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# The Social Costs of Health-related Early Retirement in Germany: Evidence from the German Socio-economic Panel

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Abstract:
This study investigates the role of stratification of health and income in the social cost of health-related early retirement, as evidenced in the German Socio-economic Panel (GSOEP). We interpret early retirement as a mechanism to limit work-related declines in health that allows poorer and less healthy workers to maximize the total discounted value of annuities received from Germany's pay-as-you-go pension system. Investments in new medical technology and better access to existing health services may help to curb the need for early retirement and thus improve efficiency, especially amid population ageing. To value the potential gains, we calibrate an intertemporal model based on ex post predictions from stratified duration regressions for individual retirement timing. We conclude that eliminating the correlation between income and health decline would delay the average age of retirement by approximately half a year, while keeping all workers in the highest of five categories of self assessed health would yield a further delay of up to three years. Had this scenario been realized during our 1992–2005 sample period, we estimate the social costs of early retirement would have been more than 20 percent lower, even without counting the direct social benefits from better health.

Keywords: Retirement timing; Health inequality; Social costs; Medical technology; Calibration

JEL classification: H55, I12, O15

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# 1 Introduction

As one of the first countries in Europe, Germany is making a serious effort to contain the costs of early retirement. It aims at raising the retirement age by making early retirement financially less attractive and by increasing the statutory retirement age from 65 to 67 years over the next two decades. Today, Germans retire much earlier: in 2006 the average retirement age was 60.9 years. Only about 33 % of workers retire at age 65 or after. This article focuses on the empirical magnitude of the social costs of *health-related* early retirement. An empirical assessment of these costs can help identify the social value of investments in new medical technology and more effective health systems that enable workers to maintain their health for longer parts of their lives. Financial incentives can only delay retirement if workers are healthy enough to work longer. According to Siddiqui (1997), the bad health of many workers reduces by approximately one fourth the effectiveness of adjustments in pension payments that are actuarially fair for the average worker in raising the average retirement age. The lower life expectancy associated with bad health reduces the net present value of a given pension annuity so that the incentive to continue working for a larger pension annuity beginning at a later age is also lower.

Previous studies have focused almost exclusively on the expenditure side of the pension system and the implied losses in terms of GDP. In this vein, Herbertsson and Orszag (2003) predict a rapid increase in the social costs of early retirement amid population aging – with output lost to early retirement reaching almost 13 % of GDP in 2010 if current participation rates of elderly workers are maintained constant in Germany. However, previous studies have tended to neglect the work-related health hazards that early retirement removes, creating a welfare benefit that lowers the net social costs. An exception is Vermeulen and Kalwij (2006) who estimate that labour force participation rates of elderly men would increase by about 12 % if everybody was in perfect health, a scenario which may help to establish an upper limit for the social costs of health-related early retirement.

The main purpose of our article is to highlight the social dimension of health-related retirement, which we think is far greater than the mere aggregation of individual costs would suggest. To this end, we investigate the timing and social cost of early retirement in the context of the health gradient, the positive correlation between income and health, as evidenced in the German Socio-economic Panel (GSOEP). The theory of the health gradient implies that at each point in time, workers' propensity to retire early on health grounds is negatively correlated with their wage income. Recent empirical studies, such as Deaton and Paxson (1998) for the US and Hostenkamp and Stolpe (2006) for Germany, show that the health gradient tends to be steeper in older age groups until about age 60, so that the opportunity costs of early retirement decline with age for those with poor health and poor earnings prospects in the labour market. Workers have different individual positions relative to the agegroup-specific health gradient that tend to persist as cohorts age because declines in health are largely related to the type of work and stratified by income class. Our measure of workers' position relative to the health gradient is obtained by the orthogonal projection of each worker's individual location in the agegroup-specific health-income diagram onto the log income axis; we then subtract the agegroup-specific median value to eliminate general aging-related changes in health or income. To construct the health gradient, we use the empirical normal transformation that Jürgens (2005; 18-19) pioneered to obtain a cardinalized health measure with the value 1.577 for *very good*, 0.431 for *good*, -0.540 for *fair*, -1.373 for *bad* and -2.196 for *very bad* self-rated health. Those at the bottom end of the health gradient are the first to face the risk of an early death if they continued to work.

