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INCENTIVE EFFECTS OF SOCIAL SECURITY ON LABOR FORCE PARTICIPATION: EVIDENCE IN GERMANY AND ACROSS EUROPE

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Participation: Evidence in Germany and Across Europe
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ABSTRACT

All across Europe, old age labor force participation has declined dramatically during the last decades. This secular trend coincides with population aging. The European social security systems therefore face a double threat: retirees receive pensions for a longer time while there are less workers per retiree to shoulder the financial burden of the pension systems.

This paper shows that a significant part of this problem is homemade: most European pension systems provide strong incentives to retire early. The correlation between the force of these incentives with old age labor force participation is strongly negative. The paper provides qualitative and econometric evidence for the strength of the incentive effects on old age labor supply across Europe and for the German public pension program.

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INCENTIVE EFFECTS OF SOCIAL SECURITY ON LABOR FORCE PARTICIPATION: EVIDENCE IN GERMANY AND ACROSS EUROPE

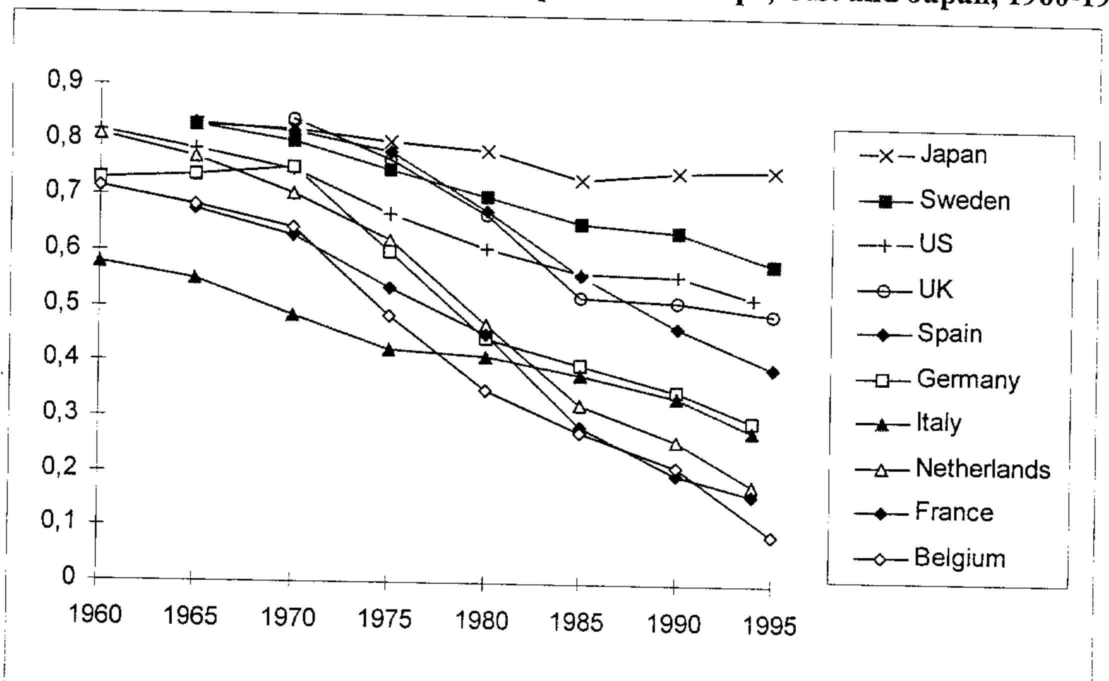
by Axel H. Börsch-Supan

1. Introduction

Pay-as-you-go (PAYG) systems dominate the old age social security programs in Europe. French, German and Italian workers rely almost exclusively on pensions financed by a PAYG system, while the Netherlands and Great Britain are exceptions in funding a substantial fraction of their future pension benefits. In the Netherlands, defined benefit plans dominate. As has been stressed quite frequently (OECD, 1988), the population aging process puts all pension systems under severe strain, but particularly defined benefit plans and most dramatically the pay-as-you-go pension systems.

At the same time, European old age labor force participation rates have declined substantially and are relatively low compared to the United States and Japan, see Figure 1.

Figure 1: Old Age Labor Force Participation in Europe, U.S. and Japan, 1960-1995

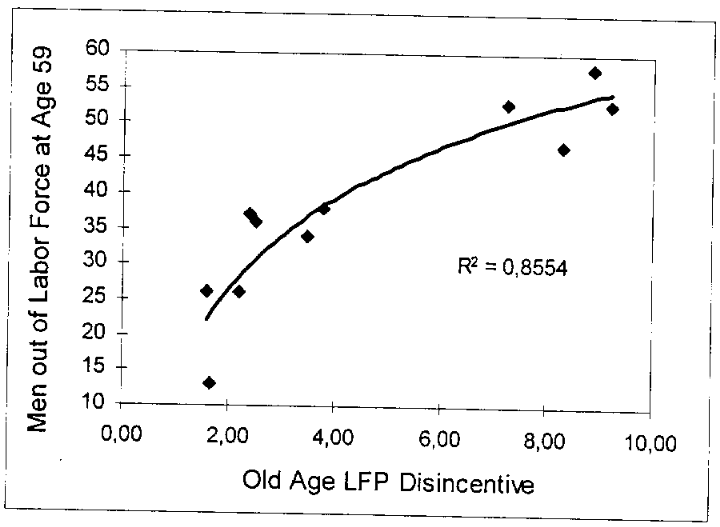


Note: Labor force participation rates of men aged 60-64. **Sources:** Belgium: Pestieau and Stjins (1998); France: Blanchet and Pelé (1998); Germany: Börsch-Supan and Schnabel (1998b); Italy: Brugiavini (1998); Japan: Oshio and Yashiro (1998); Netherlands: Kapteyn and de Vos (1998); Spain: Boldrin, Jimenez and Peracchi (1998); Sweden: Palme and Svensson (1998); UK: Blundell and Johnson (1998); U.S.: Diamond and Gruber (1998).

The decline in old age labor force participation is very pronounced in Belgium, Italy, France, the Netherlands, and Germany. All of the European countries in Figure 1 have old age labor force participation rates below Japan, and only Sweden is above the U.S. The decline in old age labor force participation amplifies the problems of financing social security in times of population aging because it implies more recipients and less contributors.

A recent volume edited by Gruber and Wise (1998) argues that the declining old age labor force is strongly correlated with the incentives created by generous early retirement provisions. Unlike defined contribution plans in funded pension systems, pay-as-you-go systems are intrinsically difficult to design in a fashion that is actuarially fair. Moreover, the early retirement provisions in most European pension systems increase this actuarial unfairness for those who retire normally or late. Pension wealth is actually decreasing when retirement is postponed past the early retirement age in all countries depicted in Figure 1. In their summary article, Gruber and Wise (1998) estimate the international correlation between aggregate measures of old age labor force participation and early retirement incentives. Figure 2 shows our variant of their results, linking the proportion of men out of labor force at age 59 with a measure of the strength of labor force disincentives after age 59, defined as the accumulated loss of pension wealth when retirement is postponed past age 59.

Figure 2: Old age Labor Force Participation Disincentives and Early Retirement



Note: The measure of old age LFP disincentives sums the losses (gains) in pension wealth relative to the average wage, when retirement is postponed by a year over all years between the applicable early retirement age in each country and age 69. A value of 10 corresponds to 10 years with a tax rate of 100%. See Gruber and Wise (1998) for similarly well-fitting alternate specifications.

Figure 2 is the starting point of this paper. Because it is dangerous to give the aggregate correlation in Figure 2 a causal interpretation, this paper uses microeconomic evidence in order to continue the above line of reasoning and to go one more step in establishing causality. It claims that older workers in Europe have responded in a very systematic and consistent way to the early retirement provisions although the provisions varied considerably in their designs over time and across European countries.

The paper uses two strands of evidence to show this claim. One the one hand, it uses international variation in pension provisions across Europe to link incentives and labor force exit decisions. Countries included are Belgium, France, Germany, Italy, Spain, Sweden and the Netherlands. This part of the analysis uses no formal econometrics. Rather, it points out „kinks" in the function that relates the present discounted value of retirement benefits to retirement age. The paper shows that these institutional factors strongly correlate with „spikes" in the departure rate from employment.

