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What happens to the husband's retirement decision when the wife's retirement incentives change?

Håkan Selin

Uppsala Center for Fiscal Studies
Department of Economics
Uppsala University
P.O. Box 513
SE-751 20 Uppsala
Sweden
Fax: +46 18 471 14 78

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Abstract. Several studies have documented a strong correlation in the timing of spouses' retirement decisions. However, considerably less is known about the causal impact of one spouse's retirement incentives on the retirement decision of the other spouse. Before, but not after, 2001 broad categories of Swedish local government workers in female dominated occupations were entitled to retire with full pension benefits already at the age of 63. In this paper, we utilize this reform – together with a micro data set covering the total Swedish population – to estimate the effect of a change in the wife's incentive on the husband's retirement behavior. We document a sharp decrease in pension benefit withdrawals among 63 year old wives in the local government sector in the years following the reform. However, we do not find any evidence of a response among husbands. This finding is at odds with some earlier results in the literature.

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[b] Uppsala Center for Fiscal Studies at the Department of Economics, Uppsala University, P.O. Box 513, SE-751 20 Uppsala, Sweden, Email: hakan.selin@nek.uu.se

1. Introduction

Several studies (e.g. Hurd 1990, An et al 2004 and Schirle 2008) have documented a strong correlation in the timing of spouses' retirement decisions. However, considerably less is known about the causal impact of one spouse's retirement incentives on the other spouse's retirement decision. Suppose that the wife's pension benefit accrual increases. If the substitution effect dominates, the wife works longer. But how is the husband's labor supply affected? While theories of family decision-making in general yield ambiguous predictions, it is an empirical issue of substantial policy relevance to find out in which direction the net effect goes, if there is any effect at all. If, for instance, complementarity in leisure is an important phenomenon, a delayed retirement age among female workers can potentially lead to an aggregate increase in labor force participation rates among elderly males.

There are at least two fundamental problems involved in estimating the causal impact of one spouse's retirement incentives on the retirement decision of the other spouse. Suppose that we would like to regress the husband's probability of retiring on the wife's benefit accrual rate. A first problem arises since individuals with correlated unobserved tastes for leisure are likely to marry each other. It is likely, in turn, that individuals with strong unobserved preferences for leisure sort into pension schemes with certain characteristics (e.g. pension schemes offering generous early pension benefits).[1] Thus, we can expect a spurious correlation between the error term and the regressor capturing the pension incentives of the wife.

A second identification problem occurs since pension incentives are typically functions of a set of observed personal characteristics that are plausibly also having direct effects on the husband's retirement decision. Suppose, for simplicity, that the wife's benefit accrual rate is a function of her age in a non-linear way. It would then be challenging to control for the direct impact of the wife's age on the husband's retirement decision in a flexible way without also destroying the identification.

The purpose of this paper is to estimate the effect of a change in pension incentives by exploiting a plausibly exogenous variation in the wife's retirement incentives. This procedure potentially overcomes these two methodological problems.

[1] In pension systems where one spouse's retirement incentive variables depends on the earnings history of the other spouse, a similar problem arises if that earnings history cannot be controlled for.

In Sweden, the workforce of local government employees is dominated by women. The old collectively agreed pension plan for this sector, which was in place prior to the turn of the millennium, provided ample incentives for broad occupational categories dominated by females to retire already at the age of 63. In contrast, the new DC plan provided incentives for the same categories of employees to work until they reached the age when they were obliged to retire. The first cohort of 63 year old workers to be exposed to the new pension plan was those born in 1938 – they turned 63 in 2001. In contrast, the collectively agreed pension plans in the private sector were not subject to any major changes during this time period. This allows us to compare retirement decisions over time of otherwise similar men who were married to women in the local government sector and in the private sector, respectively.

The main empirical analysis will be conducted on a population of men who were married to 63 year old women who participated in the labor force the preceding year, 2000-2005. Since the data contain the whole universe of Swedish taxpayers, the sample size is larger than what has often been the case in earlier studies on this topic. An individual is considered to be retired if he/she receives a positive amount of pension income. A register data source containing the total Swedish population is used in the analysis.

This study documents a sharp decrease in retirement rates among married female 63 year old local government employees in the years following the occupational pension reform. Indeed, seen over the whole post-reform period 2001-2005, the relative decrease amounts to around 23 percentage points as compared to the pre-reform year 2000. However, it is striking that there was no significant response in retirement rates among men married to women in the local government sector. In the most preferred model, which includes a full set of control variables, the point estimate of the average treatment effect on the treated is 0.004. The 99 confidence interval ranges from –0.014 to 0.022.

This result is clearly at odds with some earlier findings in the literature. In fact, earlier papers examining the interdependence between spouses' retirement decisions (e.g. Zweimüller et al 1996 and Coile 2004) have found an asymmetry in the way spouses react to each other's incentives: these papers find that husbands are sensitive to changes in their wives' incentives but that the opposite does not hold true. In one way or another, these studies suffer from the problems pointed out above (which, however, does not necessarily imply that such a causal effect does not exist in the demographic groups and countries studied). To our knowledge, this is the first quasi-experimental paper on the subject that exploits a reform affecting the retirement incentives of wives but keeps the husbands' own incentive variables at a constant level.[2] Therefore, it is interesting that we obtain a different result as compared to earlier studies.

The structure of this paper is as follows. Section 2 discusses the previous literature and Section 3 contains a brief theoretical discussion. The Swedish occupational pension system, which provides the source of exogenous variation exploited in the paper, is described in Section 4. In Section 5 we discuss the empirical specification. Section 6 provides a description of the data source and a descriptive analysis. Section 7 reports the regression results, and Section 8 concludes the paper.

2. Previous literature

Most papers on retirement behavior abstract from the interaction between spouses.[3] A plausible reason for the historical neglect of the joint nature of retirement decisions is that female labor force participation has been low in most industrialized countries in the past. As discussed by Hurd (1990), the modest female participation rates have also narrowed the range of feasible studies on small survey data sets.

To our knowledge, the only quasi-experimental study that examines cross-effects between spouses is Baker (2002). He studies the introduction of the Spouse's Allowance (SPA) to the Canadian Income Security system in 1975. SPA made age related benefits available to individuals aged 60 to 64 who were married to someone aged above 65. However, the benefits were means-tested on the basis of *family* income, creating an interaction between the two spouses' retirement incentives. Baker compares the retirement behavior of males (65 to 75) and females (60 to 64) to the retirement behavior among those who were not eligible due to the age of their spouse. He finds that there was a reduction in labor force participation rates among eligible males (aged 65 to 75), and he also detects a relative decrease in labor force participation among eligible

[2] It should be noted that the retirement incentives did in fact also change for the relevant cohorts of males. These changes affected treatment and control groups in a similar way though.

[3] See Coile and Gruber (2007) for a recent overview of empirical studies on how retirement decisions relate to financial incentives.

females. It should be noted that Baker's (2002) paper has a different focus than ours: Here, we study a reform affecting the retirement incentives of wives while keeping the husbands' own retirement incentive variables at a constant level. Thus, we can isolate the effects of the wife's incentives on the husband's decision.

Also in a reduced form, albeit not quasi-experimental, framework, Coile (2004) regresses the probability of retiring for each spouse on forward looking incentive measures for both spouses separately. Coile, who uses U.S. survey and register data from the 1992-2000 waves of the Health and Retirement Study, finds that the woman's retirement incentives have significant cross effects on her husband's retirement decision. Her estimation sample includes 6,204 observations that contain information on 1,152 unique couples 1980-1999. More specifically, when the wife's social security wealth accrual increases, the husband's probability of retiring decreases. However, she did not find the converse to be true: she did not detect any effects of the husband's incentive variables on the wife's retirement decision.

Coile's study potentially suffers from both methodological problems that were discussed in the introduction. As the independent variables, she includes incentive measures that are functions of e.g. the age of both spouses, life time earnings, current earnings and the social security benefit formula. Since these factors are most likely to have an independent impact on the retirement decision, she controls for these variables in a flexible way. Therefore, it appears as if the identification of the model rests upon the non-linear functional form of the social security benefit formula.[4] In contrast, the present study relies on identifying assumptions that are arguably more transparent.

With a special emphasis on retirement age, Zweimüller et al (1996) also study the interdependence between the two spouses' retirement decisions on an Austrian cross sectional data set of 1,886 individuals from a single year – 1983. The authors estimate a bivariate probit model for joint retirement, where the key left-hand side variables are eligibility dummies for early and regular retirement and the (imputed) earnings replacement ratios for both spouses. Even though the institutional context and the analyzed incentive measures are different, Zweimüller et al obtain results that are similar to those reported by Coile (2004): Husbands react to the wives' incentives, but not vice versa.

