

What Happens to the Husband's Retirement  
Decision when the Wife's Retirement  
Incentives Change?

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# What Happens to the Husband's Retirement Decision when the Wife's Retirement Incentives Change?

## 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, I 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. I 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, I do not find any evidence of a response among husbands. This finding is at odds with most earlier results in the literature.

JEL-Code: H550, J130, J210.

Keywords: joint retirement, retirement age, occupational pensions.

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## 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).<sup>1</sup> 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.

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<sup>1</sup> 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 couples where the wife was 63 years old and where both spouses 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 20 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 0.007. The 99 percent confidence interval ranges from -0.039 to 0.052.

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 my 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.<sup>2</sup> Therefore, it is interesting that I obtain a different result as compared to earlier studies. The main lesson is that an increase in the wife's benefit accrual rate does not necessarily translate into a change in the husband's propensity to retire.

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.<sup>3</sup> 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 my 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

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<sup>2</sup> It should be noted that the retirement incentives did in fact also change for the relevant cohorts of males since the public pension system was reformed during this period. These changes affected treatment and control groups in a similar way though.

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

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 females. It should be noted that Baker's (2002) paper has a different focus: Here, I study a reform affecting the retirement incentives of wives while keeping the husbands' own retirement incentive variables at a constant level. Thus, I 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.<sup>4</sup> 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

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<sup>4</sup> The same interpretation is made by Liebman et al (2009).

similar to those reported by Coile (2004): Husbands react to the wives' incentives, but not vice versa.

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 Mastrogiamomo (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.<sup>5</sup> 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.

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<sup>5</sup> 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

I 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

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

where  $w_s$  is cash compensation and  $E(W_s | \cdot)$  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:

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

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, I 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  $\hat{w} = w - b + m$ .<sup>6</sup> 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, I 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.<sup>7</sup>

### 3.2 *The interaction between spouses*

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

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<sup>6</sup> 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  $\frac{dC}{dh} = w - b(h, s) + \int_h^T \frac{\partial b(h, s)}{\partial h} ds$ .

<sup>7</sup> 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.<sup>8</sup> The wife maximizes  $u_f(C_f, h_f)$  subject to the life time budget constraint  $C_f = w_f h_f + \int_{h_f}^T b(h_f, s) ds$ .<sup>9</sup> While still using the notation introduced in the previous section, I 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

$$\begin{aligned} & \max_{C, h_m} u_m(C, h_m, h_f) \\ & \text{subject to } C = w_m h_m + \int_{h_m}^T b(h_m, s) ds + w_f h_f + \int_{h_f}^T b(h_f, s) ds, \end{aligned} \quad (1)$$

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, I 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*, which arises since the wife's leisure time enters the husband's utility function, 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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<sup>8</sup> 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'.

<sup>9</sup> Throughout this discussion, I neglect the possibility that the benefit formula can be dependent on the retirement behavior of the other spouse. In the Swedish system, there are few such interactions.

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’).<sup>10</sup> In this framework, spouses simultaneously solve the following constrained optimization problem

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

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

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, I 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<sup>11</sup>

### 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

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<sup>10</sup> See Vermeulen (2002) for a good survey of collective household models.

<sup>11</sup> 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.<sup>12</sup> 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.<sup>13</sup> If this employee was eligible for annual benefits from the public pension system of 60 % of her qualifying income the annual occupational pension benefit amounted to 13 % of her qualifying income; the occupational benefit was the residual amount. Hence, the early withdrawal penalties in the public pension system were irrelevant for those covered by the old

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<sup>12</sup> 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).

<sup>13</sup> 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.

collective agreement for local government workers. Rather, from the employee's perspective it was sufficient to know about the gross replacement rate.

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 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, i.e. a reduction in the gross replacement rate (see *Appendix B*). Likewise, a small adjustment was made for every month the employee postponed retirement after his/her 65<sup>th</sup> birthday.

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 Municipal 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.

Under simplified assumptions, column (1) of *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.<sup>14</sup> Most striking is the sharp fall in the benefit accrual rate at the age of 63. At this point, the individual did not earn any incremental pension rights. At this age the individual loses SEK 175,000 in pension wealth from continuing working an additional year.

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<sup>14</sup> However, for very low values of the discount rate the benefit accrual at age 62 is positive – when the discount rate is 0 the benefit accrual is SEK 8,994. In Table 1 the discount rate is set to 3%.

Table 1. Benefit accrual rates (expressed in SEK) for local government workers born in 1937 (and subject to the 63-65 rule) and in 1942

	Old plan: The 1937 cohort	New plan: The 1942 cohort
Age	(1)	(2)
61	6,341	35,955
62	-43,083	27,628
63	-174,633	18,486

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  $s-1$ . The discount rate is set to 3%. Benefit accruals are expressed in SEK in the price level of 2012. 1 EUR = 9 SEK. I assume that the individual lives until she is 86. I also assume that the individual has earned a yearly income of 5.6 price base amounts (SEK 246,400 in 2012) since the age of 25.

In 1998 a new agreement, PFA98, was signed for Swedish local government employees. The media coverage was modest.<sup>15</sup> The new PFA98 agreement came into effect on January 1, 2000 for those born in 1938 and later. The first cohort of 63 year olds to earn pension benefits from the new DC system turned 63 in 2001. Those born in 1937 and earlier were completely 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.<sup>16</sup> Accordingly, under the new system, the individual could always increase her pension wealth by postponing retirement until the age when she was obliged to retire.<sup>17</sup> Crucially, owing to the abolishment of the gross pension system the local government worker became subject to the incentives created by the public pension system. Local government workers born in 1938 and onwards were

<sup>15</sup> A search in the virtual press archive ‘Presstext’, which covers the biggest daily newspapers in Sweden, reveals that the first article mentioning ‘PFA98’ is written in the fall of year 2000.