From a social point of view, the opportunity to retire early may limit workers' private incentives to invest in the development and maintenance of their personal human capital, so that society fails to fully exploit the complementarity between health and education. Investments in new medical technology that addresses the work-related deterioration of health may prevent the health gradient from becoming steeper as workers age and turn into candidates for early retirement. To help quantify the potential welfare gains from such investments, we propose to compare the social cost of health-related early retirement in the presence of the health gradient with the hypothetical social cost in its absence. How much is at stake depends on the question whether workers' individual positions on the health gradient are very persistent and our empirical analysis pays particular attention to this issue. With persistence, workers tend to stay in their relative position as they age, whereas mobility means workers often change their position relative to the agegroup-specific health gradient. While persistence adds to the social costs, high mobility may instead call for a change in pension rules to increase flexibility and improve the private incentives for early retirees to return to work when their health recovers. Germany’s disability pension reform of 2001 can actually be interpreted as a significant step in this direction: The first disability pension may now only be granted for a limited time and is subject to review after a pre-specified number of years.

The remainder of this paper is organized as follows. In section 2, we start with a definition of the social cost of early retirement and describe alternative specifications of the counterfactual absence of the health gradient. Section 3 describes our data and presents duration regressions for the timing of early retirement by individual workers in Germany. Section 4 explains how we use these regression results to make counterfactual predictions of retirement timing, taking workers’ persistence on the health gradient into account. Section 5 finally quantifies the social costs of health-related early retirement in monetary terms. Section 6 concludes.

## 2 The Problem of Social Costs in Health-related Early Retirement

The social costs generally exceed the private costs of early retirement. While the social costs primarily include the retiring worker’s full output, the private costs are limited to lost wage income minus any potential benefits, such as pension payments, leisure time and gains in health and longevity. We ignore the possibility that pension payments add to the social costs because they may not only redistribute income but also create an excess burden as active workers’ labour supply may decline in response to rising pension system contributions. Similarly, we ignore leisure time as a direct social benefit because with impaired health it is inherently difficult to value, but we include an estimate of workers’ willingness-to-pay for the mortality-related health benefits observed after early retirement, which presumably captures much of the benefit of enjoying leisure instead of work. The health benefits do not only translate into financial gains via lower mortality and greater longevity ex post. They also represent the ex ante insurance value of the early retirement option against further declines in health that continued work would imply. This component of the social costs can only be evaluated in a dynamic context where workers' health declines in a stratified pattern with strata identified by persistent positions of workers relative to the health gradient.

In essence, we argue that systematic differences in individual retirement timing across income classes indicate social costs beyond the mere aggregation of individual costs, because the health gradient expresses a social constraint on individual human capital investments. We infer this from the human capital theory of health demand (Grossmann, 1972; Muurinen and Le Grand, 1985; Case and Deaton, 2005) in which the matching of young workers entering the labour market to available jobs puts them in different strata of health development and in effect denies many workers their full potential over the lifecycle – especially workers "selling their bodies" (Case and Deaton, 2005) in hard manual jobs. Although the opportunity for health-related early retirement provides some insurance against part of the individual costs associated with persistent health trajectories, we do not think this to be efficient. Early retirement merely *socializes* part of these costs, but neither mobilizes nor rewards the necessary investments in health and education to overcome the problem of stratification in income and health.

To assess the empirical order of magnitude of the social costs, we compare the observed loss of output and health benefits from early retirement with a counterfactual scenario in which the health gradient is absent. In fact, we specify two variants of the counterfactual scenario. One is based on the assumption that all workers enjoy the same evolution of the health status as is observed for the average worker in the highest income quintile. The other is based on the assumption of perfect health for all.

With few exceptions, the literature has neglected the possibility that early retirement may slow down or even halt work-related declines of health and that workers value these benefits. Börsch-Supan and Jürges (2006) are among the first trying to measure the intangible benefits from early retirement in terms of wellbeing, but do not assign monetary values to these gains. However, several studies (Jürges, 2005; Frijters et al., 2005; Rehfeld, 2006) show self rated health to be a good predictor for subsequent mortality, and we can use these estimates to quantify the impact of health improvements on mortality risk. In addition, there is a vast literature on the empirical willingness-to-pay (WTP) for mortality risk reductions, known as the statistical value of life, so that we can indirectly infer the value of health improvements after early retirement.