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[4] The same interpretation is made by Liebman et al (2009).

While the above mentioned works are agnostic on the exact mechanism generating the observed behavior, a related, structural, strand of literature explicitly estimates parameters of a family model. Gustman and Steinmeier (2000) adopt a non-cooperative dynamic model of family decision making and estimate a structural model where the retirement decision of one spouse depends on the retirement decision of the other spouse. On a sample of 564 U.S. couples from the time period 1969-1989, they find that the husband's retirement does not affect the retirement decision of the wife, but that the wife's retirement has a notable effect on the retirement decision of the husband. Gustman and Steinmeier (2004) exploit more recent survey data and use additional information on stated preferences on whether the interviewed spouse values leisure time together with his/her spouse or not. They find that the husband's retirement status influences the wife's retirement decision only if she values spending time in retirement with her husband. For husbands, the effect of the wife's retirement on his own retirement is roughly doubled if the husband enjoys spending time in retirement with his wife, but there is some effect even if he does not. On Norwegian data, Harnæs et al (2006) have estimated non-cooperative models of different types.

Another approach, which was proposed by Blau (1998) and further used by Mastrogiacomo (2002), is to estimate discrete choice, discrete-time retirement models. In these papers, family behavior is viewed as an outcome of a utility maximization process where the family maximizes a weighted sum of individual utility functions. In addition to taking financial incentives into account, the model also allows for past labor force behavior to affect current decisions while controlling for permanent unobserved differences across couples. The couple-specific permanent error component is treated as a random effect to be integrated out of the likelihood function.[5] On longitudinal data from the U.S., Blau (1998) finds strong associations between the labor force transition probabilities of one spouse and the labor force status of the other spouse. Interestingly, this pattern cannot be explained by financial incentives. It actually appears as if these associations are due to preferences for shared leisure.

[5] Hence, unobserved heterogeneity at the household level, which might arise due to assortative mating in the marriage market, is assumed to be uncorrelated with the regressors of interest, e.g. the incentive variables.

3. Theoretical discussion

3.1 The retirement decision

We first briefly discuss the individual's retirement decision in isolation. The total worker's compensation, $\hat{w}$, includes both remuneration in the form of cash wages, but also 'fringe benefits' such as pension wealth accruals (Lazear 1985). The total worker's compensation at age *s* can be written as

$$\hat{w}_s = w_s + E(W_s \mid r = s+1) - E(W_s \mid r = s),$$

where w_s is cash compensation and $E(W_s \mid \cdot\,)$ is the stock of expected pension wealth conditional on the individual retiring in period *r*. The pension wealth accrual can be decomposed into two components:

$$\begin{aligned} E(W_s \mid r = s+1) - E(W_s \mid r = s) &= -b_s + \left\{ \sum_{t=s+1}^{T} E(b_{s+1} \mid r = s+1) - \sum_{t=s+1}^{T} E(b_{s+1} \mid r = s) \right\} = \\ &= -b_s + m_s \end{aligned}$$

where b_s is the pension benefit that the individual foregoes while working one additional year. For simplicity, the discount rate has been set equal to 0. The second term captures the change in the sum of discounted expected yearly future pension benefits earned by the individual by working one additional year. When $m_s \geq b_s$, the pension system is said to be actuarially neutral (Queisser and Whitehouse 2006).

One way of conceptualizing the trade-off between years of retirement and the amount of life time resources (consumption) for a single individual is to define a life-time budget constraint. Suppose that the individual knows with certainty that she will live *T* years. The individual either works or is retired, and retirement is assumed to be an absorbing state. Let *R* refer to the number of years in retirement. The number of years worked is then $h = T - R$. For simplicity, we abstract from issues related to intra-period allocations and uncertainty. The individual earns an annual wage of w if she chooses to work and receives a yearly pension benefit of $b(h)$ if she retires.

If the model is formulated in continuous time, the slope of the lifetime budget constraint is given by $\hat{w} = w - b + m$.[6] In a defined benefit (DB) system, where the individual becomes eligible for full pension withdrawals at a certain age, the life time budget constraint is typically piecewise linear for a given w with kink points at ages where the benefit accrual rate changes. As in the standard static labor supply model, the optimal choice of h is given by the tangency condition that the marginal rate of substitution between C and h equals $\hat{w}$. In general, changes in $\hat{w}$ bring about a positive substitution effect and a negative wealth effect (if leisure is a normal good).

In the present paper, we analyze a reform that increased the benefit accrual rate for local government sector employees without affecting the pension wealth earned up to December 31, 1997. Since the benefit accruals from inframarginal working years are to a large extent unaffected, the income effect from the reform is likely to be small. Another issue is that the reform was unexpected for local government employees, at least in the sense that it was unknown prior to 1998. Thus, the first cohort that was treated by the reform (the 1938 cohort) knew of the reform as 60 year old workers. As discussed by Burtless (1986), unanticipated changes to retirement incentives pose special challenges in the structural modelling of retirement behavior. Since time cannot be made to run in reverse, the individual cannot retire at an earlier date even though this would have been optimal if the unanticipated change had been known at the beginning of the life cycle.[7]

3.2 The interaction between spouses

To get an idea of the basic mechanisms involved, we first consider a fairly unrealistic sequential model where the wife first maximizes life time utility without considering the optimal choices of the husband. Moreover, we assume that the husband treats the wife's lifetime earnings and pension benefits as unearned income in his own optimization

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[6] Formally, in continuous time, the budget constraint can be written as $C = wh + \int_h^T b(h,s)ds$. The slope of the budget constraint is then $\frac{dC}{dh} = w - b(h,s) + \int_h^T \frac{\partial b(h,s)}{\partial h} ds$.

[7] Partly motivated by this limitation, Stock and Wise (1990) proposed an 'option value' approach to the modelling of the individual's retirement behavior. This structural technique allows individuals to update their expectations of wages and retirement benefits.

decision.[8] The wife maximizes $u_f(C_f, h_f)$ subject to the life time budget constraint $C_f = w_f h_f + \int_{h_f}^{T} b(h_f, s)ds$.[9] While still using the notation introduced in the previous section, we now let subscript *f* refer to the wife and subscript *m* to the husband. In this setting, an increase in the benefit accrual rate will have a positive substitution effect and a negative income effect on the wife's retirement decision.

As the second mover, the husband optimizes while taking the optimal choices of his spouse as given. Suppose now that the husband solves the following problem

$$\max_{C,h_m} u_m(C, h_m, h_f)$$

$$\text{subject to } C = w_m h_m + \int_{h_m}^{T} b(h_m, s)ds + w_f h_f + \int_{h_f}^{T} b(h_f, s)ds, \qquad (1)$$

where the two latter terms on the right-hand side of the budget constraint are exogenous. Similarly to his wife, the husband knows with certainty that he will live for *T* years. Under these assumptions, the life cycle labor supply choice of the wife influences the husband's retirement decision through two channels. First, since the wife's retirement years enter the husband's utility function, the wife's labor supply decision has an externality. It is most often reasonable to believe that the husband's marginal utility from retirement years increases in the wife's retirement years (i.e. the husband enjoys spending time with his wife). Second, the wife's choice of life time income has an income effect on the husband's retirement decision.

In this paper, we analyze a situation where the wife's benefit accrual increases. Suppose that the substitution effect dominates so that the wife delays her retirement. Then, we can expect two effects going in opposite directions. The *shared leisure effect* will induce the husband to work longer. The *income effect* will lead to a reduction in working years for the husband, given that leisure is a normal good.

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[8] In the context of the standard static labor supply model where the husband first chooses his optimal supply of labor, Killingsworth (1983) named this model the 'male chauvinist model'.

[9] Throughout this discussion, we neglect the possibility that the benefit formula can be dependent on the retirement behavior of the other spouse.