<sup>16</sup> For wages above the cap of 7.5 increased price base amounts, a DB system was retained. 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 %).

<sup>17</sup> 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.

exposed to the early withdrawal penalties in the public pension system. The second column of *Table 1* shows the benefit accrual rates for local government workers born in 1942. In contrast to the accrual rates of column (1), one needs to take the public pension system into account when computing the figures of column (2). We see that post-reform accrual rates are positive at all ages. In particular, there is an enormous change in retirement incentives at the age of 63, even though accrual rates also increase for 61 and 62 year olds.

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 e.g. Sundén (2006).<sup>18</sup> This nation-wide reform affected all individuals born after 1937 included in the empirical analysis. 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.

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<sup>18</sup> 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.

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 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.

Finally, occupational pension recipients were subject to an earnings test. The main rule in the old 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.

#### *4.3 The private sector occupational pension plans*

There are two large occupational pension systems in the private sector: one for blue-collar workers and one for white-collar workers. None of these were reformed during the period of study, even though both agreements have changed after this period (primarily for younger cohorts not included in the present analysis). The pension scheme for blue-collar workers ('the SAF-LO plan') was a pure DC scheme, where the employer paid contributions amounted to 3.5% of the employee's wage. The scheme for white-collar workers ('the ITP scheme'), on the other hand, was mainly a DB system with a smaller DC component ('ITPK').

As will be discussed below in *Section 6.2*, I will define the retirement variable based on pension income receipt. To what extent is it possible to work and at the same time withdraw occupational pension benefits? As pointed out above, in the local government sector there was an earnings test preventing workers from doing this. White-collar private workers aged below 65 are not allowed to withdraw pension benefits if they work more than 8 hours per week. There are no such restrictions for blue-collar private workers. However, once a blue-collar worker starts to receive DC benefits, he cannot accumulate any new DC wealth. Under both these schemes individuals are allowed to withdraw benefits from the age of 55.



## 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). All husbands and wives included in the analysis belonged to the labor force the preceding year, i.e. they received no pension income and reported positive earnings. The 'treatment group' consists of husbands married to female *local government sector* workers who were 63 years of age and who were exposed to the new DC plan. The 'control group' consists of men married to 63 year old female *private sector* workers who were not exposed to the occupational pension reform in the local government sector. In the next section, I 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 model 2000 is the pre-reform year and 2001-2005 are the post-reform years.

The relevant difference-in-difference regression equation reads

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

for all  $i$  in couples where both spouses belonged to 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 retires year  $t$ .  $LGW_i$  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.  $\varepsilon_{it}$ , finally, is an error term. Equation (2) will be estimated in the form of a linear probability model. The key estimated interaction terms are similar if a probit model is used and interaction effects are calculated in the way proposed by Ai and Norton (2003).

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, I 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*, I 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.<sup>19</sup> Crucially, since there is an identifier for the household, I 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 already mentioned, the population of interest in this study is men married to women aged 63. Since lagged information is used to select the relevant samples, I can use the years 1999 to 2005 in the analysis. In total, there are 193,391 couples in the data 1999-2005 where the wife is aged 63. Crucially, the population is conditioned on that both spouses were part of 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.<sup>20</sup> Since the majority of 63 year old women are married to older men who receive pension income, this restriction

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<sup>19</sup> The data set was ordered from Statistics Sweden for the purpose of studying issues related to retirement savings.

<sup>20</sup> Pension income is defined as the sum of the two variables 'tpensa' and 'tprivl'. Labor income is defined as the sum of 'tjlon', 'nakte' and 'nakthb'.

reduces the sample size considerably to 26,131 couples. Some additional exclusions were made. First, I excluded husbands who were employed in the local government sector, as these husbands might be subject to the occupational pension reform. Second, I 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. These exclusions further reduced the sample to 17,626 couples. After having excluded a small number of observations with missing values on any of the key variables, the final regression sample contains 17,431 observations.<sup>21</sup>

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

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<sup>21</sup> Henceforth, unless otherwise stated, we let the term 'regression sample' denote the sample that contains the 15,518 observations for the years 2000-2005 (to be used in the main analysis) *and* the 1,913 observations for the year 1999 (to be used in the placebo analysis).

only withdrew a smaller amount of pension income. However, as described in *Section 4.3*, the different collective agreements contained rules that prevented workers and/or made it unattractive for workers to work and receive pension benefits at the same time. Moreover, individuals who received a disability pension can also be subsumed under this retirement concept prior to 2003.<sup>22</sup> As a robustness check an earnings based retirement definition will also be used (*Table 7*).

The dummy variable for being a local government employee is defined based on information from the preceding year. I also divide the studied population into six educational categories and define dummies for the educational level based on these. Finally, I 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

*Figure 1* depicts the distribution of husbands' ages. *Figure 1A* shows this distribution in the estimation sample. Remember that the estimation sample has been selected based on the criterion that both spouses were part of the labor force in the preceding year. Since there is a dramatic increase in pension benefit withdrawals among 65 year olds, there is a huge hole in the distribution starting at age 66. Therefore, it can be interesting to look at a sample where the restriction that the husband must be in the labor force in year  $t-1$  is dropped, but all other selection criteria are the same. In this sample (shown in *Figure 1B*), the mean age difference between husbands and wives is 2.9 years. This is in line with the well-known empirical regularity that husbands typically are older than their wives. In the estimation sample, on the other hand, the mean age difference is -0.89. This should be kept in mind when interpreting the estimated model parameters.

