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Employer pay policies, public transfers and the retirement decisions of men and women in Denmark

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Abstract

The empirical retirement literature measures individual responses to variations in income flows due to public transfers, private individual or employer-provided pensions. We estimate a model accounting for the incentive effects from these sources. A dynamic structural model is extended to allow both individual and employer heterogeneity. This is applied to a Danish matched panel of workers and establishments, spanning a period of reforms to a public early retirement programme. Employer-specific compensation is found to be an important determinant of work and retirement income flows. Employer effects on retirement age are only found among sub-samples where access to public transfers is limited.

JEL classification: J26; C61; C23; C25

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1. Introduction

An ageing population, longer life expectancy and earlier exit from the labour force motivate interest in the determinants of retirement. Private pensions and public transfers are of increasing importance since they finance the future consumption of a growing number of retirees for a greater proportion of their lifetime. Policy interest in early retirement is almost always motivated by the large consequences for public financing of social security (Lumsdaine and Mitchell, 1999). It is less often remarked that early

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retirement is important to employers because of the human capital endowment that older workers represent.

Pensions create incentives which have consequences for retirement behaviour. The extensive retirement literature has focused largely on the incentive effects of public transfer programmes (Gruber and Wise, 1999), somewhat less on employer-provided pensions (Gustman and Steinmeier, 1998) and more recently on personal pensions (Mitchell, 1999). The novelty of this paper is the coherent modelling of retirement incentives due to each source of pension income: personal-, employer- and publicly provided. The emphasis is on the compensation policy of different employers as a retirement determinant, while coherently modelling public transfer incentive effects and allowing individual income and retirement heterogeneity.

An employer may wish to determine the retirement age of workers for several reasons. It would allow the writing of efficient long-term labour contracts and facilitate labour force adjustments. Theories of efficient labour contracts address the agency problem of asymmetric information regarding worker effort in the presence of monitoring costs (Lazear, 1979). An efficient contract is shown to involve a back-loaded or tilted tenure-remuneration profile. This increases the expected cost of shirking because alternative employment offers lower future remuneration. An important feature of back-loaded profiles is that low-tenure workers earn less than their productivity and high-tenure workers earn more than their productivity. The employer needs to limit the length of time during which older workers are able to earn above their productivity. This can be achieved through a mandatory retirement age or the accrual profile of a defined benefit employer-provided pension plan.

Alternatively, employers may wish to induce early retirement rather than impose lay-offs when facing adverse demand shocks. Employers can carry this out by altering defined benefit pension accrual, or offering retirement plans which provide higher accrual for certain age groups for a limited time (Lumsdaine et al., 1992). Also, unexpected positive technology shocks may accelerate older workers' skill obsolescence (Bartel and Sicherman, 1993). While retraining is costly in general, older workers are at a disadvantage relative to younger cohorts, since such an investment has to be recouped over a shorter period of time.

These theories suggest that the distribution of productivity between establishments should be reflected in the distribution of retirement age between establishments. Hence establishment-specific policies are important determinants of individual retirement decisions. In particular, we expect a strong within-establishment correlation between retirement dates once other factors have been controlled for such as individual occupation, education and demographic characteristics. Nevertheless, where early retirement plans are financed by the State we would expect establishment-specific policies to have effects on individual decisions which are less pronounced or even non-existent. Indeed, if most workers are eligible for such programmes, it may be very costly for a single establishment to design incentive-compatible contracts which exhibit the desired features and survive the effect of a State funded scheme. [1]

[1] However, we may still find tilted tenure–pay profiles.

Our objective is to estimate the contributions of publicly funded programmes and employer wage policies to individual retirement incentives. A conventional structural dynamic programming model of retirement is presented. This is extended to account for heterogeneity in retirement behaviour between individuals as well as between employers. The model is estimated on a matched panel of small-to-medium sized establishments and all their employees. Information on co-workers is used to account for establishment-specific wage policies which may contribute to workers decisions to retire early. Programme effects are identified by exploiting changes in the structure of a Danish publicly funded early retirement programme. We argue that eligibility to the programme is exogenous to the retirement decision. This provides income variation between otherwise similar individuals that can be used to measure public transfer incentive effects.

The remainder of the paper is organised as follows: The next section sets up the economic model and describes the estimation method. Section 3 discusses the relevant background on retirement and pensions to motivate Denmark as of interest because of the institutional variation it provides which can inform measurement. Section 4 presents the dataset used and defines and describes the two variables of primary interest, retirement age and income. Section 5 presents, interprets and discusses the results. A final section summarises and concludes.

2. Economic model and empirical specification

The methodological consensus on retirement modelling strategy is to consider either simple option value models (Stock and Wise, 1990) or more complex structural dynamic programming models (Rust, 1995). Stern (1997) shows that structural dynamic programming models of retirement are to be preferred to option value models on the grounds of consistency with the underlying economic decision problem.[2] Furthermore, structural models are the appropriate tools for simulating policy reforms. We present a conventional structural dynamic programming model of individual retirement which is akin to Wolpin (1987) in a different context.

2.1. Economic model for individual behaviour

Our starting point is a standard discrete-time optimal stopping markov model in finite horizon. The decision to retire is final and retirement is considered to be an absorbing state. The horizon is finite because of the certainty of individual death. Retirement can occur at any moment of time before death. In this framework, age of the individual and time are equivalent concepts. We take the conventional approach in that we do not explicitly model saving behaviour.[3]

[2] Lumsdaine et al. (1992) show that both approaches have similar predictive power.

[3] See Rust and Phelan (1997) for a discussion. Rust (1995) proposes a structure which would allow saving.

The instantaneous utility of work, $u_{1,t}$ (relevant from the beginning of time period t to the end of t), depends on the income flow received during the period only. Since the earnings level may vary with experience at t, E_t, which is endogenous here we indicate that dependence, $w_t(E_t)$. We have

$$u_{1,t}(w_t(E_t)) = \alpha \ln(w_t(E_t)) + \gamma_1 + \varepsilon_\gamma(1). \tag{1}$$

Similarly, the instantaneous utility of being retired, $u_{0,t}$, depends on the flow of retirement income received during the period, b_t. This may depend on the earnings level during the last period of work, or on the complete work history until retirement. We have

$$u_{0,t}(b_t) = \alpha \ln(b_t) + \gamma_0 + \varepsilon_t(0). \tag{2}$$

We assume that $\varepsilon_t(0)$ and $\varepsilon_t(1)$ are random error terms, independently distributed across time. We further assume that both are normally distributed with zero expectation and such that the variance of their difference is equal to one.[4] This normalisation is standard in binary response models. The difference between $\gamma_0 + \varepsilon_t(0)$ and $\gamma_1 + \varepsilon_\gamma(1)$ captures the utility of leisure at any point in time. In a limited way this may reflect unobservable variations in health affecting the valuation of leisure relative to the disutility of work.