One the other hand, the paper takes Germany as an example. Germany has one of the most generous retirement systems in the world. At the very same time, Germany also faces one of the most incisive population aging processes. The proportion of persons aged 60 and older will increase from 21 percent in 1995 to 36 percent in the year 2035, when the aging process will peak in Germany. With Switzerland and Austria, this will be the highest proportion in the world in the year 2035. A formal econometric analysis exploits changes in pension regulations over time as well as cross-sectional variation in the applicable social security rules to estimate how social security regulations have influenced old age labor force participation in Germany.

The paper stresses two implications for pension policy in the face of rapid population aging. First, remaining within the framework of the pay-as-you-go pension system, early retirement can be considerably reduced by abolishing early retirement incentives. Estimates for Germany indicate that an actuarial fair retirement system will reduce retirement before age 60 by more than a third. Second, the results imply that welfare comparisons between pay-as-you-go and fully funded systems that do not model the retirement decision ignore important welfare reducing redistributions between early and normal (and late) retirees.

Strategy of this paper is to employ several different ways to identify incentive effects, using historical, international and econometric evidence, in the hope that the weight of the cumulative evidence is stronger than the sum of each single step. The reader will therefore be shuttled among different perspectives. Section 2 begins with a brief description of the German pension system and the early retirement incentives it creates. Section 3 provides informal evidence for the link between early retirement incentives and old age labor force participation in Germany, and Section 4 extends this analysis to a set of European countries. Section 5 supplies the econometric methodology that is employed in Section 6 for a formal analysis of early retirement incentives in Germany, 1984-1996. Section 7 concludes.

2. Incentives Created by the German Public Pension System

The German public pension system is particularly well-suited for a microeconometric study of incentive effects on labor force participation because it is almost universal and we do not need to account for a variety of firm pension plans that create their own incentive effects but are usually not well captured in survey data (Börsch-Supan and Schnabel, 1998a). The homogeneity arises for two reasons. First, the German public pension system is mandatory for every worker except for the self-employed and those with very small labor incomes. Because almost all German workers have been dependently employed at least at some point in their working career, almost every worker has a claim on a public pension. Second, the system has a very high replacement rate, generating net retirement incomes that are currently about 70 percent of pre-retirement net earnings for a worker with a 45-year earnings history and average life-time earnings. This is substantially higher than the corresponding U.S. net replacement rate of about 53 percent. In addition, the system provides relatively generous survivor benefits that constitute a substantial proportion of the total pension liability. As a result, social security income represents about 80 percent of household income of households headed by a person aged 65 and over, the remainder about equally divided among firm pensions, asset income, and private transfers.

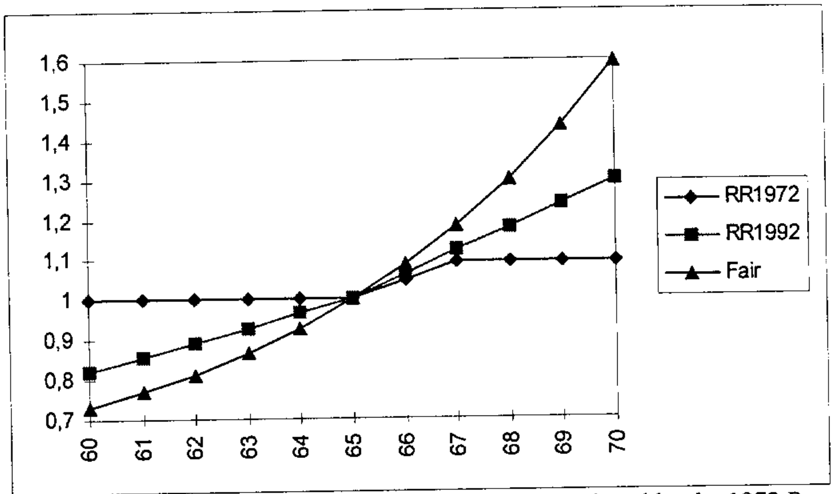
Until 1972, retirement was mandatory at age 65. In 1972, several early retirement options were introduced, „early“ defined as before age 65, the „normal“ retirement age. This gives us a first time-series handle to identify potential incentive effects. The German public pension system now provides *old-age pensions* for workers aged 60 and older and *disability benefits* for workers below age 60, that are converted to old-age pensions latest at age 65. In addition, *pre-retirement* (i.e., retirement before age 60) is possible using other parts of the public transfer system, mainly unemployment compensation. A main feature of the German old-age pensions is "flexible retirement" from age 63 for workers with a long service history. In addition, retirement at age 60 is possible for women, unemployed, and workers who cannot be appropriately employed for health or labor market reasons. This complicated system generates cross-sectional variation that gives us another handle to identify how incentive effects alter labor supply. It is noteworthy that these features were introduced as social achievements *before* unemployment started to rise in the mid-seventies. Only later, one realized that they helped to keep the unemployment rate down. 20 years after the introduction of the various early retirement options, the 1992 Pension Reform is attempting to close some of the early retirement options, see below. However, the effects will only be visible after the year 2004.

Roughly 80 percent of the budget of the German public pension system is financed by contributions, the rest by federal government revenue. Contributions are collected like a payroll tax, levied equally on employees and employers. The contribution rate in 1998 is very high: 21 percent of monthly gross income, paid in equal parts by employer and employee. The tax base for these contributions is capped at about 180 percent of average wages.

Benefits are computed on a life-time contribution basis. They are the product of four elements: (1) the employee's relative wage position, averaged over the entire earnings history, (2) the number of years of service life, (3) several adjustment factors, and (4) the average pension level. The first three factors make up the "personal pension base" which is calculated when entering retirement. Note that pensions are proportional to length of service life, a unique feature of the German pension system. The fourth factor determines the income distribution between workers and pensioners in general and is adjusted annually to net wages. Thus, productivity gains are transferred each year to all pensioners, not only to new entrants. Due to a generous exemption, social security benefits are tax free unless income from other sources is high.

Early retirement incentives are created by the (lack of) adjustment factors. Before the 1992 pension reform, there was no explicit adjustment of benefits when a worker retired earlier than at age 65, except for a bonus when retirement was postponed from ages 65 or 66 by one year, see Figure 3. Nevertheless, because benefits are proportional to the years of service, a worker with fewer years of service would get lower benefits even before the bonus. With a constant income profile and 40 years of service, each year of earlier retirement decreased pension benefits by 2.5 percent. However, this is substantially less than the actuarial adjustment which increases from about 5.5 percent for postponing retirement one year at age 60 to 8 percent for postponing retirement one year at age 65. The 1992 pension reform – not yet visible in the data – is gradually changing this by introducing retirement age-specific adjustment factors to the benefit formula. However, they will remain about 2 percent below those required for actuarial fairness.

Figure 3: Adjustment of Retirement Benefits to Retirement Age



Notes: „RR1972" denotes the adjustment factors introduced by the 1972 Pension Reform, „RR1992" symbolizes the adjustments that will be phased in by the 1992 Pension Reform, and „Fair" refers to actuarially fair adjustment factors.

The key statistic to measure the early retirement incentives exerted by the actuarially unfair adjustment factors is the change in social security wealth. If social security wealth declines because the increase in the annual pension is not large enough to offset the shorter time of pension receipt, workers have a financial incentive to retire earlier. We define social security wealth as the expected present discounted value of benefits minus applicable contributions. Seen from the perspective of a worker who is S years old and plans to retire at age R, social security wealth (SSW) is

$$SSW_S(R) = \sum_{t=R}^{\infty} YRET_t(R) \cdot a_t \cdot \delta^{t-S} \quad - \quad \sum_{t=S}^{R-1} c \cdot YLAB_t \cdot a_t \cdot \delta^{t-S},$$

with:

SSW	present discounted value of retirement benefits (=social security wealth),
S	planning age,
R	retirement age,
$YLAB_t$	labor income at age t,
$YPEN_t(R)$	pension income at age t for retirement at age R,
c_t	contribution rate to pension system at age t,
a_t	probability to survive at least until age t given survival until age S,
δ	discount factor = $1/(1+r)$.