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<sup>22</sup> 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.

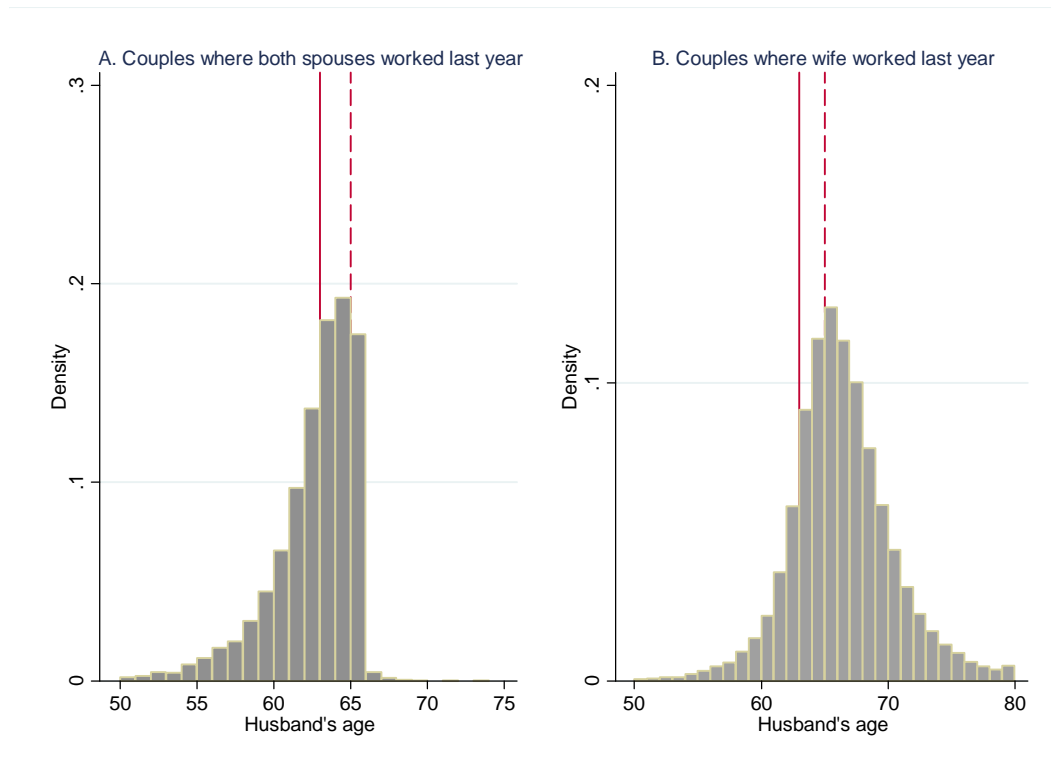


Figure 1. The age distribution of husbands married to wives aged 63 in the estimation sample. The left (solid) line depicts the wife's age (63) and the right (dashed) line depicts the normal retirement age (65).

*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 0.44. However, in 2001, the first post-reform year for 63 year olds, it had dropped to 0.31. 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 conditional 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 surprisingly well. In 2002 there is a tiny relative decrease for men married to women employed in the private sector, but otherwise it is striking how close the lines are. One may notice that the level of the conditional retirement probability on average is higher for husbands than for wives. This can be explained by the fact that around 17 percent of

the husbands are 65 years of age – 65 year olds have a retirement probability that is close to 1 (see Table A.2.).<sup>23</sup>

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<sup>23</sup> If one removes those who are aged over 64 the lines are still very parallel, even though the level of the conditional retirement probability is substantially lower. Moreover, the regression results reported in Section 7 are essentially unaffected.