These two mechanisms are present also in more general models of family labor supply behavior. A general class of models, which only assumes the outcome of the household decision making to be pareto-efficient, but nothing about the particular bargaining rule, is cooperative models of family decision making (or ‘the collective household model’).[10] In this framework, spouses simultaneously solve the following constrained optimization problem

$$\max_{C,h_I} u_I(C, h_I, h_J)$$

subject to (1) and $u_J(C, h_J, h_I) = \bar{u}_J$

for $J \neq I$. The second constraint guarantees that the outcome is pareto-efficient. In other words, there is no way of making *I* better off without making *J* worse off. In this framework, an increase in the benefit accrual rate for the wife would not only affect the husband through the shared leisure effect and the income effect. It would also change the relative bargaining power between spouses. When the second constraint is dropped, we arrive at non-cooperative bargaining models where once more an exogenous change in the benefit accrual rate will have consequences for the relative bargaining power in the household.

It should be emphasized that the reduced form empirical analysis of this paper will be insufficient to discriminate between different models of family decision making. Instead, we will be able to recover the *net effect* of a change in the wife’s incentives on the husband’s retirement decision by pursuing a transparent identification strategy.

4. The Occupational Pension System[11]

4.1 General description

In general, Swedish retirees obtain most of their pension income from the public pension system that is financed via payroll taxes. However, the occupational pension

[10] See Vermeulen (2002) for a good survey of collective household models.
[11] See also Glans (2009) for a description of the occupational pension system for local government workers. Glans does not put any emphasis on the retirement age, however.

system is a crucial complement.[12] In contrast to the U.S., where the characteristics of employer-sponsored 401 (k)-plans differ between firms, the rules governing the Swedish occupational pension system are quite uniform within a small number of large sectors of the labor market. The pension plans are determined by collective agreements that have been signed by the unions and the employers' associations. An overwhelming majority of Swedish employees (around 90 percent) are covered by collective agreements. There are four main occupational pension schemes: one for blue-collar private sector workers, one for white-collar private sector workers, one for local government employees and one for state-level government employees. Here, the focus will be on the reforms in the scheme for local government workers that occurred around the turn of the millennium. During the period of study, i.e. 1999 to 2005, there were no major changes in the private sector collective agreements.

4.2 The occupational pension reform for local government workers

The old agreement, PA-KL, which came into place in 1985, was a defined benefit (DB) plan that interacted with the old Swedish public pension system in an interesting way. As described in *Appendix B*, the old pension plan was a so-called *gross pension system*. Thus, the occupational pension plan stipulated that the sum of the annual occupational pension benefit and the annual pension benefit from the public pension system should amount to a certain fraction of the individual's qualifying income.

For a 63 year old local government female full time employee with an average wage rate, the gross replacement rate was around 73 % in 2000.[13] Hence, given that this employee was eligible for annual benefits from the public pension system of, say, 60 % of her qualifying income, the annual occupational pension benefit amounted to 13 % of her qualifying income. Thus, the occupational benefit was the residual amount.

Under the old DB plan, different occupational categories in the local public sector faced different retirement ages. The 'retirement age' could either be defined as a

[12] In 2007, pension benefits amounting to SEK 232 billion were transferred from the public pension system to retirees, whereas benefits amounting to SEK 69 billion were distributed from the occupational pension funds (Glans 2009).

[13] An average full-time wage for a female worker aged 60-64 in the local government sector in 2000 was SEK 17,700 (see SCB, http://www.ssd.scb.se/databaser/makro/Produkt.asp?produktid=AM0106). If the female worker earned the same real wage in the last five years prior to retirement, she would have a yearly qualifying income of 5.80 price base amounts (BA = 36,600 in 2000). Thus, her gross pension benefit (i.e. the sum of public pension and occupational pension benefits) would amount to 72.6 of her qualifying income.

specific point in time (the month the employee turned 65) or an interval, 63-65. At the lower end of the age interval, the employee was entitled to full retirement benefits given that he/she fulfilled the other criteria for eligibility. At the upper end of the time interval, the employee was obliged to retire unless the employer offered a prolongation. Early withdrawals before the retirement age could be made from the age of 60 under the PA-KL agreement. Early retirement was subject to a penalty.[14] Likewise, a small adjustment was made for every month the employee postponed retirement after his/her 65th birthday.

Occupational pension recipients were subject to an earnings test. The main rule in the PA-KL plan was that benefits decreased at a rate of 100% in the benefit recipient's earnings. In the post-reform PFA98 plan, the reduction rate was 73.5 %. However, the benefit recipient could get special provisions to work under both plans.

Under simplified assumptions, *Table 1* shows the benefit accrual rates at different ages under the old PA-KL plan for an individual who was subject to the 63-65 rule. It is interesting that the benefit accrual became negative already when the individual turned 62. Thus, even though the individual could increase her annual pension benefit by postponing retirement one year, the early withdrawal penalty did not fully offset the gain from receiving pension benefits the year she was 62. In other words, the system was not actuarially neutral at this age. As expected, there was a sharp fall in the benefit accrual rate at the age of 63. At this point, the individual did not earn any incremental pension rights.

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[14] The early withdrawal penalty is further described in *Appendix B.*

Table 1. Benefit accrual rates for different ages under the PA-KL plan for an individual who is subject to the 63-65 rule.

Age	Discount rate = 0.05	Discount rate = 0.02
60	2.93	4.16
61	2.77	3.82
62	-1.63	-0.98
63	-3.89	-4.01

The benefit accrual of age *s* is the difference in pension wealth from claiming retirement benefits directly after turning *s* years old and directly after turning *s+1* years old. The stock of pension wealth has been discounted back to age *s-1*. Benefit accrual rates are expressed in price base amounts (BA). In 2010, the BA is SEK 42,400 (approximately EUR 4,700). The individual is assumed to live until she is 83 and to have a qualifying income of 5.6 BA.

Broad categories of local government workers faced the 63-65 rule, for example most occupational categories belonging to the Swedish Municipal Workers' Union (Kommunalarbetareförbundet). At present, the Swedish Municipal Workers' Union organizes 512,000 members. 74 percent of the members work in the local government sector and 81 percent are women. Large occupational groups organized by the union include assistant nurses, child minders and cleaners. However, other female dominated occupations outside the area of the Swedish Municipial Workers were also subject to the 63-65 rule. For instance, hospital nurses and pre-school teachers also had the opportunity of retiring with full pension at the age of 63.

In 1998 a new agreement, PFA98, was signed for Swedish local government employees. Those born in 1937 and earlier were unaffected by the occupational reform. The implementation of PFA98 marked a move away from a DB to a funded defined contribution (DC) system. Under the new system, the employer paid pension contributions as a certain percentage of gross wages – 3.4 % – not exceeding 7.5 increased price base amounts.[15] Thus, under the new system, the individual could always increase her pension wealth by postponing retirement until the age when she was

[15] This cap roughly coincided with the ceiling of the public pension system. Above the ceiling, the employer still made a small DC contribution (1.0 %).

obliged to retire.[16] For wages above the cap of 7.5 increased price base amounts, a DB system was retained. The new PFA98 agreement came into effect on January 1, 2000 for those born in 1938 and later. Thus, the first cohort of 63 year olds to earn benefit accruals from the new DC system turned 63 in 2001.

The benefit accrual rates at the age of 63 for those born 1938 and after were, in principle, unaffected by the old PA-KL agreement. However, there was an important transitional rule in place, implying that the occupational pension reform did not change the stock of pension wealth to any considerable extent. From the age of 65, local government employees were entitled to a life annuity that was a function of the individual's employment history up to December 31, 1997. It corresponded to the annual pension benefit that the individual would have received if she had retired by December 31, 1997. If the individual was subject to the 63-65 rule under the old agreement, the life annuity was multiplied by a factor of 1.094. As the PFA98 agreement provided an opportunity to withdraw the life annuity before the age of 65, those who were subject to the 63-65 rule under PA-KL could also post-reform benefit from it when retiring before 65. Naturally, the importance of this transitional rule was declining in birth year.

It should be noted that those born in 1938 were also the first cohort to be exposed to the new Swedish public pension system, which has been described by Sundén (2006), for instance.[17] The new public pension system was phased in gradually. Those who were born in 1938 received 4/5 of their benefits from the old system and 1/5 from the new one. Each cohort then increased its share of benefits from the new system by 1/20. Benefits from the new system were distributed for the first time in 2001.

As a general rule, both the Swedish occupational pension system and the public pension system were individual-based. Thus, the stream of pension benefits received upon retirement was independent of the earnings-history of the spouse. The exception was the survivor's pension, which under some conditions could be claimed both from

[16] The DC component of the new system was a fully funded system. The first fund choice was held late 2000/early 2001. Approximately 60 % of local government employees made an active fund choice.