The accrual rate of social security wealth between age t-1 and t is

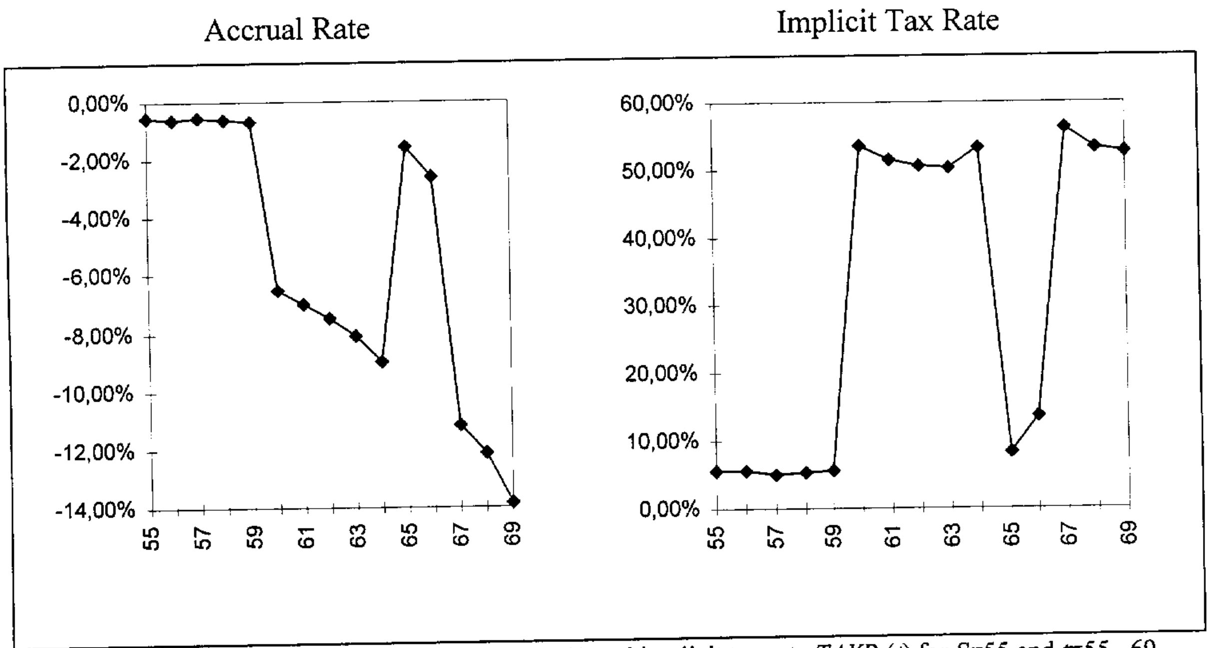
$$ACCR_S(t) = [SSW_S(t) - SSW_S(t\text{-}1)] \,/\, SSW_S(t\text{-}1).$$

A negative accrual can be interpreted as a tax on further labor force participation. We compute an implicit tax rate as the ratio of the (negative) social security wealth accrual to the net wages ($YLAB^{NET}$) that workers would earn if they would postpone retirement:

$$TAXR_S(t) = -\,[SSW_S(t) - SSW_S(t\text{-}1)] \,/\, YLAB_t^{NET}.$$

Figure 4 shows that the early retirement incentives in Germany are strong. The accrual function (left panel) has three distinctive kink points. The first kink occurs at age 60, the earliest retirement age into the public pension system without disability status. Two other kinks are generated by the bonus for postponing retirement at ages 65 and 66, interrupting the steady increase in negative pension wealth accrual.

Figure 4: Loss in Social Security Wealth When Postponing Retirement, 1972 Rules



Note: See text for definition of accrual rate $ACCR_S(t)$ and implicit tax rate $TAXR_S(t)$ for S=55 and t=55...69.

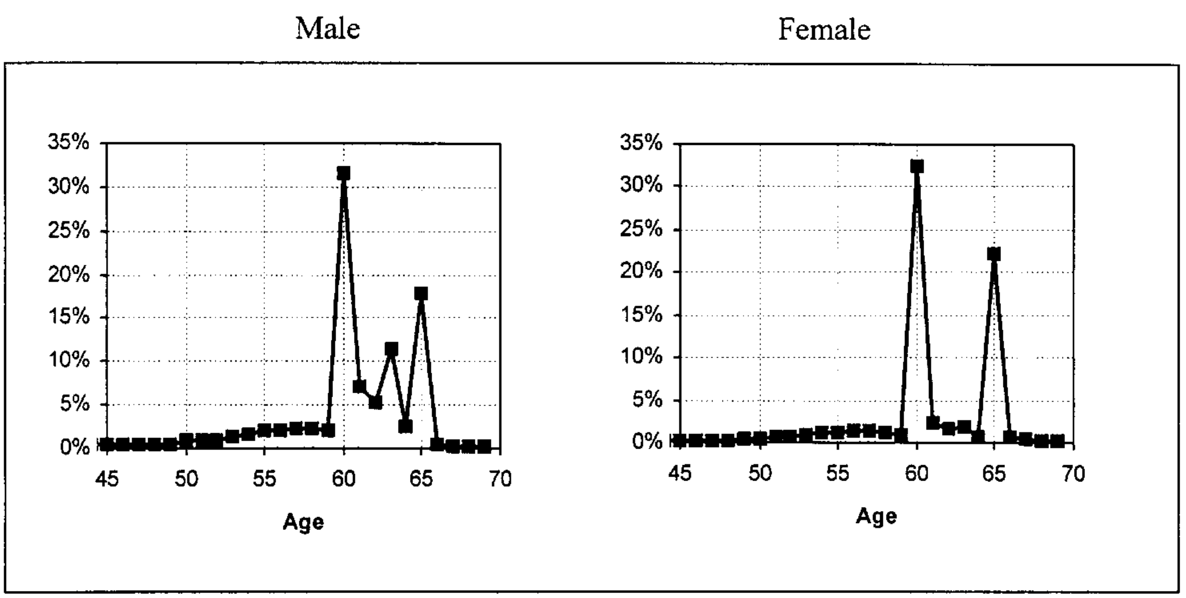
The lack of actuarial fairness of the German public pension system creates a negative accrual of pension wealth during the early retirement window at a rate reaching -9 percent when retirement is postponed from age 64 to 65. In 1995, this was a loss of about DM 22,000 (US \$ 10,500 at purchasing power parity) for the average worker. Expressed as a percentage of annual labor income, the loss corresponds to a „tax" which exceeds 50 percent.

The 1992 Pension Reform will moderate but not abolish this incentive effect. After 2004, when the 1992 reform will have fully be phased in, the negative accrual rate will reach -5 percent, corresponding to an implicit tax rate of almost 30 percent when retirement is postponed by one year at age 64.

3. Qualitative Evidence on LFP Disincentives in Germany

The labor supply disincentives for older workers created by this implicit tax are reflected in the data. Male labor force participation plunges between age 55, when it is almost 90 percent, to 38 percent at age 60. It is less than 8 percent at age 65. Figure 5 shows the cross sectional distribution of retirement ages that has its maximum at age 60, the earliest age at which retirement due to health and labor market reasons without formal claims to disability benefits is possible. The other spikes correspond to age 63 for flexible retirement and to age 65 for workers with short work histories. Quite clearly, workers take the very first legal opportunity to retire – supposedly, because they realize that a postponement inflicts a loss of social security wealth.

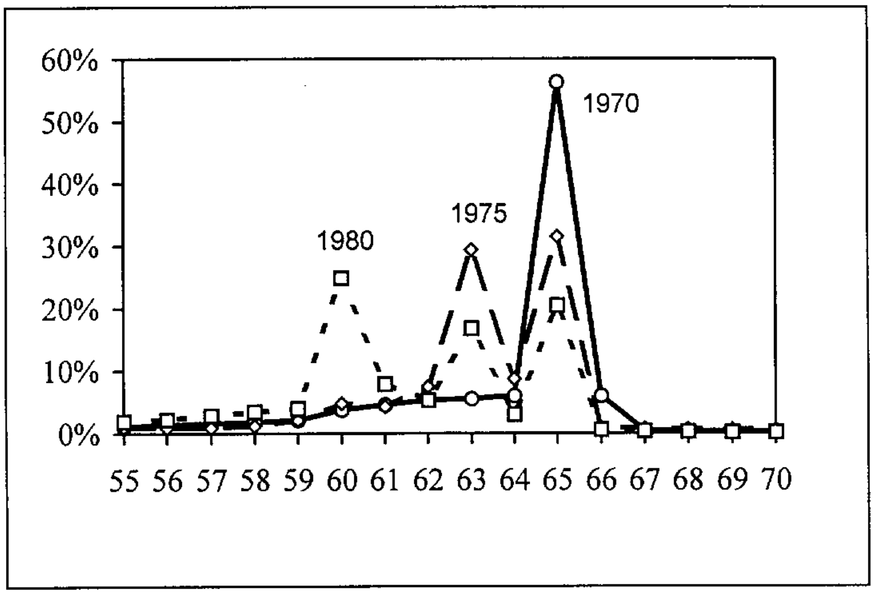
Figure 5: Distribution of Retirement Ages



Source: Verband deutscher Rentenversicherungsträger (VdR), 1997

This pattern has evolved very systematically after the introduction of the early retirement options in 1972, as shown in Figure 6. Before the 1972 Reform, the distribution of retirement ages had only one spike: in 1970, age 65 was *the* retirement age. In 1975, already two years after the 1972 Reform became effective, about half of the retirees preferred to retire earlier, and the retirement age distribution had two spikes. In 1980, another five years later, today's three-spike pattern (shown in Figure 5) had emerged.