In this vein, we calculate the social costs of early retirement as the present discounted value of labour income foregone minus the gains from mortality risk reductions. Aggregating across $N$ workers, indexed by $i$, who retire at some individual age $A_i$, the social costs $C$ can be written as

$$C = \sum_i^N \left\{ \sum_{t=A_i}^{65} \frac{Y_i}{(1+\rho)^t} \times \Pr(s_{it} \mid A_i, G, R, H) - \sum_{t=A_i}^{100} \frac{WTP}{(1+\rho)^t} \times \Delta \Pr(s_{it} \mid A_i, G, R; H) \right\} \qquad (1)$$

where $Y_i$ is the annual gross labour income[1] at age $t$, assumed constant for all future years; $\rho$ is a time preference rate equal to 3 %; $\Pr(s_{it} | A_i, G, R, H)$ is the conditional probability to survive at least $t$ more years, given survival until age $A$, gender $G$ and region of residence $R$.[2] In addition, the survivor function depends on a worker's health status $H$: Better than average age-specific health is assumed to shift the survivor function upward, as if the person were younger, while worse than average health is assumed to shift the survivor function downward, as if the person were older. Throughout, the survivor function is independent of the survival of spouses. The second sum in the curly brackets is the discounted willingness-to-pay value for mortality risk reductions, where WTP is obtained from the literature on the value of a statistical life in Germany; $\Delta \Pr(s_{it} | A_i, G, R, H)$ is the change in survival probability that can be attributed to health-related early retirement; and the reduction of mortality risk will be effective beyond the normal retirement age until a hypothetical maximum lifetime of 100 years – the last year for which age-specific mortality rates are available.

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[1] We do not include the full value of workers' output when they do not retire as a cost of early retirement. Total output will generally exceed the marginal product of labour because most workers are inframarginal. Assuming gross wages equal labour's marginal product, the output value of healthy inframarginal workers will obviously exceed their gross wage income. As noted, we also do not include the value of gains in leisure time, which have the opposite sign. The net mistake of these two omissions may be relatively small.

[2] The conditional survival probabilities are computed from the standard life tables of the German Bureau of the Census which provides this information separately for both genders, and for regions East and West. See Statistisches Bundesamt (2006).

*Constructing counterfactual scenarios*. Defining the most appropriate counterfactual scenario to determine when workers would retire in the absence of the health gradient is far from trivial. Bommier and Stecklov (2002) suggest: If equity in health is defined according to a social justice approach, “the health distribution in an ideal equitable society is one where access to health has not been determined by income or socioeconomic status.” Thus, knowing one’s health status does not help in predicting a person’s income, or vice versa. Several scenarios would be consistent with this condition. *First,* there would be no health gradient if there was no income inequality. In this vein, Deaton (2001) suggests an argument for income redistribution based on the observation that a concave relationship between individual health and income will cause greater income inequality to be associated with lower average population health. Since health is not directly transferable across individuals, inequality aversion will call for mean-preserving income transfers.

A *second* approach to define the absence of the health gradient is a distribution of health states across individuals that is independent of income. Kakwani et al. (1997), concerned with aggregate measures of income-related health inequality, have defined “avoidable” inequality as the difference between an observed health concentration index and a standardized concentration index, in which the distribution of health is age and gender standardized, to account for these “unavoidable” demographic factors. However, it is not entirely clear why the age and gender specific concentration of ill health should be unavoidable (Chotikapanich et al. 2003). Van Doorslaer and Koolman (2004) instead use the 1996 cross section of the European Community Household Panel to define “excess” inequality by relating the health concentration index in one country to the value of concentration in the country with the lowest inequality. They argue that health inequalities should primarily be addressed by health policy, not income redistribution, as the countries with the lowest income inequality are not always the countries with the lowest income-related health inequality.

