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Working Paper

Labour supply disincentive effects of old age public pensions: A case study for West Germany combining panel data and aggregate information

Diskussionspapier, No. 28

Provided in Cooperation with:

Department of Economics, University of Konstanz

Suggested Citation: Siddiqui, Sikandar (1995) : Labour supply disincentive effects of old age public pensions: A case study for West Germany combining panel data and aggregate information, Diskussionspapier, No. 28, Universität Konstanz, Forschungsschwerpunkt Internationale Arbeitsmarktforschung, Konstanz

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<https://hdl.handle.net/10419/92447>

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29. SEP. 1995 Weltwirtschaft
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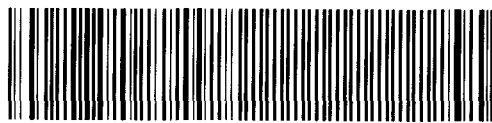
Diskussionspapier
28 - 1995

**Labour Supply Disincentive Effects of Old Age Public Pensions:
A Case Study for West Germany
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W 752 (28)



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Nr. 28

September 1995

Labour Supply Disincentive Effects of Old Age Public Pensions:

A Case Study for West Germany combining Panel Data and Aggregate Information

by Sikandar Siddiqui¹

Abstract:

In this paper, the impact of the West German pension system on the retirement decisions of elderly citizens is analyzed within the framework of a discrete-time hazard rate model deduced from a micro-economic decision rule. The model is estimated using a balanced panel of elderly West German citizens. In order to improve the precision of the estimates obtained, the data from the sample are combined with aggregate-level information on the labour force participation behaviour of the elderly. Policy-related simulation experiments based on the estimates reveal that the probability of early retirement can be reduced significantly by appropriate changes in the pension system.

JEL classifications: C32, C41, J26.

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1. Introduction

Opting for retirement is one of the most momentous decisions faced by elderly individuals. As each of the alternative retirement ages implies a specific combination of present and future amounts of income and leisure, the timing of an employee's exit from the labour force exerts considerable influence on his or her economic well-being. Seen from a macroeconomic perspective, even small variations in the average retirement age of elderly workers can result in important alterations of the aggregate supply of labour, which, in turn, are likely to bring about changes in the average wage level and the unemployment rate. Furthermore, the timing of retirement also determines when a person turns a contributor of taxes to a recipient of benefits with regard to the prevailing public pension system, which makes it a matter of high importance for fiscal and social policy issues alike. Due to the dramatic process of population ageing, which most of the industrialized countries can today be expected to undergo during the next five decades, it is the last-mentioned of these three aspects that has received most public attention in recent years.

It must not be overlooked, however, that it is the relevant pension system itself which, by means of its implicit labour supply (dis)incentives, can have a decisive impact on a person's propensity to retire early. The way in which expected labour income streams, pension benefits, and individual-specific socio-economic characteristics interact in determining retirement behaviour has therefore been the subject of numerous empirical investigations. In this context, it is noteworthy that econometric models of retirement behaviour ought to differ from conventional labour supply models (see, e.g., the survey article by LAISNEY, POHLMEIER, and STAAT, 1992) for two main reasons:

Firstly, choosing one's retirement age ought to be understood as a sequential decision process with multiple discrete alternatives (i.e. the different possible retirement ages allowed by the prevailing pension system). If a person behaves rationally in this context, he or she repeatedly compares the expected utility level associated with the option of immediate retirement with the ones of all other retirement ages still available, and chooses to quit the labour force as soon as the utility of instantaneous retirement exceeds the maximum utility attainable by postponing retirement for the first time.

Secondly, unlike temporary breaks in an individual's work history, the retirement decision is usually regarded as irreversible. In structural econometric models of retirement behaviour, this is bound to result in a problem of endogenous sample selection: Obtaining valid estimates of the parameters of a formal decision model requires that the dataset used for estimation comprises only individuals that did not retire before the beginning of the sampling period; yet on the other hand, this would imply that the selection rule applied is based upon past outcomes

of the decision process under investigation. Apparently, as POHLMIEIER (1993) points out, the only appropriate solution to this dilemma is to discard all individuals whose earliest possible year of retirement lies outside the sampling period from the dataset used for estimation, and to regard the complete time interval between a person's first potential year of retirement and the one actually chosen as the dependent variable.

In most of the early empirical works on the job exit behaviour of elderly citizens, like, e. g., the studies by BLINDER, GORDON and WISE (1980), MITCHELL and FIELDS (1984), and BURTLESS and MOFFITT (1984, 1985), the retirement decision is treated as a static choice problem with multiple discrete alternatives within the framework of a standard neo-classical model of labour supply. While approaches of this kind correctly account for the multiplicity of possible retirement ages, their usefulness in predicting alterations of individual labour force participation behaviour in response to changes in the legal and institutional setting has turned out to be doubtful. This is possibly due to the fact that they neglect the necessarily sequential nature of the retirement decision, thus disregarding the possibility that individuals might revise their plans in response to unexpected changes in income, health, or other factors influencing their preference for leisure.

An alternative modelling strategy consists in the use of continuous-time duration models in the analysis of retirement behaviour. In their analyses of retirement decisions among American citizens, DIAMOND and HAUSMAN (1984) and HAUSMAN and WISE (1984) use models of this type. Comparable studies also exist, e. g., for Finland (LILJA, 1990), Denmark (PEDERSEN and SMITH, 1991), and Germany (BÖRSCH-SUPAN and SCHMIDT, 1993). The main advantage of these approaches is that they adequately account for the problem of dynamic self-selection, which is inherent in econometric models of the job exit behaviour of elderly employees as soon as retirement is understood as an absorbing state. This is, however, counterbalanced by the fact that a model of the above-mentioned type cannot be derived from a formal utility maximization framework which appropriately describes the choice problem faced by potential retirees.

In this paper, I describe a discrete-time hazard rate model which accounts for the multiplicity of possible retirement ages, the "absorbing state" property of retirement, and the possibility of unobserved heterogeneity among individuals. As it is deduced from a rather simple utility maximization framework, its implementation with multi-wave panel data is computationally far less burdensome than the application of the more elaborate discrete dynamic programming models developed by RUST (1989) and BERKOVEC and STERN (1991), or the option value approach due to LUMSDAINE et al. (1990), and STOCK and WISE (1990). Nevertheless, it is capable of integrating possible labour (dis)incentives implied by the rules of the relevant

pension system into the empirical analysis of retirement behaviour, and can therefore be reasonably used as a basis for policy-related simulation experiments.

A second point of interest in the paper is more on the methodical side. Along the lines of recent work by IMBENS and LANCASTER (1994), it is shown how survey data can be combined with aggregate-level information on the variable(s) under investigation in order to obtain more efficient parameter estimates in micro-econometric models. The description of the corresponding econometric procedure is followed by a representation of the manner in which it can be applied to the model developed here, and of the empirical results obtained by doing so.

The paper is organized as follows: In section 2, the theoretical assumptions of the model employed are sketched out. The way in which the aggregate-level information is integrated into the estimation procedure is described in section 3. Section 4 deals with the data used and the definition of the endogenous variable. The empirical results are being discussed in section 5 along with the outcome of a simulation experiment concerning two alternative policy scenarios. A short summary (section 6) concludes the paper.