Figure 6: Distribution of Retirement Ages, Men, 1970, 1975, and 1980



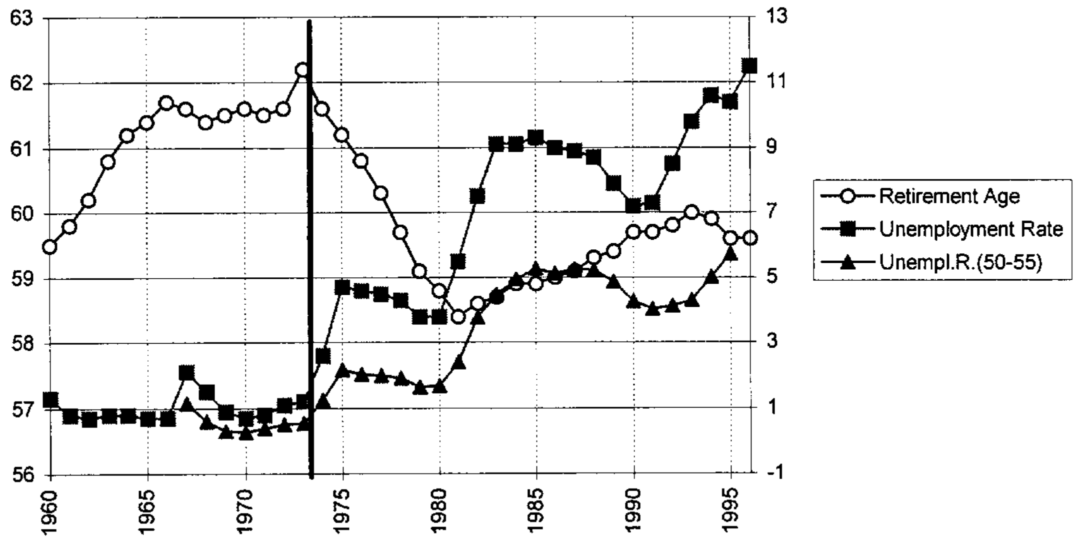
Source: VdR 1997.

One might argue that the shift in retirement age and the corresponding decline in old-age labor force participation after the 1972 Reform was mainly a demand reaction to rising unemployment rather than a supply response to early retirement incentives. We have two reasons not to believe that this was the case. First, the reform was part of the general expansion of the welfare state after the social democrats took power for the first time in the postwar period. The reform was designed at a time when unemployment was very low while the first steep increase in unemployment (1973 to 1975) happened after the 1972 Reform had already been introduced.

Second, the time-series evidence does not support the view of a demand-induced change in retirement age. Figure 7 shows the sudden decrease in the average retirement age – the average age of those who receive a public pension for the first time – after the 1972 reform. Average retirement age dropped steadily from about age 61.5 in 1971 to age 58.5 in 1981. In turn, unemployment first rose between 1970 and 1975 from 0.7 to 4.7 percent, then fell from 1975 to 1980. Therefore, there is no systematic time-series correlation between mean retirement age and unemployment between the reform and 1980. Between 1980 and

1983, the unemployment rate rose from 3.8 to 9.1 percent in a dramatic increase that eventually led to the collapse of the social democratic government. However, mean retirement age increased slightly from thereon. This positive time-series correlation between mean retirement age and unemployment is the opposite of what a demand-induced change in old age labor supply would suggest. This holds both for the general unemployment rate and the unemployment rate of men aged between 50 and 55.

Figure 7: Average Retirement Age, 1960-1995



Note: „Retirement Age" is the average age of all new entries into the public pension system. „Unemployment rate" is the general national unemployment rate. „Unempl.R.(50-55)" refers to male unemployed age 50-55.
Source: VdR 1997 and BMA 1997.

Riphahn and Schmidt (1997) confirm this view econometrically. By using panel data that combine regional, sectoral and time variation in unemployment rates and old age labor force participation, they show that the decline in the average retirement age was not a demand response to rising unemployment but rather a supply response to the introduction of early retirement incentives through the 1972 Pension Reform.

4. Pension Incentives and Retirement Age Distributions Across Europe

The qualitative analysis can be extended to other countries in Europe. In fact, the international variation in old age labor force participation provides another handle to identify disincentive effects of pension systems. The „International Social Security Project" led by

Gruber and Wise (1998) compiled comparable social security wealth calculations for a set of European countries plus Japan and the U.S. We use their calculated accrual rates and implicit tax profiles to produce a similar qualitative analysis to the one depicted in Figures 4 and 5 in the preceding section for the German case.

In all countries, labor force exits concentrate at certain ages, creating spikes in the distribution of retirement ages. These spikes can in almost all cases be identified with ages in which certain pension rules start or cease to apply. The most obvious example is the earliest age in which a worker is eligible for social security. In addition, most European pension systems have several „pathways“ to retirement, all with different eligibility ages and/or replacement rates, similar to the German example. These rules create kinks in the accrual of social security wealth. Table 2 compiles a list of the kinks in the social security wealth accrual function, and juxtaposes it to the list of spikes in the distribution of retirement ages for seven European countries plus Japan and the U.S. as reference countries. Indeed, the correspondence between kinks and spikes is quite strong.

Table 2: Kinks and Spikes

	Belgium	France	Germany
Pension wealth accrual	very sharp kink at age 60	Sharp kink at age 65, almost flat thereafter	sharp kink at age 60, reverse kink at age 65
Distribution of retirement ages	spikes at age 60, 65	sharp spike at age 60, small spike at age 65	sharp spike at age 60, small spike at age 65
	Italy	**Netherlands**	**Spain**
Pension wealth accrual	negative from age 55, kink at age 60	subsidy at age 59, sharp kink at age 60	sharp kinks at ages 60 and 65
Distribution of retirement ages	sharp spikes at ages 55,60,65	very sharp spikes at age 60	small spike at age 60, sharp spike at age 65
	Sweden	**Japan**	**U.S.**
Pension wealth accrual	almost flat until age 65	sharp kink at age 60	kinks at age 62 and 65 subsidy at 62 and 63
Distribution of retirement ages	only spike at age 65	sharp spike at age 60, small hump at age 65	spikes at ages 62 and 65

Sources: Belgium: Pestieau and Stjins (1998); France: Blanchet and Pelé (1998); Germany: Börsch-Supan and Schnabel (1998b); Italy: Brugiavini (1998); Netherlands: Kapteyn and de Vos (1998); Spain: Boldrin, Jimenez and Peracchi (1998); Sweden: Palme and Svensson (1998); Japan: Oshio and Yashiro (1998); U.S.: Diamond and Gruber (1998).

So far, the evidence on the effects of pension systems on old age labor force participation has been informal. We have accumulated three kinds of qualitative evidence: German workers appear to select the earliest retirement age legally possible (Figure 5); German workers have responded to the historical experiment of the 1972 Pension Reform is a very consistent way (Figures 6 and 7); and there is a close link between kinks in the pension accrual function and spikes in the retirement distribution both in Germany (Figures 4 and 5) and in other countries (Table 2). All this suggests that it is worthwhile to conduct a more formal analysis of early retirement incentives.

5. Econometric Estimation of Disincentive Effects: Data and Methodology

We take up the example of Germany where an econometric analysis is facilitated by the homogeneity of the retirement income provision as described in Section 2. The methodology follows the seminal work by Stock and Wise (1990). Earlier analyses of the German pension system using this framework were carried out by Börsch-Supan (1992), Schmidt (1995), Börsch-Supan and Schmidt (1996), and Siddiqui (1997). We improve on this work by exploiting more systematically the cross-sectional and time-series variation contained in the 1984-1996 German Socio-Economic Panel (GSOEP) that spans several modifications of the German pension system. This section has three parts. We will first describe the data, then the option value framework, and finally our panel data estimation procedure.