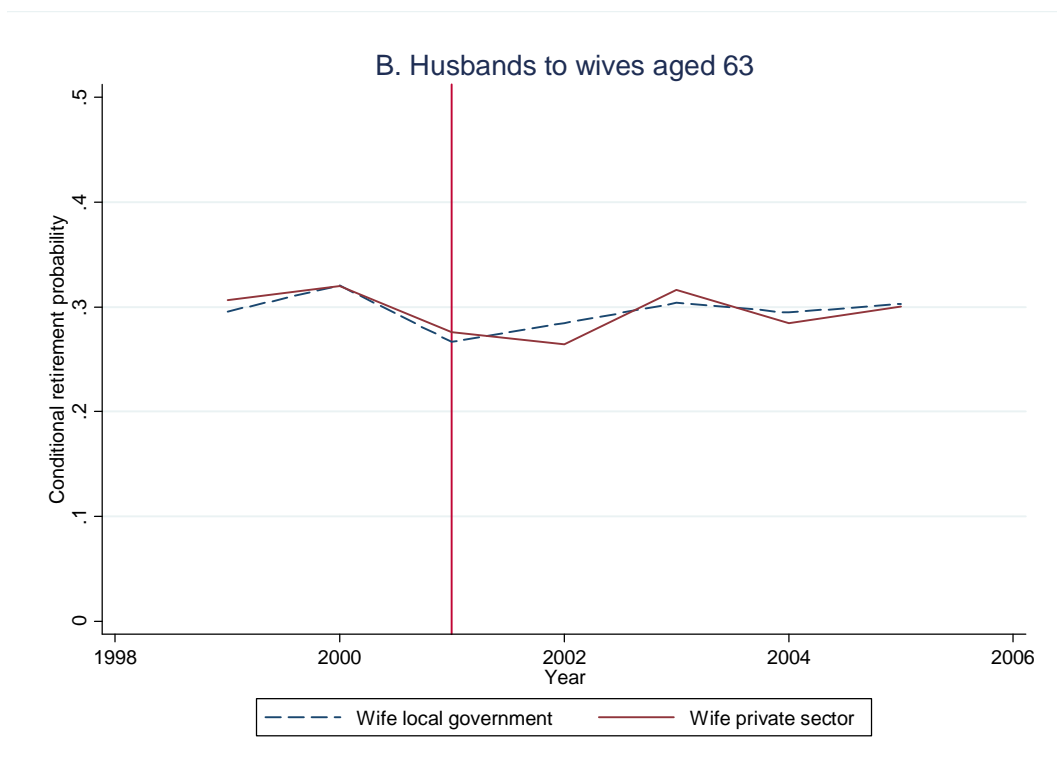
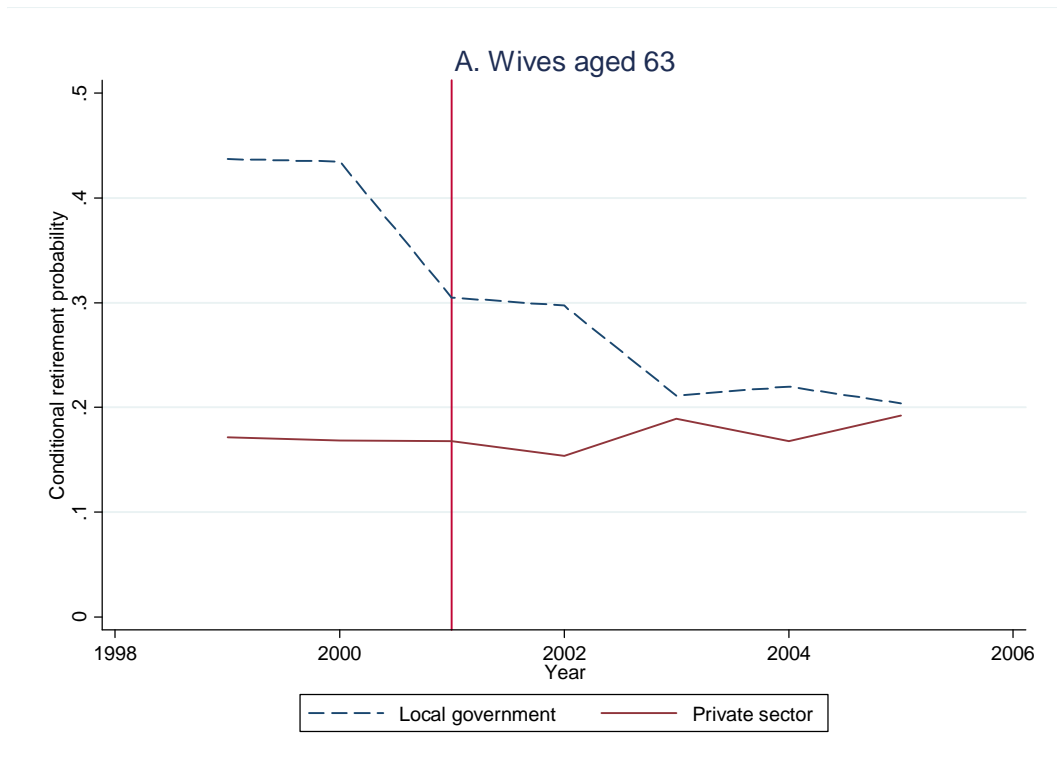


Figure 2. Conditional mean retirement probabilities for 63 year old wives and their husbands.

For the difference-in-difference approach to be valid, the group-specific fixed effect should be the same across years –  $\gamma_{LGW}$  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.<sup>24</sup> 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. Indeed, *Table 1* showed that benefit accrual rates increased also for 62 year olds, even though the magnitude of the increase was substantially larger for 63 year old local government workers. *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 see that the level is similar in 1999 and 2000 for 62 year old female local government workers, whereas there was small increase for female private sector workers 1999-2000. Accordingly, there is no graphical evidence of a clear response of 62 year old local government workers to the occupational pension reform. For other years, the lines in *Figure 3.A* are parallel, with 2004 and 2005 as exceptions. *Figure 3.B* reveals that the husbands' retirement probabilities in the two groups evolve in pretty much the same way.

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<sup>24</sup> 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.



