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Heterogene Arbeit: Positive und Normative
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Zusammenfassung:

Abstract: There are theoretical foundations which allow hypothesizing on a positive association of life expectancy or retirement age with income. If both cannot be falsified, the relationship of income and the internal rate of return of a public pension system is not straight forward. By application of a partially linear model to micro data from the German public pension system, it is found that neither life expectancy, nor retirement age is monotonously increasing in income, as measured in benefit claims. The relation of benefit claims and duration under the benefit spell (which determines the rate of return) depends on the set of covariates. Including pensions for disabled individuals, three out of four specifications exhibit a duration decreasing in benefit claims.

JEL Klassifikation : H55, I12, H23
Schlüsselwörter : Public Pension System, Life Expectancy, Redistribution, Partially Linear Model
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Longevity and Redistribution in the German Pension System

Stefan Hupfeld*

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ABSTRACT: There are theoretical foundations which allow hypothesizing on a positive association of life expectancy or retirement age with income. If both cannot be falsified, the relationship of income and the internal rate of return of a public pension system is not straight forward. By application of a partially linear model to micro data from the German public pension system, it is found that neither life expectancy, nor retirement age is monotonously increasing in income, as measured in benefit claims. The relation of benefit claims and duration under the benefit spell (which determines the rate of return) depends on the set of covariates. Including pensions for disabled individuals, three out of four specifications exhibit a duration decreasing in benefit claims.

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1. Introduction

There are good reasons to believe in the often corroborated positive relationship between income and life expectancy. The case of Germany is considered by Reil-Held (2000), who basically compares different income quantiles. Additionally, there are theoretical foundations for such a phenomenon, among the earliest the model of health capital, introduced by Grossman (1972). Such a monotonous relation would have potentially unintented consequences for a public pension scheme organized as a pay-as-you-go system, namely resulting in redistribution from poor to rich, as Diamond (2003), pp. 87 shows. However, in the following essay it is shown that the mechanism that links income and life expectancy is far more complex, at least for the participants in a public pension system. The association between life expectancy, retirement age and duration with collected benefit claims is estimated by a partially linear model, taking endogeneity of the collected benefit claims regarding retirement age into account. Though stable relationships can be discovered, they are not of the kind that life expectancy or retirement age is monotonously linked to benefit claims. Additionally, there is no unique evidence for redistribution, since not life expectancy or retirement age alone drives the rate of return from the pension system, but the duration under the benefit spell. The latter is found to be either increasing or decreasing in collected claims, depending of the specification of the econometric model.

Theory There are several strands of literature which derive second best properties of a pension system and take individual heterogeneity into account. The first to be considered is Diamond (2003), who uses a framework of optimal taxation. The dimension in which individuals may differ is (firstly) disutility of labor, which is not observable by the social planner. The individual is free to choose whether to retire in the second or third period within his three-period life[1], and this decision depends on the individual degree of disutility of labor and on the taxation on work in the second period, which is implied by the pension system. Workers with a disutility of labor beyond a certain threshold prefer to retire early, and by maximizing a utilitarian welfare function, this is optimal and should actually be induced by taxing work in the second period. However, those who prefer to retire later should benefit in two ways, namely by a higher wage in the second period *and* by higher benefits in the third period. The result can be interpreted as a tax system which is contingent on age. The second possible dimension of heterogeneity is life expectancy, and the decision to be taken by the individuals is the same as under heterogeneous disutility of labor. A fourth period[2] is introduced, and the probability of reaching the fourth period as a retiree is taken as a measure of life expectancy. Assuming a negative relation between disutility of labor and life expectancy, the welfare maximizing pension scheme is more complicated: Those with

[1] The model is reducible to two periods, as everybody will work in the first period.
[2] Again, the model is reducible by one period.

very high life expectancy will be able to take advantage of the pension scheme, because they can easily retire later, but benefit from a relatively low implicit taxation of prolonged work. This relatively low implicit taxation of work in the second period is necessary in order to provide incentives to work for those at the margin (meaning, with an intermediary life expectancy). At the same time, those individuals with very short life expectancy suffer from a relatively low rate of return, which is necessary for exactly the same reason. Diamond (2003) shows that the unequal treatment that benefits those with high life expectancy can be (partially) overcome by the introduction of a private (and perfectly discriminating) annuity market, together with lump sum payments from the public pension system.

A comparable approach in the framework of optimal taxation is taken by Cremer and Pestieau (1996) and Cremer, Lozachmeur, and Pestieau (2004). The earlier contribution starts right away with the assumption of two dimensions in which individuals may differ, namely ability and risk (of incurring a loss), which are both unobservable. Note that these dimensions can basically be transformed into the dimensions Diamond (2003) assumes. However, the authors explicitly allow for different correlation structures. One correlation structure[3] indicates that a partial social insurance system alone will already reach the second best optimum. The later contribution uses the developed framework and explicitly focusses on retirement. The two unobservable properties are productivity and health, while the latter can be translated into the already known disutility of labor. The authors find that early retirement for those "in bad health" can be second best optimal (from a utilitarian point of view) and should actually be induced by implicit taxation in the pension system.

In the above mentioned strands of literature, all heterogeneity regarding "health" or similar concepts are taken to be exogenous. Grossman (1972) proposed a model of health capital which can be augmented by (timely or costly) investments. The benefits come directly in terms of utility, and indirectly in terms of better possibilities to be productive. The time of death is also implicitly defined by the stock of health capital, and under some assumptions (namely, stressing the investment property of health capital), a positive relation between productivity and health (and therefore life expectancy) can be derived.

Empirical Insights There is some literature on the different aspects of the present analysis. For the case of life expectancy in Germany, Reil-Held (2000) finds a positive relationship between income and life expectancy. Men in the lowest income quartile must expect to be outlived by those in the highest quartile by six years, which is estimated by Kaplan-Meier and Cox-estimation, even after controlling for several socioeconomic factors. Exemplarily, Deaton and Paxson (2004) can show that income reduces the risk of mortality in the United States.

[3]High ability ⇔ low disutility of labor is associated with high risk (of longevity) ⇔ high life expectancy. This is equivalent to the correlation assumed by Diamond (2003).

Mainly concerned with policy implications, Berkel and Börsch-Supan (2004) analyze the impact of different socioeconomic variables on retirement behavior and find that people with higher income or assets actually retire *earlier* after controlling for education.[4], which is not confirmed by Schils (2005), pp. 123, who finds a positive influence of hourly wages on retirement age.[5]

As a potential link in the background, health (or health capital) ties economic capacity with life expectancy. In a more direct fashion, health status is a determinant of retirement age. The evidence on a direct causal path from economic capacity to health is mixed; Adams, Hurd, MacFadden, Merrill, and Ribeiro (2004) e. g. find a health promoting causal relationship of wealth only for certain diseases, however, some of which are found to drive mortality in a second step, such that a causal link between wealth and mortality via health can be corroborated.

The German Public Pension System The pension system in Germany is today organized as a pay-as-you-go system. Participation in (and therefore contribution to) the German public pension system is mandatory for a large fraction of individuals, the largest share of which is constituted by the dependent labor force. Contributions are paid as a percentage of the relevant monthly gross income (if it is more than EUR 400). In 2005, this fraction amounts to 19.5%, of which one half each is paid by the employer and the employee. However, the relevant income is capped at EUR 5200 per month. Other forms of employment are subject to more complex terms of participation in the public pension system, e. g. self-employed can apply for participation.

Claims against the pension system are collected in *points*. Each point earned corresponds to the contributions based on one average yearly income, which is now (2005, all figures are for Western Germany) EUR 2415 per month. An individual earning exactly the average income for 40 years will thus collect 40 points. In times of no regular employment, additional points can be earned through certain times of child care, ill health or unemployment. This rather abstract claim is transformed into a monthly pension benefit, depending on the general development of wages and employment. In the first half of 2005, each point was worth EUR 26.13. The monthly benefit is then paid as annuity to the beneficiary. Benefits are paid for several reasons, the most prominent is the old-age pension, for which one is eligible at the age of 65. Contingent on a minimum of 35 years of collecting claims, men can apply already at age 62 or 63 (depending on their year of birth) for an old-age pension.[6] They have to suffer from a discount, namely 0.3% per month earlier than 65. Yet, although the discounts are intended to be actuarially fair, on average they still provide incentives for early retirement, see Börsch-Supan and Schnabel (1998). Under certain contingencies, benefits are paid to

[4]This phenomenon supports the conjecture of an income effect at the lower tail of the income distribution, see sections 5 and 6.

[5]To be precise, Schils (2005) finds that hourly wages decrease the hazard into retirement.

[6]Different, but similar rules apply to women.

individuals incapable of working as well as to surviving spouses and orphans. The legal basis for these rules can be found in SGB VI (German Social Security Code No. 6).

Organization of the Essay Section 2 will state the hypotheses explicitly. The data set is introduced with some descriptive statistics in section 3. Section 4 then introduces the estimation technique, namely the partially linear model, and covers related issues such as bandwidth choice and confidence bands. The implementation of the suggested methodology and the results are presented in section 5, and section 6 summarizes and concludes.

2. The Hypotheses

The main concern of this work is on two hypotheses and a corollary. For ease of exposition, they are clearly stated below.

Hypothesis 1 (*Age*) There is a positive relationship between socioeconomic status and life expectancy among the participants of the German public pension system. The socioeconomic status is measured by collected benefit claims (in points), whereas death of the individual is directly observed. The positive association leads *ceteris paribus* to redistribution to those with high benefit claims, since the internal rate of return of the pension system increases with life expectancy.

Hypothesis 2 (*Retirement Age*) There is a positive relationship between socioeconomic status and retirement age, which is due to the following reasons:

(i) Low income (and therefore low benefit claims) are associated with worse health. Individuals who collected relatively few claims hereby are forced to retire earlier due to their inability to work.

(ii) Relatively bad health (or low income) may be associated with high disutility of labor. Individuals with low benefit claims deliberately choose to retire earlier.

(iii) Following hypothesis 1 (*Age*) and despite discounts for early retirement, individuals with low benefit claims may choose to offset this disadvantage of shorter life expectancy by retiring earlier.

Corollary (*Duration*) If both hypotheses 1 and 2 cannot be rejected, the aggregate dependence of the duration under the pension benefit spell on socioeconomic status is not clear, and therefore it is not clear how (or if) the internal rate of return from the pension system is related to the collected benefit claims.

3. The Data

On the basis of the report of a commission[7] installed by the German Federal Ministry of Education and Research, the administrators of Germany's social security system were obliged to improve the cooperation among scientists and the various institutions of social security. The German public pension system, represented by the former Federation of German Pension Insurance Institutes (VDR, now: Deutsche Rentenversicherung Bund), began to publish their data in 2005. The data in use here is the collection of pension discontinuations due to death of the beneficiary[8], beginning in 1993 and ending in 2003. The data set contains a 10% stratified sample (based on the federal states) of all pensions that were discontinued, which adds to a total of 951,560 observations and 31 variables, described in table 1.

The data is not based on individuals, but on the pension, as the individual is not the main subject of interest for the pension system. Sometimes both concepts coincide, but accounting for benefits paid to widows and orphans, an individual may receive more than one pension at a time. All these double payments are excluded. Pensions paid to the insured himself cover both old-age pensions and disability pensions. The latter are transferred to old-age pensions, once the beneficiary is eligible. Both are not distinguishable in the data set. Additionally, the value of a benefit claim increases by the potential payment to widows and orphans.

Descriptive Statistics In the following, some descriptive statistics of the data set are presented. In table 2, mean and standard deviation are given for the variables of interest, firstly for the whole (raw, but corrected for missing values) data, and secondly for the adjusted data set as it is finally used for the estimation. At the bottom, the descriptive statistics are given for variables which are not contained in the data set, but which can be easily constructed. Histograms for age, retirement age, duration, and collected claims (based on the adjusted data set) are presented in figures 1 to 4.