A *third* approach to define the absence of the health gradient is to assume everybody had the same health status, irrespective of whether that status was excellent, fair or bad. Vermeulen and Kalwij (2006), for example, assess the importance of being healthy for individual retirement behaviour by quantifying a counterfactual scenario where everybody is in perfect health and labour market participation rates increase by more than 12 %age points for German men, and about 8 %age points for women. We think equal health for all is the least ambiguous scenario to quantify. By contrast, constructing counterfactual scenarios on the basis of no income inequality or on the basis of statistical independence of health and income is more difficult as these assumptions are specific only at the population level, but do not specify how counterfactual health states should be assigned to individual workers.

# 3 Estimating Individual Retirement Behaviour in Germany

*Construction of the health gradient*. The present study is based on an *unbalanced* panel drawn from 13 consecutive annual waves of the GSOEP starting in 1992 when our preferred health measure was first collected. We use the variable *self rated health status*, a subjective measure of health based on answers to the WHO recommended question “How would you describe your current health?” on a five point ordinal rating scale from *very good* coded as 1 to *very bad* coded as 5. As respondents’ self assessment of overall health is likely to include physical as well as mental aspects, we prefer this measure over more objective, yet often incomplete measures of health. The latter are inherently difficult to assess comprehensively, but can be useful for specific conditions. For example, Larsen and Gupta (2004) show that depression is an important factor in the timing of retirement, especially among women.

To identify the exact beginning of workers’ retirement, we choose the month when a person first declares him- or herself retired, as this is the most exact variable available in the GSOEP. This variable, moreover, may be less affected by the justification bias that might arise if respondents understated their health in order to justify early retirement in a socially acceptable way. In the GSOEP, respondents are not asked to give a reason for retirement and health and retirement variables are not recorded in the same questionnaire, so that conscious misreporting seems unlikely.

The second variable needed to construct the health gradient is income. The GSOEP provides continuous measures, recorded in Euros, of nominal individual labour income and of total household income[3]. We transform these variables using the OECD consumer price deflator so that real income is expressed in prices of 2001. For regression analyses, we use the natural logarithm of the variable *post-government* income, which is defined as the combined total household income of all household members after taxes and transfers. In addition, we run regressions in which household income is replaced by equivalent income, i.e. adjusted for equivalent household size by calculating the ratio of household income to the modified OECD scale, the sum of 0.5, the number of children times 0.3 and the number of other household members times 0.5.

Our regressions also include variables that were significant explanatories of retirement timing in previous studies: the natural logarithm of imputed rent for owner occupied housing, to control for wealth effects; the number of years in education, to measure a person's attainment beyond the legally required minimum; a dummy variable indicating whether a person is a public sector employee, to account for differences in the pension systems for private and public employees; and dummies for gender, region of residence, marital status and unemployment.

As the decision to retire is essentially a dynamic inter-temporal decision problem, the natural empirical approach is duration analysis. However, duration analysis cannot account for longitudinal weighting, using person-specific time-varying weighting factors. To account at least for differences in initial sampling probabilities across subgroups, we use the cross sectional weights of the first year that an individual is observed to weight the subsequent observations.[4] This might lead to a slight over-weighting of individuals that participated from the start in 1992, relative to individuals of subsamples included later, since we cannot examine differences in panel attrition rates across subgroups. Selection bias, due to the lower attrition of workers in good health, may render the estimated hazard to retire on health grounds too small.

Another problem often found in duration analysis is unobserved heterogeneity. To account for groupwise unobserved heterogeneity, we estimate stratified duration models that allow the baseline hazard to differ between strata, defined on the basis of worker's quintile position relative to the agegroup-specific health gradient. In the context of

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[3] In the so-called CNEF equivalent file.
[4] This procedure was recommended by Dr. Martin Kroh of the DIW, Berlin in personal communication.

parametric duration analysis, this means that the shape parameter $p$ is allowed to differ between strata but the distribution is confined to be of the Weibull family. This allows for time-varying heterogeneity due to unobserved persistent health shocks between, but not within strata, so that we can explore some basic implications of the human capital model of health demand, identify differences in retirement timing between workers in different strata and make out-of-sample predictions of the mean time to retirement.