2. Basic Theoretical Assumptions

In the model which is set up here, it will be assumed that the utility level U which individual i can enjoy at time t if he (or she) chooses to retire at time $r_i \geq t$ can be expressed as a function of the expected present discounted value of all present and future income streams (which will be denoted by $Y_{ii}(r_i)$ here and in the following), the expected length of his (her) retirement period, $L_{ii}(r_i)$, and a disturbance term $\varepsilon_{ii}(r_i)$. For the alternative-specific utility level $U_{ii}(r_i)$ which is associated with retirement at time r_i , the following functional form is specified:

$$U_{ii}(r_i) = \ln Y_{ii}(r_i) + \beta_{ii} \ln L_{ii}(r_i) + \varepsilon_{ii}(r_i) \quad (1)$$

The parameter β_{ii} is a taste parameter which measures the relative intensity of an individual's preference for leisure and is allowed to vary among individuals as well as over time. It is assumed to be a linear function of a vector γ of unknown parameters and of a set of exogenous variables, which are collected in the vector z_{ii} and represent certain socio-economic characteristics of the individuals observed:

$$\beta_{ii} = z_{ii}' \gamma \quad (2)$$

$\pi_i(s|t)$ denotes the individual-specific probability that a person who was alive at time t is still alive at time $s \geq t$, and S_i is the last period in which $\pi_i(s|t)$ is positive. It is assumed that individuals earn a labour income before retiring and pension benefits from the date of retirement onwards. In this case, the expected present discounted value of present and future income levels, $Y_{it}(r_i)$, can be expressed as follows:

$$Y_{it}(r_i) := \sum_{s=t}^{r_i-1} E_t(y_{is}) \cdot \pi_i(s|t) \cdot (1+\rho)^{t-s} + \sum_{s=r_i}^{S_i} E_t(b_{is}(r_i)) \cdot \pi_i(s|t) \cdot (1+\rho)^{t-s} \quad (3)$$

In this equation, y_{is} denotes an individual's labour income at time s , the variable b_{is} stands for the amount of retirement benefits received by the respective person, E_t is the expectations operator given the information set of time t , and ρ is a discount factor, which is set to 0.03 for the sake of simplicity. (The amount of assets held by an individual is not included as an argument of the utility function within the framework of this model because the corresponding statements in the German Socio-Economic Panel, which is the data base used here, were found to be rather imprecise and incomplete). The expected length of a person's retirement period is defined as

$$L_{it}(r_i) = \sum_{s=r_i}^{S_i} \pi_i(s|t) \quad (4)$$

Let t_{i0} denote the period in which person i first has the opportunity to retire and r_i^{max} the latest possible year of retirement for the same person. (For reasons that will become obvious later, t_{i0} is assumed to be the year of the corresponding individual's 58th birthday). This implies that an individual who has not retired until the beginning of period t can either opt for immediate retirement or plan to retire at one of the $(r_i^{(max)} - t)$ remaining future years of retirement. Thus, an individual can be said to have the choice between a total number of $(r_i^{(max)} - t_{i0} + 1)$ different retirement ages at time t_{i0} . These alternatives are henceforth numbered in consecutive order beginning with zero. In order to retain relative computational simplicity while adding some stochastic flexibility to what otherwise would be a restrictive multinomial logit or independent probit model, a so-called "one-factor" decomposition (AMEMIYA, 1985, p. 323f.) is chosen as to the stochastic error terms $\varepsilon_{it}(r_i)$:

$$\varepsilon_{it}(r_i) = \delta_{r_i-t_{i0}} \cdot v_i + u_{it}(r_i) \quad (5)$$

The error components $u_{it}(r_i)$ and v_i are assumed to be normally distributed with mean zero; the "factor loadings" δ_0 to $\delta_{r_i^{(max)}-t_{i0}}$ are unknown parameters which have to be estimated.

In order to be able to identify the elements of the parameter vector γ , however, the following normalizations are introduced:

$$E[u_{is}(r_i) v_i] = 0 \text{ for all } r_i, \quad (6)$$

$$E[u_{is}(r_i)^2] = 0.5, \quad (7)$$

$$E[u_{is}(r_i) u_{is'}(r_i')] = 0 \text{ for all } r_i, r_i' \text{ if } s \neq s', \quad (8)$$

$$E[v_i^2] = 1, \quad (9)$$

and

$$\delta_0 = \delta_1 = 1. \quad (10)$$

Whenever the number of panel waves, T , falls short of the number of possible retirement ages or if not all retirement ages are observed with sufficient frequency, even further equality constraints as to the covariance structure of the error terms have to be added.

As mentioned in the introduction, an individual is assumed to retire at time t if this is the period in which the utility level associated with the option of immediate retirement exceeds the maximum utility level that can be achieved by postponing retirement for the first time. Let \bar{r}_i denote the year of retirement that has actually been chosen by individual i (which is observable for the econometrician if it lies within the sampling period) and r_i^* be the corresponding person's optimal year of retirement (which is a discrete random variable from the econometrician's point of view because it is a function of the unknown parameters and the stochastic components of the individual's utility function). Then, the probability that this individual retires at time t given he/she has not retired until $t-1$ and conditional on v , can be expressed as follows:

$$\begin{aligned} & \Pr(r_i^* = t \mid r_i^* > t-1; v) \\ &= \prod_{r=t+1}^{r_i^*(\max)} \Phi \left[\ln Y_{it}(t) - \ln Y_{it}(r) + z_{it}' \gamma \cdot [\ln L_{it}(t) - \ln L_{it}(r)] + [\delta_{t-t_0} - \delta_{r-t_0}] \cdot v \right] \end{aligned} \quad (11)$$

Here and in the following, $\Phi(\cdot)$ denotes the cumulative distribution function and $\phi(\cdot)$ the probability density function of the standard normal distribution. It follows from equation (11) that the unconditional probability of individual i retiring at time t is

$$Pr(r_i^* = t) = \int_{-\infty}^{\infty} Pr(r_i^* = t \mid r_i^* > t-1; \nu) \cdot \prod_{s=t_{i0}}^{t-1} [1 - Pr(r_i^* = s \mid r_i^* > s-1; \nu)] \phi(\nu) d\nu \quad (12)$$

(The probability of a person retiring before t_{i0} is zero by definition).

As the term $\Delta_i := (r_i - t_{i0})$ measures the length of the time interval between the earliest possible retirement age and the one actually chosen, this model can be understood as a duration model in which time is measured in discrete intervals. Seen from this angle, the right-hand side of equation (11) equals a hazard rate (conditional on ν) with regard to the transition from work to retirement. The one-factor structure which was chosen as to the error terms allows for a large variety of correlation patterns among the hazard rates referring to different retirement ages; it can thus be seen as a solution to the problem of spurious duration dependence, which is typical of duration models in which unobserved heterogeneity is neglected.

It becomes obvious that the unknown parameters of the decision model described here can be estimated by the method of maximum likelihood. Suppose that the individuals in the sample are observed during a period of T consecutive years. In this case, there will be some persons who did not choose to retire during the sampling period although they were already entitled to retire. Taking this phenomenon (which can be thought of as a kind of right-hand-side censoring if this model is understood as a duration model in discrete time) into account, the contribution of individual i to the sample likelihood function (Λ_i) can be expressed as follows:

$$\Lambda_i = \begin{cases} Pr(r_i^* = \bar{r}_i) & \text{if } \bar{r}_i \leq T \quad (\Leftrightarrow \text{observed}) \\ \int_{-\infty}^{\infty} \prod_{s=t_{i0}}^T [1 - Pr(r_i^* = s \mid r_i^* > s-1; \nu)] \phi(\nu) d\nu & \text{if } \bar{r}_i > T \quad (\Leftrightarrow \text{unobserved}) \end{cases} \quad (13)$$

Then, the log-likelihood function for an entire sample of N individuals simply is

$$\ln \Lambda = \sum_{i=1}^N \ln \Lambda_i \quad (14)$$

and can be maximized by means of conventional numerical optimization techniques. Yet it has to be taken into consideration that the above-mentioned objective function might exhibit multiple local optima. It is therefore highly recommendable to combine standard gradient ascent methods with relatively robust steplength determination algorithms, like the "golden section" method due to BRENT (1972), when trying to optimize $\ln \Lambda(\cdot)$. Furthermore, once a (local) optimum has been found, one should examine it for neighbouring (and possibly superior) optima by means of random and/or line search methods, as described, e.g., by PINSKE

(1991). The fact, however, that even a combination of these strategies, as it was applied in the application presented here, does not necessarily lead to the detection of the global optimum with absolute certainty, remains an unresolved problem.