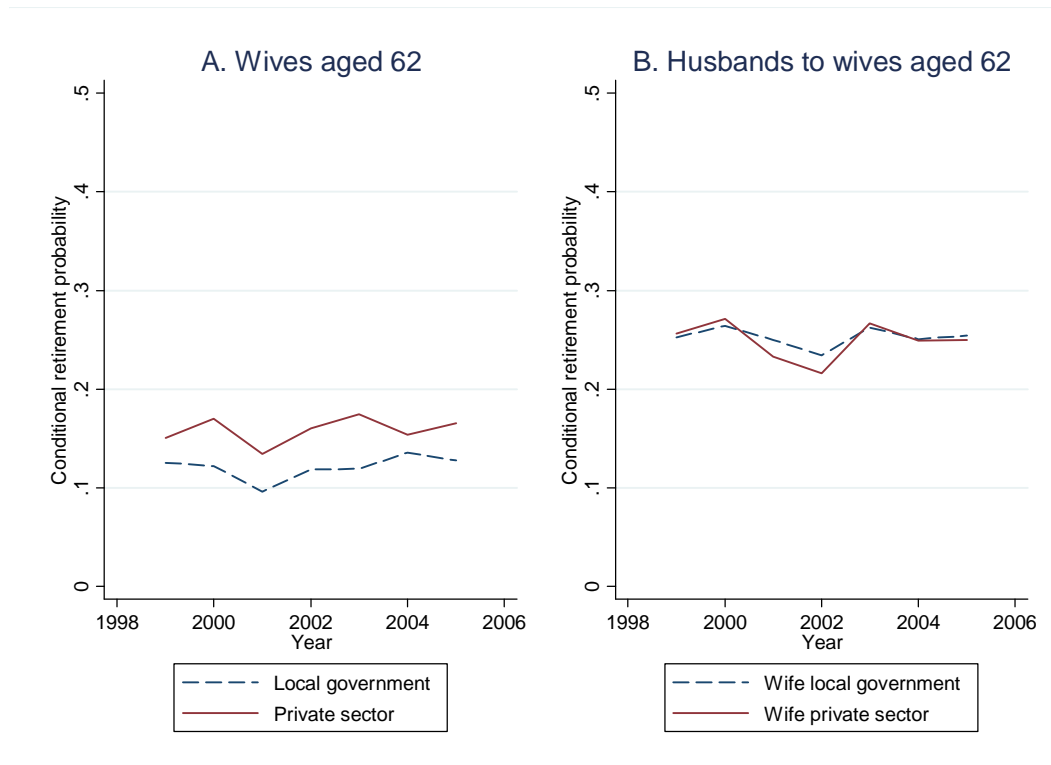


Figure 3. Conditional mean retirement probabilities for 62 year old wives and of husbands married to 62 year old wives, 1999-2005.

### 6.3 The correlation between the two spouses' retirement decisions

Before turning to the regression analysis it can be fruitful to examine the prevalence of joint retirement in Sweden. More specifically, we look at the correlation between the two spouses' retirement probability, conditional on that the two spouses were in the labor force year  $t-1$ . *Table 2* illustrates the correlation coefficient between the wife's retirement and the husband's retirement in the data 1999-2005 for the ages 60-65. In general, the retirement decision of spouses of age  $s$  are most strongly correlated with the other spouse's retirement decision if the other spouse is also aged  $s$ .<sup>25</sup> The grey shaded column of *Table 2* is of special interest as it reflects a large subset of the estimation sample. In couples where both spouses are aged 63, the correlation coefficient is 0.19.

<sup>25</sup> The only exception is that 62 year old women's retirement decisions are more correlated with 64 year old men's retirement decisions than 62 year old husbands' retirement decisions.

Table 2. Correlation coefficients for the husbands' and wives' retirement decisions

		Wife's age					
		60	61	62	63	64	65
Husband's	60	0.1785	0.1005	0.0853	0.067	0.0458	0.0786
age		(8528)	(4516)	(2294)	(1141)	(554)	(273)
	61	0.0882	0.2225	0.14	0.0804	0.0383	0.0972
		(9481)	(6477)	(3348)	(1686)	(798)	(377)
	62	0.0964	0.1752	0.1867	0.1394	0.1963	0.0218
		(8499)	(6966)	(4705)	(2383)	(1102)	(558)
	63	0.0798	0.1228	0.1693	0.19	0.1817	0.004
		(6375)	(6161)	(5002)	(3157)	(1498)	(724)
	64	0.092	0.1111	0.1163	0.1494	0.2177	0.0499
		(4292)	(4620)	(4318)	(3350)	(1914)	(1028)
	65	0.0399	0.0246	0.0269	0.0384	0.0229	0.1586
		(2781)	(3234)	(3404)	(3030)	(2122)	(1339)

The table reports correlation coefficients for the husband's retirement and the wife's retirement, 1999-2005. The number of observations is in parenthesis. The sample is selected in the same way as the estimation sample, i.e. both spouses should be in the labor force year t-1, none of the spouses should be employed in the central government sector and the husband should not be employed in the local government sector. The grey shaded column reflects a subset of the estimation sample.

## 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 3* 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 (2), the homogenous

treatment effect is  $ATET = \frac{\sum_{t=2001}^{2005} \beta_t N_t}{N_{PR}}$ , 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 3*, the estimate of the ATET is  $-0.202$ . The heteroskedasticity robust standard error is rather small – the 99 percent confidence interval ranges from  $-0.26$  to  $-0.15$ . 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 20 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 estimate of column (2) is identical to the estimate reported in column (1). The robust standard error is, however, somewhat larger when the control variables are added.

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. These results are, of course, fully consistent with *Figure 2.A*. 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. Once more, the estimated treatment effects are fairly similar in the regressions with and without control variables. A  $\chi^2$  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 0.82.

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.