The objective of the individual is to maximise his or her expected lifetime utility over all possible future states:

$$\operatorname*{Max}_{\{d_t\}_{t=0,\dots,T}} \mathrm{E}_0 \left[\sum_{t=0}^{T} \beta^t (d_t u_{1,t}(w_t) + (1 - d_t) u_{0,t}(b_t)) \right], \tag{3}$$

where d_t is the strategy in t, $\{d_t\}_{t=0,\dots,T}$ is the sequence of strategies over the lifetime. Since this problem has an optimal stopping structure, the strategy in t is of the reservation value type. In our case, the reservation utility is such that the value of retiring now is equal to the value of working an additional period. In a finite horizon although the reservation utility cannot be expressed analytically, it is possible to compute it using backward recursion. In the present context this is relatively straightforward. Let us denote $V_t(d_t; w_t, b_t, E_t, \varepsilon_t)$ the value of taking the decision d_t given the state variables. In t the value of taking decision d_t is

$$\begin{aligned} V_t(d_t; w_t, b_t, E_t, \varepsilon_t) = {} & d_t(\alpha \ln(w_t) + \gamma_1 + \varepsilon_t(1)) \\ & + (1 - d_t)(\alpha \ln(b_t) + \gamma_0 + \varepsilon_t(0)) \\ & + \beta E_t \left[\max_d \{V_{t+1}(d; w_{t+1}, b_{t+1}, E_{t+1}, \varepsilon_{t+1})\} \mid d_t \right]. \end{aligned} \tag{4}$$

[4] In a model with two alternatives this is without consequences.

Given the assumptions we made earlier about the error term distribution it is easy to compute the last term for $d_t = 1$:

$$
\begin{aligned}
&\mathrm{E}_t\left[\max_d \{V_{t+1}(d; w_{t+1}, b_{t+1}, E_{t+1}, \varepsilon_{t+1})\} \mid d_t = 1\right] \\
&\quad = \bar{V}_{t+1}(0; E_t) \\
&\qquad + \Phi[\bar{V}_{t+1}(1; E_t + 1) - \bar{V}_{t+1}(0; E_t)](\bar{V}_{t+1}(1; E_t + 1) - \bar{V}_{t+1}(0; E_t)) \\
&\qquad + \phi[\bar{V}_{t+1}(1; E_t + 1) - \bar{V}_{t+1}(0; E_t)],
\end{aligned} \tag{5}
$$

where $\Phi[\cdot]$ stands for the normal cumulative distribution function and $\phi[\cdot]$ stands for the normal probability distribution function, and where $\bar{V}_t(\cdot;\cdot)$ stands for the expectation of $V_t(\cdot;\cdot)$ given the relevant history up to, but not including, t. The last term arises in the calculation of the expectation of the maximum of two normal random variables (for a detailed discussion see Wolpin, 1987). Since retirement is an absorbing state, the value of retiring is expressed as a discounted sum of expected flows of income in retirement:

$$
\begin{aligned}
V_t(0; w_t, b_t, E_t, \varepsilon_t) &= \alpha \ln(b_t) + \gamma_0 + \varepsilon_t(0) + \mathrm{E}_t\left[\sum_{\tau=t+1}^{T} \beta^{\tau-t}(\alpha \ln(b_\tau) + \gamma_0 + \varepsilon_\tau(0))\right] \\
&= \alpha \ln(b_t) + \gamma_0 + \varepsilon_t(0) + \sum_{\tau=t+1}^{T} \beta^{\tau-t}(\alpha\, \mathrm{E}_t[\ln(b_\tau)] + \gamma_0).
\end{aligned} \tag{6}
$$

If we assume, as we do in our empirical model, that out-of-work income flows are constant and are equal to $\ln(b_t)$ after retirement at date t until death at date T, the previous expression simplifies to

$$
V_t(0; w_t, b_t, E_t, \varepsilon_t) = \frac{1 - \beta^{T-t+1}}{1 - \beta}(\alpha \ln(b_t) + \gamma_0) + \varepsilon_t(0). \tag{7}
$$

Substituting (5) into (4) gives the value of not retiring in t

$$
\begin{aligned}
&V_t(1; w_t, b_t, E_t, \varepsilon_t) \\
&\quad = \alpha \ln(w_t) + \gamma_1 + \varepsilon_t(1) + \beta\{\bar{V}_{t+1}(0; E_t) \\
&\quad + \Phi[\bar{V}_{t+1}(1; E_t + 1) - \bar{V}_{t+1}(0; E_t)](\bar{V}_{t+1}(1; E_t + 1) - \bar{V}_{t+1}(0; E_t)) \\
&\quad + \phi[\bar{V}_{t+1}(1; E_t + 1) - \bar{V}_{t+1}(0; E_t)]\}.
\end{aligned} \tag{8}
$$

The last two expressions allow backward computation of the value of retirement and the value of working one more year. Hence the strategy in t is such that

$$
d_t = \begin{cases} 1 & \text{if } V_t(1, w_t, b_t, E_t, \varepsilon_t) > V_t(0, w_t, b_t, E_t, \varepsilon_t), \\ 0 & \text{otherwise.} \end{cases} \tag{9}
$$

2.2. Introducing heterogeneity

It is important to introduce some observable or unobservable heterogeneity into the model outlined above. In particular, this allows for correlation between observations within groups which are observed and are potentially relevant for the decision process being modelled. In our case the data allows us to control for both individual and establishment group effects. Following Berkovec and Stern (1991), we endow the stochastic component of income flows with a random error component structure:

$$\ln(w)_{t,f,i} = x_{t,f,i}\beta_w + \eta_{t,f,i}(1) + \hat{\xi}_f\delta_w + \omega_i,$$

$$\ln(b)_{t,f,i} = z_{t,f,i}\beta_b + \eta_{t,f,i}(0) + \hat{\xi}_f\delta_b + \theta\omega_i, \tag{10}$$

where the subscripts t, f, and i indicate time, establishment and individual, respectively. The error term $\eta_{t,f,i}(\cdot)$ is assumed to be independently and identically distributed across time, establishments and individuals, the term ξ_f describes the establishment-specific effect, and the term ω_i describes the individual-specific effect. $x_{t,f,i}$ and $z_{t,f,i}$ are vectors of covariates that determine the in-work and out-of-work income flows, and β_w and β_b are conformable vectors of parameters. The expectation operator in the conditional value functions is understood to be conditional on the value of these terms.