3. Combining Survey Data and Aggregate-Level Information

Basic Theoretical Considerations

Up to this point, it has been implicitly assumed that the sample used for estimation is the only available source of information on the economic relationship under examination. However, as census reports often provide (almost) exact information on moments of economic variables with respect to the underlying population, this need not be the case. Being based upon previous literature on the use of prior information in econometrics (e.g. THEIL and GOLDBERGER, 1961) and the combination of datasets (ARRELANO and MEGHIR, 1992, LUSARDI, 1992, and ANGRIST and KRUEGER, 1992), a recent study by IMBENS and LANCASTER (1994) has shown that, under certain conditions, "micro" and "macro" data can be combined in a Generalized Method of Moments (GMM) framework to improve parameter estimates in microeconomic models. In the first subsection of this chapter, the method introduced in the last-mentioned paper is summarized. (The notation used here differs from the one of the previous chapter in order to highlight the general applicability of the method described). The manner in which it is applied to the model introduced in chapter 2 is dealt with in the second subsection of this chapter.

Let Y and X denote random vectors with joint density $f(Y, X)$. The conditional density of Y for given X is assumed to be a function of the unknown parameter vector θ_0 . Thus we have

$$f(Y, X) = f(Y|X; \theta_0) \cdot r(X) \quad (15)$$

In this equation, $r(X)$ denotes the (unknown) density function of X . Furthermore, let

$$\{y_i, x_i\}, \quad i = 1, \dots, N \quad (16)$$

be a random sample of independent observations on Y and X . (For notational convenience, the vectors y'_1, \dots, y'_N and x'_1, \dots, x'_N will be gathered in the matrices y and x , respectively).

Provided that some very general regularity conditions (see, e.g., THEIL, 1971, ch. 8) are fulfilled, θ_0 can then be estimated consistently by the method of maximum likelihood, which amounts to maximizing the objective function

$$L_N(y, x, \tilde{\theta}) := \sum_{i=1}^N \ln f(y_i | x_i; \tilde{\theta}) \quad (17)$$

with regard to the parameter vector $\tilde{\theta}$.

Since it fulfils the necessary condition for an optimum of $L_N(\cdot)$,

$$N^{-1} \cdot \sum_{i=1}^N \frac{1}{f(y_i, x_i, \tilde{\theta})} \cdot \frac{\partial f(y_i, x_i, \tilde{\theta})}{\partial \tilde{\theta}} \stackrel{!}{=} 0, \quad (18)$$

the resulting Maximum Likelihood estimator $\hat{\theta}_N$ can also be understood as a method of moments estimator.

Now suppose that, in addition to the random sample already mentioned, the expectation κ^* of a given function $\kappa(Y, X)$ is known, for example, from corresponding compilations of aggregate-level data. ($\kappa(\cdot)$ is assumed to be a column vector of dimension $(M \times 1)$). In this case, the condition

$$E[h(Y, X)] := E[\kappa(Y, X) - \kappa^*] = 0 \quad (19)$$

holds, and, as a consequence, the following equality is valid:

$$\begin{aligned} E[h(Y, X)] &:= E_X \left[E_Y [h(Y, X) | X] \right] \\ &= E_X \left[\int h(z, X) \cdot f(z | X; \theta_0) dz \right] \\ &= E_X [\bar{g}(X, \theta_0)] = 0, \end{aligned} \quad (20)$$

with

$$\bar{g}(X, \theta_0) := \int_z h(z, X) \cdot f(z | X; \theta_0) dz$$

The condition set up in equation (19) can thus be taken into account in the estimation of θ_0 by adding the following $(M \times 1)$ vector of stochastic equality constraints:

$$\text{p lim}_{N \rightarrow \infty} \left(N^{-1} \sum_{i=1}^N \bar{g}(x_i, \tilde{\theta}) \right) = 0 \quad (21)$$

As mentioned above, the (unconstrained) maximum likelihood estimator $\hat{\theta}_N$ can also be understood as a specific GMM estimator. A constrained estimate $\hat{\theta}_N^{(c)}$ calculated under inclusion of the stochastic restriction (21) can thus be computed by minimizing the function

$$\begin{aligned} \Xi_N(y, x, \tilde{\theta}, \bar{S}) &:= \left[\xi_N(y, x, \tilde{\theta}) \right]' \cdot \bar{S} \cdot \left[\xi_N(y, x, \tilde{\theta}) \right] \\ \text{with } \xi_N(y, x, \tilde{\theta}) &:= \sum_{i=1}^N \begin{bmatrix} \left[\frac{\partial \ln f(y_i, x_i, \tilde{\theta})}{\partial \tilde{\theta}} \right] \\ \left[\bar{g}(x_i, \tilde{\theta}) \right] \end{bmatrix} \end{aligned} \quad (22)$$

with respect to $\tilde{\theta}$. In (22), \bar{S} either stands for an arbitrary, symmetric, non-stochastic and positive-definite matrix with dimension $(k+M) \times (k+M)$, or for a sequence of matrices converging almost surely to such a matrix. It follows straightforwardly from the general theory of extremum estimators (see, e.g. GOURIÉROUX and MONFORT, 1989, pp. 224-228) that the asymptotic distribution of $\hat{\theta}_N^{(c)}$ can be written as follows:

$$\sqrt{N} \left(\hat{\theta}_N^{(c)}(\bar{S}) - \theta_0 \right) \xrightarrow{d} N(0, \Omega(\bar{S}))$$

with

$$\begin{aligned} \Omega(\bar{S}) &:= \left[\left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \tilde{\theta}} \right]' \cdot \bar{S} \cdot \left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \tilde{\theta}} \right] \right]^{-1} \\ &\quad \cdot \left[\left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \tilde{\theta}} \right]' \cdot \bar{S} \cdot V(\xi(y, x, \theta_0)) \cdot \bar{S} \cdot \left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \tilde{\theta}} \right] \right] \\ &\quad \cdot \left[\left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \tilde{\theta}} \right]' \cdot \bar{S} \cdot \left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \tilde{\theta}} \right] \right]^{-1} \end{aligned} \quad (23)$$

If \bar{S} were set to

$$S^* := V(\xi(y, x, \theta_0))^{-1}, \quad (24)$$

or to a sequence of matrices converging almost surely to $V(\xi(y, x, \theta_0))^{-1}$, the asymptotic covariance matrix of $\hat{\theta}_N^{(c)}(S^*)$ would simplify to

$$\Omega(S^*) := \left[\left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \theta} \right]' \cdot V(\xi(y, x, \theta_0)) \cdot \left[E \frac{\partial \xi(y, x, \theta_0)}{\partial \theta} \right] \right]^{-1} \quad (25)$$

For any other symmetric, positive-definite matrix $\tilde{S} \neq S^*$, the difference $\Omega(\tilde{S}) - \Omega(S^*)$ is a positive-semidefinite matrix (for a proof, see, e.g., DAVIDSON and MACKINNON, 1993, p. 597). $\hat{\theta}_N^{(c)}(S^*)$ is thus termed the optimal GMM estimator of θ_0 . Yet as S^* itself is a function of the unknown vector θ_0 , such an optimal GMM estimation is infeasible. However, as any matrix (or sequence of matrices) that meets the above-mentioned requirements can be used to obtain consistent estimates of θ_0 , one can utilize a two-step estimation procedure to obtain an asymptotically optimal GMM estimator. The first step consists of finding a consistent (but asymptotically inefficient) estimate $\hat{\theta}_N^{(1)}$, e.g. by setting \tilde{S} to an identity matrix of dimension $(k+M)$, or by minimizing the criterion function

$$\Xi_N^{(1)}(y, x, \tilde{\theta}, \cdot) := -\frac{1}{N} \left[\sum_{i=1}^N \ln f(y_i, x_i, \tilde{\theta}) + \sum_{i=1}^N \bar{g}(x_i, \tilde{\theta})' \cdot \bar{g}(x_i, \tilde{\theta}) \right] \quad (26)$$

with respect to $\tilde{\theta}$. The resulting first-stage estimate $\hat{\theta}_N^{(1)}$ can then be used to estimate the covariance matrix of $\xi(\cdot, \theta_0)$ by

$$\hat{V} := \frac{1}{N} \cdot \sum_{i=1}^N \xi(y, x, \hat{\theta}_N^{(1)}) \cdot \xi(y, x, \hat{\theta}_N^{(1)})' \quad (27)$$