The raw data set is adjusted in the following way: Firstly, all observations containing missing values concerning the date of birth and nationality are abandoned. Secondly, only pensions that are based on own claims are considered, so especially all widower's or orphan's pensions are discarded. There remain only those where the individual that paid contributions coincides with the individual that benefits directly from the pension. Finally, only male individuals who never worked in the former GDR are considered. Female workers who retired between 1993 and 2003 seem to exhibit a career pattern which is not comparable to the one of male workers. However, including the few (roughly 10,000) female pensioners who worked at least 25 years does not alter the later results substantially. People who worked in GDR and received benefit

[7]See: Kommission zur Verbesserung der informationellen Infrastruktur zwischen Wissenschaft und Statistik (2001).

[8]The "SUF Demographie Rentenwegfall 1993-2003 Versicherte, Witwer- und Witwen".

claims after reunification are excluded, as the variance of wages in the former GDR was relatively low and cannot be easily compared to wages paid in the former FRG. All individuals who worked less than 25 years and contributed to the pension system are excluded, as they do not represent the usual career pattern. However, the results presented later are robust against the inclusion of individuals who worked less. The adjusted data set contains 48,765 observations. Most of the dropped observations are females and/or pensioners who worked less than 25 years, see table 3 for a detailed description of the difference between raw and adjusted data.

4. Methodology

The major focus of this work is the relationship of two variables, namely age, retirement age, or duration under the benefit spell as dependent variable and collected claims as explanatory variable. Yet, there are more variables involved, which should be included in the analysis, e.g. a dummy for the type of health insurance, and the months spent in ill-health and unemployment.

Denote the respective dependent variables by the scalar y_i (with i being the individual observation), the main explanatory variable (a scalar again) is denoted by z_i, and the additional control variables can be found in the vector x_i, which is of dimension k. In order to infer the nature of the relationship between z_i and y_i, it is convenient to estimate this relation non-parametrically, circumventing an imposed linear or higher polynomial structure. However, it is generally not convenient to estimate the influence of all x_i non-parametrically as well, due to the so-called curse of dimensionality. This means that the requirement of observations increases exponentially with the number of regressors, or put differently, the approximation error increases more than proportionally if the number of observations is held constant, but the dimension of the regressor matrix is increased.[9] In order to inspect the combined influence of more than one variable graphically, one is bound to two regressors that enter the model non-parametrically anyhow.

Partially Linear Model Accounting for the trade-off between imposed structure and the necessity of additional control variables, a partially linear model is applied. This has the following form:

$$y_i = f(z_i) + x_i'\beta + \epsilon_i \tag{1}$$

The function f is not known, only smoothness is assumed. The parameter vector β is not known either. In order to approach an estimation technique, rewrite the partially linear model in terms of expectations, conditional on z_i:

[9] See e. g. Yatchew (2003), p. 17.

$$E(y_i|z_i) = f(z_i) + E(x_i|z_i)'\beta \tag{2}$$

The conditional expectations are now estimated non-parametrically, e. g. by fitting a local polynomial. Denote the estimates by

$$\begin{aligned}
\widehat{E(y_i|z_i)} &=: m_y(z_i) \\
\widehat{E(x_{1i}|z_i)} &=: m_{x1}(z_i) \\
&\vdots \\
\widehat{E(x_{ki}|z_i)} &=: m_{xk}(z_i)
\end{aligned}$$

and

$$(m_{x1}(z_i) \quad \ldots \quad m_{xk}(z_i))' =: m_x(z_i). \tag{3}$$

The partially linear model in terms of conditional expectations of equation 2 can then be rewritten as

$$y_i - m_y(z_i) = [x_i - m_x(z_i)]'\beta + \epsilon_i, \tag{4}$$

and β can be estimated by least squares. Denoting the estimate by $\widehat{\beta}$ and using equations 1, 2 and the definition 3, an estimate for $f(z_i)$ is finally obtained by

$$\widehat{f}(z_i) = m_y(z_i) - m_x(z_i)'\widehat{\beta} \tag{5}$$

However, note that the elements of the partially linear model can only be identified under two restrictions[10], namely

$$E(\epsilon_i|z_i, x_i) = 0 \tag{6}$$

and the absence of a constant in the parametric regressor vector x_i. The first condition will be violated once y_i and z_i are endogenous variables. The latter is due to the fact that $f(z)$ is left unspecified, such that any constant term in x_i can not be distinguished from a shift of $f(z)$.

[10] See e. g. Pagan and Ullah (1999), p. 198.

Endogeneity Suppose that z_i is endogenous in the sense that the identifying restriction in equation 6 is violated. However, there exists a variable $\widetilde{z}_i$ which does satisfy this restriction and which is associated with the original z_i by

$$z_i = \widetilde{z}_i\theta + u_i. \tag{7}$$

Under the further assumption of linearity of $E(\epsilon_i|z_i, u_i) = u_i'\rho$, the relationship among the residuals can be expressed by

$$\epsilon_i = u_i'\rho + \nu_i, \tag{8}$$

where ϵ_i is the residual of the model in equation 1. The endogeneity-adjusted model can then be written[11] as

$$y_i = f(z_i) + u_i'\rho + x_i'\beta + \nu_i \tag{9}$$

and finally be estimated semi-parametrically as well, with u_i entering the model as additional parametric regressor. As u_i is not directly observable, it has to be replaced by an estimate, namely the residual of equation 7 estimated by least squares. The significance of IV-residual in the final partially linear estimation will falsify the hypothesis of exogeneity of the respective regressor.

Local Polynomial Estimation and Bandwidth Choice The non-parametric estimates $m_x(z_i)$ and $m_y(z_i)$ still desire choices to be taken that influence the shape of the analyzed relation, namely

(i) The degree p of the local polynomial: The most prominent choices are $p = 0$ (which leads to the Nadaraya-Watson estimator) or $p = 1$. Consider the non-parametric estimator $m_y(z)$, for example[12]. The estimator minimizes one of the following terms, depending on either $p = 0$ or $p = 1$ (the extension to polynomials of higher order is straight forward):[13]

$$p = 0: \qquad m_y^{(0)}(z_i) = \arg\min_{m^{(0)}} \sum_{i=1}^{n} \left(y_i - m_y^{(0)}(z_i)\right)^2 K\left(\frac{z_i - z}{h}\right) \tag{10}$$

$$p = 1: \qquad m_y^{(1)}(z_i) = \arg\min_{m^{(1)},\beta} \sum_{i=1}^{n} \left(y_i - m_y^{(1)}(z_i) - (z_i - z)\beta\right)^2 K\left(\frac{z_i - z}{h}\right) \tag{11}$$

[11]See Yatchew (2003), pp. 87 and Blundell and Duncan (1998).
[12]The following applies to the estimation of $m_x(z)$ as well.
[13]Compare e. g. Pagan and Ullah (1999), pp. 93.

The estimator $m_y^{(\cdot)}(z)$ therefore either fits a local constant or a local line around x_i, weighting the neighboring observations around x_i with the kernel function $K(\cdot)$. Note that (whenever applicable), the respective objective functions are also minimized with respect to the local β. In the case of $p = 1$, the asymptotic bias of the estimated function is zero, which is not the case under $p = 0$, compare Mittelhammer, Judge, and Miller (2000) , pp. 622. As Loader (2004) shows, the asymptotic bias will vanish whenever the degree of the polynomial is odd, and especially the bias at the boundaries of the data set will decrease, compared to the Nadaraya-Watson estimator.

(ii) The weighting kernel $K(\cdot)$: There are a couple of possibilities for the weighting scheme, the most prominent being the Gaussian kernel and the Epanechnikov kernel:

$$K^{\text{GAUSS}}\left(\frac{z_i - z}{h}\right) = 2\pi^{-\frac{1}{2}} \exp\left(\frac{-\left(\frac{z_i-z}{h}\right)^2}{2}\right) \tag{12}$$

$$K^{\text{EPAN}}\left(\frac{z_i - z}{h}\right) = \begin{cases} \frac{3}{4}\left(1 - \left(\frac{z_i-z}{h}\right)^2\right), & \left|\frac{z_i-z}{h}\right| < 1 \\ 0, & \text{else} \end{cases} \tag{13}$$

The latter proves to be the efficient one, see e. g. Pagan and Ullah (1999), p. 28. Using the Gaussian kernel, no observation (no matter how far from x_i) ever receives a weight of zero, which may cause some computational burden with large data sets.[14]

(iii) The bandwidth h used with the weighting kernel: A bandwidth chosen to be too high will leave the estimate "over-smoothed" and potentially ignore specific patterns, whereas an under-smoothed estimate may hide the pattern of interest behind erratic components, leading in the limit (as $h \to 0$) to an exact replication of the unfitted data. This phenomenon is known as the bias-variance-tradeoff[15]. Nowadays the standard procedure of choosing the bandwidth is cross-validation. Thereby the mean integrated square error ($MISE$) is asymptotically and indirectly minimized by minimizing the cross-validation function

[14] However, a large number of zero weights may yield a different computational difficulty, namely due to singular matrices; note that the local β-vector in equation 11 can be estimated by weighted least-squares, compare Loader (2004):

$$\widehat{\beta} = (z'Wz)^{-1}z'Wy,$$

where W is a diagonal matrix with the respective kernel weights on the main diagonal. The matrix $z'Wz$ may be singular for certain outcomes of the kernel weights, and thus not invertible.

[15] Compare e. g. Yatchew (1998).

$$CV(h) = \frac{1}{n}\sum_{i=1}^{n}\left(y_i - m_{y,-i}(z_i, h)\right)^2 \tag{14}$$

with respect to h, where $m_{y,-i}(z_i, h)$ is the leave-one-out estimator, i. e. the estimator for y_i based on the whole data except the i-th observation. This basically means that the local polynomial fit has to be estimated n times for each potential bandwidth within a discrete set of candidates.

Alternatively, the optimal bandwidth can be approximated by a rule of thumb, which may be advisable while using large data sets. Compare e. g. Pagan and Ullah (1999), p. 103, who propose the bandwidth to be of the order of magnitude $n^{-1/5}$. There are several rule-of-thumb or "plug-in" methods which specify the bandwidth more explicitly, e. g. Loader (1999) and Loader (2004), who propose the optimal bandwidth to be

$$h = \left(\frac{\sigma^2(b-a)^2\int K(v)^2 dv}{n\left(\int v^2 K(v)dv\right)^2\int m''(x)^2 dx}\right)^{-1/5}, \tag{15}$$

where σ^2 is the error variance, $m''(x)$ is the second derivative of the estimated function, and a and b are the lower and upper bounds of x. Using a first stage or pilot estimate, the error variance can be estimated by

$$\widehat{\sigma^2} = \frac{1}{n - 2\nu_1 + \nu_2}\sum_{i}^{n}(y_i - m(x_i))^2, \tag{16}$$

with ν_1 and ν_2 adjusting the degrees of freedom (see Loader (2004) for the computation). The second derivative $m''(x)$ of the estimate is obtained by fitting a local *quadratic* to the data, hence by solving

$$m_y^{(2)}(z_i) = \arg\min_{m^{(1)},\beta_1,\beta_2}\sum_{i=1}^{n}\left(y_i - m_y^{(1)}(z_i) - (z_i - z)\beta_1 - (z_i - z)^2\beta_2\right)^2 K(\cdot)\,. \tag{17}$$

An estimate for the second derivative is then given by $2\widehat{\beta}_2$. Using the Gaussian density function as the weighting kernel, we have

$$\int K(v)^2 dv = 0.282 \quad (18)$$

$$\left(\int v^2 K(v) dv\right)^2 = 1. \quad (19)$$

However, there remains a pilot bandwidth to be chosen, as the respective $\widehat{\beta}_2$ and $m''(x)$ are sensitive to the bandwidth as well.

Conditional Moment Estimation with Dummies Once the respective column in the matrix of controls consists of a dummy variable, the propensity score can easily be estimated by a Probit model. Admittedly, such a binary choice model assumes a parametric structure.