Instead of integrating out the establishment-specific effect (the marginal approach) we condition on an out-of-sample measurement of the establishment-specific effect. Since our dataset is exhaustive of employees within sampled establishments, we can use all workers *except* our sub-sample of potential retirees and estimate a simple within-establishment OLS model to identify the establishment effect. This generates a distribution of establishment unobserved heterogeneity terms (one for each establishment), say $\hat{\xi}_f$.

δ_w and δ_b are parameters which measure the impact of establishment-specific effects, $\hat{\xi}_f$, on in-work and out-of-work income flows. If they are equal, then there is no within establishment correlation of retirement ages. A positive value of δ_w means that establishments which pay their employees who have not reached the age of potential retirement above average (in which case $\hat{\xi}_f > 0$) will pay those who have reached the age of potential retirement above average. The individual-specific effect, ω_i, accounts in a conventional way for unobserved individual heterogeneity in retirement behaviour. Therefore, it accounts for a possible serial correlation between income paid in different periods to the same individual. Moreover, a θ different from unity allows for individual retirement decisions to depend on the individual-specific effect ω_i. Indeed, the decision rule at any point in time compares the differences between income flows in-work and out-of-work. That difference depends on ω_t if and only if $\theta = 1$.

2.3. Likelihood specification

Given the stochastic assumptions we have made previously (in particular conditional independence), the value in t of each possible alternative for an individual i in firm f is

$$V_{t,f,i}(0) = \alpha(z_{t,f,i}\beta_b + \hat{\xi}_f\delta_b + \theta\omega_i) + \gamma_0 + \varepsilon_{t,f,i}(0) + \alpha\eta_{t,f,i}(0)$$
$$+ \sum_{\tau=t+1}^{T} \beta^{\tau-1}\{\alpha(z_{\tau,f,i}\beta_b + \hat{\xi}_f\delta_b + \theta\omega_i) + \gamma_0\}, \tag{7$'$}$$

$$V_{t,f,i}(1) = \alpha(x_{t,f,i}\beta_w + \hat{\xi}_f\delta_w + \omega_i) + \gamma_1 + \varepsilon_{t,f,i}(1) + \alpha\eta_{t,f,i}(1) + \beta\,\mathrm{E}[\max\{V_{t+1,f,i}(0), V_{t+1,f,i}(1)\}]. \tag{8'}$$

As usual with multi-regime models, because of partial observability the correlation between income flows is not identified and therefore set to zero. Given the firm and individual fixed effects, we have

$$\begin{pmatrix} \varepsilon_{t,f,i}(1) + \alpha\eta_{t,f,i}(1) - \varepsilon_{t,f,i}(0) - \alpha\eta_{t,f,i}(0) \\ \eta_{t,f,i}(0) \\ \eta_{t,f,i}(1) \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \\ 0 \end{pmatrix}, \begin{pmatrix} 1 & \sigma_{10,0} & \sigma_{10,1} \\ \sigma_{10,0} & \sigma_0^2 & 0 \\ \sigma_{10,1} & 0 & \sigma_1^2 \end{pmatrix}\right), \tag{11}$$

independently across individual, years and firms, where the two-digit subscripts in the variance–covariance matrix refer to terms involving a difference in value function between two alternatives, whereas the single-digit numbers refer to income flows.

The sample of N individuals is composed of N_1 individuals who retire at some date $t_{1,i}$ [5] with completed spells of participation, and N_+ individuals who have an uncompleted spell of participation of length l_i. To simplify notation in what follows we drop the firm subscript (this is warranted since the firm-specific component is among the regressors). The contribution to the likelihood of each sub-sample is

$$L_1 \propto \prod_{i=1}^{N_1}\prod_{\tau=1}^{t_{1,i}-1}\left\{\frac{1}{\sigma_1}\phi\left[\frac{e_{\tau,1,i}}{\sigma_1}\right]\Phi\left[\frac{e_{\tau,10,i}}{\sqrt{1-\rho_{10,1}^2}} - \frac{\rho_{10,1}e_{\tau,1,i}}{\sigma_1\sqrt{1-\rho_{10,1}^2}}\right]\right\} \left\{\frac{1}{\sigma_0}\phi\left[\frac{e_{t_{1,i},0,i}}{\sigma_0}\right]\left(1-\Phi\left[\frac{e_{t_{1,i},10,i}}{\sqrt{1-\rho_{10,0}^2}} - \frac{\rho_{10,0}e_{t_{1,i},0,i}}{\sigma_0\sqrt{1-\rho_{10,0}^2}}\right]\right)\right\}, \tag{12}$$

$$L_+ \propto \prod_{i=1}^{N_+}\prod_{\tau=1}^{l_i}\left\{\frac{1}{\sigma_1}\phi\left[\frac{e_{\tau,1,i}}{\sigma_1}\right]\Phi\left[\frac{e_{\tau,10,i}}{\sqrt{1-\rho_{10,1}^2}} - \frac{\rho_{10,1}e_{\tau,1,i}}{\sigma_1\sqrt{1-\rho_{10,1}^2}}\right]\right\}, \tag{13}$$

where $e_{t,k,i} = \ln(y)_{t,k,i} - \ln(\bar{y})_{t,k,i}$ with $\ln(y)_{t,1,i} = \ln(w)_{t,i}$, and $\ln(y)_{t,0,i} = \ln(b)_{t,i}$, $e_{10,t,i} = \bar{V}_{t,i}(1) - \bar{V}_{t,i}(0)$. $\ln(\bar{y})_{t,k,i}$ stand for the unconditional mean of the relevant income flow at time t for individual i. If income flows are not observed during part of the history the unconditional probabilities of participation/retirement can be substituted into these

[5] Note that the subscript i indicates that these individuals may have different retirement ages.

expressions, provided that non-observability is not correlated with the unobservables of the model.