In the second stage of the estimation procedure, the expression \tilde{S} in equation (22) is either set to \hat{V}^{-1} or to a sequence of matrices \hat{S} of which the limiting value converges in probability to S^* as N tends to infinity. The resulting objective function $\Xi_N(y, x, \tilde{\theta}, \hat{V}^{-1})$ or $\Xi_N(y, x, \tilde{\theta}, \hat{S})$ is then, again, minimized as to $\tilde{\theta}$. The limiting distribution of the resulting estimator $\hat{\theta}_N^{(c)}(\hat{V}^{-1})$ or $\hat{\theta}_N^{(c)}(\hat{S})$ for $N \rightarrow \infty$ equals the one of the optimal GMM estimator $\hat{\theta}_N^{(c)}(S^*)$ in equation (25). If $\xi_N(\cdot)$ is defined as in (22), S^* can be written as

$$S^* = \left[\begin{array}{cc} E \left[\left(\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right) \left(\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right)' \right] & E \left[\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \cdot \bar{g}(x, \theta_0)' \right] \\ E \left[\bar{g}(x, \theta_0) \cdot \frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right] & E \left[\bar{g}(x, \theta_0) \cdot \bar{g}(x, \theta_0)' \right] \end{array} \right]^{-1} \quad (28)$$

As to the expression $E \left[\bar{g}(x, \theta_0) \cdot \frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right]$ one obtains

$$\begin{aligned}
& E \left[\bar{g}(x, \theta_0) \cdot \frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right] \\
&= E_x \left[E_{y|x} \left(\bar{g}(x, \theta_0) \cdot \frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right) \right] \\
&= E_x \left[\bar{g}(x, \theta_0) \cdot E_{y|x} \left(\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right) \right] \tag{29} \\
&= 0 \\
&\text{because the equality } E_{y|x} \left(\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right) = 0 \text{ holds by assumption.}
\end{aligned}$$

Equation (28) thus simplifies to

$$S^* := \begin{bmatrix} E \left[\left(\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right) \left(\frac{\partial \ln f(y, x, \theta_0)}{\partial \theta} \right)' \right]^{-1} & 0 \\ 0 & E[\bar{g}(x, \theta_0) \cdot \bar{g}(x, \theta_0)']^{-1} \end{bmatrix} \tag{30}$$

Because of equations (24) and (25), the covariance matrix of the optimal GMM estimator can be rewritten as

$$\Omega(S^*) = \left[-E \frac{\partial^2 \ln f(y, x, \theta_0)}{\partial \tilde{\theta} \partial \tilde{\theta}'} + E \left(\frac{\partial \bar{g}(x, \theta_0)}{\partial \tilde{\theta}} \cdot \bar{g}(x, \theta_0) \cdot \bar{g}(x, \theta_0)' \cdot \frac{\partial \bar{g}(x, \theta_0)}{\partial \tilde{\theta}'} \right) \right]^{-1} \tag{31}$$

It is obvious that the difference between the asymptotic covariance matrix of the unconstrained maximum likelihood estimator and $\Omega(S^*)$ is a positive-definite matrix. Thus it can be concluded that the inclusion of macroeconomic information in microeconomic models increases the precision of the parameter estimates obtained.

Implementation for the Retirement Model

The estimation procedure just described can be applied directly to the model introduced in section 2 if the econometrician has access to some information about age-specific retirement probabilities or labour force participation rates within the population from which the sample used for information constitutes a random draw. Fortunately, this condition is fulfilled to a certain degree in the case studied here. One important source of macro level information in this context are the tabulations of the age distribution of the individuals receiving old-age pensions for the first time, which are annually published by the Association of Trustees to the Public

Pension System in Germany (Verband Deutscher Rentenversicherungsträger, VDR). As the date at which a person first receives old age pension benefits (virtually) always follows the one at which he or she left the labour force, these tabulations can be used to calculate age-specific "survival" (=non-retirement) probabilities for males.

With females, things are considerably more complicated. Due to certain peculiarities of the German Pension system regarding the valuation of pre-retirement years of work and childrearing, it cannot be taken for granted that women who left the labour force before their sixtieth birthday are entitled to receive old-age pension benefits from that date onwards. It would thus be at least potentially misleading to use the VDR's statistics regarding the age distribution of females when first receiving pension benefits as a basis for the computation of age-specific retirement probabilities. However, the labour force participation rates published by Germany's Federal Bureau of Statistics, are ordered by age categories and can therefore be utilized in this context.

A major problem associated with the combination of individual- and aggregate-level data in microeconomic models is that the "micro" dataset in use is not necessarily representative for the population as a whole. As IMBENS and LANCASTER (1994, p. 663) point out, erroneously assuming that the individual-level information in use is compatible with the corresponding macro data is a form of misspecification and can lead to incorrect inference. The authors suggest a simple test procedure based on the central limit theorem to examine whether micro and macro data are compatible. Using the notation introduced above, the corresponding test statistic becomes

$$\tau_c := N \cdot (\bar{\kappa} - \kappa^*)' \cdot \hat{V}_\kappa^{-1} \cdot (\bar{\kappa} - \kappa^*)$$

with

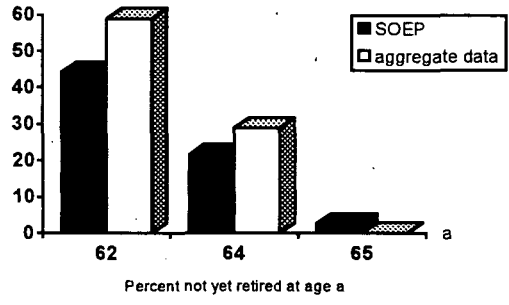
$$\bar{\kappa} := N^{-1} \cdot \sum_{i=1}^N \kappa(y_i, x_i) \quad (32)$$

and

$$\hat{V}_\kappa := N^{-1} \cdot \sum_{i=1}^N (\kappa(y_i, x_i) - \bar{\kappa}) \cdot (\kappa(y_i, x_i) - \bar{\kappa})'$$

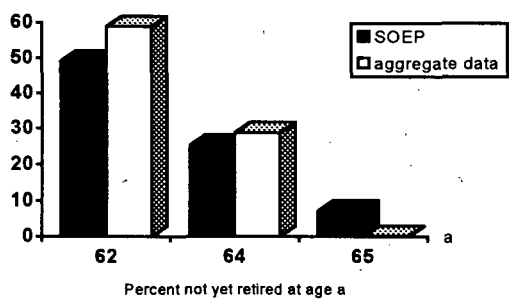
In principle, the test described above consists essentially of a comparison of sample averages and population means. As N goes to infinity, the asymptotic distribution of τ_c is $\chi^2(\dim(\kappa^*))$. In the case studied here, the hypothesis that the sources of micro and macro data used are compatible is not rejected for any of the two subsamples (one of them including and one excluding foreigners) used for estimation. The results of a comparison of micro and macro information for the retirement model are summarized in figures 1 to 3.