Confidence Interval around $\widehat{f}(z_i)$ Though inference is generally not a big issue due to the sheer size of the data set, confidence bands around the estimated function $\widehat{f}(z_i)$ can be bootstrapped. The procedure used for the bootstrapped confidence interval is borrowed from Yatchew (2003), p. 161. First, an over-smoothed and under-smoothed estimate $\underline{\widehat{f}}(z_i)$ and $\widehat{\widehat{\overline{f}}}(z)$ are produced, using $0.95h$ and $1.05h$. Based on the under-smoothed estimate, the residuals $\widehat{\epsilon_i} = y_i - \underline{\widehat{f}}(z_i)$ are calculated. From all $\widehat{\epsilon_i}$, the new errors ϵ_i^B are drawn with replacement, and the bootstrap data set is constructed by

$$y_i^B = \widehat{\widehat{\overline{f}}}(z_i) + \epsilon_i^B. \quad (20)$$

Based on this bootstrap sample and the original bandwidth h, a new estimate of the nonparametric function $\widehat{f_B}$ is calculated. The drawing of ϵ_i^B and the subsequent estimation of $\widehat{f_B}$ is repeated several times, and the 95% confidence interval finally lies between the 0.025% and the 0.975% quantile of the empirical distribution of all $\widehat{f_B}$.

5. Implementation and Results

General Specification In order to capture a variety of potential influences, a total of five different specifications is estimated. The first specification only consists of a local linear fit without further explanatory variables, whereas the following four variants contain different sets of X.

It can be argued that the sum of claims (hence, the PSEGPT variable) as main explanatory variable is endogenous with respect to the retirement age and therefore also with respect to the duration under the benefit spell. So a proper instrument has to be found, which is not driven by the end of the working career. If the claims which are gathered beyond a certain age are excluded, the sum of these adjusted claims would not be affected by the retirement age, and would therefore be exogenous. Hence, an instrument z_{adj} for z can be constructed by

$$z_{\text{adj}} = z - \frac{z}{t} t_{52}, \tag{21}$$

where t is the sum of years in which claims have been collected, and t_{52} is the sum of years in which claims are collected beyond the age of 52. The correlation between the two variables is larger than 0.999.

The following computations are based on a local polynomial of degree $p = 1$, used in all estimations of conditional expectations. In order to avoid singular matrices, the weighting scheme used in the local regressions is Gaussian. Often, the local polynomial regression is not performed on all observations, but only on a grid of a few tens' or hundreds of data points. However, this faster procedure is not feasible once the results (the estimated conditional expectations) are used again, as is done here in a least squares estimation, and where the full set of observations is necessary. However, all conditional moments $E(X|z) = m_x(z_i)$ are the same in the following regressions, so they only have to be estimated once. For the estimation of confidence intervals around duration (following specification 4), the distribution and its respective quantiles of the estimated function $\widehat{f}(z_i)$ is based on 100 bootstrap samples, and the 95% confidence band around $\widehat{f}(z_i)$ is simply based on the 0.025% and 0.975% quantile. The number of resamples is relatively low, which is, however, due to the computational burden of data-driven procedures.

Specification 0 As a benchmark, only the pure nonparametric relationship of life expectancy, retirement age or duration with the sum of claims or the sum of adjusted claims is considered. The estimated equation reduces to

$$y_i = f(z_i) + \epsilon_i. \tag{22}$$

The results are given in figures 5 to 7. The pilot bandwidths are chosen to be

$$h_{\text{pilot}} = 5n^{-1/5} = 0.577. \tag{23}$$

Specification 1 The first partially linear estimate contains the smallest set of covariates. The model as implemented can be written as follows:

$$y_i = f(z_i) + \beta_1 x_{1i} + \beta_2 x_{2i} + \beta_3 x_{3i} + \epsilon_i, \tag{24}$$

where the variables of the regressions are organized as follows:

- y_i Age at death, measured in years and months (the latter as fractions of a year) **or** Retirement Age (in years) **or** Duration under the benefit spell (in years)
- z_i Collected benefit claims (in points)
- x_{1i} Type of health insurance (dummy: $1 =$ private, $0 =$ public)
- x_{2i} Time, in which claims have been earned due to ill-health and/or usage of certain types of rehabilitation (in months)
- x_{3i} Time of unemployment (in months)

For the case of y_i being retirement age or duration, the model is augmented with the residual of the following first stage regression,

$$z_i = \widetilde{z}_i\theta + u_i, \tag{25}$$

with z_i being the collected claims PSEGPT and $\widetilde{z}_i$ being the adjusted claims.

Specification 2 The second specification contains all variables as in specification 1, with one addition: A dummy is introduced, which indicates whether the year of retirement was 1992 or later. In this year, actuarially fair discounts were introduced for those who chose early retirement. Ceteris paribus, retirement age is expected to rise.

Specification 3 The third specification contains again all variables as in specification 1, with additional cohort dummies. All individuals are divided into 10 year cohorts, starting in 1900. The first cohort is used as reference group, so the estimated equation is augmented with five additional dummies.

Specification 4 The last specification contains all variables together, namely those of specification 1, 2, and 3.

Unfortunately, a measure for education is not available. Though not in included in the data set, an indicator for the highest degree of education has been surveyed in recent years, yet, it consists of 98% missing values. In the present data set an indicator for the profession is provided (the variable BFKL); however, more than 93% are so-called reassessments, missing, or are simply coded as "unskilled workers".

Results: Partially Linear Regression of Age on Benefit Claims Compare table 8 and figures 5 and 8 for the results of the partially linear model with age as dependent variable.

In the local linear estimation without additional regressors, a monotonously positive relationship can only be discovered for individuals with claims beyond approximately 38 points. However, note that beyond 70 points there very few observations, such that the majority of the observations can be found in the U-shaped part of the estimated function. This pattern can be found again in all partially linear estimation results, however, less pronounced. A private health insurance might be taken as an additional indicator for high income, as only high income individuals are allowed to be privately insured. However, conditional on higher income, a private health insurance is also an indicator for good health, as private health insurers can discriminate their customers regarding their health status, such that it might not be worthwhile for those in bad health to buy private insurance. Accordingly, the existence of a private health insurance is associated with higher life expectancy. Also unemployment, which is either an indicator for low productivity, or even exerts direct pressure on an individual's status (regarding health, income and social status in general), lowers life expectancy in all estimated specifications. Time spent in ill-health also lowers life expectancy, except in specification 3. Including cohort effects does not change the main results; yet, the coefficients of the respective cohort dummies cannot be directly interpreted in terms of rising life expectancy through the years (which does not show up in the results as well). The interpretation has be conditioned on the fact that there are no birth cohorts considered, but rather a cohort of individuals who have died; so a cohort effect is always conditional on reaching at least the year 1993, no matter when the individual was born, which explains the negative coefficient of the youngest cohort (those born in 1950 and later).

The U-shaped relation between collected claims and life expectancy especially around the mean is astonishing and calls for explanation. Low life time earnings in terms of collected claims against the public pension system can either be explained by low yearly earnings, or by a short contribution period. An artifact of the German public pension system is that certain professions which are traditionally performed self-employed are eligible to opt out of the public pension system. Among those professions are some which are typically well-paid, such as advocates, medical doctors or careers in the public service. Individuals in these professions contribute more by chance than deliberately to the public pension system in some (usually early) stages in their career, such that they appear to have had a low income, but actually belong to the better (and best) educated and earning individuals, though still included in the lowest $PSEGPT$ region. However, this conjecture cannot be corroborated, as we consider only individuals with at least 25 years of work in a job where contributions to the public pension system are mandatory. So a second conjecture utilizes the relatively rigid hourly wages in

Germany, especially at the lower end of the income distribution. An individual with low life time earnings may have worked fewer hours per week than an individual with intermediary life time earnings, while facing a comparable hourly wage. So the increase in life time earnings from 10 to 20 points, e. g., is not due to an increase in productivity, but due to an increase in hours worked, which by itself does not add to life expectancy.

To sum up, hypothesis 1 (Age) can only be corroborated for observations with more than 38 collected points of benefit claims and must be rejected in general.

Results: Partially Linear Regression of Retirement Age on Benefit Claims This estimation comes potentially twofold, as the data set offers different possibilities to interpret the retirement age. Firstly, retirement age is the age at which an individual receives his first benefit payment, which can be either the old-age pension or a disability pension. Secondly, the age at entry into the "actual" pension is observed, which constitutes the benefit payments that are paid until the end of the pensioner's life, and which can again be either one of both named types. In some cases, both pensions might coincide, but not necessarily in general, such that an individual might receive two or even more pension types during his life as a beneficiary. In order to capture the meaning of hypothesis 2, the first concept is preferred (this applies to the estimation regarding duration as well). Consider a worker who becomes unable to work at the age of 55 and receives the respective pension benefit. At the age of 65, the benefits are changed into the old-age pension; so if the second concept of retirement age were applied, the reaction to an incentive (or necessity) for people with ill-health to retire early could only be observed partially.

The results of these estimations can be found in the figures 6 and 9, as well as in table 9. The final construction of the nonparametric (local linear) relationship between retirement age and benefit claims yields again a strongly pronounced U-shaped pattern: A monotonously increasing relationship can only be observed for those individuals who collected more than 38 points of benefit claims, such that in specification 0, hypothesis 2 must generally be rejected.

Including additional regressors, the shape of the relationship changes. Although a monotonously increasing relationship can still not be found, the estimated minimum retirement age is at approximately 20 points and increasing thereafter. Individuals with private health insurance retire earlier, which is not in line with expectations. The introduction of discounts for those who retire early, however, resulted in adaption of the individuals, which retired later after the legislation in 1992. Yet, the reaction is smaller as measured by Berkel and Börsch-Supan (2004), e. g., while the difference crucially depends on the different data set and sub-sample. Time spent in ill-health lowers the retirement age, which justifies hypothesis 2, whereas unemployment actually increases the retirement age. This might be due to institutional factors, depending on the possibilities of substitution between early retirement and unemployment at the end of a career. Also possible is a dominating income effect at the lower end of the income distribution, such that certain individuals cannot afford to retire earlier as they do not want to face *any* discount. Finally, even in the partially linear specifications, hypothesis 2 must be rejected in general, although the upward sloping relationship dominates and encompasses the majority of individuals (compare again the histogram of claims, figure 4).

Results: Partially Linear Regression of Duration under the Benefit Spell on Benefit Claims The results of the estimations can be found in figures 7 and 10, as well as in table 10. The local linear estimate exerts a strong, inverted U-shaped pattern, which does not appear in any of the partially linear specifications.

First of all, the IV-residuals appear to have a coefficient significantly different from zero, such that the assumption of endogeneity of PSEGPT is justified also with duration as left-hand variable. Specifications 1, 3, and 4 yield a downward sloping curve, meaning that there is no redistribution from poor to rich, as the duration and therefore the rate of return decreases slightly with income. Yet, specification 2 indicates exactly the opposite, namely a slightly increasing relationship between duration and collected claims. A private health insurance is associated with a longer duration. Time spent in ill-health is positively related to duration, off-setting the corresponding lower life expectancy. Hence, paying pensions to early retirees who are potentially in bad-health enhances their rate of return. Unemployment lowers the duration, which may be due to institutional factors. A dramatic decrease in the duration is caused by the introduction of discounts for early retirees - however, this decrease is (in absolute terms) larger for retirees with more than the average amount of collected claims, such that ceteris paribus the status of poorer individuals was relatively improved.[16] For specification 4, a confidence interval is constructed by the method proposed in section 4; see figure 11.

To sum up, there is neither clear evidence for an upward sloping relationship between claims and duration, nor for a downward sloping one. However, all specifications indicate that the duration is not orthogonal with respect to collected claims.

Robustness Note that all patterns are not an artifact of the estimation technique; the general shape is corroborated by fitting a global polynomial of degree three by least squares, and by replacing the local linear estimates of the conditional moments by lowess or Nadaraya-Watson estimates. Plots of the distribution of the dependent variables by PSEGPT-quantiles are not monotonously ordered, which also indicates a more complex relation. Even the inclusion of women with at least 25 years of contributions does not alter the major results; however, the number of women in the analyzed cohorts with such a career pattern is comparably low ($\sim$ 10,000).