*Results of duration regressions*. Table 1 reports our estimates of the Weibull model, stratified on workers' quintile position relative to the agegroup-specific health gradient, using the lowest quintile as the base category. Positive coefficients indicate that higher covariate values are associated with a greater hazard to retire, and vice versa. A prior comparison of the log likelihoods from stratified and non-stratified regressions, available from the authors on request, showed that stratification significantly improves the explanatory power of the model. All dummy variables for quintile position turned out highly significant with the expected signs. The scale and the shape of the hazard function are significantly different from the base category in each of the other strata. The regressions also show that most of the variation can be explained by workers' quintile position, and that variation within strata is satisfactorily accounted for by the other significant covariates, as our continuous measure of workers' position relative to the health gradient becomes insignificant once quintile position is controlled for.

Our stratified Weibull regressions confirm that the hazard of early retirement is highest in the lowest quintile position relative to the health gradient. The estimated effect of moving from the lowest to the second lowest quintile is particularly large. In models using unadjusted household income, the hazard to retire is smallest in the highest quintile, whereas in models using income per equivalent household member, the hazard is smallest in the second highest quintile. Based on the Akaike information criterion, models using income per equivalent household member appear to perform best.

The estimated coefficients for the other covariates are also robust across different model specifications and have the expected signs. More specifically, we find that public-sector employment, unemployment for males and eastern residence for females significantly increase the hazard to retire, consistent with the view that financial variables are important determinants of retirement behaviour at the individual level. However, continuous measures of household income are always insignificant in our stratified regressions, so that their impact seems to be fully accounted for by the strata identifying variable. Summing up, we find that health, education and financial variables have a significant effect on the retirement behaviour of men and women.

# 4 Predicting Individual Retirement Behaviour

Based on ex post predictions of hazard rates and mean time to retirement, this section examines the persistence of workers' positions relative to the agegroup-specific health gradient, using Markov chains, and quantifies counterfactual scenarios in which the health gradient is absent. Based on our stratified regression estimates with household income and health as separate covariates, the hazard function is

$$\hat{h}\left(t_j \middle| x_j\right) = \exp(\beta_0)\exp(\ln(p)) \cdot t_j^{\exp(\ln(p))-1} \exp\begin{pmatrix} +\beta_1 srhs + \beta_2 public + \beta_3 unempl + \beta_4 edu \\ +\beta_5 sex + \beta_6 married + \beta_7 east + \beta_8 birthyear \end{pmatrix} \quad (2)$$

where $\beta_o$ and ln*(p)* vary by quintile position. Household income enters our analysis via the definition of quintile positions relative to the agegroup-specific health gradient. Assuming constant covariate values, the mean time to retirement is the expected value of survival time, or the first moment of the distribution of survival time, conditional on covariate values $x_j$: $\mu_{Tj} = E\left(T_j \middle| x_j\right) = \int_0^\infty f\left(t \middle| x_j\right)dt = \int_0^\infty S\left(t \middle| x_j\right)dt$. We use these predictions to calibrate welfare effects of health-related early retirement, without attempting to account for time-varying covariates at the individual level.[5] The implied prediction error will be greater, the more the main covariates of interest in our analysis, workers' health and their individual position on the health gradient, change in reality.

Assuming a health status or quintile position that is *worse* than the true value would result in an under-prediction of the mean time to retirement, because better health as well as a better individual position on the health gradient both lower the predicted hazard to retire in each point in time, thus increasing predicted time at work. By

[5] According to personal communications with Dr. Freitag-Wolf at the Institute of Medical Informatics and Statistics of Kiel's University Hospital, sophisticated programming could be employed to estimate how covariate values change with time at the population level and in principle these changes could be incorporated in our predictions; yet this would be beyond the scope of our present study, and variations in the rate of covariate changes at the individual level could still not be accounted for.

contrast, assuming a health status or quintile position that is *better* than the true value results in an over-prediction of the mean time to retirement. To obtain a realistic corridor for the estimated time to retirement of each individual, we perform two sets of ex-post predictions. In the first, we set all covariates constant at the values observed at age 40, when workers' time at risk to retire is set to begin. In the second, we set the covariate values constant as of the moment of retirement; in virtually all cases, workers' health is better at age 40 than at the moment they retire. Table 3 shows the mean time to retirement across all observations for these two scenarios. The difference is about 155 days. Expected retirement age is 59.02 years when covariates are set constant at age 40. It is 58.60 when covariates are assumed constant as of the moment of retirement for each worker. Our ex post predictions are thus very close to the actual average retirement age for the entire sample, which is 58.9 years.[6]