The likelihood of the model with individual-specific effects can be derived directly from the previous expressions since given the specific effect the error structure is identical to the error structure without. The contribution to the likelihood of one individual belonging to the first group becomes

$$L_{1,i} \propto \int_{-\infty}^{+\infty} \prod_{\tau=1}^{t_{1,i}-1} \left\{ \frac{1}{\sigma_1} \phi\left[\frac{e_{\tau,1,i}(\omega)}{\sigma_1}\right] \Phi\left[\frac{e_{\tau,10,i}(\omega,\theta\omega)}{\sqrt{1-\rho_{10,1}^2}} - \frac{\rho_{10,1} e_{\tau,1,i}(\omega)}{\sigma_1\sqrt{1-\rho_{10,1}^2}}\right]\right\}$$
$$\left\{\frac{1}{\sigma_0}\phi\left[\frac{e_{t_{1,i},0,i}(\theta\omega)}{\sigma_0}\right]\left(1-\Phi\left[\frac{e_{t_{1,i},10,i}(\omega,\theta\omega)}{\sqrt{1-\rho_{10,0}^2}} - \frac{\rho_{10,0} e_{t_{1,i},0,i}(\theta\omega)}{\sigma_0\sqrt{1-\rho_{10,0}^2}}\right]\right)\right\}$$
$$\frac{1}{\sigma_\omega}\phi\left[\frac{\omega}{\sigma_\omega}\right] \mathrm{d}\omega, \tag{14}$$

where we make clear how the terms $e_{\tau,1,i}(\omega)$, $e_{t_{1,i},0,i}(\theta\omega)$ and $e_{t_{1,i},10,i}(\omega,\theta\omega)$ depend on the individual specific effect ω, either through the in-work income flow or through the out-of-work income flow.

A similar expression can be obtained for the contributions to the likelihood of individuals belonging to the second group. In practice, the integral can be computed at relatively low cost using quadrature methods.

3. Retirement and pensions in Denmark

Projections for all OECD countries indicate that the elderly proportion of the population will increase dramatically over the next 40 years. Denmark is no exception with the ratios of those aged 65 plus to those aged 15–64 projected 0.22, 0.32, 0.38, respectively, in 2000, 2020, 2030. [6] Of the countries for which comparable information is available, Denmark had the highest participation rate for this group in 1960 (90%), and one of the largest subsequent reductions to 1992 (45%). However, all European countries experience a qualitatively similar evolution (ILO, 1995).

Public transfers to pensioners constituted 10% of Danish GDP in 1994 (Statistics Denmark, 1995). This compares with the OECD average of 9%. Twenty-five per cent of pension transfers were in the form of early retirement benefits. Other countries can expect to learn something from Danish retirement experience since the behaviour of its elderly workers is similar to that elsewhere, and publicly financed early retirement programmes are of similar size to those existing in other countries. [7]

[6] OECD member country averages are 0.21, 0.30, 0.38, respectively, for 2000, 2020, 2030 (OECD, 1995).

[7] Many OECD nations with the exception of Denmark are discussed in OECD (1995).

Consumption by the elderly in Denmark is largely financed by public transfer programmes. Provision for retirement through private pensions is empirically much less important. Supplementary private pensions in the form of tax-subsidised company or individual pension schemes are widespread with 54% of the work-force contributing (Pedersen and Smith, 1995). However, they are not well developed with a large majority far from maturity and currently representing only a small fraction of total pension income. [8]

The publicly financed old age pension is available at 67, covers everyone regardless of previous attachment to the labour market, and was worth 40% of average production worker (APW) earnings in 1992. However, the average age of retirement was then 61.5 (Danish Social Commission, 1993). Hence, early retirement is the preferred exit route from the labour market for a large proportion of elderly workers.

3.1. Out-of-work transfer programmes

The importance of non-work amongst the population of potential working age (17–66) in Denmark has increased from 11% in 1980 to 19% in 1991. Public transfer programmes primarily finance consumption for those who are out-of-work: Unemployment Insurance Benefit (7% of the population), Invalidity Benefit (7%), Post-Employment Wage (3%) and Social Assistance (2%). [9] It is informative to briefly compare the rules and generosity of each of these transfer programmes, in order to motivate our focus on the Post-Employment Wage as the programme of primary interest for measuring incentives.

Social Assistance (Bistandshjælp) is the lowest level of income support, which is means-tested and available to those who are neither working nor insured against unemployment. It is worth about 24% of (APW) earnings. Unemployment Insurance Benefit (Arbejdsløshedsforsikring) is 90% of former wage up to a ceiling (60% of APW). It is available to those insured during at least 6 of the previous 12 months and currently actively seeking work, and lasts for up to 30 months. Provision for those not working due to disability or long-term sickness is through the Invalidity Benefit (Førtidspension) programme. It is available to those aged over 17 according to three levels of assessed ability to work on medical and social criteria (46% APW).

The Post-Employment Wage (Efterløn, hereafter PEW) is an early retirement programme where remuneration is the same as for Unemployment Insurance Benefit, but after 30 months, instead of support being withdrawn, it is reduced to 80% of its former level until age 66. Eligibility is restricted to those aged 60–66 who have contributed to an unemployment insurance fund.

It is important to reemphasise the relative incentives for early exit that the various public transfer programmes create. PEW (for those eligible) is as generous as

[8] See Ministry of Finance (1995) for some measures. Simply computing the ratio of active to receiving members of private pension funds: 20/1 gives an idea of their immaturity. Mean 1993 receipt (conditional on receiving) was DKK 64,000.

[9] This breakdown understates the relative importance of post-employment wage for our elderly sample of interest. After a large influx when the programme was introduced, participation soon stabilised at 100,000.

Unemployment Insurance Benefit without the job search test, and dominates after 30 months, where Social Assistance becomes the alternative. The lowest level of Invalidity Benefit, which may be regarded as a substitute to the extent that it may be given on social grounds, is much less financially rewarding. In summary, PEW is the most attractive programme, and relative to employment it is most appealing to low wage earners.

3.2. Eligibility to the Post-Employment Wage programme

The Post-Employment Wage programme was introduced in 1979. Current members of all Danish Unemployment Insurance Funds were able to participate in the programme and retire between ages 60 and 66 inclusive, as soon as they had been a member of a fund for at least 5 of the last 10 years.[10] The eligibility requirement for the number of years of membership was increased in 1980 (to 10 of the last 15), 1985 (to 15 of the last 20) and 1990 (to 20 of the last 25 years). Membership of a fund is not mandatory in Denmark, but 80% of private sector workers are members.