Figure 1
**Comparison of Micro and Macro Level Information
 for West Germany
 (German Males only)**



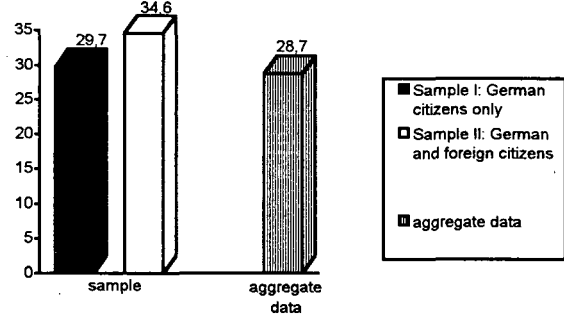
χ^2 test statistic (d. f. = 3) = 0.52198
 p - value: 0.9140

Figure 2
**Comparison of Micro and Macro Level Information
 for West Germany
 (German and Foreign Males)**



χ^2 test statistic (d. f. = 3) = 1.1465
 p - value: 0.7659

Figure 3
**Comparison of Micro and Macro Level Information
 for Females in West Germany**



(labour force participation rates of females previously employed;
 age group: 60 to 64)

	χ^2 test statistic (d. f. = 1)	p-value
German females only	$3.979 \cdot 10^{-4}$	0.98408
German and foreign females	0.03676	0.84796

4. The Data

The definition of "retirement"

In previous micro-econometric studies of retirement behaviour, a variety of different definitions of the endogenous variable have been used. To a certain extent, this is due to the fact that the questions of main interest within this field vary considerably among researchers. Another reason is that the informational content of the datasets in use differs from case to case. The striking differences among the results of previous empirical investigations concerning the labour force exit behaviour of elderly Americans can, in part, be attributed to these factors. HURD and BOSKIN (1984), for instance, who use the "Retirement History Survey" (RHS) as a data base, only consider a person as retired if he or she has left the labour force finally and irrevocably. Contrary to that, in a study by BURTLESS and MOFFITT (1984), which is based on PSID data, an elderly employee's retirement age is defined as the age in which a sudden and pronounced decline in the individual's number of working hours occurs. If the incentive effects of firm pension plans are being analyzed, it is most appropriate to equate a person's date of retirement with the period when he or she first receives pension benefits, as it is done by STOCK and WISE (1990).

In an empirical study of retirement behaviour in West Germany conducted by BÖRSCH-SUPAN (1992) using the 1984 cross-section of the SOEP, the period of retirement is defined as the year in which the number of a person's working hours per week first falls short of fifteen. However, as "retirement" is understood as an absorbing state in the model described in section 2, it seems more reasonable in this context to equate a person's date of retirement to the moment in which he/she first declares himself/herself retired. This is possible with data from SOEP because it contains a very detailed set of information on a person's present and past labour force status in which "retirement" is a separate category.

One of the main goals of the empirical study described here is to explain the effect of Germany's pension system on retirement behaviour. Nevertheless, it is assumed here that individuals can opt for retirement up to two years before they reach the age of 60, which is the minimum age for the receipt of old-age pension benefits in this system. The reason for this assumption is that, under the rules of the German system of unemployment insurance, unemployed individuals who are older than 58 can receive full unemployment benefits without having to declare explicitly that they are willing to work as long as they are ready to apply for the receipt of pension benefits at the earliest possible age (see BUNDESMINISTER FÜR ARBEIT UND SOZIALORDNUNG, 1993, p. 111). As an evaluation of the labour force histories of all SOEP respondents born between 1915 and 1925 reveals, less than 3% of all persons in this category who were either unemployed or out of the labour force at the age of 58 returned to

employment afterwards. This indicates that setting the minimum retirement age ("retirement" being defined as a state in which a person does not intend to return to paid employment) to 58 is a reasonable choice.

Exogenous variables

The intensity of an individual's preference for leisure, which is measured by the parameter β within the framework of the model described in chapter 2, is assumed to be influenced mainly by the person's current health status and educational background. In addition to these variables, the dummy variable PUBLIC, which indicates whether a person is or was a public sector employee, is included in order to control for the influence of possible differences between the pension systems applying to public and private sector employees, respectively. From a psychological point of view, it might be interesting to find out whether married persons exhibit a higher preference for leisure than others; therefore, a corresponding dummy variable (MARRIED) is taken account of. Finally, in order to examine in how far macroeconomic conditions on the labour market influence retirement behaviour, the gender-specific unemployment rate in the corresponding age group, UR, is included in the set of regressors. Definitions and descriptive statistics of the explanatory variables used are summarized in tables 1 and 2.

Table 1.: Definitions of explanatory variables in β_{it}

D_DISAB	Degree of disability with regard to the requirements of the corresponding person's job. $0 \leq D_DISAB \leq 1$.
PUBLIC	Dummy variable; PUBLIC = 1 if a person is or was formerly employed in the public sector.
CH_MAL	Dummy variable; CH_MAL = 1 if a person suffers from a chronic malady.
UR	Gender-specific unemployment rate in the age group a person belongs to.
FOREIGN	Dummy variable; FOREIGN = 1 if a person is not a German citizen.
MARRIED	Dummy variable indicating whether a person is married.
HAUPTS	Dummy variables for the highest grade in general education achieved;
REALS	HAUPTS = 1 corresponds to 9 years, REALS to 10 years, and ABI
ABI	to 13 years of schooling.
UNI	Dummy variable; UNI = 1 if a person has a university degree.
FEMALE	Dummy variable for gender.

Table 2. : Descriptive statistics of the exogenous variables in β_{it}

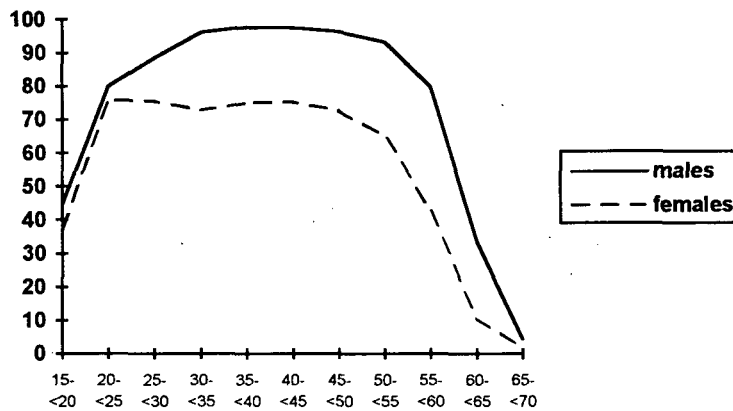
Variable	Mean	Standard Deviation	Minimum	Maximum
D_DISAB	0.1111	0.2475	0.0000	1.0000
PUBLIC	0.1462	0.3533	0.0000	1.0000
CH_MAL	0.3956	0.4890	0.0000	1.0000
UR	11.7500	2.0700	8.2300	17.5300
FOREIGN	0.2811	0.4498	0.0000	1.0000
MARRIED	0.6495	0.4772	0.0000	1.0000
HAUPTS	0.3533	0.4780	0.0000	1.0000
REALS	0.0856	0.2797	0.0000	1.0000
ABI	0.0578	0.2333	0.0000	1.0000
UNI	0.0389	0.1933	0.0000	1.0000
FEMALE	0.3011	0.4590	0.0000	1.0000

Selection of Sample

In order to implement the model described in section 2, it is necessary to confine the sample to the persons whose earliest possible year of retirement lies within the sampling period. This implies that in case described here only individuals born between 1926 and 1933 were included in the dataset used for estimation. As the pension systems that apply to most of the self-employed in Germany differ significantly from the one which is relevant for employees, persons who report to have been self-employed in one of the years of the the sampling period were excluded from the sample. The same applies to individuals whose statements about their labour force status were found to be self-contradictory or incomplete.

The significant differences between the labour force participation patterns of men and women (see figure 1 for a comparison) make it difficult to decide in how far female employees should be included in the sample.