As workers' quintile position is defined relative to agegroup-specific health gradients, whose slope increases with age, there is bound to be persistence over time. And the more persistent workers' positions, the more reliable can be our long-term predictions of retirement timing. Put differently: Since workers' quintile position is estimated to have the biggest single effect on the retirement hazard, as shown in the bottom panel of Table 1, our prediction error will increase with workers' propensity to move from one quintile to another. To obtain an empirical measure of persistence, we estimate a discrete-time Markov chain and compare the size of entries on the main diagonal with off-diagonal entries. The states of the Markov chain are defined by quintile positions so that our estimates give information on workers' intra-distributional mobility between quintiles on the health gradient. We use fractile Markov chains where every discrete state contains the same number of individuals of the population. However, the sample of our distribution of workers' across quintiles is not exactly equal because the different sampling probabilities of each individual have been taken into account using the person expansion factors provided by the GSOEP.

*Results of Markov chain estimation*. Table 2 presents estimates of first-order time-stationary transition probabilities over periods of one and five years for the population

[6] Note that the Deutsche Rentenversicherung (2007) reports an official average retirement age in Germany of 60.9 years in 2006, reflecting the fact that the average retirement age increased slightly during the observation period.

40 plus.[7] The first panel gives the one-step annual transition matrix, whose $(i, j)$ entry is the conditional probability that an individual has transited from quintile $i$ to quintile $j$ after one year. Persistence is greater in quintile one and five, exceeding 68 % and 76 % respectively. Yet considerable intra-distributional mobility that is evident in the three middle quintiles, where the entries on the main diagonal fall below 50 %, leads to a long run stationary distribution, reported in the first panel along with the sample distribution, that does not show any concentration of probability mass at the extremes.

The second panel gives estimates of first-order five-year transition matrices, where the entries on the main diagonal are lower, indicating much lower persistence over longer periods of time. Nonetheless, the one-year transition estimates are qualitatively confirmed. To illustrate consistency of our short- and long-run estimates, we iterate the one-year transition matrix five times and report this in the third panel. The entries on the main diagonal are between 8 and 25 percentage points smaller than the corresponding direct estimates of the five-year transition matrix. This comparison suggests that the true persistence may be higher than estimated by first-order Markov chains.

The stationary distribution was observed after 25 iterations, with further changes in transition probabilities below 0.002%. However, we can only claim stationarity relative to the underlying true state space if suitable measures of the inter-quintile range do not change over time. Since we apply fractile Markov chains based on quintile positions, the boundaries between states are not fixed over time, but are determined endogenously for each agegroup. If the agegroup-specific distribution becomes wider in the course of workers' life, this alone would reduce the number of observed transitions between health states, so that we may underestimate the true mobility unless we take changes in the entire distribution into account. To this end, the bottom section of Table 2 shows the development of the boundaries between the first and the second quintile as well as the range of quintiles two, three and four. All increase until the five-year agegroup centred

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[7] Ignoring any information about transition probabilities which may be contained in the initial probability distribution and assuming the transition probabilities to be invariant with respect to time as well as gender, we use the Maximum Likelihood estimator $p_{ij} = h_{ij} / h_i$, where $h_{ij}$ denote the observed frequency of transitions from state $i$ to state $j$ and $h_i = \sum_j h_{ij}$, to determine the transition probabilities. The transition matrix can then be used to calculate a long run stationary distribution using the Chapman-Kolmogorov equation as described in Osaki (1992).

in 58 and decline thereafter, suggesting the variance of the distribution increases as workers age until they reach the typical age of retirement. Regressing the inter-quintile ranges on age using OLS for agegroups 28 to 58 revealed average effects of moving to a higher agegroup of 2.3 to 3.8 %.[8]

*Quantifying counterfactual scenarios*. Figure 2 illustrates the background to our counterfactuals by showing the mean health status of early retirees and of the highest income quintile by agegroup. Younger early retirees tend to have particularly bad health. By contrast, the average worker in the highest quintile on the income distribution can stays above the mean of the overall health distribution (at 0 on the horizontal axis) until about age 50.