The response in retirement entries should, of course, be viewed in relation to the magnitude of the incentive change. I have made back-of-the-envelope calculations of the elasticity of the retirement decision with respect to the one minus the implicit tax rate, where I use information from *Table 3*, column (4), and *Table 1*.<sup>26</sup> When evaluated at an earnings level corresponding to the average wage in the local government sector, the retirement elasticity with respect to one minus the implicit tax rate is -0.22. Thus, even though the treatment effect is estimated to be -0.25 the implied retirement elasticity is not terribly large. The reason is that the change in benefit accrual rates for 63 year olds between 2000 and 2005 is huge. Since individual pension incentives cannot be calculated for each individual in the sample this elasticity figure should, however, be seen as a rough approximation of the ‘true’ elasticity.

<sup>26</sup> The elasticity is defined as

$$\eta_{retirement} = \frac{\beta_{2005}}{RETIRED_{2000}^{LGW\_WIFE}} \bigg/ \frac{\Delta(1 - \tau^{LGW\_WIFE}) - \Delta(1 - \tau^{PS\_WIFE})}{(1 - \tau_{2000}^{LGW\_WIFE})}, \quad \text{where}$$

$RETIRED_{2000}^{LGW\_WIFE}$  is the retirement probability of local government wives in year 2000 and  $\tau$  is the implicit tax caused by the pension system. The implicit tax rate at age 63 is defined as minus the benefit accrual divided by annual earnings, i.e.  $\tau = -\frac{\{E(W_{63} | r = s + 1) - E(W_{63} | r = s)\}}{wh}$ .  $\tau^{LGW\_WIFE}$  is

the implicit tax rate for a local government worker under the assumptions in *Table 1*.  $\tau^{PS\_WIFE}$  is the implicit tax rate for a blue-collar private sector worker, where all other assumptions are the same as in *Table 1*. The change in the implicit tax between 2000 and 2005 in the latter group is small.

Table 3. The wife's retirement status.

	(1)	(2)	(3)	(4)
Treatment	-0.202	-0.202		
2001-2005	(0.021)***	(0.023)***		
Treatment 2001			-0.129 (0.027)***	-0.127 (0.029)***
Treatment 2002			-0.123 (0.026)***	-0.128 (0.028)***
Treatment 2003			-0.245 (0.026)***	-0.240 (0.027)***
Treatment 2004			-0.214 (0.025)***	-0.218 (0.026)***
Treatment 2005			-0.256 (0.024)***	-0.254 (0.026)***
Control variables	No	Yes	No	Yes
<p>The regression sample consists of couples where the wife is 63 years old and where both spouses 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 15,518. 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.</p>				

### 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 4*.

Again, 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 -0.001 and the heteroskedasticity robust standard error is approximately of the same magnitude as in the regressions for females, 0.022. When covariates are added (column 2), the point estimate of the treatment effect changes sign and becomes positive, but is still very small, 0.007. The standard error of the point estimate is now reduced as the

inclusion of covariates decreases the variance in the estimated residuals. The 99 percent confidence interval ranges from -0.039 to 0.052. 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. None of these estimates are statistically different from 0. I therefore conclude that the husbands of those wives who were affected by the occupational pension reform in the local government sector did not change their retirement behavior due to reform, despite the large response of the wives. This finding is surprising given the well-documented correlation in the timing of spouses' retirement decisions. The results are also at odds with the results in e.g. Zweimüller et al (1996) and Coile (2004) who both find that husbands are sensitive to changes in their wives' incentives.

### 7.3 Placebo tests

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, I will exploit the fact that our data set allows us to define the dependent variable also for 1999, an earlier pre-reform year.<sup>27</sup> 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:

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

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 (3) should not be significantly distinct from 0. From *Table 5* we can infer that it is indeed the case that the coefficients for the

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<sup>27</sup> 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.

interaction term are close to zero (and very insignificant). Thus, consistent with the graphical evidence provided in *Figure 2.A*, the placebo estimates are close to zero both for wives and husbands.

Table 4. The husband's retirement status.

	(1)	(2)	(3)	(4)
Treatment	-0.001	0.007		
2001-2005	(0.022)	(0.018)		
Treatment 2001			-0.011	0.011
			(0.029)	(0.021)
Treatment 2002			0.018	0.012
			(0.028)	(0.021)
Treatment 2003			-0.015	0.003
			(0.028)	(0.022)
Treatment 2004			0.006	-0.000
			(0.027)	(0.021)
Treatment 2005			-0.002	0.009
			(0.026)	(0.021)
Control variables	No	Yes	No	Yes
<p>The regression sample consists of couples where the wife is 63 years old and where both spouses 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 15,518. 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.</p>				

Table 5. Placebo regressions

	Wives		Husbands	
	(1)	(2)	(3)	(4)
Treatment 2000	0.003	0.015	0.020	0.005
	(0.029)	(0.030)	(0.030)	(0.023)
Control variables	No	Yes	No	Yes
<p>The regression sample consists of couples where the wife is 63 years old and where both spouses were employed in the previous year, 1999-2000. Linear probability model. The regression equation follows equation (3) in the main text. Robust standard errors are in parenthesis. * denotes significance at 10%, ** at 5% and *** at 1%. The number of observations is 3,923. 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.</p>				

#### 7.4 Further robustness checks

One potential concern of the analysis is related to the dynamics of the retirement response. There are, indeed, two problems at hand. First, in *Section 4.2* we noted that retirement incentives for local government workers born in 1938 changed already when they were 62 years old. Thus, even though the graphical analysis did not indicate any substantial response of the 1938 cohort at the age of 62, it is desirable to examine this further in a regression analysis.