[16] Less than average PSEGPT is observed for 24,928 observations, 23,837 observations have higher claims. The respective coefficients for the introduction of actuarial fairness in 1992 are in specification 2: -7.702 for the "poor" and -9.306 for the "rich", in specification 4: -5.026 for the "poor" and -7.697 for the "rich". The differences are significant.

The results are also relatively robust against the selection of the sub-sample used in the estimation. If the desired minimum of contribution is reduced to 5 years and including women, the results shown in appendix B are obtained. The respective estimates are based on 115,114 observations. The four specifications of the partially linear model remain the same, except of the introduction of sex as an additional dummy variable. The relationship of life expectancy and retirement age with benefit claims does not change in its general shape, however, the local linear estimate of duration (given benefit claims) is now monotonously increasing. The patterns derived from the partially linear estimation are again very similar to the ones shown in appendix A, only that now all four specifications yield a *decreasing* relationship of duration and benefit claims. Out of 84 coefficients which are estimated, 23 change (significantly) their sign due to the change of the sub-sample.

6. Summary

Life expectancy is not monotonously increasing in income (as measured by claims against the public pension system), neither is retirement age. The relationships we find are nevertheless robust. Especially around the mean of the collected claims, the relation between life expectancy and collected claims is U-shaped, which calls for an explanation. It can be shown that this pattern can be observed for the estimated influence of collected claims on retirement age as well, and that this pattern is neither artefact of the data set, nor of the estimation procedure. Similar (however, also rougher) findings can be derived by fitting a global polynomial and by survival analysis.

A conjectured explanation can be found in the rigidity of hourly wages in lowest income groups: Most jobs which give rise to mandatory contributions to the pension system are subject to collective wage bargaining. It can be argued that especially the lowest wages are higher than wages which would arise on a perfect labor market. If an individual belongs to the very bottom of the distribution of benefit claims (but still has worked for more than 20 years), this is due to relatively low hourly wages *and* few weekly hours of work. In order to move to the right on the claims distribution, the increase in hourly wages is potentially small or close to zero; yet, individuals simply work more hours, which does not add to life expectancy. Only beyond a certain level (when full-time jobs are reached), higher income is associated with higher hourly wages and not with more work time, which then increases life expectancy. Unfortunately, this conjecture cannot be analyzed directly, as hours of work are not surveyed in this setting.

The estimated relationship between retirement age and benefit claims also shows a U-shaped pattern. Although the non-actuarial discounts seem to encourage early retirement at the margin, this might not be true for those at the bottom of the income distribution: Every discount, even if it is relatively low, is not affordable due to a required (or desired) minimum level of the pension benefits. Therefore, individuals at the bottom of the income distribution choose to retire late, although they might suffer from high disutility of labor. Only individuals in the intermediary sections of the income distribution can actually afford to retire early, whereas those at the top have a low disutility of labor (and are in good health), and accordingly retire late.

Finally, in three of four specifications the duration under the benefit spell is negatively associated with claims, and only in one specification a positive relationship can be found. A general mechanism which links income to a higher rate of return (or less implicit taxation) from the public pension system can therefore not be corroborated. However, since the distributions of the duration and the collected claims are not independent, the pension system is not redistributionally neutral.

Yet, there are two major factors which have not been considered yet. Firstly, the whole analysis is conditional on *reaching* retirement age. Clearly, many individuals die without ever having received benefits, and their fraction might depend on the size of collected benefits. However note that all individuals are included who receive disability benefits, which weakens the counter argument. Secondly, pensions paid to widows and orphans are excluded, although their potential payment constitutes an asset to the (altruistic) contributor. Unfortunately, within the used data set, their value cannot be assessed individually.

A. Results including Men with at least 25 Years of Contribution

Variables in Data Set		
Name	**Description**	**Characteristics** (if not self-explanatory)
SK	Label of Dataset	always 90
JA	Year of Report	
CASE	Case Number	
FMSD	Marital Status	
GBJAVS	Year of Birth (Insurant)	
GBMOVS	Month of Birth (Insurant)	
GEVS	Sex (Insurant)	
SAVS	Nationality (Insurant)	
BFKL	Profession	2 Digits
BLAND	Residence (Federal State)	
REGBEZ	Residence (District)	
GBJABC	Year of Birth (Beneficiary)	
GBMOBC	Month of Birth (Beneficiary)	
GEBC	Sex (Beneficiary)	
SABC	Nationality (Beneficiary)	
RTBE1	Year of 1st Pension Payment	
ZTPR1	Year of Actual Pension Payment	
RTWF1	Year of Discontinuation	
RTWF2	Month of Discontinuation	
AT	Type of Health Insurance	i. e., private or public insurance
EKAH	Own Income if widowed	joint payment of own pension and widower's pension
ZLKL1	Number of Children	if parenting is assessed, up to 5
FRGLD	Foreign Contributions	especially former GDR
PSEGPT	Sum of Collected Claims	in points
WIRTZQ	Bonus for Children if widowed	in points
BYVL	Full Years of Contribution	up to 45
AUAZ	Ill-Health	up to 48 (months)
AJAZ	Unemployment	up to 120 (months)
FRGMO	Foreign Contribution	up to 45 (years)
POP	Population Code	own or widower's pension
SATZ	Type of Dataset	always 2

Table 1: Variables **FDZ-RV-SUFRTWF93VWITD** to **FDZ-RV-SUFRTWF03VWITD**

Descriptive Statistics				
Variable	**Raw Data**		**Adjusted Data (if feasible)**	
	Mean/Fractions	**St. Dev.**	**Mean/Fractions**	**St. Dev.**
FMSD	not married: 8.11%			
	married: 18.93%			
GBJAVS	1917.87	11.57		
GEVS	female: 37.23%			
SAVS	Germany: 95.67%			
	EU-15 1.74%			
	Italy: 0.85%			
	Turkey: 0.24%			
BLAND	West: 76.78%		West: 91.86%	
	East: 18.99%		East: 1.77%	
	foreign: 3.59%		foreign: 3.05%	
GBJABC	1918.93	10.76	1934.70	5.80
GEBC	female: 61.97%		female: 0.00%	
SABC	Germany: 95.76%		Germany: 90.93%	
	EU-15 1.76%		EU-15 2.60%	
	Italy: 0.83%		Italy: 1.32%	
	Turkey: 0.22%		Turkey: 0.76%	
RTBE1	1978.47	11.45	1993.00	5.54
ZTPTR1	1987.78	9.20	1995.70	2.78
AT	private: 4.50%		private: 10.80%	
	public: 86.23%		public: 78.59%	
FRGLD	none: 94.73%		none: 100.00%	
	former GDR: 0.44%			
PSEGPT	32.53	19.12	43.22	13.05
BYVL	5.24	11.99	38.11	5.98
AUAZ	0.57	2.97	2.99	6.39
AJAZ	1.25	8.16	9.24	20.32
FRGMO	0.33	3.08	0.06	1.16
POP	insurant: 30.80 %		insurant: 100%	
Generated Variables:				
AGE	80.38	11.31	64.84	6.10
RET. AGE	59.54	8.15	57.75	5.25
DUR	20.32	11.18	7.10	5.63
Ret. $\geq$ 1992			74.41%	
Cohort 1900-1909			0.10%	
Cohort 1910-1919			0.27%	
Cohort 1920-1929			18.03%	
Cohort 1930-1939			63.99%	
Cohort 1940-1949			15.53%	
Cohort >1950			2.08%	

Adjustment of Data Set		
Variable	**Discarded Obs.**	**Total Number of Obs.**
POP	if POP = 2 (Widow or Orphan)	262,465
GBJABC	if Missing	463
GBMOBC	if Missing	600
GEBC	if Female	549,782
SABC	if Missing	1,309
RTBE1	if Missing	42,582
ZTPTR1	if Missing	51,521
FRGLD	if FRGLD = 13 (GDR)	3,662
FRGLD	if FRGLD = else (foreign)	39,417
PSEGPT	if PSEGPT $\geq$ 86	1,532
BYVL	if BYVL $<$ 25 (Contribution)	773,632

The number of observations is reduced to 48,765.

Table 3: Adjustment of Data Set

Expectations, conditional on PSEGPT Estimates by Probit		
Variable	**Constant**	**Coefficient**
Health Insurance	-7.699	-7.946
Actuarial Fairness ('92)	0.396	0.006
Cohort 1910-1919	-0.748	-49.512
Cohort 1920-1929	0.075	-7.102
Cohort 1930-1939	-0.656	-12.627
Cohort 1940-1949	-1.180	-7.162
Cohort $>$ 1950	-2.948	-9.998
IV Estimation Step 1 Dependent Variable: PSEGPT		
Variable	**Coefficient**	**St. Error**
PSEGPT-ADJ	1.010	0.000

Table 4: First Stage Results

Life Expectancy Partially Linear Estimates				
	Specification			
	1	**2**	**3**	**4**
Health Insurance	0.000	0.001***	1.692***	2.048***
	(0.000)	*(0.000)*	*(0.162)*	*(0.156)*
Actuarial Fairness	—	-5.463***	—	-2.696***
(1992)		*(0.062)*		*(0.065)*
Cohort 1	—	—	-2.253***	-1.821***
(1910-1919)			*(0.200)*	*(0.190)*
Cohort 2	—	—	12.082***	10.467***
(1920-1929)			*(0.370)*	*(0.364)*
Cohort 3	—	—	7.008***	6.380***
(1930-1939)			*(0.371)*	*(0.351)*
Cohort 4	—	—	-1.238***	-1.376***
(1940-1949)			*(0.362)*	*(0.336)*
Cohort 5	—	—	-6.736***	-7.119***
($>$ 1950)			*(0.174)*	*(0.168)*
Ill-Health	-0.074***	-0.009***	0.030***	-0.004*
	(0.004)	*(0.004)*	*(0.003)*	*(0.003)*
Unemployment	-0.009***	-0.004***	-0.002***	-0.001
	(0.001)	*(0.001)*	*(0.001)*	*(0.010)*
R^2	0.006	0.162	0.539	0.569

$n = 48,765$, Newey-West Standard Errors in Paranthesis

*** denotes significance on 99% level, ** on 95% level, and * on 90% level.

Table 5: Life Expectancy — Results of Partially Linear Estimation

Retirement Age

Partially Linear Estimates

	Specification			
	1	**2**	**3**	**4**
IV-Residual	15.848***	15.840***	15.583***	15.384***
	(0.074)	*(0.070)*	*(0.074)*	*(0.073)*
Health Insurance	-0.001***	-0.001***	-0.336***	-0.325***
	(0.000)	*(0.000)*	*(0.025)*	*(0.025)*
Actuarial Fairness	—	0.017	—	0.219***
(1992)		*(0.031)*		*(0.031)*
Cohort 1	—	—	-1.201***	-1.221***
(1910-1919)			*(0.036)*	*(0.038)*
Cohort 2	—	—	2.681***	2.831***
(1920-1929)			*(0.072)*	*(0.079)*
Cohort 3	—	—	2.980***	3.037***
(1930-1939)			*(0.068)*	*(0.072)*
Cohort 4	—	—	3.052***	3.004***
(1940-1949)			*(0.071)*	*(0.072)*
Cohort 5	—	—	-1.641***	-1.693***
(> 1950)			*(0.068)*	*(0.068)*
Ill-Health	-0.010***	-0.010***	-0.017***	-0.016***
	(0.003)	*(0.003)*	*(0.003)*	*(0.003)*
Unemployment	0.012	0.012***	0.011***	0.011***
	(0.074)	*(0.001)*	*(0.001)*	*(0.001)*
R^2	0.839	0.839	0.859	0.859