Based on our corridor of ex post predictions, Table 3 summarizes the two alternative counterfactual scenarios consistent with the absence of the health gradient, defined as no correlation between health and income, that we use to assess the social costs of health-related early retirement. When – in our first counterfactual scenario – we assume equal access to existing health care and medical technology and all workers are able to maintain their health as individuals in the highest quintile on the income distribution, each early retiree is assigned the age-specific mean health of the highest income quintile. In this situation, retiring workers' health still declines with age, as can be seen by the development of the dark bars in Figure 2. But investment in better medical technology may lower the rate of health decline below the rate currently observed in the highest income quintile. In our second counterfactual, we assume – as the first-best case – health does not decline at all with age so that all workers are able to maintain the state of health classified as very good on the ordinary scale in the GSOEP.

In Table 3, the first panel compares our ex post predictions with the two counterfactual predictions for 40-year old workers, the second panel reports the potential gain from eliminating the health gradient and the third panel reports the potential gain from eliminating health decline. It appears that most working years are lost because workers'

8 In addition to our earlier finding that the cross sectional measure of the health gradient becomes steeper over the course of people's lives, we find that persistence in the individual position on the health gradient increases with age as well, when comparing estimates for the entire adult population with those adults above 40 years of age.

health deteriorates with age. Only between 0.32 and 0.63 working years would be saved per early retiree in our first, but between 2.63 and 3.04 years would be saved per early retiree in our second counterfactual.

## 5 Quantifying the Social Costs

To complete our calibration of the social costs of health-related early retirement, we now quantify the health improvements that can be attributed to early retirement, relate these health improvements to subsequent mortality risk reductions and changes in survival probabilities and then briefly describe how we assign monetary valuations to the gains from early retirement. To identify the relevant health improvements empirically, we would ideally like to compare each worker's health after retirement with a situation where that worker had not retired early. Aggregated across all early retirees, we would then have an estimate of the improvements in population health from early retirement. However, we do not observe the counterfactual and thus base our estimate on the nearest comparable scenario in the dataset, simplifying a methodology developed by Börsch-Supan and Jürges (2006). More specifically, we ask to what extent is the effect on health more favourable for those taking early retirement compared to those retiring at the normal retirement age. Consistent with non-health-related retirement pathways that become available after age 60, we divide the observations into two groups: all retirees who are younger than 60, but at least 40 years old at the moment of retirement entry and the "normal" retirees defined as women above 60 years and men above 61, but in any case below 70 years of age.

Figure 3 shows the development of mean health of normal and early retirees after retirement. We can see that early retires are on average a lot unhealthier than normal retirees, and controlling for age would reveal even larger differences. More importantly, the graph indicates that retirement has a positive effect on health only for early retirees. Average health of normal retirees continues to decline, whereas the average early retiree enjoys health improvements over the course of two years after retirement. Only in the third period after retirement does the health of early retirees start to decline again, but still at a lower rate than that observed for normal retirees.

We use this comparison of normal and early retirees to obtain an estimate of the effect of early retirement on health. To be conservative, we model the health improvement as a level effect, namely as the difference between the average health changes of normal and early retirees after two years. The effect is slightly larger for women than for men; women who retire early can improve their health measured on the cardinal health scale by 0.193 points on average, while men can improve their health by 0.181. We next assume the size of the effect to be linearly decreasing with age and obtain the slope of the linear relationship by comparing the average health improvements at the mean age of early retirees with retirees at age 60 (the omitted age category) where assume health changes due to retirement are zero. Since male early retirees are on average about a year older than female early retirees, the relationship between age and health improvements from early retirement is steeper for men.

*Relating health improvements to mortality risk reductions*. Bad health lowers the probability to survive at every age so that people who suffer from bad health die younger (Frijters et al., 2005 and Rehfeld, 2006). Using a Cox proportional hazard model, Jürges (2005) finds that the hazard to die is about 48 % higher for men in bad health compared to men in fair health, controlling for other relevant covariates recorded in the GSOEP.