The effect of the programme rules and reforms described is to create three distinct groups according to date of birth and membership status. The three PEW eligibility groups are: (1) eligible from age 60 because of a long membership history, (2) eligible from some age between 60 and 66 once sufficient membership duration is acquired, (3) never eligible because of non-current membership or membership history too short to be increased sufficiently before age 67. Consider, for example, the case of an individual aged 58 in 1980 having been a fund member for the last: (1) 18 years, may retire on PEW in 1982 at age 60, (2) 6 years, may retire on PEW in 1984 at age 62, (3) 4 years may not retire on PEW.

Changes in eligibility criteria for the PEW programme form the basis for identification of the effects of transfer programmes on individual income and retirement. Calculating individual eligibility requires knowledge of unemployment insurance fund membership back to 1965. We do not observe membership prior to the sample period, which only spans 1980–1991. However, we do know the employment history for each person during 1964–1979, thanks to the history of mandatory pension contributions. Fund membership is imputed during these missing years by using employment history together with fund membership status in 1980. The assumption is that membership status in 1980 reflects membership during the period 1965–1979. This can be justified by the fact that very few older workers (only 3% of those aged over 44) change membership status.

Unemployment insurance fund contributions and benefits are a function of part-time (PT) and full-time (FT) employment. PEW entitlement for a history of PT employment of a given length is worth about half that of FT. Furthermore, the unemployment insurance system and pension system distinguish between PT and FT work according to the same rules. Hence, conditional on the assumption regarding the equivalence of mandatory pension contributions and insurance fund membership, we are able to impute to each worker an age at which he or she becomes eligible to the PEW. In our sample, 89% of men are potentially eligible, 82% of FT women and 61% of PT women.

[10] Prior to 1979 no pension benefits were associated with unemployment insurance fund membership.

4. Data description and empirical strategy

The dataset used in estimation is a sample drawn from the Statistics Denmark Integrated Database for Labour Market Research (IDA). This database is unique among the Danish administrative registers in that it contains a longitudinally consistent link between all establishments and all employees.[11] The sample that we have access to contains 5% of private sector establishments with 5–500 employees in week 48 of 1980. These establishments are followed until 1991, with replacement representative of establishment births into the size class. This sample is therefore representative of small-to-medium sized private sector Danish establishments in each annual cross section. To these establishments are matched all persons employed in week 48 each year who are aged 15–75. Each person is followed over time as long as they are employed at a sampled establishment, and for one additional year should they separate from the establishment.

For the analysis of retirement we consider sub-samples of older workers, where it is safe to assume that the next labour supply decision is retirement, rather than movement to another establishment. Men are sampled from age 55 and women from 50, after which the annual rate of between-establishment mobility is below 5%.[12]

The most important question in an empirical retirement study is the definition and identification of retirees. Retirement is defined as a transition to long-term (at least 1 year) non-work for older workers. Non-work, for us, requires that three conditions are fulfilled for a year: (1) zero mandatory pension contributions which are only made in-work; (2) below 20% of median labour earnings in the year; (3) no attachment to an employer.[13] Consequently, those unemployed for a full calender year, those in receipt of invalidity benefit and pensioners with limited labour market attachment will all be labelled as retired.

Hazards of moving from work to non-work are presented in Fig. 1 for each of the three samples. Spikes in the retirement hazard occur at ages 60 and 67, which correspond to the ages of first potential eligibility to the Post-Employment Wage early retirement programme and the National Old Age Pension, respectively.[14]

Transitions between several other combinations of origin and destination state are informative, but not presented because their age dimension is unimportant: Transitions from non-work to work for men and women and between PT and FT for women are negligible. We conclude that non-work is an absorbing state. This justifies

[11] An establishment is a business operation with a distinct physical address. A legal business entity may contain several different establishments. It is the continuity of the physical rather than the legal unit which has undergone the most thorough consistency checks. The establishment is the unit by which we shall group employees in this analysis. See Leth-Sørensen (1995) for more details on IDA, and Abowd and Kramarz (1999) for a survey of other analyses using matched employer–employee data.

[12] Description of other explanatory variables used in the analysis is relegated to the appendix (Table 3).

[13] This definition is similar to that used in other retirement studies using Danish register data, for example Pedersen and Smith (1996).

[14] These retirement hazards differ from those found in the population as a whole because of the absence of public sector employees from our sample. Junior and senior public sector workers face mandatory retirement ages of 67 and 70, respectively. Otherwise the hazards correspond closely.

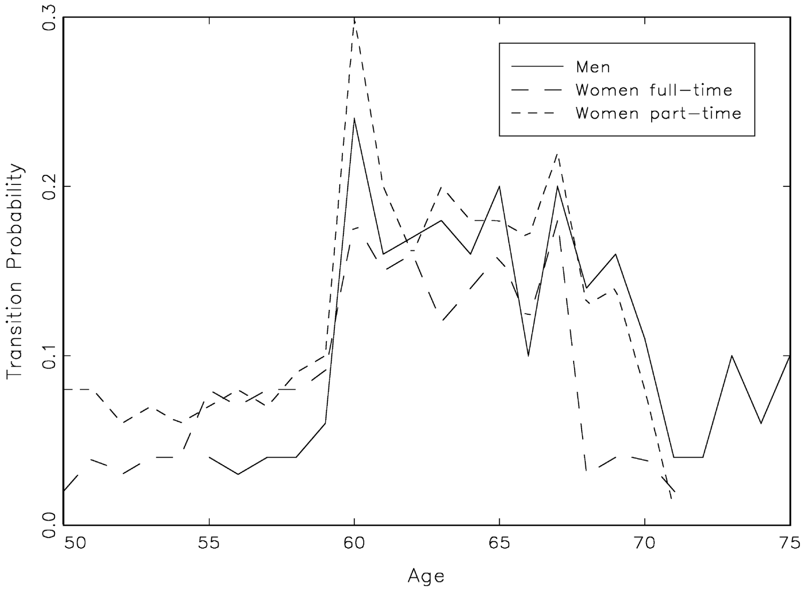

Fig. 1. Transitions to non-work.

defining retirement as a transition to non-work. Furthermore, it is appropriate to split the female sample by origin state and estimate separately. In summary, all inference is conditional on working in a given establishment at age 50 for women and 55 for men. Further, for women it is conditional on employment status then, either full-time or part-time.