Data

The German Socio-Economic Panel (GSOEP) is an annual panel study of some 6000 households and some 15000 individuals. Its design closely corresponds to the U.S. Panel Study of Income Dynamics (PSID). The panel started in 1984; 13 waves through 1996 are currently available. Since 1990, the West German panel was augmented by an East German sample. We will not use the East German subpanel, however, since retirement patterns in the East are still very much dominated by the difficult transition from a command to a market economy. Response rates and panel mortality are comparable to the PSID. The GSOEP data provide a detailed account of income and employment status. The data is used extensively in Germany, and the increasing interest in the U.S. prompted the construction of an English-language user file available from Richard Burkhauser and his associates at Syracuse University. Burkhauser (1991) reports on the usefulness of the German panel data and provides English-language code books as well as an internationally accessible GSOEP version.

Our subsample covers all persons aged 55 through 70 in West Germany for which earnings data is available. This includes 1639 individuals. We construct an unbalanced panel of these individuals. We obtain 8474 observations, i.e., average observation time is slightly more than 5 years. The panel is left-censored as we include only persons who have worked at least one year during our window in order to reconstruct an earning history. There is only little right censoring due to missing interviews. Of the 1639 individuals, 678 have no transitions, 609 have a single transition from employment to retirement, and 352 individuals have more complex histories with at least one reverse transition. 35 percent are female, and the most frequent retirement age is age 60.

The option value model

We capture the economic incentives provided by the pension system by the option value to postpone retirement (Stock and Wise, 1990). This value captures for each retirement age the trade-off between retiring now (resulting in a stream of retirement benefits that depends on this retirement age) and keeping all options open for some later retirement date (with associated streams of first labor, then retirement incomes for all possible later retirement ages). Consequently, the option value for a specific age is defined as the difference between the maximum attainable consumption utility if the worker postpones retirement to some later year minus the utility of consumption that the worker can afford if the worker would retire now.

Let $V_t(R)$ denote the expected discounted future utility at age t if the worker retires at age R. Let R*(t) denote the optimal retirement age if the worker postpones retirement past age t, i.e., $\max(V_t(r))$ for r>t. With this notation, the option value is

$$G(t) = V_t(R^*(t)) - V_t(t)$$

Since a worker is likely to retire as soon as the utility of the option to postpone retirement becomes smaller than the utility of retiring now, retirement probabilities should depend negatively on the option value.

We specify the expected utility as follows:

$$V_t(R) = \sum_{s=t}^{R-1} \mathrm{u}(YLAB_s) \cdot a_s \cdot \delta^{s-t} + \alpha \sum_{s=R}^{\infty} \mathrm{u}(YRET_s(R)) \cdot a_s \cdot \delta^{s-t}$$

with	$YLAB_s$	Labor income at age s, s=t...R-1,
	$YRET_s(R)$	Retirement income at age s, s>R,
	R	Retirement age,
	α	Marginal utility of leisure, to be estimated,
	a	Probability to survive at least until age s,
	δ	Discount factor, set at 3 percent.

To capture the utility from leisure, utility during retirement is weighted by $\alpha>1$, where $1/\alpha$ is the marginal disutility of work. We will estimate this parameter by grid search, see below. We apply a very simple utility function by identifying consumption with income. Preliminary estimates with an isoelastic utility function, $u(Y) = Y^{\gamma}$, yield a γ coefficient that is not significantly different from one. The discount factor δ is assumed to be 3 percent. Other discount factors in the range between 1 and 6 percent yield qualitatively similar results.

Retirement income depends on retirement age according to the adjustment factors and on previous labor income by the benefit formula explained in Section 2. The option value captures therefore the economic incentives created by the pension system and the labor market. The option value is computed for every person, using the applicable pension regulations and individual earning histories. Additional private pension income is ignored because it represents only a very small proportion of retirement income as described before. Note that due to the linear utility function, the second part of $V_t(R)$ corresponds closely to the pension wealth $SSW_t(R)$ defined in Section 2.

Econometric estimation method

The variable to be explained is old age labor force status. Because Germany has very few part-time employees, we model only two states – fully in labor force and fully retired – unlike the competing risk analysis of Sueyoshi (1989). The definition of „retired" is problematic, although less so in Germany than in other countries (for the U.S., see Rust, 1990). Retirement definitions commonly employed in the literature include the retirement status self-reported by the respondent, the fact that there are few work hours, or the receipt of retirement benefits, among other definitions. We use the first concept, and include pre-retirement, mainly financed by a mixture of unemployment compensation and severance pay, in our definition of retirement.

Our main explanatory variable is the option value described in the previous subsection. The other explanatory variables are the usual suspects: an array of socio-economic variables such as gender, marital status, wealth (indicator variables of several financial and real wealth categories) and a self-assessed health measure. We do not use the legal disability status as a measure of health since this is endogenous to the retirement decision.

We link the explanatory variables to the dependent variable by a panel probit model with a parametrized correlation pattern over time. The model can be interpreted as a semi-nonparametric hazard model for multiple spell data permitting unobserved heterogeneity and state dependence. It is non-parametric in the sense that the model does not impose a functional form on the duration in a given state. Fairly flexible hazard rate models of retirement have been estimated by Sueyoshi (1989) and Meghir and Whitehouse (1997), however, not in combination with an option value describing the incentives to retire. Parametric hazard rate models for German data have been estimated by Schmidt (1995) and Börsch-Supan and Schmidt (1996).

Inserting the option value in a regression-type model is much less computationally involved and more practical than the estimation procedure employed by Stock and Wise (1990) which in turn much closer approximates the underlying dynamic programming structure (Rust and Phelan, 1997), see Lumbsdaine, Stock and Wise (1992), and at the same time generates robust estimates of the average effects of the option value on retirement.

Nevertheless, discrete choice models for panel data are necessarily computationally involved because the space of possible outcomes is so large. For I discrete choices and T panel periods, there are I^T different choice sequences $\{i_t\}, t = 1,\ldots,T$. Although not all of these I^T choice sequences are observed in our data, there are sufficiently many complex choice sequences in the data to warrant a departure from a simple one-spell hazard model. We structure the choices by cross-sectional utility maximization:

(1) $i=i_t$ is chosen $\Leftrightarrow u_{it}$ is maximal element in $\{u_{jt}, j=1,..,I\}$,

where the utility of choice i in period t is the sum of a deterministic utility component which depends on the vector of observable variables X_{it} and a parameter vector β to be estimated, and a random utility component ε_{it}. We assume that this random component is normally distributed. In our case, the number of choices equals two and all relevant information is the difference between the two states, u_{1t}-u_{2t}. For simplicity and in a slight abuse of notation, we normalize the utility of the second state to 0 and suppress the index i, denoting the difference u_{1t}-u_{2t} by u_t. Finally, we specify the deterministic utility of the first state as a linear function

(2) $u_t = X_t\beta + \varepsilon_t$.

The probability of a choice sequence $\{i_t\}$, expressed by condition (1) for all periods t=1 through T is therefore a T-dimensional normal probability. If utilities are correlated over time, which is likely in panel data, these probabilities are difficult to evaluate. Examples for intertemporal linkages are random effects and autocorrelation in ε_t. In the language of duration models, random effects capture unobserved population heterogeneity, and autocorrelated errors capture elements of state dependency. More formally, a random-effect structure is imposed by specifying

$\varepsilon_t = \alpha + \nu_t$ for ν_t i.i.d.

This yields an equicorrelation structure of the covariance matrix of ε, denoted by M, parametrized by the standard deviation of α. This structure allows for a factorization of multidimensional probability that makes computation relatively easy (Mofitt, 1987). An autoregressive error structure can be incorporated by specifying

$$\varepsilon_t = \rho \cdot \varepsilon_{t-1} + \nu_t \quad \text{for } \nu_t \text{ i.i.d.}$$

The resulting familiar AR(1) structure of M does not permit an easy factorization of the T-dimensional probability. The two error structures can be combined by specifying

$$\varepsilon_t = \alpha + \eta_t \quad \text{where} \quad \eta_t = \rho \cdot \eta_{t-1} + \nu_t \quad \text{for} \quad \nu_t \text{ i.i.d.}$$

This amounts to overlaying the equicorrrelation structure with the AR(1) structure. It should be noted that this model can only be identified in reasonably long panels, and formally if $T \geq 3$ and $\rho < 1$.