[17] The idea behind the Swedish pension reform, which was announced in the early 1990's and finally legislated in 1998, was to design a fiscally sustainable system with clear links between contributions and benefits. In addition, the Swedish policy makers had the ambition of enhancing work incentives. In the old system, benefits were defined based on the individual's earnings during the 15 best years of employment. For full eligibility, the individual was required to have worked 30 years. In the new system, pension benefits are based on life time income.

the public pension system and the occupational pension system. In fact, the rules for the survivor's pension were similar in the old PA-KL plan and in the new PFA98 plan. Therefore, the survivor's pension does not pose any threat to identification in this paper.

5. Empirical model

The idea of this paper is to estimate the effect of the pension reform on the husbands' behavior. The empirical model will allow us to recover the average treatment effect on the treated (ATET). The 'treated population' consists of husbands married to female *local government sector* workers who are 63 years of age and who belonged to the labor force in the preceding year. The 'untreated population' consists of men married to female *private sector* workers who are 63 years of age and who belonged to the labor force in the preceding year. In the next section, we will further discuss the definitions of these concepts. The 1938 cohort of female local government workers – the first cohort to be exposed to the new DC plan – turned 63 in 2001. In the main difference-in-difference regressions, we will use 2000 as the pre-reform year and 2001-2005 as the post-reform years.

The relevant regression equation reads

$$RETIRED_{it}^{husband} = \sum_{t=2001}^{2005} \beta_t \left(LGW_i \times YEAR_t \right) + \gamma_{LGW} LGW_i + \sum_{t=2001}^{2005} \gamma_t YEAR_t + \delta\ X_{it} + \varepsilon_{it} \tag{2}$$

for all i with wives in the labor force in the year $t-1$, where i is an individual index and t is a time index. $RETIRED_{it}^{husband}$ is an indicator variable that takes the value of 1 if the husband is retired. LGW_i is an indicator that is 1 if the wife was employed in the local government sector in year $t-1$. $YEAR_t$ is a year t specific-fixed effect. X_{it} is a vector of family characteristics that also includes a constant term. ε_{it}, finally, is an error term. Equation (2) will be estimated in the form of a linear probability model. Note that no constraint is imposed on the husband's labor force participation in $t-1$.

For two reasons, one might expect the treatment effects to be heterogeneous for the years 2001-2005. First, if norms adjust slowly in response to the new pension plan, one can expect a delayed response among the treated wives and, as a consequence, a delayed response among husbands. Second, as described in *Section 4*, there was a transitional rule in place post-reform that affected the stock of pension wealth more for older than for younger cohorts.

Therefore, we also report estimates of the treatment effect for each specific year. The crucial identifying assumption is that the outcome variable evolved in the same way in the treated group (husbands married to wives in the local government sector) as in the non-treated group (husbands married to wives in the private sector) in the absence of a pension plan reform. Below, in *Section 7.3*, we will further discuss the identifying assumption.

6. Data issues

6.1 Data source, sample selection and variables

This study exploits a register data source that covers the total Swedish population for the years 1998 to 2005.[18] Crucially, since there is an identifier for the household, we observe married couples. Furthermore, a number of demographic characteristics are included in the data. Importantly, there is information on sector affiliation for all years. From the sector affiliation code, it is possible to observe whether the individual is employed in the local government sector, the central government sector or in the private sector.

As stated in the introduction, the population of interest in this study is men married to women aged 63. The population is also conditioned on the wife belonging to the labor force in the preceding year. It was necessary to impose this constraint since information on sector affiliation was only available in the data for those who actually did work. Here, labor force participation is defined as having positive labor earnings and no pension income.[19] Since lagged information is used to select the populations, we can use the years 1999 to 2005 in the analysis. Some additional exclusions were made. First, we excluded husbands who were employed in the local government sector, as

[18] The data set was ordered from Statistics Sweden for the purpose of studying issues related to retirement savings.

[19] Pension income is defined as the sum of the two variables ‘tpensa’ and ‘tprivl’. Labor income is defined as the sum of ‘ttjlon’, ‘nakte’ and ‘nakthb’.

these husbands might be subject to the occupational pension reform. Second, we also omitted couples where either the husband or the wife was affiliated with the central government sector. This was done since a new collective agreement, which affected retirement incentives, came into place during the period of study (2003) in this sector. After these exclusions, the regressions sample contains 68,354 observations.[20]

It should be noticed that even though it is possible to identify individuals as local government employees, it is not feasible to observe eligibility for the 63-65 rule. Thus, the treatment group also includes wives who faced a retirement age of 65 in the pre-reform system. However, all wives in the local government sector went from a DB to a DC plan regardless of retirement age. Note also that the individual's stock of social security wealth cannot be calculated as the data set both lacks information on earnings histories and information on specific retirement age rules.

The main advantage of register based data – as compared to smaller self-reported survey data sets – is that some types of measurement errors are, in principle, absent and that the entire population can be studied. On the other hand, as discussed by Hallberg (2008) for instance, a disadvantage in the context of studying retirement decisions is that the tax registers provide data on a yearly, rather than monthly, basis. Hence, we typically observe pension income and earnings, but it is not possible to tell whether the individual has been retired on a part time basis or on a full time basis for a shorter period of time. In our data, the age variable is defined based on the age of the individual as of December 31. Accordingly, some of the '63 year olds' have been aged 62 and have faced the marginal pension incentives applying to 62 year olds during 11 months of the year.

In this paper, we define the indicator for retirement in the following way: Individuals who receive a positive amount of pension income will be classified as retired.[21] Naturally, this is quite a generous definition, which also includes those who

[20] Henceforth, unless otherwise stated, we let the term 'regression sample' denote the sample that contains the 59,895 observations for the years 2000-2005 (to be used in the main analysis) *and* the 8,459 observations for the year 1999 (to be used in the placebo analysis).

[21] Due to the strict earnings testing of occupational benefits (described in *Section 4.2*), the benefit withdrawal should be a good proxy for the retirement decision of 63 year old local government employees. For private sector employees, i.e. the husbands in our analysis, the situation is somewhat more complicated. Private sector blue-collar workers were allowed to work and receive occupational pension benefits at the same time from the age of 55. White-collar private sector workers were only allowed to work 8 hours a week if they withdrew pension benefits before they turned 65. As reported in *Section 7.2* below, we have also run regressions where the husband's earnings is the dependent variable.

only withdrew a smaller amount of pension income. Moreover, individuals who received a disability pension can also be subsumed under this retirement concept prior to 2003.[22]

The dummy variable for being a local government employee is defined based on information from the preceding year. We also divide the studied population into six educational categories and define dummies for the educational level based on these. Finally, we also construct 21 dummy variables for the county of residence. Descriptive statistics for all variables used in the empirical analysis are reported in Appendix A.

6.2 Graphical analysis

It is a well-known empirical regularity that husbands are typically older than their wives. In the estimation sample, the mean age difference between husbands and wives is 2.9 years. *Figure 1* depicts the distribution of husbands' ages, where the husbands are married to wives who are 63 years of age. The left (solid) vertical line depicts the wife's age 63, and the right (dashed) vertical line displays the normal retirement age 65. The median age of husbands married to women aged 63 is 66 in the sample. For ages exceeding 64, virtually all husbands in the population claim at least some pension benefits and are therefore classified as retired. This can be seen in *Table B2* of Appendix B. Accordingly, the action in the sample occurs to the left of the dashed line of *Figure 1*.

[22] Disability pension income was classified as sickness insurance benefits (labor income) from 2003 and onwards, but as pension income prior to 2003. Probably for that reason, there is a large increase in labor income 2003 and a decrease in pension income that year.