$n = 48,765$, Newey-West Standard Errors in Paranthesis

*** denotes significance on 99% level, ** on 95% level, and * on 90% level.

Table 6: Retirement Age — Results of Partially Linear Estimation

Duration Partially Linear Estimates				
	Specification			
	1	**2**	**3**	**4**
IV-Residual	-5.164***	-0.692***	-10.737***	-4.424***
	(0.115)	*(0.085)*	*(0.147)*	*(0.147)*
Health Insurance	0.000	0.002***	1.063***	0.721***
	(0.000)	*(0.000)*	*(0.124)*	*(0.075)*
Actuarial Fairness	—	-8.868***	—	-6.939***
(1992)		*(0.052)*		*(0.084)*
Cohort 1	—	—	-1.405***	-0.754***
(1910-1919)			*(0.159)*	*(0.101)*
Cohort 2	—	—	8.938***	4.180***
(1920-1929)			*(0.314)*	*(0.231)*
Cohort 3	—	—	3.889***	2.093***
(1930-1939)			*(0.310)*	*(0.201)*
Cohort 4	—	—	-2.865***	-1.361***
(1940-1949)			*(0.278)*	*(0.157)*
Cohort 5	—	—	-3.098***	-1.479***
(> 1950)			*(0.130)*	*(0.088)*
Ill-Health	0.174***	0.076***	0.090***	0.057***
	(0.005)	*(0.004)*	*(0.004)*	*(0.004)*
Unemployment	-0.019***	-0.010***	-0.013***	-0.009***
	(0.001)	*(0.001)*	*(0.001)*	*(0.001)*
R^2	0.162	0.541	0.449	0.601