Figure 1
Age-specific labour force participation rates of males and females in West Germany, 1991



As the model described in chapter 2 only applies to individuals who did not already leave the labour force a long time before the assumed minimum retirement age, it would clearly be misleading to draw all women from the original dataset regardless of their past employment histories into the subsample used for estimation. On the other hand, the rather steady growth in female labour force participation that West Germany has experienced during the past three decades suggests that it would be equally inexpedient to exempt all women from the analysis conducted here. The compromise advocated in this paper is to include only those females in the

relevant subsample who reported to have participated in the labour force for at least two years after their 55th birthday.

Income forecasts

Individual-specific forecasts of net labour income levels were computed on the basis of a classical, MINCER (1974)-type earnings function. In order to avoid the problem of a possible selectivity bias in the earnings estimates, the model was extended to a "type-two-Tobit" model (see AMEMIYA, 1985, p. 385-389) by including a separate selection equation. The results, in turn, can be combined to panel estimates by means of a minimum distance method described by GOURIÉROUX and MONFORT (1989, pp. 385-387). In the particular case discussed here, this was done separately for balanced panels of 1.633 male and 1.985 female SOEP respondents born between 1925 and 1959. The high degree of variation in the number of working hours which can be observed among female employees was taken into account by using the log of the hourly net wage rate as the left-hand side variable in the earnings equation for women, and by specifying a selection equation which allows for three separate categories (non-employment, part-time and full-time work) among females. The corresponding estimation results and descriptive statistics of the variables used can be obtained from the author on request.

Because of the complexity of the German pension law, the computation of hypothetical pension benefits (b_{it}) for persons who did not retire during the sampling period proved to be a rather difficult task. In Germany, an individual's pension level is a function of both the person's number of years of service and of all past and present labour income streams. (For details, see, e.g., CASMIR, 1990). Fortunately, the SOEP contains rather detailed information on the labour force histories of the individuals, so that at least the number of years in which a person paid social security contributions could be computed with adequate accuracy. Together with the estimation results from the earnings equation, these estimates made it possible to evaluate a person's claims on the pension system at least in an approximate manner. (The mean absolute forecast error of this approximation with regard to the monthly pension income levels of persons already retired proved to be less than DM 100).

5. Empirical Results

The model set up in chapter 2 was estimated on two samples. Sample I consists only of German citizens, whereas in sample II, foreigners are included as well. It cannot be told unequivocally which of these two datasets leads to more reliable results when being used for estimation: On one hand, as immigrants are deliberately oversampled in the SOEP database, including foreigners raises the sample size by almost one third. Yet on the other hand, many immigrants can be reasonably assumed to have claims on the pension systems of their home countries, which are likely to differ enormously both among nations and individuals. As the estimates obtained do not permit any conclusive inferences about which kind of sample selection is more adequate, the corresponding results will simply be placed side by side here.

The estimation procedure for combining micro and macro data described in chapter 3 raises the question in how far unconstrained maximum likelihood estimates and GMM estimates calculated under inclusion of the stochastic constraints described in equation (21) differ in magnitude and precision. That is why the results from the ML estimation as well as of both steps of the two-step procedure described in section 3 are shown in this paper (see table 3 and tables A.1 and A.2 in the appendix). It can be seen that, while the qualitative implications of the estimates obtained remain the same for virtually all the explanatory variables under consideration, some of the estimated coefficients (especially the constant term and the one referring to the age- and gender-specific unemployment rate UR) differ considerably in absolute value. This should, however, not be regarded as too surprising since the size of each of the two samples in use is small compared to the population of some 4 million potential new retirees in the 1983 to 1990 period. The efficiency gains resulting from the combination of "micro" and "macro" information in the context of this model are considerable, as a comparison of the asymptotic standard errors of the coefficient estimates in tables 3 and A.1 makes obvious. The following discussion of the estimation results will therefore, unless indicated otherwise, refer to the estimates obtained from the second stage of the procedure described in chapter 3.

The estimation results obtained provide strong evidence for the importance of a person's health status for retirement behaviour. The coefficients referring to the two health-related variables in the dataset, "degree of disability" and "presence of chronic complaints", both bear the expected positive sign, indicating that, on average, a bad health status increases the probability of early retirement. While being far from surprising, this finding has an important implication: If raising the average retirement age is a goal of public policy, measures directed at improving employees' health and removing working conditions that are potentially hazardous in this respect can be expected to be highly effective in this context. (Only for the sake of completeness, though, it should be mentioned that the increase in longevity brought about by

an improvement of individuals' health might limit the financial relief this involves for the prevailing pension system).

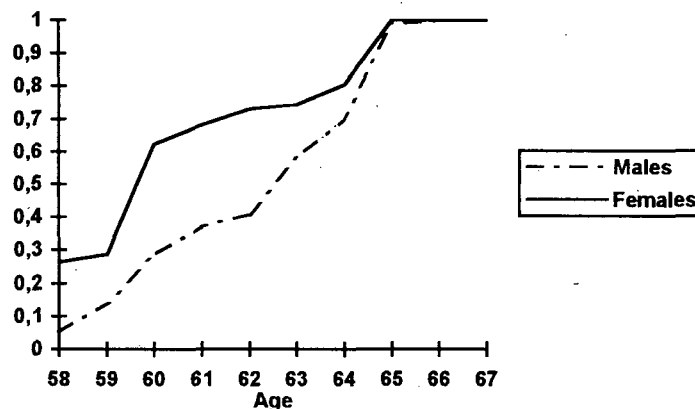
Table 3:
Parameter Estimates for combined micro and macro information
 (Second-stage estimates; GMM with asymptotically optimal weighting matrix)

Variable	Sample I: German citizens		Sample II: Germans and foreign citizens	
	Estimate	asy. t-statistic	Estimate	asy. t-statistic
CONSTANT	16.6093	10.163	10.1065	6.638
D_DISAB	3.7129	4.876	6.0105	7.568
HAUPTS	2.5526	5.980	2.4764	6.652
REALS	2.4794	4.400	2.8880	5.465
ABI	-0.2882	-0.324	-0.1113	-0.094
UNI	3.6638	2.799	3.5811	2.666
MARRIED	-0.0951	-0.233	0.8294	2.463
PUBLIC	-2.8973	-5.753	-3.1625	-5.886
CH_MAL	0.6774	2.076	1.2117	4.353
FEMALE	10.5907	16.393	9.1611	20.599
FOREIGN	-	-	1.0942	2.860
UR	-1.2820	-7.582	-0.7341	-5.048
δ_2	-3.2369	-2.573	-4.2536	-4.143
δ_3	1.1200	5.731	-3.2017	-3.146
δ_4	1.7501	13.087	1.6803	15.627
δ_5	2.1948	19.109	2.1232	22.824
δ_6	-4.5677	-3.671	-5.7966	-5.592
Mean $\hat{\beta}$	6.5296		7.3489	
Mean Log Likelihood (constrained)	-1.3122		-1.2854	
# of cases	635		900	

Another important result is that, on average, foreigners have a higher propensity to retire early than their German colleagues. A plausible explanation for this phenomenon can be found in an empirical study by STEINER and VELLING (1993), who find that persons who wish to re-migrate to their country of origin often do so within a short time after retirement, which suggests that the motive of return migration can be expected to lead foreigners to retire earlier than would otherwise be expected. It should be noted, however, that at least some of the elderly "guest workers" currently living in Germany have no intention to leave their "host country" after having retired. This gives rise to the supposition that the probable presence of discrimination against foreigners on the labour market, which can result in a higher disutility of labour or in untimely retirement following involuntary unemployment, can also contribute to an explanation of the above-mentioned finding.