Classical evaluation of the resulting panel-data probit choice probabilities is computationally infeasible since the number of operations increases exponentially with the length of the panel, T. Approximation methods, such as the Clark approximation (Daganzo, 1981) or its variant proposed by Langdon (1984), are tractable - their number of operations increases quadratically in T - but they remain unsatisfactory since their relatively large bias cannot be controlled by increasing the number of observations. Rather, we simulate the choice probabilities by drawing pseudo-random realizations from the underlying error process.

The most straightforward simulation method is to simulate the choice probabilities by frequencies obtained by counting how often the event (1) appears when a large number of ε_t are randomly drawn (Lerman and Manski, 1981). However, in order to obtain reasonably accurate estimates of small choice probabilities, a very large number of draws is required. That results in unacceptably long computer runs. The occurrence of small choice probabilities is more likely in long panels because they generate more complex choice sequences.

We use instead the smooth simulated maximum likelihood estimation (SSML) method proposed by Börsch-Supan and Hajivassiliou (1993). This algorithm is very quick, depends continuously on the parameters β and M, and is currently the most efficient method for panel data probit models, see the surveys Hajivassiliou, McFadden and Ruud (1996) and Stern (1998). We sketch this method – as it applies to our application – in the remainder of this section.

The choice probability of (1) corresponds to a truncated normal probability where the truncation occurs at the left at $a = X_t\beta$. Simulating probit choice probabilities thus corresponds to drawing truncated normal variates. It is straightforward to draw univariate truncated normal variates by applying the integral transform theorem: If u is drawn from a univariate standard uniform distribution, $0<u<1$, then

$$(3) \qquad e = \Phi^{-1}\{[1-\Phi(a)]\cdot u + \Phi(a)\}$$

is distributed left-truncated normal with truncation point a. We use the notation $e = N(0,1)$ s.t. $a \le e \le \infty$. Φ denotes the univariate standard normal cumulative distribution function. Note that e is a continuously differentiable function of the truncation parameter a. This continuity is essential for computational efficiency.

The multivariate case is more difficult, see Geweke (1991). In our case, we have to draw a T-dimensional vector of unobserved utilities ε_t. Let Γ be the lower diagonal Cholesky factor of the covariance matrix M of the unobserved utility vector ε such that $\Gamma\Gamma' = M$. We begin by drawing sequentially an auxiliary vector e of T univariate left-truncated normal variates

$$e = N(0, I) \text{ s.t. } a \le \Gamma e \le \infty,$$

where $a = (a_1,\ldots a_T)$ is now a vector of T truncation points with $a_t = X_t\beta$. Because Γ is triangular, the restrictions on e are recursive:

$$(4) \qquad \begin{aligned} e_1 &= N(0,1) \quad \text{s.t. } a_1 \le \gamma_{11}e_1 \le \infty \Leftrightarrow a_1/\gamma_{11} \le e_1 \le \infty, \\ e_2 &= N(0,1) \quad \text{s.t. } a_2 \le \gamma_{21}e_{1l} + \gamma_{22}e_2 \le \infty \Leftrightarrow (a_2 - \gamma_{21}e_1)/\gamma_{22} \le e_2 \le \infty, \end{aligned}$$

etc. until $t=T$.

Hence, each e_t, $t=1..T$, in (4) can be drawn using the univariate formula (3).

Finally, define $\varepsilon=\Gamma e$. Because $\Gamma\Gamma' = M$, ε has covariance matrix M and is subject to $a \le \Gamma e \le \infty \Leftrightarrow a \le \varepsilon \le \infty$. Therefore, the probability for a choice sequence $\{i_{tn}\}$ of observation n is, in an expected value sense, the probability that a draw of a T-dimensional vector of truncated normal variates $e_r = (e_{r1},\ldots,e_{rT})$ falls in the set of intervals given by (4):

$$\begin{aligned} \widetilde{P}_r(\{i_{tn}\}) &= \Pr(a_1/\gamma_{11} \le e_{r1} \le \infty)\cdot \Pr[(a_2-\gamma_{21}e_{r1})/\gamma_{22} \le e_{r2} \le \infty \mid e_1] \cdot \ldots\ldots \\ &= [1-\Phi(a_1/\gamma_{11})]\cdot[1-\Phi((a_2-\gamma_{21}e_{r1})/\gamma_{22})]\cdot\ldots, \end{aligned}$$

To increase precision, the choice probability is approximated by the average over R replications:

$$(5)\quad \widetilde{P}(\{i_{tn}\}) = \frac{1}{R}\sum_{r=1}^{R}\widetilde{P}_r(\{i_{tn}\}).$$

which are inserted into the SSML likelihood function

$$(6)\quad \widetilde{L}(\beta, M) = \prod_{n=1}^{N}\sum_{r=1}^{R}\widetilde{P}_r(\{i_{tn}\}).$$

Börsch-Supan and Hajivassiliou (1993) prove that $\widetilde{P}$ is an unbiased estimator of P. Because the likelihood function (6) depends nonlinearly on the simulated probabilities, $\widetilde{L}$ carries a small sample bias that vanishes only if R goes to infinity. In practice, however, are the choice probabilities well approximated by (5) already for reasonably small number of replications, independent of the true choice probabilities, such that the bias becomes quickly unnoticeable. Because the generated draws as well as the resulting simulated probabilities depend continuously and differentiably on the parameters β in the truncation vector a and the covariance matrix M, conventional optimization methods can be used to solve the first order conditions for maximizing the simulated likelihood function, and the computational effort in the simulation increases nearly linearly with panel length.

6. Econometric Estimation of Disincentive Effects: Results

Estimation results are presented in Table 3 below. Dependent variable is labor force status „retired“ which includes formal retirement as well as the various forms of pre-retirement. A positive coefficient indicates that the explanatory variables increases the probability of retirement. In addition to the option value, health, and an array of socio-economic variables, we include a full set of age dummies to non-parametrically capture all other unmeasured effects on the retirement decision that are systematically related to age, such as social customs. The reference category is age 65.

We estimate three models. Model 1 has i.i.d. errors. Model 2 corrects for unobserved heterogeneity by a random effect whose standard deviation is reported at the bottom of Table 3. Finally, Model 3 adds an autoregressive error component to Model 2.

Even the simple i.i.d. model fits the data well. The pseudo-R^2 - one minus the ratio of the likelihood at the estimated parameters over the likelihood at zero - is 60.5 percent. Introducing random effects increases the log likelihood significantly, the pseudo-R^2 increases to 68.7 percent. The additional inclusion of an autoregressive component is also statistically significant, the pseudo-R^2 now rises to 69.4 percent. The prediction success is about 87 percent for all three models. The sample distribution is 30 percent in the labor force and 70 percent retired.

Table 3: Multiperiod probit model of the retirement decision

	MODEL 1: i.i.d.		MODEL 2: RAN		MODEL 3: RAN+AR1	
Variable	**Parameter**	***t-Stat***	**Parameter**	***t-Stat***	**Parameter**	***t-Stat***
Option value	-0.0041	*-7.45*	-0.0087	*-9.30*	-0,0096	*-8,17*
Health	-0.1385	*-10.1*	-0.1190	*-4.90*	-0,0991	*-3,71*
Female	-0.1246	*-1.81*	-0.2238	*-1.29*	-0,4172	*-1,91*
Married	-0.0328	*-0.43*	0.0887	*0.45*	0,0337	*0,15*
Education	0.1035	*0.98*	0.5822	*2.23*	0,4303	*1,30*
Civil servant	-10.010	*-10.0*	-15.226	*-10.0*	-19,777	*-10,0*
Firm pension	-2.6465	*-5.62*	-4.1637	*-5.57*	-4,3530	*-4,91*
Life insurance	-0.1719	*-2.83*	-0.2341	*-2.17*	-0,1781	*-1,44*
Stocks/bonds	0.0720	*0.97*	-0.0719	*-0.51*	-0,0776	*-0,51*
Real estate	-0.6970	*-8.08*	-1.0468	*-6.50*	-1,0152	*-5,55*
Owner occup.	0.2666	*4.16*	0.2270	*1.45*	0,2441	*1,33*
Age≤59	-3.2770	*-21.2*	-6.2911	*-22.1*	-6,3849	*-17,1*
Age=60	-2.4725	*-15.8*	-4.8332	*-18.6*	-4,9591	*-15,2*
Age=61	-1.2705	*-8.83*	-2.6248	*-12.0*	-2,7235	*-10,7*
Age=62	-0.9559	*-6.65*	-1.9526	*-9.23*	-2,0195	*-8,54*
Age=63	-0.8527	*-5.84*	-1.8068	*-8.52*	-1,8860	*-8,23*
Age=64	-0.2487	*-1.61*	-0.4544	*-2.11*	-0,4817	*-2,22*
Age=66	0.9891	*4.56*	1.6842	*5.67*	1,7242	*5,62*
Age=67	0.6708	*3.21*	1.3595	*4.51*	1,3682	*4,41*
Age≥68	0.8810	*4.89*	1.9073	*7.13*	1,9398	*6,45*
Constant	2.2246	*13.4*	3.3728	*10.8*	3,4100	*9,35*
RAN			2.3848	*14.8*	2,2869	*10,3*
AR1					0,6903	*7,05*
Log likelihood	-2321.32		-1836.91		-1744.98	
Individuals	8474		1639		1639	
Max. periods	1		13		13	
Observations	8474		8474		8474	

Notes: The dependent variable is the dummy variable „retired". T-statistics are based on robust standard errors. The log likelihood value at zero parameter values is 5873.7. 100 Replications.