$n = 48,765$, Newey-West Standard Errors in Paranthesis

*** denotes significance on 99% level, ** on 95% level, and * on 90% level.

Table 7: Duration — Results of Partially Linear Estimation

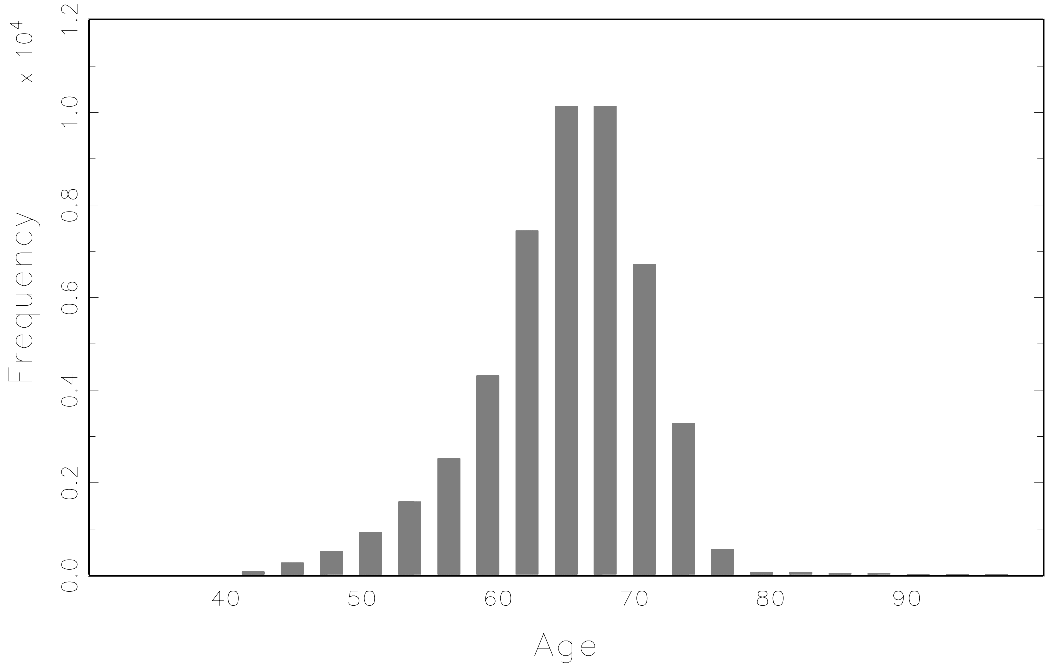


Figure 1: Histogram of Age

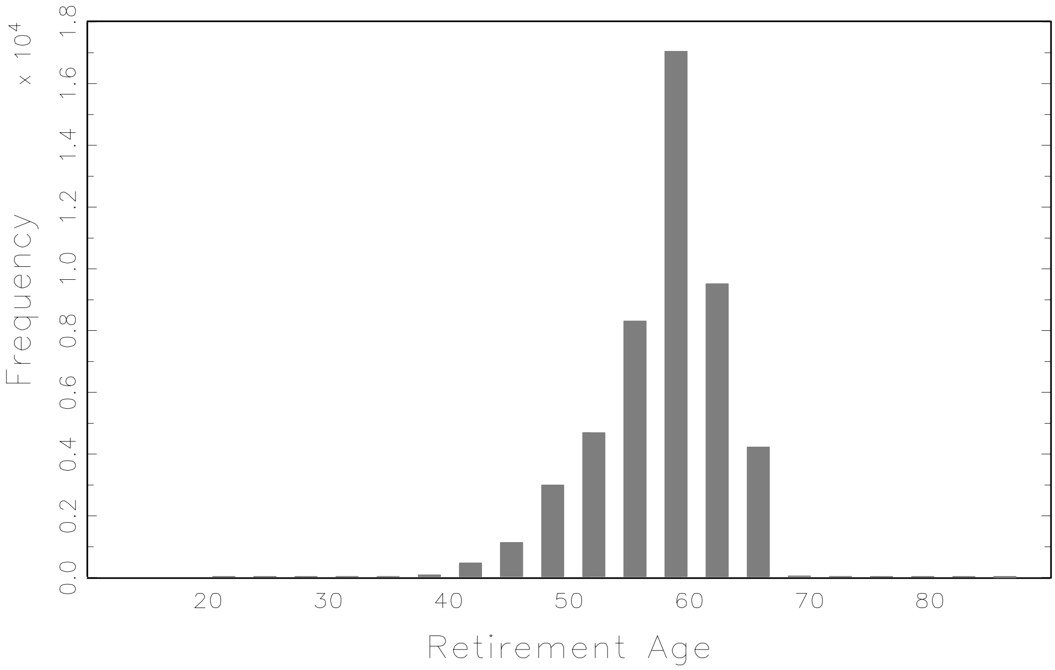


Figure 2: Histogram of Retirement Age

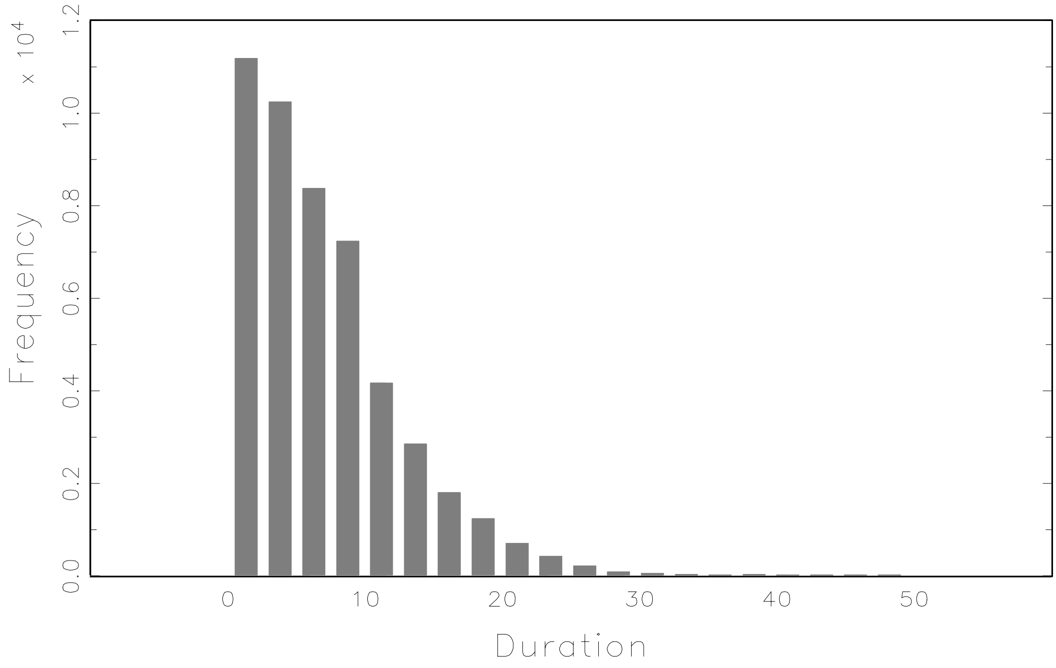


Figure 3: Histogram of Duration

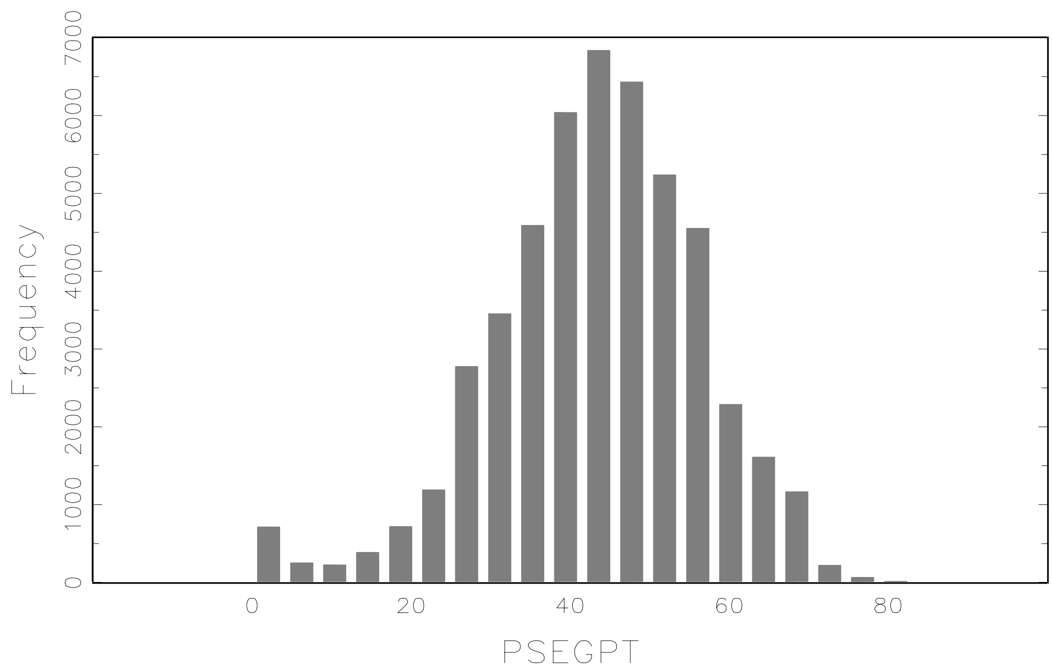


Figure 4: Histogram of PSEGPT

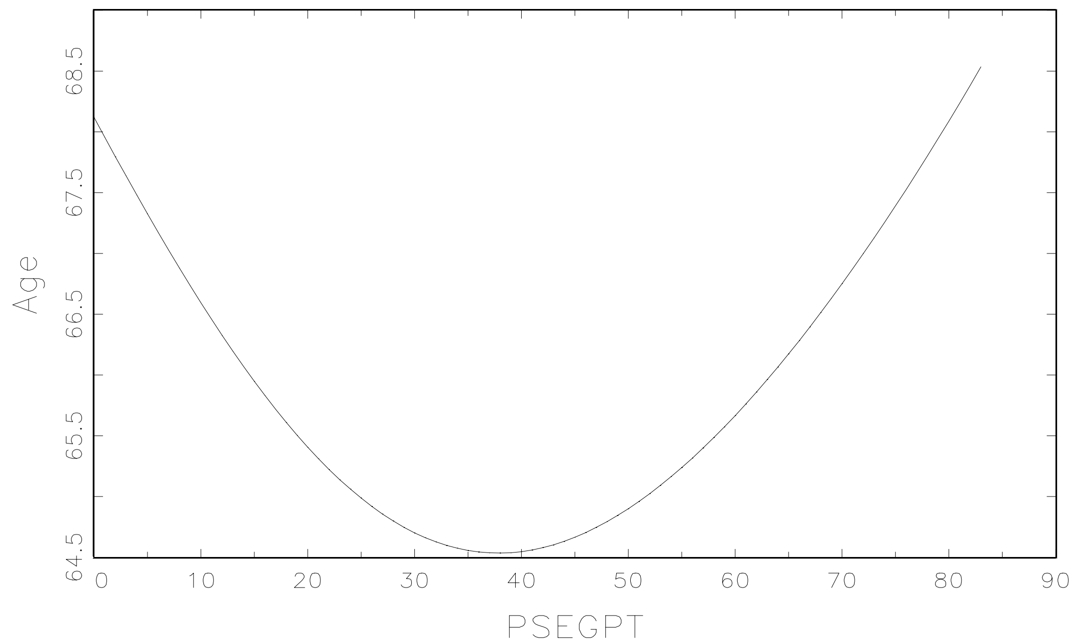


Figure 5: Local Linear Estimate of Age on PSEGPT, $h = 21.265$

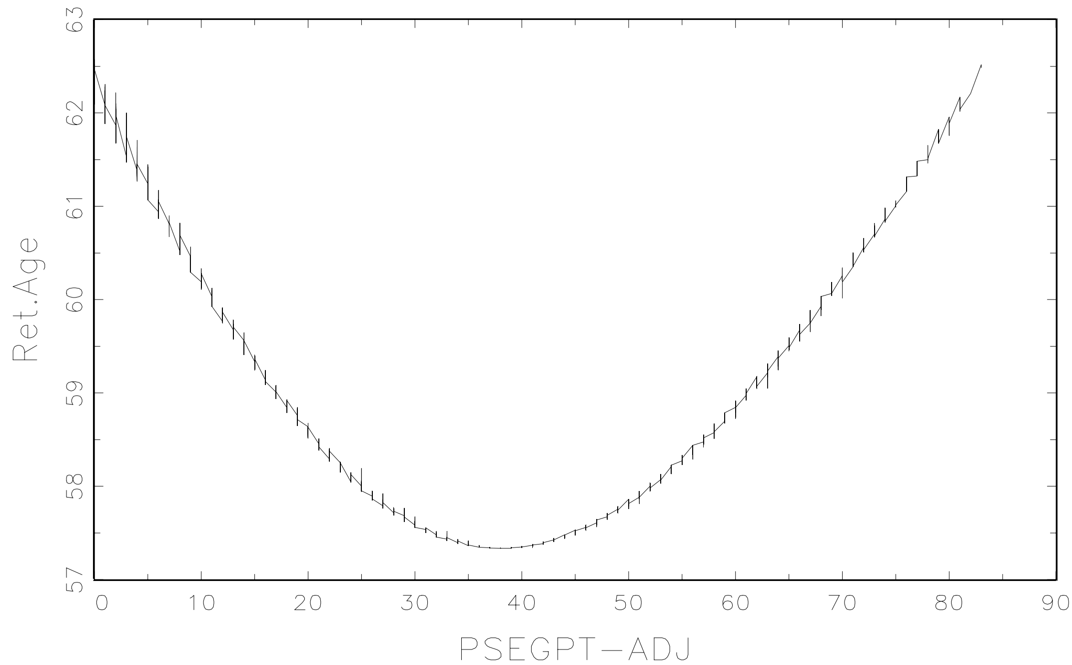


Figure 6: Local Linear Estimate of Retirement Age on PSEGPT-ADJ, $h = 21.192$

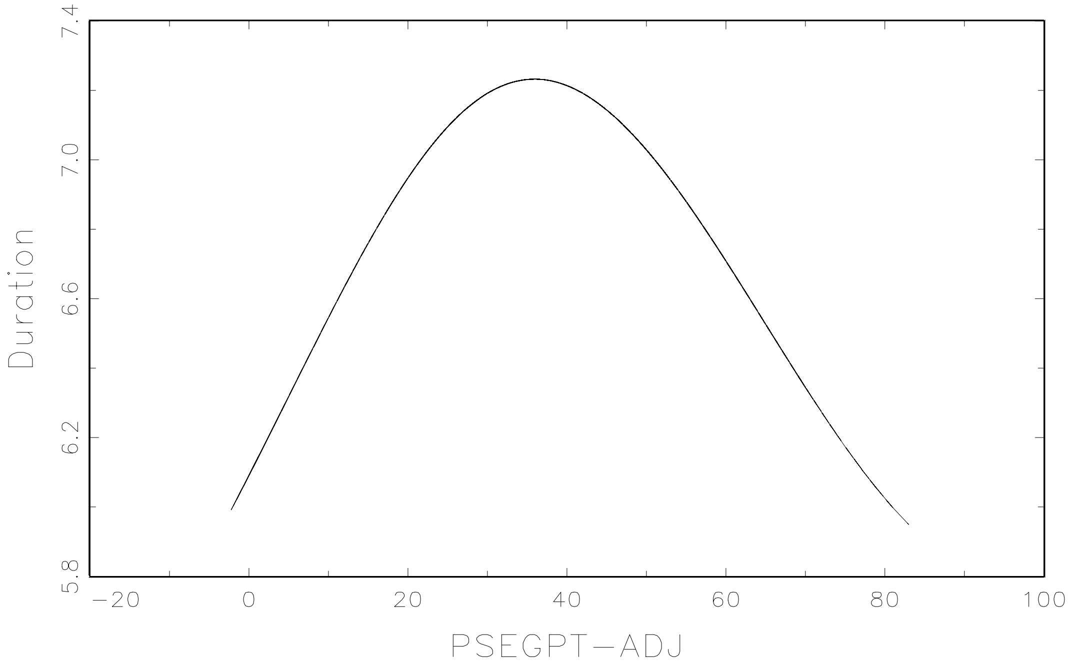
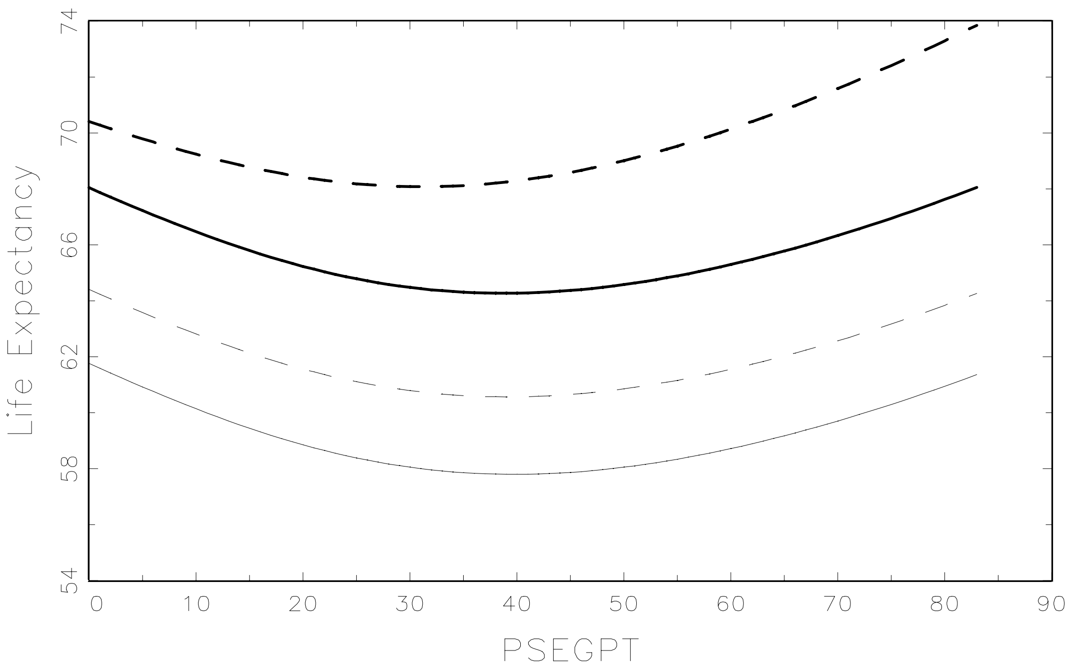
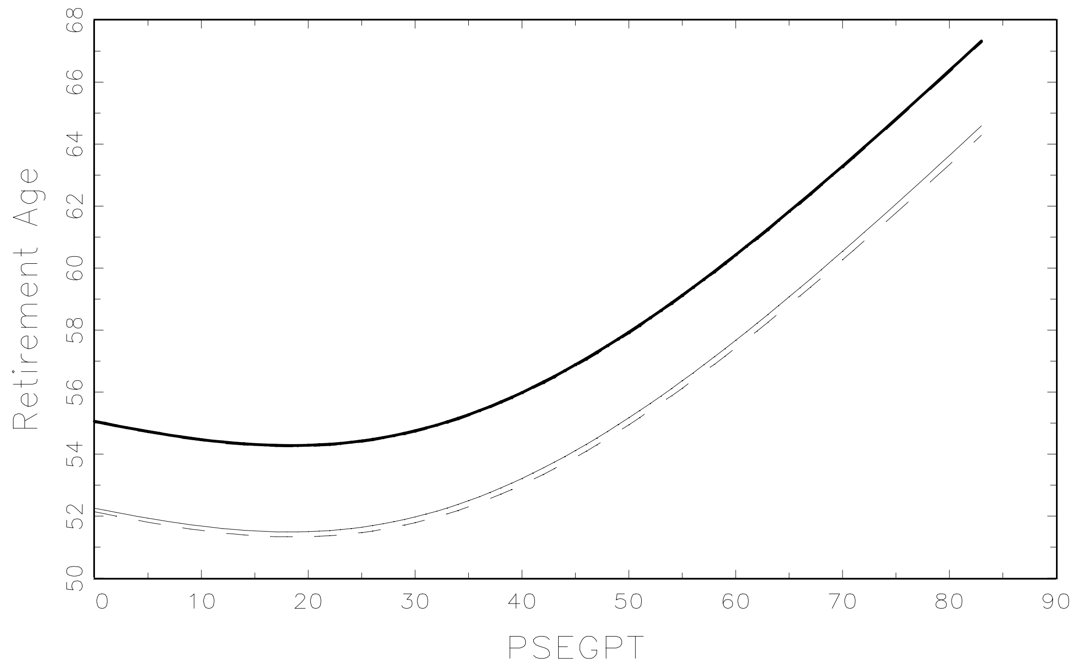


Figure 7: Local Linear Estimate of Duration on PSEGPT-ADJ, $h = 20.776$



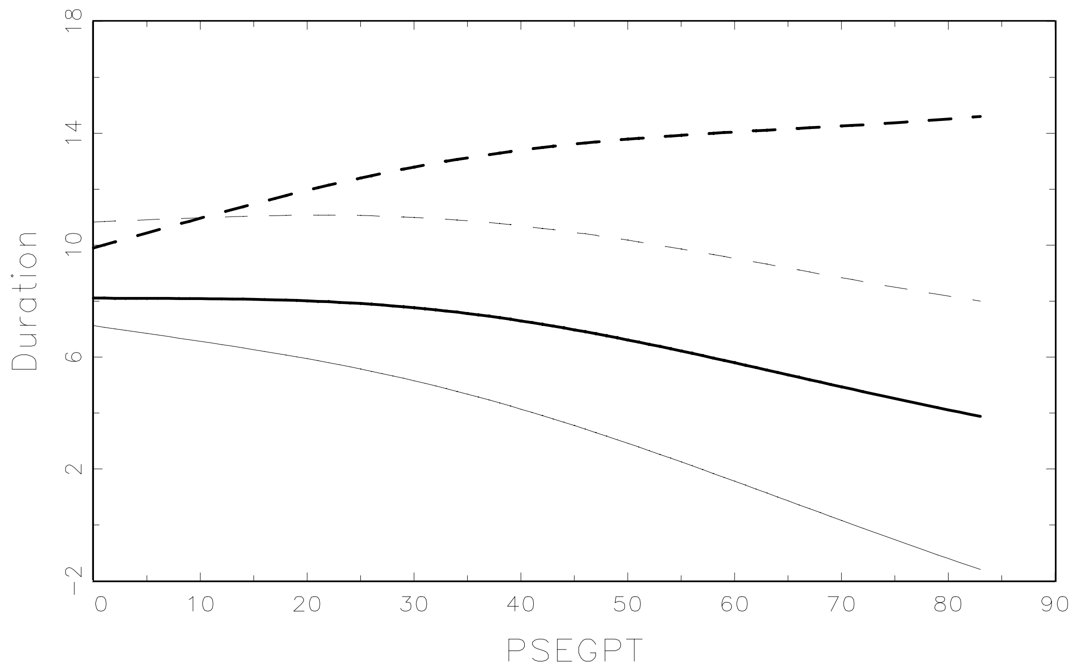
Specification 1: Solid (Bold), Specification 2: Dashed (Bold), Specification 3: Solid (Thin), Specification 4: Dashed (Thin)