An equally important question in a study of retirement incentives is the measurement of income. Real net income is computed from administrative tax records.[15] Our sampling frame only allows us to observe at most 1 year of retirement income. This provides a good indication of future expected but unobserved flows, since we know from other datasets that income exhibits a large fall on retirement, which quickly stabilises at a much lower level than in-work (on average 50%).

Establishment-specific effects in income are obtained from two auxiliary regressions on co-workers. To provide establishment effects for women an OLS log real net income equation is run for the whole IDA sample of workers excluding women aged 50 and above. Similarly, to obtain establishment effects for men a regression is run on a sample excluding men aged 55 and over. Table 4 in the appendix presents estimates from these regressions.

[15] Deductions and total taxable income are recorded consistently throughout the sample period. Net income is calculated by applying the average tax rate per gross income group (see Statistics Denmark, 1991/1993/1995b).

5. Results and discussion

Our aim is to estimate retirement preference parameters recovered from the structural model. Income flows in-work and out-of-work are the main economic determinants of interest, and our aim is to pay careful attention to modelling these flows. Specifically, we want to measure the effects of two income determinants: Establishment heterogeneity and public transfer programme eligibility. The former is the novelty of this paper, and the latter provides identifying variation in income flows. Hence we, respectively, contribute to the personnel economics literature and address the simultaneity problem faced by purely structural (as opposed to natural experimental) approaches.

The model described in Section 2 is estimated separately on three IDA sub-samples: Men aged over 54, FT women aged over 49 and PT women aged over 49. We assume everyone dies before age 100 and retires before 76. Following Berkovec and Stern (1991) we allow for the possibility of death before age 100. This makes the discount rate dependent on time and gender (g) as follows: $\beta^*_{gt} = \beta\pi_{gt}$, where β is a constant time preference element and π_{gt} reflects the mortality rate.[16] Income flows are explained by occupation dummies, region dummies, measured labour market experience for the in-work income equation, and age at retirement and benefit eligibility for the retirement income equation. The difference between γ_0 and γ_1 depends linearly on years of education and age.

Note that the occupation and region dummies are excluded from the description of preferences. Note further that it is, in principle, possible to perform a test for the existence of non-pecuniary job-specific unobserved heterogeneity. In our context, this would amount to testing whether the occupation dummies must be included in the characterisation of preferences. In practice, their exclusion is important for identification of parameters α and β. Relaxing the exclusion restrictions, as is implicit in the definition of the test, would make the identification of α and β tenuous and the proposed test would have low power. For this reason we do not perform such a test.

The three equations presented in Table 1 are the retirement decision (top), in-work income (middle) and retirement income (bottom). Consider first the retirement decision. α represents the instantaneous utility of income. This parameter is only estimated up to scale, but its size relative to other parameters in the retirement decision indicates that income is much more important for men than for women, and slightly more important for FT than PT women. β is the rate of time preference. The parameter is estimated quite precisely for males at 0.459. However, for females it is estimated much less precisely, and may lie in the range 0–0.5. This characterises very myopic behaviour. It is rare in the retirement literature to freely estimate discount rates, and most studies are content to assume 0.97.[17] The parameters of the income process, which are of primary interest in our model, are qualitatively similar if the discount rate is fixed at 0.97. It is reassuring that discount rates are not so pivotal as to drive all other inference.

[16] Mortality rates are obtained from Statistics Denmark (1975/1985/1995a). Over this period cohort effects were negligible for our purposes.

[17] See the Lumsdaine and Mitchell (1999) retirement literature survey and references there in.

Table 1
Retirement decisions and income flows model estimates (standard errors)

Parameter	Women part-time	Women full-time	Men
α	1.093 (0.235)	1.591 (0.228)	2.824 (0.247)
β	0.085 (0.209)	0.204 (0.195)	0.459 (0.045)
Leisure utility $\gamma_0 - \gamma_1$			
Constant	−1.882 (0.617)	−3.131 (0.659)	−1.005 (0.369)
Education/100	−0.896 (0.975)	−2.902 (0.948)	1.349 (0.328)
Age/100	2.981 (0.859)	5.908 (1.048)	3.078 (0.522)
Married no kids	0.222 (0.066)	0.050 (0.040)	−0.065 (0.017)
Married with kids	0.212 (0.073)	0.117 (0.063)	0.015 (0.020)
Cohabiting	0.328 (0.153)	0.052 (0.085)	0.028 (0.032)
Work log net income			
Constant	4.210 (0.015)	4.739 (0.012)	5.078 (0.009)
δ_w	0.544 (0.030)	0.350 (0.019)	0.741 (0.015)
Experience/100	1.541 (0.053)	0.519 (0.031)	−0.303 (0.020)
Retirement log net income			
Constant	1.521 (0.322)	2.602 (0.226)	4.118 (0.111)
δ_b	0.323 (0.132)	0.102 (0.109)	0.744 (0.026)
Ret. age/100	2.633 (0.501)	2.113 (0.391)	0.256 (0.177)
Eligibility	0.570 (0.061)	0.319 (0.037)	0.139 (0.010)
Covariance matrix			
σ_1^2	0.054 (0.001)	0.025 (0.001)	0.029 (0.001)
σ_0^2	0.260 (0.052)	0.229 (0.009)	0.049 (0.003)
$\rho_{10,1}$	−0.105 (0.126)	0.558 (0.035)	0.658 (0.012)
$\rho_{10,0}$	0.441 (0.187)	0.935 (0.005)	0.332 (0.079)
σ_ω	0.416 (0.002)	0.313 (0.001)	0.266 (0.001)
θ	1.024 (0.047)	0.857 (0.035)	0.933 (0.009)
Log-likelihood	4074	5977	15834
# observations	1914	1924	4939

Note: Also included but not presented in the table are 4 occupation dummies and 7 region dummies.

However, it does call into question the widely maintained assumption in other studies that potential retirees are quite so forward looking in their dynamic behaviour.