Second, we cannot *a priori* rule out that husbands reacted to the change in the wives' incentives by retiring before the wife retires. In order to examine this hypothesis I construct a new regression sample that includes all couples where both partners were in the labor force in period  $t-2$ . I then define a binary outcome variable that takes the value of 1 if the individual either retires in period  $t$  or in  $t-1$  (assuming that retirement is an absorbing state). With these modifications I run the same regressions defined by equation (2) above. I now allow for a response, both of husbands and wives, the year before the wife turns 63. The number of observations increases from 15,518 to 21,793. It is now impossible to perform placebo tests as 1998 is the first year of data and the sample is selected based on labor force status in year  $t-2$ .

*Table 6*, column (1), shows that the common treatment dummy, which now is estimated to be -0.208, is very close the treatment effect estimate in the main model,



where the corresponding point estimate was -0.202. The year-specific treatment effect estimates of column (2) are also close to those of column (4) of *Table 3*. The same holds true for the parameters reflecting the husbands' behavior. The treatment effect for 2001-2005 reported in column (3) is estimated to be 0.013 instead of 0.007 in the baseline model (column 2, *Table 4*).

Note that this exercise also serves as a robustness check of the way in which I select the regression sample. Until now I have defined the sample based on labor force status in year  $t-1$ . When running regressions with the standard binary outcome variable (which is 1 if the individual retires in period  $t$ ) on the larger sample defined based on information in  $t-2$  I obtain results that are similar (not reported).

Table 6. The probability to retire in year  $t$  or in year  $t-1$

	Wives		Husbands	
	(1)	(2)	(3)	(4)
Treatment 2001-2005	-0.208 (0.020)***		0.013 (0.016)	
Treatment 2001		-0.136 (0.026)***		0.022 (0.020)
Treatment 2002		-0.109 (0.026)***		0.013 (0.020)
Treatment 2003		-0.262 (0.025)***		0.029 (0.020)
Treatment 2004		-0.257 (0.025)***		-0.005 (0.020)
Treatment 2005		-0.237 (0.024)***		0.010 (0.019)
Control variables	Yes	Yes	Yes	Yes

The regression sample consists of couples where the wife is 63 years old and where both spouses were employed in year  $t-2$ , 2000-2005. The dependent variable is an indicator variable that takes on the value of 1 if the individual retired in year  $t$  or  $t-1$ . Linear probability model. Robust standard errors are in parenthesis. \* denotes significance at 10%, \*\* at 5% and \*\*\* at 1%. The number of observations is 21,793. 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.

Finally, one might wonder what happens if one instead uses an earnings based outcome variable. In the analysis above we defined an individual as being retired if he/she received a positive amount of pension income in year  $t$ . Since earnings will be positive the year an individual retires an earnings based definition necessarily needs to be based on earnings information as of year  $t+1$ . One candidate is to impose an arbitrary earnings limit for  $t+1$ . In what follows, I let the outcome variable take on the value of 1 in period  $t$  if the individual had earnings below 1.75 price base amount in  $t+1$ .<sup>28</sup> The average male retirement rate in the sample now slightly falls from 0.29 to 0.25.

To achieve comparability with the baseline results I will define the regression sample in the same way as before: those who had positive earnings and received zero pension income in year  $t-1$  are considered as being in the labor force in year  $t-1$  and are, hence, included in sample. However, since the dependent variable now builds on information from period  $t+1$  I will not be able to use data for 2005, the most recent year in the data set. As can be seen from *Table 7* the results are roughly similar to those in the main analysis.

If one instead sets the earnings cut-off limit to 0, which is considerably less arbitrary, there is a dramatic drop in retirement rates in the sample. The average male retirement then falls from 0.29 to 0.12. Accordingly, this retirement definition is very restrictive; it assigns some of those who we would consider as being genuine retirees to the working population. The regression results, which can be provided upon request, are now *qualitatively* similar to those in the main analysis. The female response is, however, approximately half as large as in the main analysis.

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<sup>28</sup> This cut-off limit was used by Glans (2009). According to this definition, the individual with a median wage needs to work approximately a quarter of a year to be included in the sample. A 'price base amount' is a Swedish administrative concept primarily used in the social insurance system. In 2012 a price base amount approximately amounts to EUR 5,000.

Table 7. The probability to retire: earnings based retirement definition

	Wives			Husbands		
	(1)	(2)	(3)	(4)	(5)	(6)
Treatment 2001-2004	-0.191 (0.024)***			-0.012 (0.019)		
Treatment 2001		-0.133 (0.031)***			0.006 (0.024)	
Treatment 2002		-0.132 (0.030)***			-0.032 (0.023)	
Treatment 2003		-0.239 (0.028)***			0.001 (0.023)	
Treatment 2004		-0.238 (0.027)***			-0.018 (0.022)	
Placebo treatment 2000			0.004 (0.032)			0.029 (0.025)
Control s	Yes	Yes	Yes	Yes	Yes	Yes

The regression sample consists of couples where the wife is 63 years old and where both spouses were employed in the previous year, 2000-2005. Linear probability model. In columns 1, 2, 4 and 5 the regression equation follows equation (2) in the main text. In columns 3 and 6 the regression equation follows equation (3) in the main text. Robust standard errors are in parenthesis. \* denotes significance at 10%, \*\* at 5% and \*\*\* at 1%. The number of observations is 12,068 (3,923) in columns 1,2,4 and 5 (columns 3 and 6). 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.