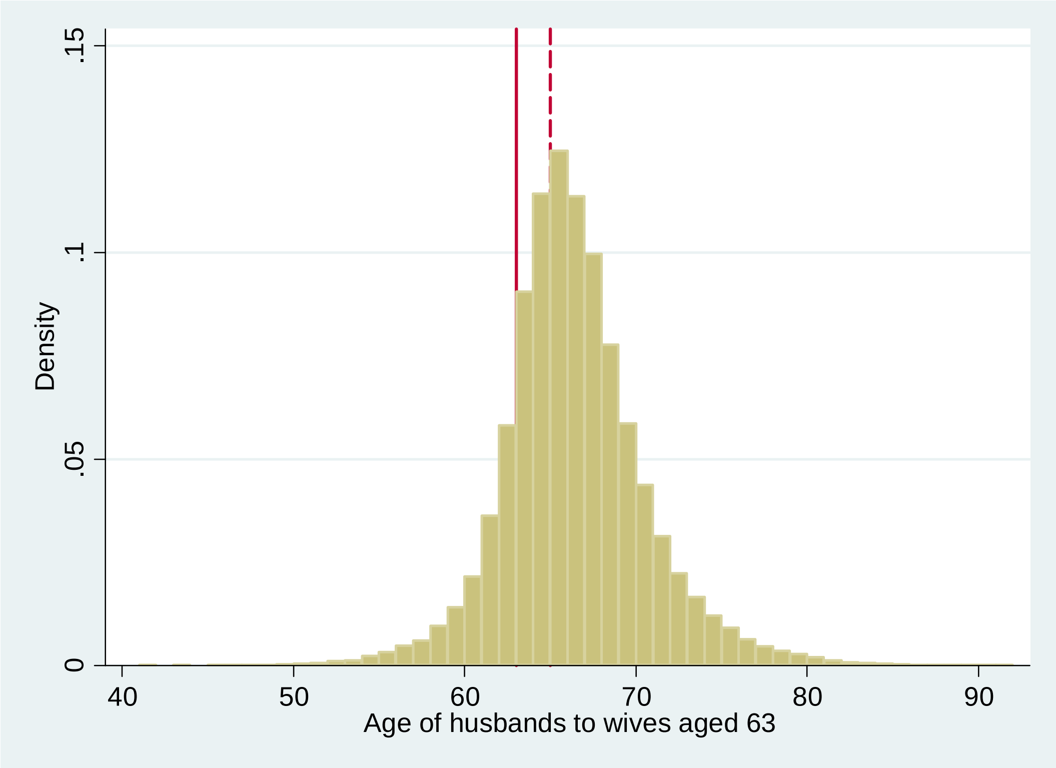

Figure 1. The age distribution of husbands married to wives aged 63 in the estimation sample. The left (solid) line depicts the wife's age (63) and the right (dashed) line depicts the normal retirement age (65).

Figure 2.A illustrates the probability of retiring for married females, conditional on being in the labor force in the preceding year. Among the 63 year olds in 2000, which was the last cohort to be exposed to the old DB plan, the mean conditional retirement probability was 0.47. However, in 2001 it had dropped to 0.35. Apparently, this is a sharp decline. In 2003, there was a new decrease in the conditional retirement probability and from 2003 and onwards, the conditional retirement rates almost converged in the local government sector and the private sector.

Figure 2.B shows the evolution of retirement rates of men who were married to women aged 63 in the local government sector and the private sector, respectively. From ocular inspection, we note that the two lines track each other fairly well. There is a slight increase between 2001 and 2002 among husbands married to local government sector wives, whereas there is a decrease in the other group.

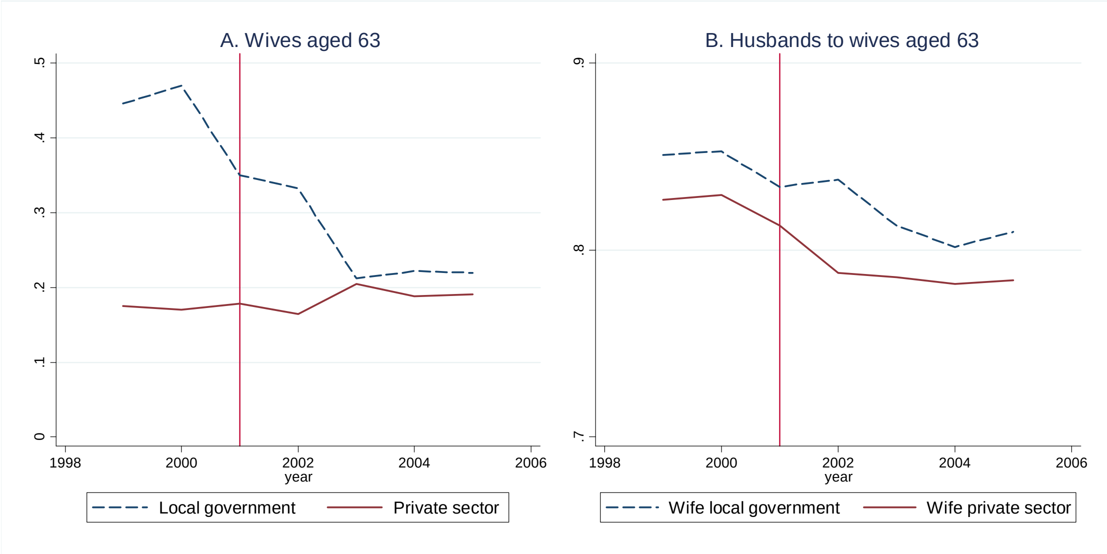

Figure 2. Conditional mean retirement probabilities for 63 year old wives and unconditional mean retirement probabilities for husbands 1999-2005. Note the difference in scales!

For the difference-in-difference approach to be valid, the group-specific fixed effect should be the same across years – γ_{LGW} of equation (2) should be time invariant. Thus, we want the composition of the treatment and control groups to be unchanged by the reform. If there were a large response among 62 year old women to the occupational reform, a duration model would be called for.[23] Of particular interest is the fact that those who were aged 63 in 2001 were exposed to the new DC plan already in the year 2000 as 62 year olds. *Figure 3.A* depicts the evolution of the retirement rate for 62 year old women, conditional on being part of the labor force as 61 year olds. We infer that the decrease in retirement rates was modest in 1999-2000 among the 62 year olds. Moreover, the conditional retirement rates have been fairly stable throughout the time period. Interestingly, the level of retirement rates was always lower in the local government sector than in the private sector. *Figure 3.B* reveals that the husbands' retirement probabilities in the two groups evolve in pretty much the same way.

[23] At present, the empirical analysis includes individuals born 1936-1942. Since the untreated cohorts of 1936 and 1937 can only be traced back to 1998, the possibilities of conducting an interesting duration analysis are nonetheless limited here. See An et al (2002) for a duration model of joint retirement.

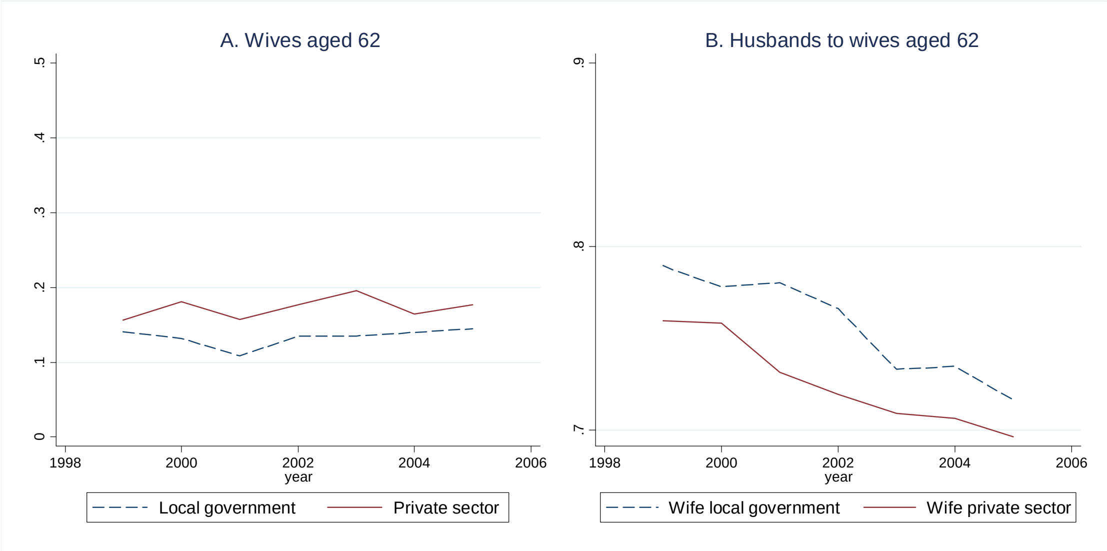

Figure 3. Conditional mean retirement probabilities for 62 year old wives and unconditional mean retirement probabilities for husbands 1999-2005. Note the difference in scales!