Our most important results relate to the coefficients of the option value. In all models, they are statistically highly significant. The signs are negative, as expected, indicating a lower probability of retirement when the option value to postpone retirement is large. Quite noticeable is the lack of any spikes in the pattern of age dummies. In this sense, retirement behavior is correctly described by the option value, the main economic incentive for retirement.

Taking account of the intertemporal correlations in the panel appears to be very important. The numerical value of the option value coefficient is severely underestimated in the i.i.d. model. We will use simulation exercises below to illustrate the size of magnitude of the estimated coefficients.

The other economic incentives for retirement, the wealth variables, are only partially significant. The GSOEP data does not contain levels of wealth, only indicators whether certain portfolio components – firm pension, life insurance, stock and bonds, and real estate – are present. There are many missing values, here coded as „not present". In general, presence of financial and real wealth decreases the retirement probability. This is not particularly plausible for the presence of a firm pension. However, significant firm pensions are rare in Germany and usually indicate higher valued jobs in which retirement may occur later for reasons not related to the firm pension per se.

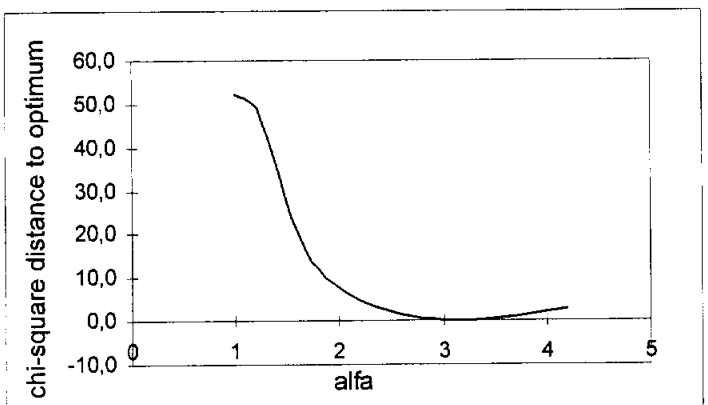
The pattern of age dummies reflects the obvious: older workers are more likely retired than younger ones. It is important to measure the option value with the age dummies included in order to purge its estimated coefficient from all other non-economic effects. The omission of age dummies about triples the estimated coefficient of the option value. Most other socio-demographic variables are not significant. The important differences in social security regulations between men and women (women can retire at age 60 if have at least 15 years of retirement insurance history, while men need 35 years to retire at age 63, unless they claim disability) appears to be fully captured by the option value. Marital status and education is also insignificant. We did not correctly model the retirement subsystem for civil servants. They are actually treated as if they were part of the standard social security system which is not the case. Civil servants are required to work longer than other employees, with a fairly rigid retirement age at 65, although claims to disability are frequent as well as early retirement due to downsizing of the civil service sector. We find a corresponding rather strong negative coefficient, indicating later retirement for civil servants.

Very interesting are the estimated coefficients of the health variable. It is coded 0 for „very poor“ to 10 for „excellent“. As expected, the coefficients are negative. Less healthy workers retire earlier. In the i.i.d. model, health is more significant than the option value. However, as soon as unobserved population heterogeneity is accounted for, this changes, and the estimated coefficient becomes somewhat smaller. This shows once more the importance to account for intertemporal linkages. Note that we did not use disability status because it is an endogenous variable. The desire for early retirement may prompt workers to seek disability status, and frequently the employer helps in this process to alleviate restructuring. Until recently, disability status was granted for labor market reasons without a link to health.

The estimation results in Table 3 are based on 100 replications. There is no appreciable parameter change if the number of replications is increased, showing that the small sample bias of the SSML estimator has gone away at this number of replications (see Appendix).

Finally, there is one additional parameter in the model that does not fit in the simple linear utility framework of the probit model, namely the inverse marginal disutility of work, α. Rather than implementing a nonlinear utility function, we adapted a more pragmatic approach. Since a grid search over α yielded a rather flat likelihood with an optimum at 3.13, we used this value for all estimated models (see Figure 8). This value means that a dollar that has to be earned by work is valued at only about a third of a dollar that is given as a public transfer through the retirement system. This value is somewhat higher than estimates for the U.S. but not implausible for Germany with an arguably higher preference for leisure.

Figure 8: Gridsearch over the Inverse Marginal Disutility of Work (α)

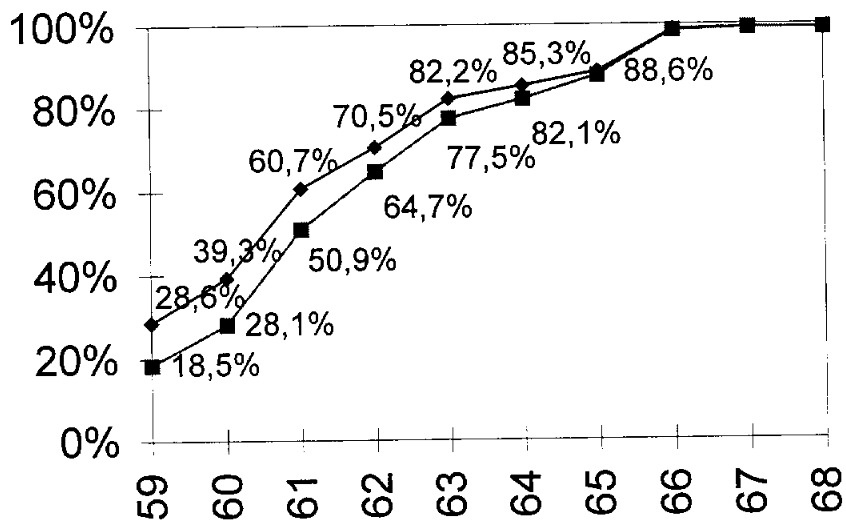


Note: Based on Model 1.

Simulations

The estimated coefficients can be used to compute the effect of the non-actuarial benefit adjustments on retirement age. Introducing an actuarially fair benefit formula induces a shift of the cumulative distribution to the right, see Figure 9: Early retirement ages (until age 65) are less frequent than before. The effects are most powerful for very early retirement, i.e., retirement before the official window period through disability or pre-retirement schemes. A shift to an actuarial fair system would cause retirement at ages 59 and below to drop from 28.6 percent to less than 18.5 percent.

Figure 9: Actual and Predicted Cumulative Distribution of Retirement Age



Note: Percent retired at given age. Upper graph: current rules (baseline). Lower graph: actuarially fair adjustment factors. Microsimulation based on estimated coefficients of Model 3, Table 3.

7. Summary and Conclusions

The strategy of this paper was to use several different strands of evidence to argue that pension incentives are an important cause for the decline in, and the relatively low level of, old age labor participation in Europe. It was motivated by the hope that the weight of the cumulative evidence is stronger than the sum of each single step. We exploited international and German historical and cross-sectional variation in social security regulations. Pension incentives, visible as kinks in the accrual function, correspond closely to spikes in the distribution of retirement ages in the countries we inspected. We looked at the German case in more detail. The German case is interesting because population aging in Germany is particularly strong and while Germany nevertheless has one of the most generous and rigid PAYG systems. Moreover, because of the ubiquitousness of the German public pension system, microeconometric analysis is much easier than, say, in the U.S. or the UK with their very heterogeneous firm pensions. We did three exercises. We looked at the kinks and spikes in more detail; we looked at the historical response to the 1972 pension reform with its sudden introduction of various early retirement options; and we estimated a formal econometric retirement model based on option value analysis. All three exercises yielded the same result, corresponding to the results of the international comparisons: Workers have responded consistently and strongly to the economic incentives to retire earlier. It is interesting to note that they have done so before unemployment became widespread in the beginning of the 1980s. Policy-induced early retirement accounts for about a third of early retirement, thus a substantial fraction of total pension expenditures.