The difference $\gamma_0 - \gamma_1$ indicates instantaneous utility of leisure. For men, higher education is associated with greater utility of leisure (retirement). In other words, given a constant income stream, a more educated man would retire earlier than one with less education. Whereas for women the opposite is true. Though for PT women, education is insignificant. Age increases the disutility of work for everyone. For marital status effects, the reference group contains singles. PT women living with a partner have a higher utility of leisure relative to utility of work than do PT women living alone. Married FT women with children living at home have significantly higher utility of leisure than other FT women. Men without children have a higher utility of work than other groups.

Consider now the income equations.[18] Age of retirement in the out-of-work net income equations is significant for women but not for men. For women, working an extra year increases retirement income by 2%, whereas for men the increase is only 0.2%. Since older men have often contributed more towards retirement income security, the financial return to delaying retirement is incrementally lower than for women. As predicted by life-cycle theory, the return to experience is economically small although statistically significant. It is much more important as one considers samples with weaker labour market attachment. Since we have an accurate *measure* of experience, FT women are earlier along their experience-earnings profile than men, and PT women are even earlier.

Eligibility for the public early retirement scheme (PEW) increases out-of-work net income significantly for all samples. This is worth 13.9%, 31.9% and 57.0% of retirement income, respectively, for men, FT women and PT women. PEW appears relatively generous for women, all else equal. Indeed, women's income streams are lower than men's, and since PEW has an absolute maximum regardless of gender, it contributes in greater proportion.

Establishment effects, δ, are highly significant both in-work and out-of-work. Hence establishment unobserved heterogeneity is an important determinant of pay, but less so for older workers. This should be expected as observables explain more of older workers pay. The establishment effect on retirement comes through the difference between the two equations ($\delta_w - \delta_b$). This is significant for women, but not so for men. Consequently, establishment pay policy accounts for some female retirement behaviour, but not for males. Unobserved individual heterogeneity is important in describing individual retirement decisions to the extent that θ is different from one and therefore individual effects in the income processes do not cancel each other out. Only for PT women is θ not significantly different from unity.

Introducing individual-specific effects has the expected consequence of reducing the variance of the transitory component of the income flow error terms. From the estimates, we calculate that individual-specific effects account for around 75% of the variance of in-work incomes. Somewhat more of the out-of-work income variance remains unattributed, as is to be expected, especially for women. Note that we have already explained a large proportion of establishment heterogeneity with the inclusion of the firm-specific effect $\hat{\xi}_f$.

The correlations between each income flow and the differences in utility values are positive irrespective of the type of income considered. A structure similar to a Roy model (Heckman and Honoré, 1990), where occupation choice is based on a

[18] Dummy variables for occupation and region are included in estimation, but not presented in Table 1 for the sake of brevity. There is a similar structure of net income in-work and net income out-of-work, according to the hierarchy of skill groups which are maintained into retirement. For men, both the size and significance of skill and regional differentials are maintained into retirement. For FT women the similarity is weaker, and for PT weaker still.

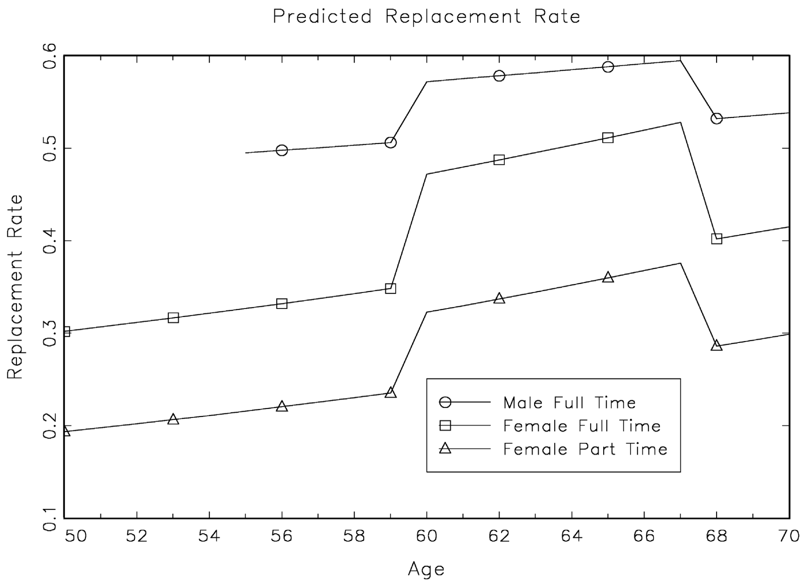

Fig. 2. Predicted replacement rates by age.

comparison of (discounted) earnings only, would lead the two correlations between the difference in values and the income equations to have opposite signs (see Eq. (12)). The structure we present here allows the choice to depend on an additional source of unobserved heterogeneity, $\varepsilon_t(1) - \varepsilon_t(0)$, which allows for more complex correlation patterns. Alternatively, it may well reflect the mispecification of the income flow equations. For example, the omission of a variable in both income flow equations and utility may have such an effect.

Predicted replacement rates from the model are presented in Fig. 2. They exhibit a similar pattern between samples whereby the hump represents the PEW eligibility window. Women benefit relatively more from the PEW, but the male replacement rate is uniformly higher at any age.

A permanent change of 1% in the replacement rate has clear effects on retirement probabilities, as presented in Fig. 3. Elasticities decline with age for all samples, but for men from a much higher level and therefore more steeply. Note that these elasticities reflect a change in the replacement rate relative to that observed in the data. The observed replacement rate for men is significantly higher than for women. The difference in slope between the male and female profiles is due to men's higher instantaneous marginal utility of income and lower rate of time preference.

A summary of model predictions is presented in Table 2. We cannot meaningfully compare expected retirement age with observed because of sample selection, whereby we do not observe everyone up until age of retirement. Men are expected to retire later than women, but among women, it is the part-timers who retire later. This is consistent with part-time women relative to full-timers having a higher return to experience and

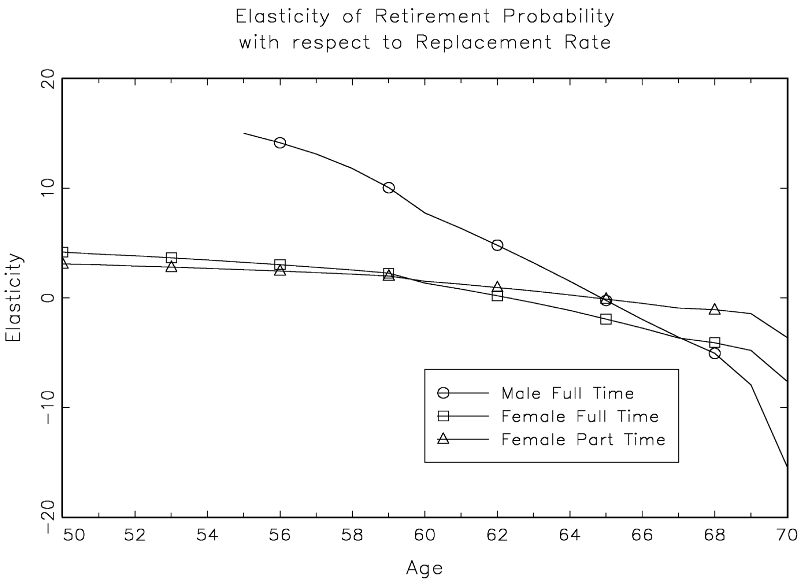

Fig. 3. Retirement elasticities by age.