7. Regression results

7.1 The wives' own response

The main purpose of this study is to examine the husbands' response to the pension reform in the local government sector. However, as a first step we examine the wives' own response to the shift in the occupational pension plan. *Table 2* reports the estimate of the average treatment effect on the treated (ATET) from an OLS difference-in-difference regression, where the wife's retirement status is regressed on a constant, the treatment dummies, a dummy for the local government sector and a full set of year dummies. Columns (1)-(2) report a treatment effect that is assumed to be homogenous across the whole period 2001-2005. Using the notation of equation (1), the homogenous treatment effect is $ATET = \sum_{t=2001}^{2005} \frac{\beta_t N_t}{N_{PR}}$, where N_{PR} is the total number of observations in the estimation sample 2001-2005. Thus, the ATET for all years is a weighted average of the year-specific estimated treatment effects.

In the specification without covariates, reported in column 1 of *Table 2*, the estimate of the ATET is –0.223. The heteroskedasticity robust standard error is small – the 99 percent confidence interval ranges from –0.249 to –0.196. According to this estimate, the probability of withdrawing retirement benefits among 63 year old wives, who were in the labor force as 62 year olds, decreased by 22 percentage points in the local government sector relative to the private sector during the post-reform period. The same qualitative pattern was noticed by Glans (2009), even though he did not focus on 63 year olds.

If the treatment dummy is orthogonal to the error term, we expect the estimate of the ATET to be unaffected by the inclusion of control variables. Intuitively, if the addition of control variables has a large impact on the estimate, there might be some concern that *unobserved* characteristics are also correlated with the regressor of interest. Column (2) reports results from a regression where a large number of control variables have been added to the regression equation. Interestingly, the treatment effect estimates are almost identical. In fact, a χ^2 test does not reject the null hypothesis that the treatment effect estimates in column (1) and column (2) are equal, the p-value being 0.155.

When the estimated treatment effects are allowed to be heterogeneous across years (columns 3 and 4), it turns out that the response does indeed differ across post-treatment years. In particular, the relative decrease in retirement rates in the local government sector – as compared to the pre-reform year 2000 – was considerably lower in 2001-2002 than in 2003-2005. This is clearly in line with the graphical evidence presented in *Figure 2.A*. Once more, the estimated treatment effects are fairly similar in the regressions with and without control variables. A χ^2 test does not reject the null hypothesis that the coefficient vectors containing the 5 coefficient estimates reported in column (3) and column (4) are equal at any reasonable level of significance, the p-value being 0.139.

These results are in line with findings in a large body of literature that estimates the effect of the individual's own retirement incentives on the probability of retiring or the retirement hazard. In fact, recent studies that exploit policy discontinuities for identification (Liebman et al 2009 and Manoli and Weber 2011) indicate that females are more responsive than males at the retirement margin.

Table 2. The wife's retirement status.

	(1)	(2)	(3)	(4)
Treatment	-0.223	-0.228		
2001-2005	(0.010)***	(0.011)***		
Treatment 2001			-0.128	-0.135
			(0.013)***	(0.014)***
Treatment 2002			-0.130	-0.140
			(0.013)***	(0.014)***
Treatment 2003			-0.292	-0.291
			(0.013)***	(0.013)***
Treatment 2004			-0.266	-0.271
			(0.012)***	(0.013)***
Treatment 2005			-0.271	-0.275
			(0.012)***	(0.013)***
Control variables	No	Yes	No	Yes

The regression sample consists of 63 year old wives who were employed in the previous year, 2000-2005. Linear probability model. The regression equation follows equation (2) in the main text. Robust standard errors are in parenthesis. * denotes significance at 10%, ** at 5% and *** at 1%. The number of observations is 59,895. All regressions include an intercept, a dummy for the local government sector and a full set of year dummies. The control variables are 17 dummies for the husband's age, 5 dummies for the wife's level of education, 5 dummies for the husband's level of education, 20 region dummies and interactions between year dummies and educational dummies.

7.2 The husbands' response

We now turn to the husbands' response, which is the main focus of the paper. The results for the husband's retirement status are displayed in *Table 3*. Remember that the male sample, in contrast to the female sample, is not conditioned on the individual being part of the labor force in the previous year.[24]

Once more, columns (1)-(2) report a treatment effect that is assumed to be homogenous across the whole period 2001-2005. Interestingly, column 1 reports a treatment effect estimate which is not statistically distinct from 0. The point estimate is

[24] As can be seen below in *Table 4*, the results are by and large robust to conditioning the regression sample on the husband also being in the labor force in the previous year. However, this inflates the standard errors somewhat since the number of observations is reduced.

0.005 and the heteroskedasticity robust standard error is approximately of the same magnitude as in the regressions for females, 0.009. When covariates are added (column 2), the treatment effect estimate is virtually unaffected. The standard error of the point estimate is now further reduced as the inclusion of covariates decreases the variance in the estimated residuals. The 99 confidence interval ranges from –0.014 to 0.022. This result suggests that a change in the wife's pension incentive does not affect the retirement decision of the husband.

Columns (3)-(4) show the estimated year-specific treatment effects. With one exception, these estimates are not statistically different from 0. The exception is 2002. The specification without covariates (column 3) generates a treatment effect for 2002 of 0.027, which is statistically significant at a level of 5 percent. If this estimate were to be interpreted as a behavioral response to the occupational pension reform, it would suggest that an increase in the benefit accrual rate for wives would lead to an *increase* in the husbands' probability of retiring.

We are, however, reluctant to interpret the 2002 estimate in this way. First, it should be noticed that the estimate is reduced from 0.027 to 0.016 when the full set of covariates is included in the regression equation (column 4). The p-value for the t-statistic then increases from 0.015 to 0.081. This raises some doubts about the robustness of this finding.[25] Second, there is no good reason why we would observe this effect in 2002, but not in any other year. The wives' response is estimated to be the largest for 2003-2005.

[25] Actually, the inclusion of covariates has a larger impact on the 2002 estimates also for the wives' response (see columns (3)-(4) of *Table 2*) as compared to the other post-treatment years. When testing for parameter equality for each pair of estimated year-specific treatment effects for wives, it turns out that the null hypothesis of parameter equality can only be rejected at any reasonable level of significance for 2002. If we do the same exercise for husbands, we cannot reject parameter equality in any post-treatment year, including 2002.

Table 3. The husband's retirement status.

	(1)	(2)	(3)	(4)
Treatment	0.005	0.004		
	(0.009)	(0.007)		
Treatment 2001			-0.003	0.008
			(0.011)	(0.009)
Treatment 2002			0.027	0.016
			(0.011)**	(0.009)*
Treatment 2003			0.004	0.008
			(0.011)	(0.009)
Treatment 2004			-0.005	-0.006
			(0.011)	(0.009)
Treatment 2005			0.002	-0.001
			(0.011)	(0.009)
Control variables	No	Yes	No	Yes

The regression sample consists of 63 year old wives who were employed in the previous year, 2000-2005. Linear probability model. The regression equation follows equation (2) in the main text. Robust standard errors are in parenthesis. * denotes significance at 10%, ** at 5% and *** at 1%. The number of observations is 59,895. All regressions include an intercept, a dummy for the local government sector and a full set of year dummies. The control variables are 17 dummies for the husband's age, 5 dummies for the wife's level of education, 5 dummies for the husband's level of education, 20 region dummies and interactions between the year dummies and the educational dummies.

All this suggests that husbands did not respond to the change in their spouses' incentives. This finding is surprising given the well-documented correlation in the timing of spouses' retirement decisions. The results are also at odds with the findings in e.g. Zweimüller et al (1996) and Coile (2004) which both find that husbands are sensitive to changes in their wives' incentives.

What can explain the discrepancy between the results in this study and earlier work in the area? A feature of this study, which is somewhat novel as compared to previous studies, is that the estimation sample consists of couples where the wife is 63 years old and participated in the labor force in the preceding year. These restrictions were imposed as a consequence of the structure of the occupational pension plan reform and the structure of the data.[26] As women tend to marry men who, on average, are older than themselves, the estimation sample is dominated by males who are fairly old in a retirement context – the mean age of husbands is 65.9. As is apparent from *Table A.2* of Appendix A, virtually all men withdraw pension benefits at the age of 65.

Even though almost all males aged 65 and above received pension benefits, they could still report positive earnings. Therefore, we also ran the same set of regressions as reported in *Table 3* where we replaced the probability of retiring with the husband's earnings (and the log of husband's earnings). The estimated parameters from these regressions also reflect a continuous dimension of the labor supply decision. It turned out that none of the ATET estimates (neither the homogenous treatment effects nor the heterogeneous treatment effect estimates) were significantly different from 0 at a level of 10 percent.