Figure 8: Partially Linear Estimates of Age on PSEGPT



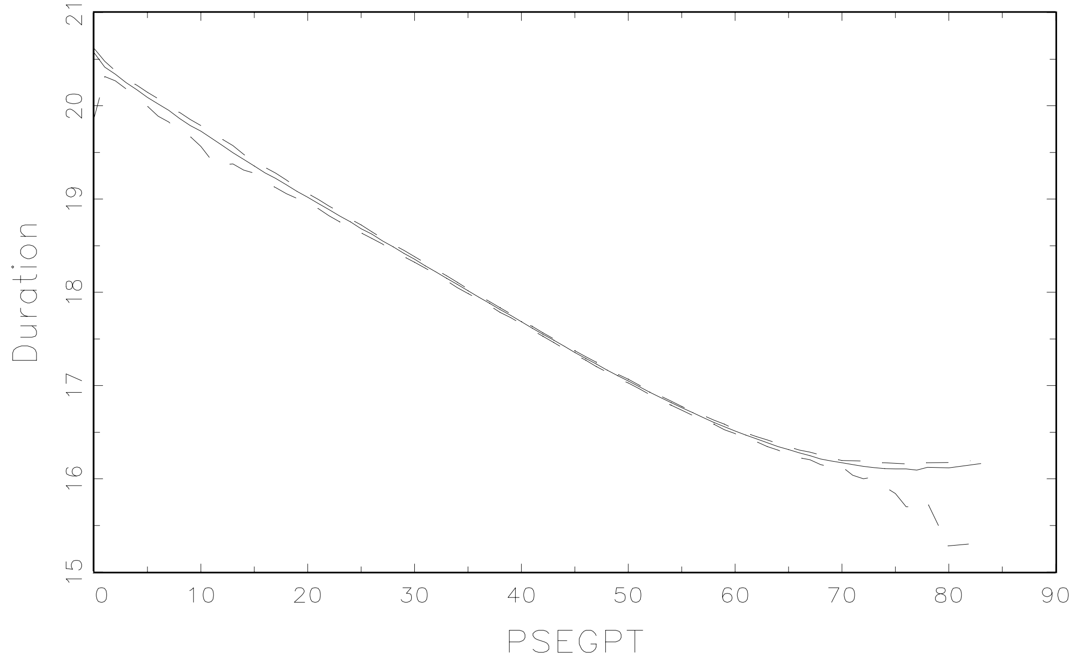
Specification 1: Solid (Bold), Specification 2: Dashed (Bold), Specification 3: Solid (Thin), Specification 4: Dashed (Thin)

Figure 9: Partially Linear Estimates of Retirement on PSEGPT



Specification 1: Solid (Bold), Specification 2: Dashed (Bold), Specification 3: Solid (Thin), Specification 4: Dashed (Thin)

Figure 10: Partially Linear Estimates of Duration on PSEGPT



Based on model specification 4. Due to computation time, the current confidence interval is constructed by using a 10% random sub-sample only (with $n = 4,948$) as 0.025 and 0.975 quantiles around the median of the estimates.

Figure 11: Confidence Interval around Duration

B. Results including Women and Men with at least 5 Years of Contribution

Life Expectancy

Partially Linear Estimates

	Specification			
	1	**2**	**3**	**4**
Health Insurance	0.252***	-1.762***	-0.556***	0.021
	(0.057)	*(0.076)*	*(0.050)*	*(0.048)*
Actuarial Fairness	—	-0.345***	—	-3.380***
(1992)		*(0.016)*		*(0.059)*
Cohort 1	—	—	5.890***	5.705***
(1910-1919)			*(0.183)*	*(0.159)*
Cohort 2	—	—	10.447***	9.441***
(1920-1929)			*(0.099)*	*(0.100)*
Cohort 3	—	—	5.610***	5.405***
(1930-1939)			*(0.096)*	*(0.088)*
Cohort 4	—	—	0.002***	0.002***
(1940-1949)			*(0.000)*	*(0.000)*
Cohort 5	—	—	-16.859***	-16.411***
($>$ 1950)			*(0.141)*	*(0.128)*
Ill-Health	0.097***	0.083***	0.023***	-0.013***
	(0.004)	*(0.004)*	*(0.002)*	*(0.002)*
Unemployment	-0.035***	-0.036***	-0.005***	-0.006***
	(0.001)	*(0.001)*	*(0.001)*	*(0.001)*
Sex	0.023***	-0.166***	2.633***	2.487***
	(0.005)	*(0.007)*	*(0.076)*	*(0.066)*
R^2	0.011	0.019	0.609	0.638

Including Woman and at least 5 Years of Contribution. $n = 115,114$, Newey-West Standard Errors in Paranthesis

*** denotes significance on 99% level, ** on 95% level, and * on 90% level.

Table 8: Life Expectancy — Results of Partially Linear Estimation

Retirement Age

Partially Linear Estimates

	Specification			
	1	**2**	**3**	**4**
IV-Residual	20.418***	20.526***	8.197***	17.061***
	(0.074)	*(0.077)*	*(0.063)*	*(0.060)*
Health Insurance	0.198***	-0.352***	-0.208***	-0.329***
	(0.030)	*(0.030)*	*(0.022)*	*(0.022)*
Actuarial Fairness	—	-0.094***	—	1.394***
(1992)		*(0.006)*		*(0.030)*
Cohort 1	—	—	0.664***	0.908***
(1910-1919)			*(0.054)*	*(0.043)*
Cohort 2	—	—	-0.486***	0.293***
(1920-1929)			*(0.036)*	*(0.035)*
Cohort 3	—	—	0.234***	0.562***
(1930-1939)			*(0.031)*	*(0.027)*
Cohort 4	—	—	0.000?	0.000?
(1940-1949)			*(0.000)*	*(0.000)*
Cohort 5	—	—	-8.671***	-9.185***
(> 1950)			*(0.087)*	*(0.086)*
Ill-Health	0.026***	0.023***	-0.005***	0.002
	(0.002)	*(0.002)*	*(0.002)*	*(0.002)*
Unemployment	0.012***	0.012***	0.010***	0.010***
	(0.000)	*(0.000)*	*(0.000)*	*(0.000)*
Sex	0.018***	-0.033***	-0.595***	-0.421***
	(0.003)	*(0.003)*	*(0.027)*	*(0.023)*
R^2	0.771	0.772	0.847	0.852

Including Woman and at least 5 Years of Contribution. $n = 115,114$, Newey-West Standard Errors in Paranthesis

*** denotes significance on 99% level, ** on 95% level, and * on 90% level.

Table 9: Retirement Age — Results of Partially Linear Estimation

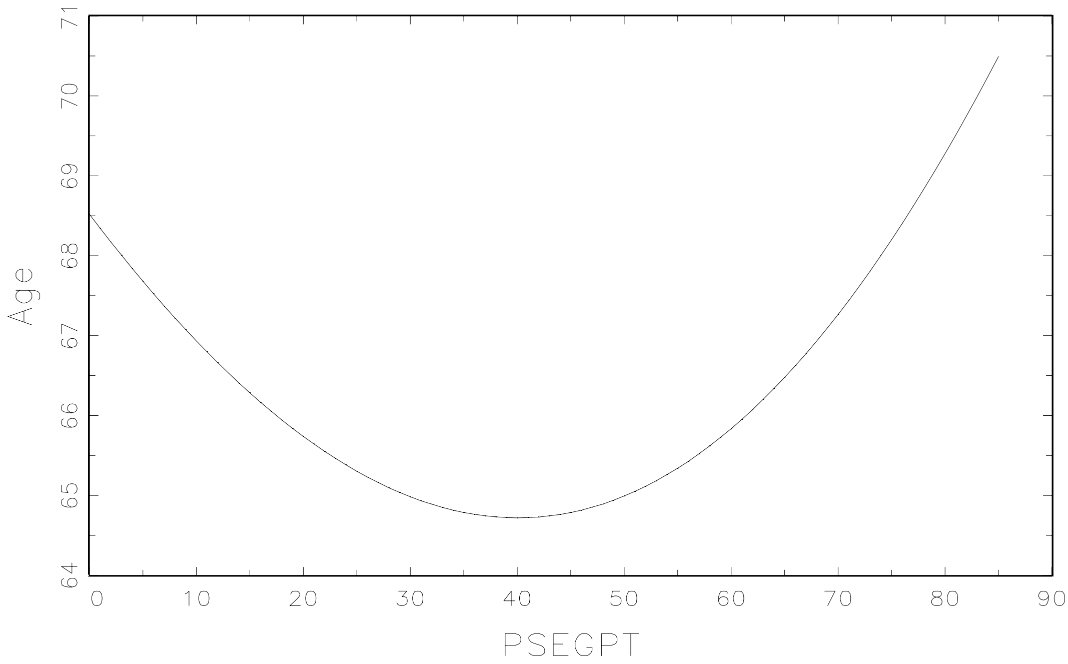
Duration
Partially Linear Estimates

	Specification			
	1	**2**	**3**	**4**
IV-Residual	-4.996***	-4.272***	-10.946***	-3.445***
	(0.096)	*(0.095)*	*(0.140)*	*(0.101)*
Health Insurance	-0.282***	-3.997***	-1.201***	-0.403***
	(0.034)	*(0.054)*	*(0.045)*	*(0.030)*
Actuarial Fairness	—	-0.633***	—	-9.202***
(1992)		*(0.012)*		*(0.063)*
Cohort 1	—	—	3.998***	2.387***
(1910-1919)			*(0.169)*	*(0.078)*
Cohort 2	—	—	8.431***	3.290***
(1920-1929)			*(0.124)*	*(0.087)*
Cohort 3	—	—	3.771***	1.603***
(1930-1939)			*(0.110)*	*(0.063)*
Cohort 4	—	—	0.001***	0.002***
(1940-1949)			*(0.000)*	*(0.000)*
Cohort 5	—	—	-5.949***	-2.557***
($>$ 1950)			*(0.094)*	*(0.048)*
Ill-Health	0.161***	0.139***	0.083***	0.037***
	(0.004)	*(0.003)*	*(0.003)*	*(0.002)*
Unemployment	-0.022***	-0.023***	-0.009***	-0.007***
	(0.001)	*(0.001)*	*(0.001)*	*(0.001)*
Sex	-0.026***	-0.374***	2.407***	1.263***
	(0.096)	*(0.005)*	*(0.071)*	*(0.035)*
R^2	0.090	0.134	0.347	0.597

Including Woman and at least 5 Years of Contribution. $n = 115,114$, Newey-West Standard Errors in Paranthesis

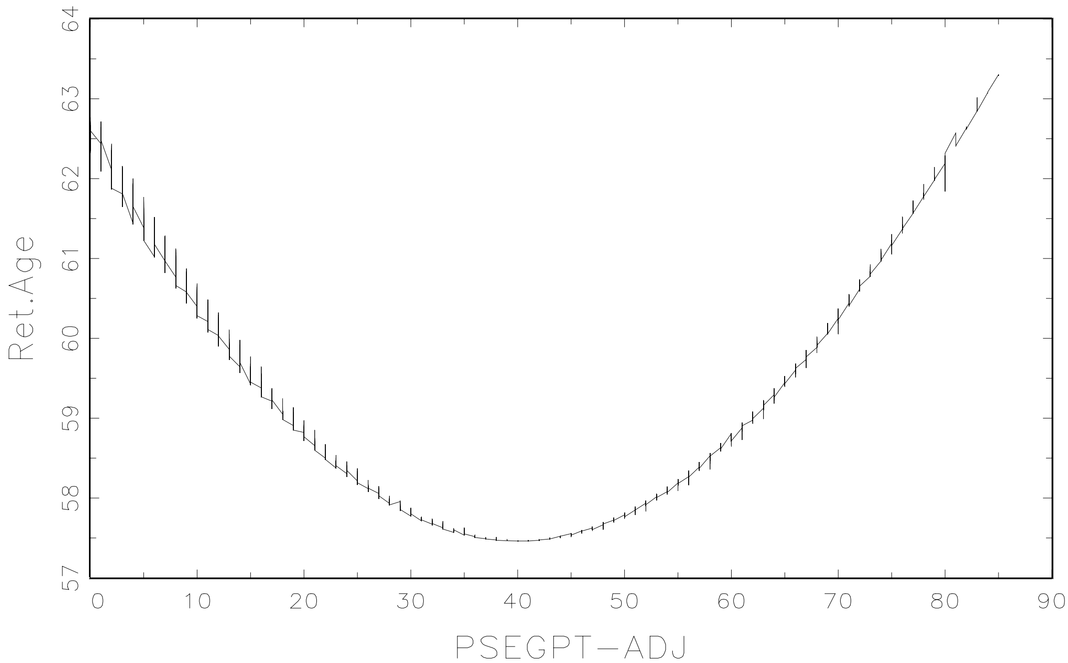
*** denotes significance on 99% level, ** on 95% level, and * on 90% level.