Furthermore, the results obtained with both subsamples in use indicate that females formerly at work are much more likely to opt for early retirement than males. In fact, as figure 4 demonstrates, the estimated probability of retiring at the earliest possible age is almost five times as high for women as it is for men. The obvious explanation is that women in employment very often have to carry a twofold burden of paid work on one hand and domestic work and childrearing on the other, which probably entails, on average, an earlier exhaustion of the physical ability to meet the demands of one's respective job.

Figure 4 Estimated cumulative retirement probabilities



Interpreting the coefficients referring to the effect of educational achievement turns out to be a rather difficult task. This can be attributed to the fact that due to the lack of sufficiently reliable information on individual asset holdings in the database in use, these parameters capture two

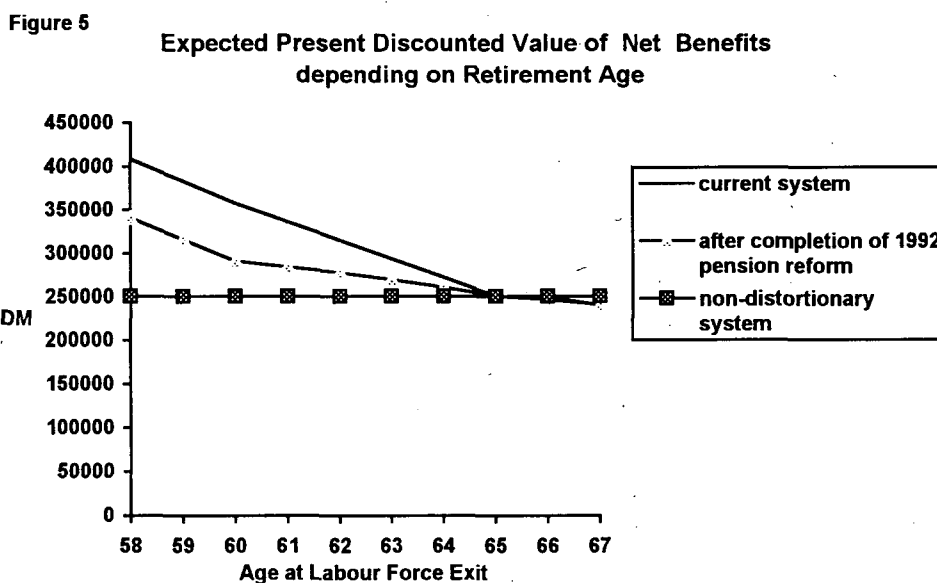
effects simultaneously: On one hand, a person's educational status can reasonably be expected to be positively correlated with wealth, which, from a life-cycle point of view, would imply that highly educated people tend to leave the labour force earlier than others because they can afford to. On the other, individuals with a relatively low degree of educational achievement are more likely to be laid off when (or even before) having reached the minimum retirement age. The estimation results obtained do not permit an unequivocal conclusion as to which of these two effects dominates the other.

Unobserved wealth effects may also be one of the reasons why the preference for early retirement appears to be significantly below average among public sector employees, who are, on average, paid significantly lower wages than their private sector colleagues. Another possible explanation for this phenomenon is that, for the vast majority of the public sector employees in Germany, the probability of involuntary early retirement following a dismissal by the employer is virtually reduced to zero by the relevant legal regulations. But, of course, also readers who simply believe the average disutility of labour to be considerably lower in the public sector than in the private one can regard this result as a confirmation of their verdict.

What appears to be highly implausible at first sight is the estimate obtained for the coefficient belonging to the age- and gender-specific unemployment rate, UR. One would normally expect that untimely retirement resulting from layoffs occurs more often in times of high unemployment, and thus assume that the corresponding coefficient takes a positive value. It has to be kept in mind, though, that unemployment may influence the propensity to retire through a second channel, as well: In times of high unemployment, people still at work probably tend to feel more pessimistic about future income prospects and therefore are more inclined to postpone their exit from the labour force. (In Germany, the annual growth rate of pension benefits is linked to the one of wages through an adjustment mechanism arranged accordingly). Thus, one possible explanation for the negative sign of the estimated coefficient referring to UR is that the latter of these two effects over-compensates the former on total. What also might help to explain this finding is that, due to a number of peculiarities of the German labour market legislation, by no means all cases in which persons had to retire before the age of sixty following a dismissal by their employer are registered in the official unemployment figures. For example, one possible "exit route" frequently chosen by people with health-related problems who would otherwise become unemployed (as it is defined in the official statistics) is to apply for disability pensions, which, according to ZIMMERMANN (1992), are much more generously granted in times when (re-)employment prospects are dull (see RIPHAHN, 1994, for a thorough theoretical and empirical analysis of disability retirement). Another means by which individuals aged 58 or 59 could be "bought off" the labour market was the so-called "Precipitous Retirement Act", which was in effect from 1985 to 1988 in West Germany and prescribed that under certain conditions, individuals laid off as a

consequence of personnel reductions in their firms could claim a specific early retirement allowance until being entitled to regular old age pensions. (A more detailed summary of the legal framework is provided by LAMPERT, 1988).

Being deduced from a micro-economic decision model in which the individuals are assumed to balance the gain in income attainable by postponing retirement against the increase in leisure implied by opting for an early exit from the labour force, the results of the investigation presented here can be used as a basis for simulation experiments aimed at forecasting the manner in which assumed changes of the current pension system affect the frequency distribution of retirement ages. This is done in the following for two different cases. The first experiment relates to Germany's most recent Pension Reform Act, which was enacted on January 1, 1992 but of which the main constituents will only become effective after the turn of the century. One of the main goals of this new law is to reduce the financial attractiveness of early retirement by introducing an extra deduction of 0.3% from annual pension benefits for every month that lies between the date when a person first receives them and his/her 65th birthday. The second policy scenario of which the simulated outcomes are presented here is the introduction of a hypothetical, "non-distortionary" social security system in which the expected present discounted value of net transfer payments (pensions plus unemployment benefits minus taxes) is independent of the retirement age chosen. In figure 5, the pension system prevailing during the sampling period is contrasted with the one coming into force after the completion of the 1992 pension reform and with the supposed nondistortionary system. This is done by comparing the age-specific expected present discounted values of the net transfer payments, computed under the assumptions of chapter 2 for a person aged 58 and with average earnings, corresponding to each of these three systems.



If, as it is frequently done, the expected present discounted value of net benefits is understood as a measure for the relative financial attractiveness of different possible retirement ages, one can conclude from this comparison that the financial incentives in favor of early retirement which have been typical of the prevailing pension system will be significantly reduced, but not completely eliminated by the 1992 Pension Reform Act.

By replacing the original values of the income variables $Y_{it}(r_i)$ with the ones corresponding to the two alternative policy régimes studied here and subsequently re-estimating the age-specific retirement probabilities on the basis of these newly generated arguments, an estimate of the effect these policies can be obtained. In figures 6 and 7, the results of such an experiment are summarized.