What would happen to the results in the main analysis if we also conditioned the estimation sample on the husband being part of the labor force in the preceding year? This restriction dramatically reduces the sample size, and the mean male age in the sample falls to 62.11 years. Columns (1)-(2) of *Table 4* show results for wives and husbands, respectively, under this sample restriction. The ATET for wives is now 2.6 percentage points lower in absolute terms, but still roughly of the same magnitude. The point estimate of the ATET for husbands is still close to 0 – it is 0.007. Thus, our main results are, by and large, robust to conditioning the estimation sample on the husband's labor force participation in the preceding year.

A concern might still be that the results are now potentially driven by the fact that the estimation sample is dominated by males who are younger than their wives. Therefore, we make an even more ambitious sample limitation. In addition to the sample restriction that husbands were in the labor force the previous year, we now only include couples where both spouses were aged 63 in the estimations. The results are reported in columns (4)-(5) of *Table 4*. Even though we started out with the total Swedish population, this exercise leaves us with a sample size of 2,799. The ATET for females is −0.217, as compared to −0.228 in the main specification. Moreover, the

[26] The reform had the largest impact for 63 year olds. Data on occupational sector affiliation are only available if the individual is employed.

ATET estimate for males is close to 0 – the point estimate being 0.008. Unsurprisingly, the standard errors are larger in columns (4)-(5).

Table 4. Alternative sample restrictions and correlations between the retirement decisions of the husband and the wife.

	Sample conditioned on the husband's participation last year			Sample conditioned on the husband's participation last year and the husband being aged 63		
	Wife's retirement	Husband's retirement	Husband's retirement	Wife's retirement	Husband's retirement	Husband's retirement
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment	-0.202	0.007		-0.217	0.008	
	(0.023)***	(0.018)		(0.055)***	(0.051)	
Wife's retirement			0.094			0.172
			(0.007)***			(0.021)***
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Observations	15,518	15,518	15,518	2,799	2,799	2,799

All regression samples consist of 63 year old wives who were employed in the previous year, 2000-2005. Linear probability model. The regression equation follows equation (2) in the main text. Robust standard errors are in parenthesis. * denotes significance at 10%, ** at 5% and *** at 1%. All regressions include an intercept, a dummy for the local government sector and a full set of year dummies. Control variables are 17 dummies for the husband's age, 5 dummies for the wife's level of education, 5 dummies for the husband's level of education, 20 region dummies and interactions between the year dummies and the educational dummies.

The robustness checks reported in *Table 4* indicate that the surprising zero-result for males is not driven by the non-standard age-composition of the sample. A more interesting source of discrepancy between the present study and previous research is that we exploit a reform affecting the retirement incentives of wives while keeping the relative retirement incentives of husbands at a constant level. By utilizing exogenous variation in the wives' retirement incentives, we circumvent endogeneity problems that might have plagued the previous literature.

The correlation between the husband's and the wife's retirement decisions is indeed present also in our estimation sample. If one replaces the treatment dummies (i.e.

$\sum_{t=2001}^{2005} \beta_t \left(LGW_i \times YEAR_t \right)$) of equation (2) with a dummy for the wife's retirement, we get the partial correlation between the husband's and the wife's retirement decisions. Since more than 80 % of the husbands in the main estimation sample are already retired, this partial correlation is rather low in the sample used in the main analysis, 0.035 (robust standard error 0.003). *Table 4* reports the same coefficient for the two more narrowly defined subsamples that were used in the robustness analysis. In the subsample consisting of couples where both spouses are aged 63 and were part of the labor force in the preceding year, the partial correlation was as large as 0.17.

7.3 Placebo test and further robustness analysis

Once more, the crucial identifying assumption is that the outcome variable would have evolved in the same way in the treated group (husbands married to wives in the local government sector) as in the non-treated group (husbands married to wives in the private sector) in the absence of a pension plan reform. To obtain a view on the validity of this identifying assumption, we will exploit the fact that our data set allows us to define the dependent variable also for 1999, an earlier pre-reform year.[27] Therefore, we can perform 'placebo-type' regressions where we assume that the pension plan reform instead affected the 1937 cohort (those who turned 63 in 2000), but not the 1936 cohort (those who turned 63 in 1999). In reality, both cohorts were unaffected by the reform. The following regression will be performed on data for 1999 and 2000:

$$RETIRED_{it}^{husband} = \beta_{placebo} \left(LGW_i \times YEAR_{2000} \right) + \gamma_{LGW} LGW_i + YEAR_{2000} + \delta \; X_{it} + \varepsilon_{it} \quad (3)$$

The placebo tests for wives, with and without controls, are reported in columns (1)-(2) of *Table 5*. If the pre-reform trends were the same for the two groups, the coefficient for the interaction term in equation (2) should not be significantly distinct from 0. *Table 5* reveals that the retirement rates actually rose somewhat in the 'treated' group as compared to the 'non-treated' group 1999-2000. However, the increase is about an order of magnitude smaller than the estimated decrease in retirement rates in the main analysis. This is, by the way, consistent with the graphical evidence provided

[27] The data set contains data for the years 1998-2005. But since we need to condition on labor force participation in the preceding year, we cannot include the retirement decision in 1998 in the analysis.

in *Figure 2.A*. It is comfortable, however, that the placebo tests for husbands (columns 3-4) generate treatment effect estimates that are very close to 0.

Table 5. Placebo regressions.

	(1)	(2)	(3)	(4)
Treatment 2000	0.027	0.035	0.001	0.004
	(0.014)**	(0.014)**	(0.011)	(0.009)
Control variables	No	Yes	No	Yes

The regression sample consists of 63 year old wives who were employed in the previous year, 1999-2000. Linear probability model. The regression equation follows equation (2) in the main text. Robust standard errors are in parenthesis. * denotes significance at 10%, ** at 5% and *** at 1%. The number of observations is 17,231. All regressions include an intercept, a dummy for the local government sector and a full set of year dummies. The control variables are 17 dummies for the husband's age, 5 dummies for the wife's level of education, 5 dummies for the husband's level of education, 20 region dummies and interactions between year dummies and educational dummies.

A potential concern, highlighted by the placebo test exercise, is that retirement rates among 63 old females were at a temporarily high level in 2000 in the local government sector. Ideally, one would like to have access to a large number of pre-reform years to rule out this possibility. The best we can do here is to use data from both available pre-reform periods, 1999 and 2000, in the difference-in-difference regressions.[28] When the additional information from 1999 is exploited, we still obtain results that are consistent with the main analysis. The estimated ATET for males (females) now amounts to 0.006 (–0.210) with a robust standard error of 0.005 (0.008) in the specification that includes the full set of control variables.

8. Concluding discussion

In Sweden, the workforce of local government employees is dominated by women. The old collectively agreed pension plan for this sector, which was in place prior to the turn of the millennium, provided ample incentives for broad occupational categories dominated by females to retire already at the age of 63. In contrast, the new DC plan

[28] The estimating equation then becomes $RETIRED_{it}^{husband} = \sum_{t=2001}^{2005} \beta_t \left(LGW_i \times YEAR_t \right) + \gamma_{LGW} LGW_i + \sum_{t=2000}^{2005} \gamma_t \, YEAR_t + \delta \; X_{it} + \varepsilon_{it}$

provides incentives for the same categories of employees to work until they reach the age when they are obliged to retire. In this paper, we have exploited this reform to study how a change in the wife's retirement incentive affects the retirement decision of the husband.

We make two major findings. First, there was a clear reduction in retirement rates among 63 year old female local government workers after the reform. Seen over the whole post-reform period 2001-2005, the decrease was 23 percentage points relative to private sector workers. Second, it was *not* the case that husbands married to local sector employees changed their retirement behavior in comparison to those who were married to wives who were private sector employees. The point estimate in the main specification is 0.004. The 99 percent confidence interval ranges from –0.014 to 0.022. Both key findings are robust to including a host of control variables and alternative sample restrictions.

The results of this study are consistent with the analysis of Blau (1998) who also found strong associations between the labor force transition probabilities of one spouse and the labor force status of the other spouse, but that the pattern could not be explained by financial incentives. Still, the results are at odds with some other earlier studies (e.g. Coile 2004 and Zweimüller et al 1996) which found significant cross-effects of the wife's incentives on the husband's behavior. While comparisons of studies conducted on data from different countries should be made with care, it should be noted that, in contrast to previous studies in the literature, we have exploited a transparent and plausibly exogenous source of variation in the wife's incentives. As far as we know, the present paper is the first quasi-experimental study that is able to isolate the cross-effect from the wife's retirement incentive.