Table 2
Retirement age modelled outcomes

Sample	Expected	Elasticity with respect to permanent wage
Men	64.2	0.443
Women full-time	61.6	0.181
Women part-time	63.5	0.136

less eligibility to PEW. As a consequence of mens' higher instantaneous marginal utility of income and lower rate of time preference, the elasticity of retirement age with respect to a permanent increase in wages is greater for men than for women. PT women have a higher instantaneous marginal utility of income than full-timers and thus have a lower elasticity.

6. Summary and conclusions

We have analysed the effect of between-establishment variation in remuneration policy on individual worker retirement ages. A conventional structural dynamic programming model is extended to allow employer unobserved heterogeneity to influence retirement decisions through income streams in-work and out-of-work. These employer effects are identified by exploiting a matched panel of Danish small-to-medium sized establishments and all their employees. Exogenous eligibility to a public early retirement programme identifies the effect of individual income flows on retirement.

Several strong assumptions have been imposed in order to obtain our results, and it is important to restate them as caveats to any conclusions drawn. Exogeneity of programme eligibility at the time of the reform cannot be tested within sample since both programme and dataset began in 1979/80. Measurement of eligibility is indirectly through mandatory pension contributions and subject to errors to the extent that workers may change insurance fund membership status before 1980. We have ignored potentially important retirement determinants (for example, household decisions, health and saving) in order to focus on adding employer heterogeneity to a relatively simple model of individual decisions.

Notwithstanding these caveats, pay policy is a potentially important instrument through which employers may influence workers retirement ages, together with mandatory retirement rules. The potential for this instrument is made clear since employer effects on income in-work are found to carry over strongly into out-of-work income. In the presence of efficient labour contracts, tilted tenure-remuneration profiles should be accompanied by employer-specific retirement policies, according to the between-employer productivity distribution. However, only for women do we find evidence of a direct establishment effect on retirement age, after controlling for observables. The relatively attractive public early retirement programme, which is available to the majority of male workers from the age of 60, may disguise or reduce the importance of establishment retirement policies. This transfer is not means-tested against private unearned income, and so employer-effects on non-work income remain, though they are not important enough to generate employer-effects on retirement age. The finding that employer effects on retirement age are only found among sub-samples where access to a public programme is relatively limited is consistent with the idea that generous public benefits make it too expensive for employers to write optimal incentive contracts.

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Appendix

Description of other explanatory variables used in the analysis is shown in Table 4. Auxilliary regression estimates are presented in Table 4.

Table 3
Descriptive statistics means (standard deviations)

Variable	Women PT	Women FT	Men
Education/100	0.103 (0.019)	0.104 (0.020)	0.108 (0.022)
Experience	16.730 (5.614)	20.548 (6.454)	38.611 (7.434)
Firm effect	−0.060 (0.126)	−0.025 (0.116)	−0.002 (0.111)
Age at dataset entry	54.797 (5.234)	53.607 (4.343)	57.985 (4.048)
Age at dataset exit	57.976 (5.593)	56.313 (4.773)	61.144 (4.336)
Ret. age/eligible	58.254 (4.145)	58.208 (4.260)	61.585 (3.079)
Ret. age/ineligible	60.850 (4.826)	62.100 (4.790)	65.046 (4.337)
Duration from 55/55/50	8.976 (5.593)	7.313 (4.777)	7.144 (4.336)
Eligible/retire	382	576	1687 —
Ineligible/retire	140	70	197 —
Married with children	0.541	0.449	0.578
Married no children	0.255	0.185	0.235
Cohabiting	0.024	0.047	0.038
# observations	1914	1924	4939

Table 4
Auxiliary within-establishment income regression model estimates (standard error), mean (standard deviation)

Variable	Excluding women 50+		Excluding men 55+	
	Estimate	Descriptive	Estimate	Descriptive
Log(income)		4.8705 (0.4394)		4.8181 (0.4647)
Intercept	3.8463 (0.0110)		3.6255 (0.0118)	
Full-time	0.3843 (0.0023)	0.8776 (0.3277)	0.4300 (0.0022)	0.8436 (0.3632)
Male	0.1990 (0.0018)	0.6602 (0.4736)	0.2256 (0.0018)	0.5575 (0.4966)
Education	0.0207 (0.0003)	11.2033 (2.1465)	0.0251 (0.0004)	11.1936 (2.1340)
Experience	0.0343 (0.0001)	15.5234 (11.7995)	0.0530 (0.0003)	12.7100 (8.1968)
Experience2	−0.0005 (0.0000)	380.2043 (534.496)	−0.0012 (0.0000)	228.7325 (254.412)
Management	0.1718 (0.0026)	0.0951 (0.2933)	0.1596 (0.0030)	0.0782 (0.2686)
Skilled	−0.0496 (0.0022)	0.1960 (0.3969)	−0.0367 (0.0024)	0.1786 (0.3830)
Unskilled	−0.0944 (0.0021)	0.2348 (0.4239)	−0.0756 (0.0022)	0.2269 (0.4188)
Others	−0.2095 (0.0030)	0.0653 (0.2472)	−0.1928 (0.0029)	0.0839 (0.2772)
log(size)	0.0081 (0.0024)	3.8691 (1.1998)	0.0121 (0.0026)	3.8403 (1.2034)
$\bar{R}^2$/# obs.	0.5877	188304	0.6142	176163
# groups	1514	1899	1516	1890

Note: Also included but not presented in the table are 11 year dummies and 7 region dummies.

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