The responsiveness of the choice of retirement age to the incentives offered by pension systems has strong policy impacts. The public pension systems do not use retirement-age-dependent adjustments of pension benefits as policy instruments for balancing the budget of the pension system, they even yield incentives to retire early, most frequently at the earliest legal retirement age. While this has alleviated the unemployment problem in the short run, is was an expensive way to do so, and it will hurt in the long run. Rather than awarding later retirement to moderate the labor supply disincentives created by quickly rising social security taxes, social security regulations across the world have encouraged early retirement, thus aggravating the imbalance between the number of workers and pensioners in times of population aging.

The responsiveness to actuarially unfair pension regulations also shows that there is widespread redistribution between from late to early retirees. This increases the deadweight burden generated by contributions to PAYG pension systems, a crucial parameter in computations that try to design welfare-improving transitions from PAYG to fully funded pension systems (e.g., Feldstein and Samwick, 1996; Kotlikoff,1996). The estimates of the option value model can be used to quantify these order of magnitude for such deadweight burden calculations.

References

Blanchet, D., and L.-P. Pelé, 1998, Social Security and Retirement in France, NBER Working Paper No. 6214, Cambridge, Mass.

Blundell, R., and P. Johnson, 1998, Pensions and Retirement in the United Kingdom, NBER Working Paper No. 6154, Cambridge, Mass.

Boldrin, M., S. Jimenez and F. Peracchi, 1998, Social Security and Retirement in Spain, NBER Working Paper No. 6136, Cambridge, Mass.

Börsch-Supan, A., 1992, Population Aging, Social Security Design, and Early Retirement, *Journal of Institutional and Theoretical Economics 148*, 533-557.

Börsch-Supan, A., and Hajivassiliou, V., 1993, Smooth Unbiased Multivariate Probability Simulators for Limited Dependent Variable Models. *Journal of Econometrics* 58, 347-368.

Börsch-Supan, A., und R. Schnabel, 1998a, Social Security and Declining Labor Force Participation in Germany, *American Economic Review 88.2*, 173-178.

Börsch-Supan, A., und R. Schnabel, 1998b, Social Security and Retirement in Germany, in: J. Gruber and D.A. Wise (eds), *International Social Security Comparions*, University of Chicago Press: Chicago, in press.

Börsch-Supan, A., and P. Schmidt, 1996, Early Retirement in East and West Germany, mimeo, forthcoming in: R. Riphahn, D. Snower and K. Zimmermann, eds., *Employment Policy in the Transition to Free Enterprise: German Integration and Its Lessons for Europe*, London.

Brugiavini, A., 1998, Social Security and Retirement in Italy, NBER Working Paper No. 6155, Cambridge, Mass.

Bundesministerium für Arbeit und Sozialordnung (BMA), 1997, *Statistisches Taschenbuch*, Bonn: Bundespresseamt.

Burkhauser, R., 1991, An Introduction to the German Socio-Economic Panel For English Speaking Researchers, mimeo, Syracuse University

Daganzo, C., 1981, *Multinomial Probit*, New York: Academic Press.

Diamond, P., and J. Gruber, 1997, Social Security and Retirement in the U.S., NBER Working Paper 6097, forthcoming in: Gruber, J., and D.A. Wise (eds), *International Social Security Comparisons,* University of Chicago Press: Chicago.

Feldstein, M., und A. Samwick, A., 1996, The Transition Path to Privatizing Social Security, NBER Working Paper, Cambridge, Mass.

Geweke, J., 1991, Efficient Simulation from the Multivariate Normal Distribution Subject to Linear Inequality Constraints and the Evaluation of Constraint Probabilities, *Computer Science and Statistics: Proceedings of the Twenty-Third Symposium on the Interface*, 571-78.

Gruber, J., and D.A. Wise (eds), 1998, *International Social Security Comparions,* University of Chicago Press: Chicago, forthcoming.

Hajivassiliou, V., D. McFadden and P. Ruud, 1996, Simulation of Multivariate Normal Rectangle Probabilities and their Derivatives: Theoretical and Computational Results, *Journal of Econometrics 72(2),* 85-134.

Kapteyn, Arie, and Klaas de Vos, 1998, Social Security and Retirement in the Netherlands, NBER Working Paper No. 6135, Cambridge, Mass.

Kotlikoff, L.J., 1996, Simulating the Privatization of Social Security in General Equilibrium, NBER Working Paper, Cambridge, Mass.

Langdon, M.G., 1984, Methods of Determining Choice Probability in Utility Maximising Multiple Alternative Models, *Transportation Research* B18.

Lerman, S., and Manski, C., 1981, On the Use of Simulated Frequencies to Approximate Choice Probabilities, in: Manski, C., and McFadden, D. (eds.), *Structural Analysis of Discrete Data with Econometric Applications*, Cambridge, Mass.: MIT-Press.

Lumbsdaine, R.L., J.H. Stock, and D.A. Wise, 1992, Three models of Retirement: Computational Complexity versus Predictive Validity, in D.A. Wise, ed., *Topics in the Economics of Aging*, Chicago: University of Chicago Press, 16-60.

Meghir, C. and E. Whitehouse, 1997, Labour Market Transitions and Retirement of Men in the UK, *Journal of Econometrics 79*, 327-354.

Moffitt, T.E., 1987, Life Cycle Labor Supply and Social Security: A Time-Series Analysis, in: G. Burtless (ed.), *Work, Health, and Income for the Elderly,* Washington, D.C.: The Brookings Institution.

OECD, 1988, *Ageing Populations: The Social Policy Implications*, OECD: Paris.

Oshio, T., and N. Yashiro, 1998, Social Security and Retirement in Japan, NBER Working Paper No. 6156, Cambridge, Mass.

Palme, M., and I. Svensson, 1998, Social Security and Retirement in Sweden, NBER Working Paper No. 6137, Cambridge, Mass.

Pestieau, P., and J.-P. Stjins, 1998, Social Security and Retirement in Belgium, NBER Working Paper No. 6169, Cambridge, Mass.

Riphahn, R.T., and P. Schmidt, 1997, Determinanten des Ruhestandes: Lockt der Ruhestand oder drängt der Arbeitsmarkt? Jahrbücher für Wirtschaftswissenschaften, 48(1).

Rust, J., 1990, Behavior of Male Workers at the End of the Life Cycle: An Empirical Analysis of States and Controls, in: D.A. Wise (ed.) *Issues in the Economics of Aging*, Chicago: University of Chicago Press.

Rust, J., and C. Phelan, 1997, How Social Security and Medicare Affect Retirement Behavior in a World of Incomplete Markets, *Econometrica 65(4)*, 781-831.

Schmidt, P., 1995, *Die Wahl des Rentenalters - Theoretische und empirische Analyse des Rentenzugangsverhaltens in West- und Ostdeutschland*, Lang, Frankfurt.

Schnabel, R., 1997, Internal Rates of Return of the German Pay-As-You-Go Public Pension System, SFB 504 Discussion Paper: University of Mannheim.

Siddiqui, S., 1997, The Pension Incentive to Retire: Empirical Evidence for West Germany, *Journal of Population Economics 10(4),* 463-86.

Stern, S., 1997, Simulation-Based Estimation, *Journal of Economic Literature 35,* 2001-39

Stock, J.H., and Wise, D.A., 1990, The Pension Inducement to Retire: An Option Value Analysis, in: D.A. Wise (ed.) *Issues in the Economics of Aging*, Chicago: University of Chicago Press, 1990, 205-30.

Sueyoshi, G.T., 1989, Social Security and The Determinants of Full and Partial Retirement: A Competing Risk Analysis, NBER Working Paper 3113, Cambridge.

Verband deutscher Rentenversicherungsträger (VdR), 1997, *Die Rentenversicherung in Zeitreihen*, Frankfurt am Main.

Appendix:

Simulated maximum likelihood parameter estimates by number of replications

Number of replications	10	20	50	100	200	500	1000
Coefficient of option value	-0,0077	-0,0076	-0,0089	-0,0096	-0,0094	-0,0095	-0,0095
Random effect	1,82	1,96	2,18	2,29	2,26	2,26	2,25
Autocorrelation	0,69	0,69	0,71	0,69	0,70	0,70	0,70

Notes: See Table 3, Model 3.