Finally, when reflecting on the policy implications of this paper, it should be kept in mind that we have studied a pension reform that unexpectedly altered the benefit accrual of the wife at a late stage of her working career. The cross-effects might be more important if the family's planning horizon is the whole working career. Thus, the absence of a short-run effect of the wife's retirement incentives on the husband's retirement decision does not necessarily rule out longer-term effects.

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Appendix A. Descriptive statistics.

Table A1. Descriptive statistics. Regression sample.

	Local government sector		Private sector		Total	
	Mean	Std	Mean	Std	Mean	Std
Husband retired	0.83	0.38	0.80	0.40	0.81	0.39
Wife retired	0.31	0.46	0.18	0.39	0.25	0.43
Husband's age	66.04	4.17	65.77	4.22	65.91	4.20
Wife - education level 1	0.21	0.41	0.40	0.49	0.30	0.46
Wife - education level 2	0.06	0.23	0.12	0.32	0.09	0.28
Wife - education level 3	0.43	0.50	0.37	0.48	0.40	0.49
Wife - education level 4	0.01	0.09	0.01	0.12	0.01	0.11
Wife - education level 5	0.29	0.45	0.10	0.29	0.19	0.40
Wife - education level 6	0.00	0.05	0.00	0.04	0.00	0.05
Husband - education level 1	0.38	0.49	0.40	0.49	0.39	0.49
Husband - education level 2	0.04	0.20	0.05	0.22	0.05	0.21
Husband - education level 3	0.37	0.48	0.38	0.48	0.37	0.48
Husband - education level 4	0.02	0.15	0.02	0.15	0.02	0.15
Husband - education level 5	0.15	0.36	0.12	0.32	0.13	0.34
Husband - education level 6	0.01	0.09	0.01	0.08	0.01	0.09
Stockholm	0.13	0.34	0.17	0.38	0.15	0.36
Upppsala	0.03	0.18	0.02	0.16	0.03	0.17
Södermanland	0.03	0.18	0.03	0.18	0.03	0.18
Östergötland	0.05	0.21	0.05	0.22	0.05	0.21
Jönköping	0.05	0.21	0.05	0.21	0.05	0.21
Kronoberg	0.03	0.16	0.02	0.15	0.03	0.16
Kalmar	0.03	0.18	0.03	0.18	0.03	0.18
Gotland	0.01	0.08	0.01	0.08	0.01	0.08
Blekinge	0.02	0.15	0.02	0.14	0.02	0.14
Skåne län	0.11	0.32	0.14	0.35	0.13	0.34
Hallands län	0.04	0.20	0.04	0.20	0.04	0.20
Västra Götaland	0.18	0.39	0.17	0.38	0.18	0.38
Värmland	0.04	0.19	0.03	0.17	0.03	0.18
Örebro	0.03	0.18	0.03	0.17	0.03	0.18
Västmanland	0.03	0.18	0.03	0.18	0.03	0.18
Dalarna	0.03	0.18	0.03	0.17	0.03	0.18
Gävleborg	0.04	0.19	0.03	0.17	0.04	0.18
Västernorrland	0.03	0.18	0.03	0.16	0.03	0.17
Jämtland	0.01	0.12	0.01	0.11	0.01	0.11

Table A.1 continued						
Västerbotten	0.03	0.18	0.02	0.14	0.03	0.16
Norrbotten	0.03	0.16	0.02	0.15	0.02	0.15
Year 1999	0.12	0.33	0.13	0.33	0.12	0.33
Year 2000	0.12	0.33	0.14	0.34	0.13	0.33
Year 2001	0.13	0.33	0.13	0.34	0.13	0.34
Year 2002	0.14	0.35	0.13	0.34	0.14	0.34
Year 2003	0.15	0.36	0.13	0.34	0.14	0.35
Year 2004	0.16	0.36	0.16	0.37	0.16	0.36
Year 2005	0.18	0.38	0.18	0.38	0.18	0.38
Number of observations	33,320		35,034		68,354	
Note that 'the regression sample' both includes the 2000-2005 sample that is used in the main analysis and the 1999 sample that is used in the placebo test.						

Table A2. Mean retirement status and number of observations by the age of the husband.

Age of husband	Number of observations	Mean		Age of husband	Number of observations	Mean
41	2	0.00		70	2994	1.00
43	2	0.00		71	2147	1.00
45	5	0.00		72	1526	1.00
46	6	0.17		73	1141	1.00
47	16	0.00		74	824	1.00
48	14	0.14		75	632	1.00
49	25	0.08		76	439	1.00
50	37	0.08		77	324	1.00
51	46	0.09		78	251	1.00
52	82	0.09		79	194	1.00
53	84	0.12		80	145	1.00
54	157	0.09		81	91	0.99
55	224	0.14		82	61	1.00
56	337	0.18		83	41	1.00
57	417	0.20		84	34	1.00
58	662	0.23		85	27	1.00
59	972	0.22		86	5	1.00
60	1479	0.29		87	8	1.00
61	2479	0.38		88	11	1.00
62	3977	0.48		89	4	1.00
63	6190	0.57		90	4	1.00
64	7815	0.64		91	4	1.00
65	8520	0.99		92	2	1.00
66	7766	1.00				
67	6812	1.00				
68	5314	1.00				
69	4005	1.00				
Total				68354	0.81	68354

Note that 'the regression sample' both includes the 2000-2005 sample that is used in the main analysis and the 1999 sample that is used in the placebo test.

Appendix B

The total benefit level in the gross pension system

As described in *Section 4.2,* the pension plan for local government workers that was in place prior to 2000, PA-KL, stipulated that the sum of the public pension benefit (*PB*) and the occupational pension benefit (*OB*) should be equal to a certain fraction, γ, of the individual's qualifying income (*QI*). Thus, the OB was determined as a residual amount:

$$TB = \gamma \times QI = PB + OB\,,$$

where *TB* denotes the total annual pension benefit. The local government worker was eligible for full pension benefits given that he/she had reached the retirement age and had worked at least 30 years since the age of 28 in the sector. The *QI* was defined as an average of the five best earnings years out of the of the seven most recent earnings years. *Table B.1.* shows the replacement rate γ as a function of the qualifying income, *QI*.

Table B1. Formula for gross pension in the public sector

Qualifying income (QI)	Total gross replacement rate in the interval	Total gross replacement rate at the upper end of the interval, γ
0 – 1 BA	96 %	96 %
1 BA -2.5 BA	78.5 %	85.5 %
2.5 BA – 3.5 BA	60 %	78.2 %
3.5 BA – 7.5 BA	64 %	70.6 %
7.5 BA – 20 BA	65 %	67.1 %
20 BA – 30 BA	32.5 %	55.6 %
30 BA -	0	-
BA = price base amount. In 2010 BA is SEK 42,400 (approximately EUR 4,700).		

The level of the public pension benefit, *PB*, was determined in the following way

$$PB = \underbrace{0.6 \times APP \times \min(\frac{N_{PP}}{30}, 1)}_{Supplementary\ pension\ ('ATP')} + \underbrace{\omega \times BA}_{Basic\ pension}$$

where *APP* is the average pension points earned during the 15 best earnings years. At maximum, the individual could earn 6.5 pension points each year. 1 pension point amounted to 1 price base amount, *BA*. N_{PP} is the number of years that the individual has earned pension points. ω is a parameter that determines the level of basic pension. ω was 96 % for singles and 78.5 % for married individuals.[29]

Early and late withdrawals in the gross pension system

Early withdrawals before the retirement age could be made from the age of 60 under the PA-KL agreement. Early retirement was subject to a penalty. *Table B2* shows the monthly adjustment of the gross pension benefit (*TB*) in percent. Suppose that the employee, under the old agreement, was subject to the 63-65 rule, i.e. she was eligible for full pension benefits by the month she turned 63. If she withdrew pension benefits already from the month she turned 60, the gross pension benefit was reduced by

$$12 \times 0.5\% + 12 \times 0.45\% + 12 \times 0.35\% = 15.6\%.$$

Note, however, that the individual is unaffected by the early withdrawal penalty in the public pension system.

Table B2. Adjustments for early/late withdrawals (in % of gross pension) under PA-KL.

Age	60	61	62	63	64	65	66
Retirement age 65	-0.5	-0.5	-0.4	-0.3	-0.3	0.1	0.1
Retirement age 63-65	-0.5	-0.45	-0.35	0	0	0.1	0.1

[29] See Palme and Svensson (2004) for a more detailed description of the old Swedish pension system.

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