Figure 6
Estimated Impact of Policy Changes on Cumulative Retirement Probabilities (Males only)

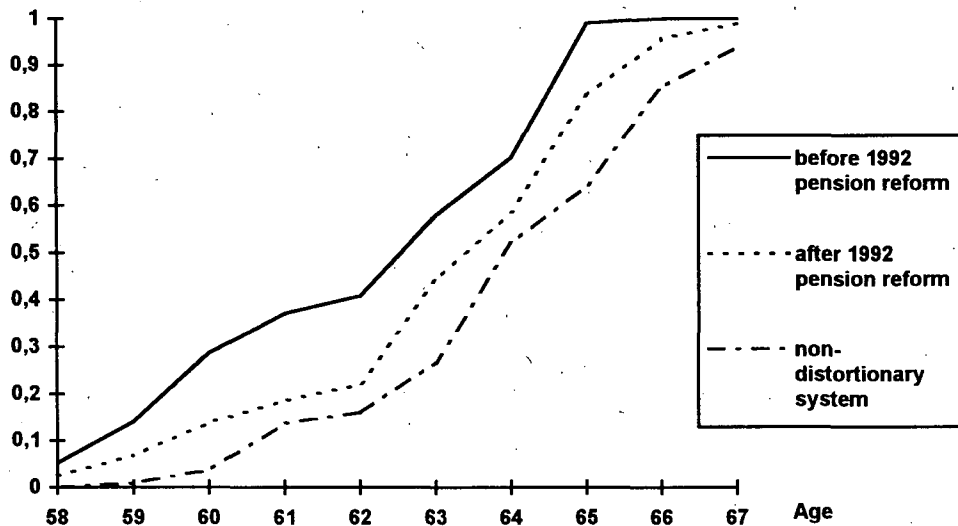
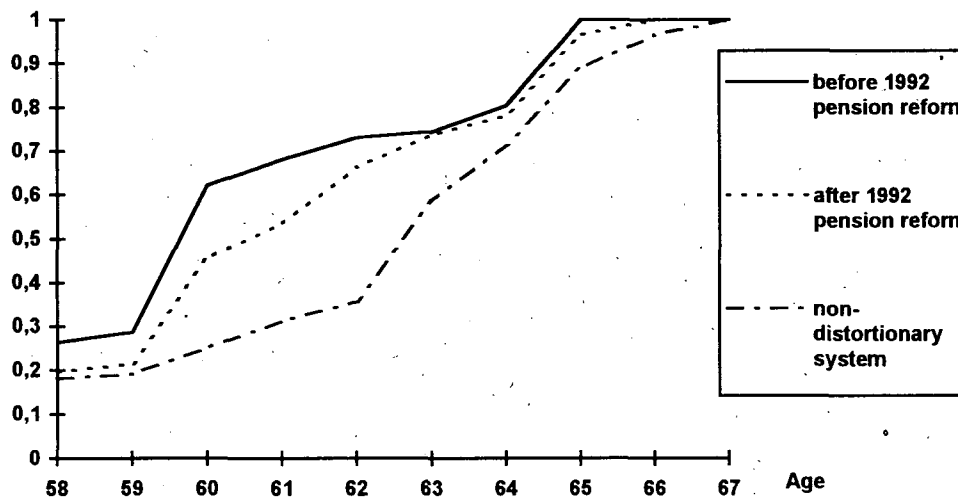


Figure 7
Estimated Impact of Policy Changes on Cumulative Retirement Probabilities (Females only)



Changes in the incentive structure of the underlying pension scheme can obviously exert considerable influence on the age-specific retirement probabilities. A replacement of the current German pension scheme by the non-distortionary system described above raises the estimated average retirement age by 1.9 years for men and 1.7 years for women previously at work (see table 3 below). The question in how far such an increase might help to compensate the additional financial burden levied on Germany's social security system by the process of population ageing that is expected to take place during the next five decades, however, cannot be answered at this stage.

	Males	Females
Current System	62.5	60.9
after 1992 Pension Reform	63.5	61.5
Non-Distortionary System	64.4	62.6

A word of caution ought to be said with regard to the simulation exercise of which the results are gathered here: As can be concluded from LUCAS' (1976) critique of econometric policy evaluation, the validity of its outcomes depends critically on the correctness of the assumption that the true parameters of the decision models set up in section 2 are independent of the rules governing the current labour market and pensions policy. This problem undoubtedly is of high importance, yet there does not appear to be a manifest solution to it.

6. Summary and Conclusions

In this paper, the job exit behaviour of elderly West German employees was analyzed empirically using a sequential, multi-period decision model for multiple discrete alternatives. In order to improve the precision of the estimates obtained, the information contained in the panel dataset in use was combined with aggregate-level information on the frequency distribution of retirement ages and on female labour force participation using the approach developed by IMBENS and LANCASTER (1994). The estimation results reveal that health-related factors play a key role in determining the age a person chooses for the transition from work to retirement.

Education, too, does matter, but the influence a person's educational status has on the probability of an early exit from the labour force appears to follow a rather complicated, non-linear pattern. The supposition that, because of the re-migration motive and - possibly - the presence of discrimination problems, elderly foreigners have, on average, a tendency to quit the labour force earlier than their German colleagues is also strongly supported by the estimates. A policy-related simulation experiment reveals that a removal of the financial incentives for early retirement implied by the current German pension scheme would probably result in a significant increase in the average retirement age.

The main advantage of the approach presented in this paper is that it adequately reflects the sequential nature of the retirement decision and the multiplicity of possible retirement ages, while also taking account of the possibility of unobserved heterogeneity among individuals. One major point of criticism, of course, is that its relative computational simplicity necessitates the introduction of rather restrictive assumptions concerning the utility maximization framework it is based upon. Another drawback might be that it does not explicitly account for the possibility of rationing on the labour market. In spite of the fact that it is doubtful whether the informational content of the database used here is high enough to solve this problem in this particular case, the author believes it is the last-mentioned aspect that deserves most attention in future empirical research on retirement behaviour.

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Appendix:

Table A.1: Unconstrained Parameter Estimates

<i>Variable</i>	<i>Sample I: German citizens only</i>		<i>Sample II: Germans and foreigners</i>	
	Estimate	asy. t-statistic	Estimate	asy. t-statistic
CONSTANT	4.3067	2.800	2.4201	2.018
D_DISAB	3.0667	4.068	4.2987	5.496
HAUPTS	2.4804	7.243	1.6801	4.841
REALS	2.3772	5.147	1.9139	4.069
ABI	1.2423	1.156	0.8655	1.012
UNI	2.2242	1.860	1.3233	1.420
MARRIED	-0.2635	-0.776	0.2667	1.004
PUBLIC	-2.3514	-5.728	-2.1325	-6.803
CH_MAL	1.2210	4.294	1.2962	4.800
FEMALE	2.7295	4.723	2.0047	4.074
FOREIGN	-	-	0.7346	2.704
UR	-0.2665	-1.967	-0.0306	-0.290
δ_2	-0.6621	-1.424	-0.5448	-1.322
δ_3	0.1535	-0.349	-0.0397	-0.100
δ_4	0.3071	0.568	0.2157	0.515
δ_5	1.1829	8.169	1.1401	8.894
δ_6	0.3082	0.436	0.0381	0.061
Mean $\hat{\beta}$	3.8780		4.7938	
Mean Log Likelihood	-1.07091		-1.15221	
# of cases	635		900	

Table A.2:
Parameter Estimates for combined micro and macro information
 (First-stage estimates; GMM with sub-optimal weighting matrix)

<i>Variable</i>	<i>Sample I:</i> <i>German citizens</i>		<i>Sample II:</i> <i>Germans and foreign citizens</i>	
	Estimate	asy. t-statistic	Estimate	asy. t-statistic
CONSTANT	9.1652	7.320	4.8372	4.372
D_DISAB	3.5872	4.256	4.7997	5.921
HAUPTS	2.4816	5.928	1.8282	5.737
REALS	2.3396	4.063	1.9821	3.722
ABI	0.9129	1.109	0.7461	0.837
UNI	2.5919	2.200	1.6120	1.510
MARRIED	-0.2142	-0.522	0.3806	1.393
PUBLIC	-2.5036	-5.088	-2.2724	-4.774
CH_MAL	1.0453	3.191	1.1908	4.382
FEMALE	5.3599	10.652	3.9809	10.603
FOREIGN	-	-	0.7908	2.545
UR	-0.6883	-5.686	-0.2691	-2.663
δ_2	-1.3844	-2.448	-0.8278	-2.236
δ_3	-0.7433	-1.395	-0.2551	-0.702
δ_4	1.1083	7.614	-0.1502	-0.420
δ_5	1.4012	11.219	1.3682	12.565
δ_6	-1.3724	-2.245	-0.8061	-1.843
Mean $\hat{\beta}$	4.5752		5.1727	
Mean Log Likelihood (constrained)	-1.1043		-1.1672	
# of cases	635		900	