## 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 provides incentives for the same categories of employees to work until they reach the age when they are obliged to retire. In this paper, I have exploited this reform to study

how a change in the wife's retirement incentive affects the retirement decision of the husband.

I 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 20 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 0.007; the 99 percent confidence interval ranges from -0.039 to 0.052. Both key findings are robust to alternative specifications.

The main lesson from this paper is that an increase in the wife's benefit accrual rate does not necessarily translate into a change in the husband's propensity to retire. The results are broadly 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, I have exploited a transparent and plausibly exogenous source of variation in the wife's incentives. As far as I know, the present paper is the first quasi-experimental study that is able to isolate the cross-effect from the wife's retirement incentive. This study highlights the importance of separating the correlation between spouses' retirement decisions from the causal effect of one spouse's incentives on the other spouse.

Finally, when reflecting on the policy implications of this paper, it should be kept in mind that I 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.

Table A1. Descriptive statistics. Regression sample.

	Local government sector		Private sector		Total	
	Mean	Std	Mean	Std	Mean	Std
Husband retired	0.30	0.46	0.29	0.46	0.29	0.46
Wife retired	0.28	0.45	0.17	0.38	0.23	0.42
Husband's age	62.12	2.80	62.10	2.91	62.11	2.86
Wife - education level 1	0.19	0.40	0.39	0.49	0.29	0.46
Wife - education level 2	0.05	0.22	0.11	0.32	0.08	0.28
Wife - education level 3	0.43	0.50	0.37	0.48	0.40	0.49
Wife - education level 4	0.01	0.10	0.01	0.12	0.01	0.11
Wife - education level 5	0.31	0.46	0.10	0.31	0.20	0.40
Wife - education level 6	0.00	0.05	0.00	0.05	0.00	0.05
Husband - education level 1	0.33	0.47	0.36	0.48	0.35	0.48
Husband - education level 2	0.06	0.23	0.06	0.24	0.06	0.24
Husband - education level 3	0.43	0.50	0.42	0.49	0.43	0.49
Husband - education level 4	0.03	0.17	0.03	0.16	0.03	0.17
Husband - education level 5	0.13	0.34	0.12	0.32	0.12	0.33
Husband - education level 6	0.01	0.10	0.01	0.09	0.01	0.10
Stockholm	0.14	0.35	0.19	0.39	0.17	0.37
Uppsala	0.03	0.17	0.02	0.15	0.03	0.16
Södermanland	0.04	0.19	0.03	0.18	0.03	0.18
Östergötland	0.05	0.21	0.05	0.22	0.05	0.21
Jönköping	0.05	0.22	0.05	0.22	0.05	0.22
Kronoberg	0.02	0.15	0.02	0.15	0.02	0.15
Kalmar	0.03	0.18	0.03	0.17	0.03	0.18
Gotland	0.01	0.08	0.00	0.07	0.01	0.07
Blekinge	0.02	0.14	0.02	0.13	0.02	0.14
Skåne län	0.13	0.34	0.15	0.36	0.14	0.35
Hallands län	0.04	0.19	0.04	0.20	0.04	0.20
Västra Götaland	0.19	0.39	0.17	0.38	0.18	0.39
Värmland	0.03	0.16	0.02	0.15	0.03	0.16
Örebro	0.03	0.18	0.03	0.17	0.03	0.17
Västmanland	0.04	0.19	0.04	0.19	0.04	0.19
Dalarna	0.03	0.17	0.03	0.17	0.03	0.17
Gävleborg	0.04	0.19	0.03	0.17	0.03	0.18
Västernorrland	0.03	0.17	0.02	0.16	0.03	0.16
Jämtland	0.01	0.09	0.01	0.09	0.01	0.09

Table A.1 continued						
Västerbotten	0.03	0.16	0.02	0.13	0.02	0.14
Norrbottn	0.02	0.12	0.01	0.11	0.01	0.12
Year 1999	0.10	0.31	0.11	0.32	0.11	0.31
Year 2000	0.11	0.31	0.12	0.33	0.12	0.32
Year 2001	0.12	0.32	0.12	0.33	0.12	0.32
Year 2002	0.13	0.34	0.14	0.34	0.13	0.34
Year 2003	0.16	0.36	0.14	0.34	0.15	0.35
Year 2004	0.18	0.38	0.17	0.38	0.18	0.38
Year 2005	0.20	0.40	0.20	0.40	0.20	0.40
Number of observations	8,302		9,129		17,431	
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
41	2	0.00
43	2	0.00
45	5	0.00
46	5	0.00
47	15	0.00
48	11	0.00
49	23	0.00
50	33	0.03
51	43	0.02
52	75	0.01
53	71	0.00
54	143	0.02
55	197	0.05
56	287	0.05
57	344	0.05
58	525	0.06
59	784	0.07
60	1141	0.10
61	1686	0.13
62	2383	0.16
63	3157	0.20
64	3350	0.20
65	3030	0.97
66	78	0.63
67	24	0.50
68	7	0.29
69	6	0.50
71	2	0.00
73	1	1.00
74	1	0.00
Total	17431	0.29

## 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,  $\gamma$ , 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  $\gamma$  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, $\gamma$
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

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

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.  $\omega$  is a parameter that determines the level of basic pension.  $\omega$  was 96 % for singles and 78.5 % for married individuals.<sup>29</sup>

*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

<sup>29</sup> See Palme and Svensson (2004) for a more detailed description of the old Swedish pension system.