Table 10: Duration — Results of Partially Linear Estimation



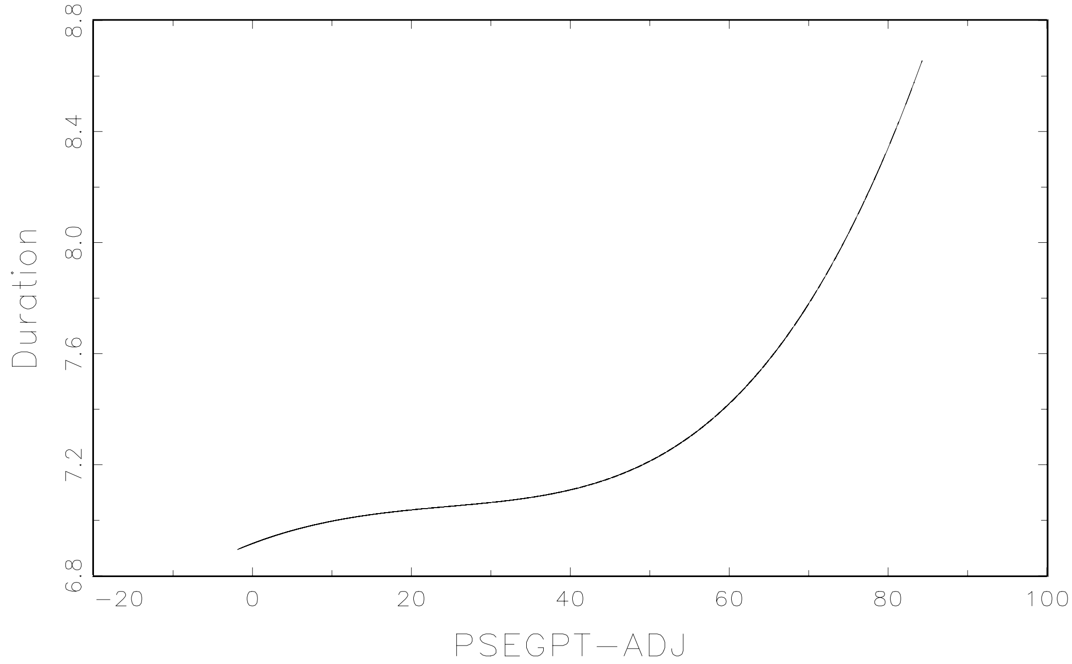
Including Women and at least 5 Years of Contribution.

Figure 12: Local Linear Estimate of Age on PSEGPT, $h = 32.768$



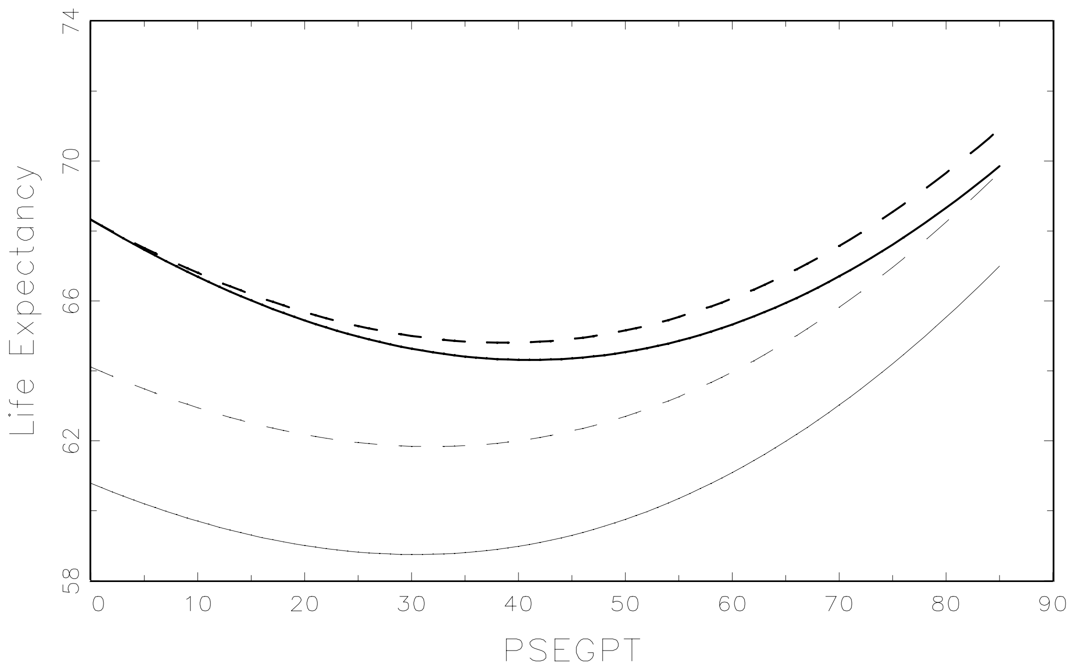
Including Women and at least 5 Years of Contribution.

Figure 13: Local Linear Estimate of Retirement Age on PSEGPT-ADJ, $h = 27.387$



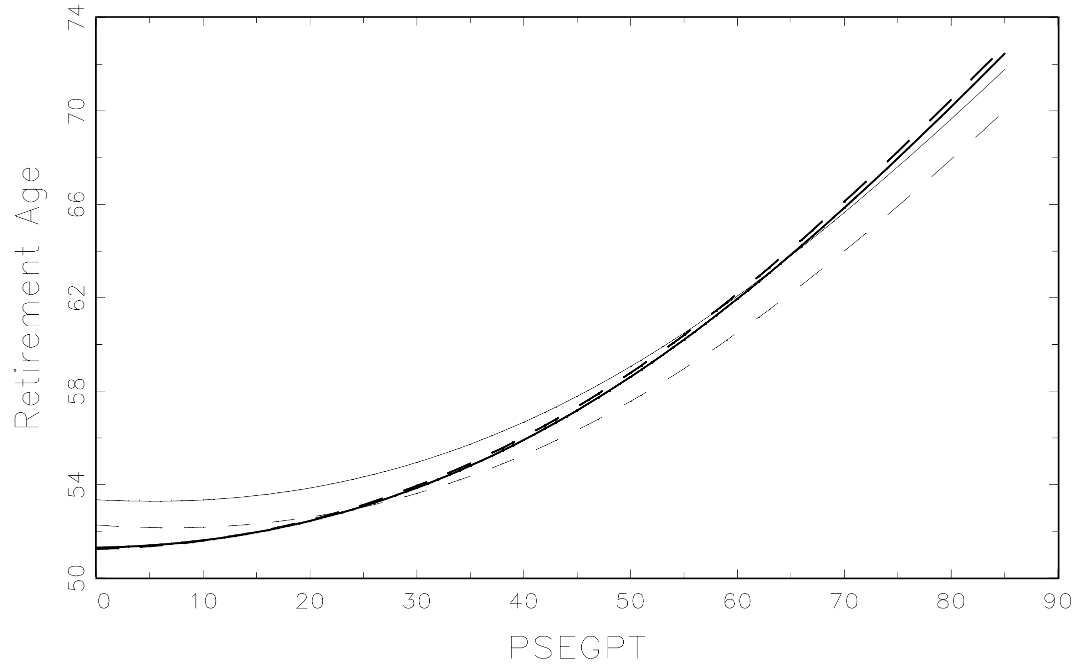
Including Women and at least 5 Years of Contribution.

Figure 14: Local Linear Estimate of Duration on PSEGPT-ADJ, $h = 30.827$



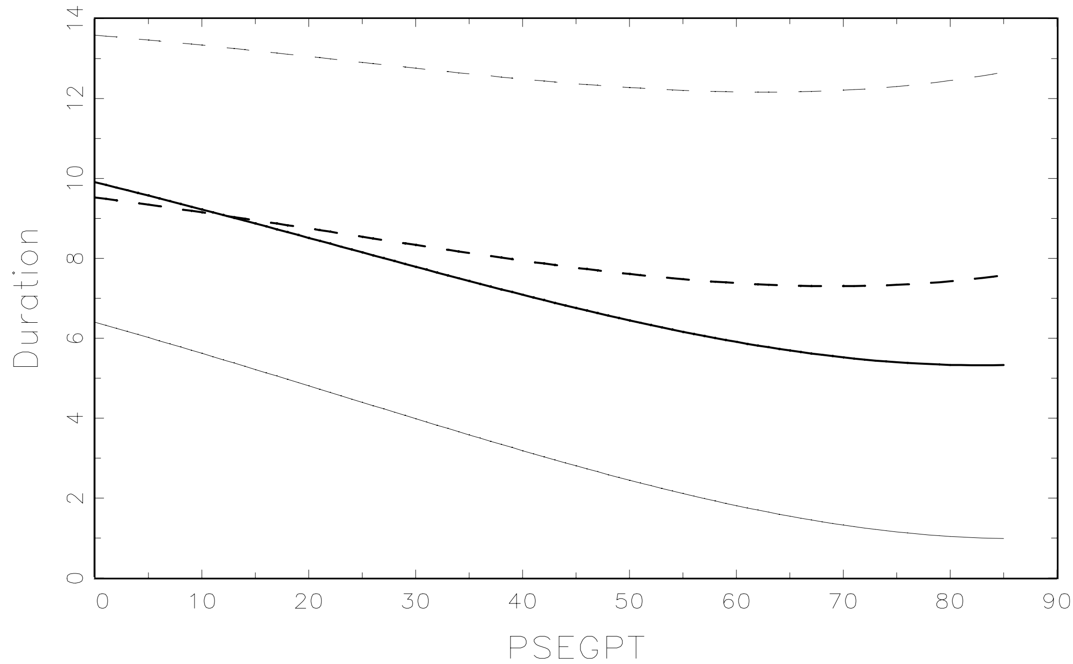
Including Women and at least 5 Years of Contribution. Specification 1: Solid (Bold), Specification 2: Dashed (Bold), Specification 3: Solid (Thin), Specification 4: Dashed (Thin)

Figure 15: Partially Linear Estimates of Age on PSEGPT



Including Women and at least 5 Years of Contribution. Specification 1: Solid (Bold), Specification 2: Dashed (Bold), Specification 3: Solid (Thin), Specification 4: Dashed (Thin)

Figure 16: Partially Linear Estimates of Retirement on PSEGPT



Including Women and at least 5 Years of Contribution. Specification 1: Solid (Bold), Specification 2: Dashed (Bold), Specification 3: Solid (Thin), Specification 4: Dashed (Thin)

Figure 17: Partially Linear Estimates of Duration on PSEGPT

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