The estimated mortality hazard relative to those of men and women in fair health are 0.503* for men and 0.759 for women in very good health, 0.726** for men and 0.812 for women in good health, 1.481*** for men and 1.412*** for women in bad health, and 2.518*** for men and 2.479 *** for women in very bad health.[9] This suggests, for example, that a health improvement of 0.834 in our continuous measure reduces male workers' relative mortality hazard from bad health to the level of a worker in fair health. Assuming a piecewise linear relationship between health and mortality, a health improvement of 0.181 would reduce the hazard rate to die by 0.087. The magnitude of these effects is very similar in the case of women. To use this information in the calculation of social costs, we have to multiply the relative mortality hazards with a baseline hazard. Since Jürges' (2005) semi-parametric Cox model does not specify a baseline hazard function, we use standard life tables from the German federal statistics

[9] The number of stars indicates the level of significance, 1, 5 or 10 %.

office to construct the baseline hazard to die. In doing so, we make three technical assumptions: First, the hazard ratios are applicable to the entire population although Jürges' (2005) analysis used only the West German sub-sample. Second, for women in good and very good health the estimates of the proportional hazard to die are economically relevant although they are not significantly different from the baseline category in a statistical sense.[10] Third, we multiply the proportional hazard rates by a constant factor to make the average health status the baseline category as Jürges (2005) defined the middle health category as his reference category. Figure 4 shows that this results in a parallel upward shift of the proportional hazards dependent on health and that relative changes in health states are unaffected by this adjustment.

To obtain health-specific probabilities to survive, we multiply the proportional hazard to die conditional on health with the probability to die at a certain age and calculate the complement: $1-\{\ \Pr(die|A,G,R)\times Hazard(die|H)\ \}=\Pr(s|A,G,R,H)$. The probabilities to survive conditional on age, gender and region are taken from the standard life tables of the years 2003/05 (Statistisches Bundesamt, 2006), which may lead to a slight underestimation of age-specific mortality rates because life expectancy increased during the observed period. However, the difference between years 1992 and 2002 is significant only for males in eastern Germany (Eisenmenger, 2005). To account for health-specific mortality after retirement, we make the additional assumption that mortality always increases with age. A worker who is in a bad health status at retirement is assumed to *age* like an average person facing the same (low) survival probability at a much higher age.

*Willingness-to-pay for mortality risk reductions*. If retirement leads to a mortality reduction for early retirees, what should this be worth to society in monetary terms? Estimates of the value of a statistical life (VSL) can be derived from tradeoffs between money and fatality risks and range between 0.5 and 16 million US dollar for a statistical life in the US – with a median of 5 million and an income elasticity is around 0.5 to 0.6 (Viscusi, 1993). A reasonable value for Germany will be lower than most US estimates. However, the literature for Germany is surprisingly scarce. Miller (2000) estimates a

[10] Including only the effect for women in bad health would lead to a downward bias of the health state specific probability to survive.

value of 3.2 million Euros for a statistical life in Germany. The European Commission (2000) suggests an upper limit of 2.5 million Euros for a VSL in environmental cost benefit analysis. Krewitt and Friedrich (2000) notes that estimates for the VSL tend to be higher in medical than in environmental settings and suggests 3 million Euros as a reasonable value for the EU 15 countries. We apply this value and ignore the possibility that the value of a statistical life may vary with age.[11]

*Social cost estimates*. All costs are counted in the year that the individual enters retirement. We first calculate the present value of the future costs and benefits of early retirement for each individual for all years following retirement. The first additive term represents the present value of the lost labour income from the entry into early retirement until the pensioner reaches the statutory retirement age.[12] This involves double discounting – reflecting workers' time preference and the conditional probability to survive, which is below 100 %. The second term includes the discounted gains from future mortality risk reductions from the entry into early retirement until age 100. Since we define early retirees as between 40 and 60 years of age, mortality risk reductions through early retirement can only be assigned to this group of retirees. In the last step, the social costs accrued for an individual are multiplied with the individual-specific person expansion factor and aggregated across all individuals to obtain the social costs at the population level. Here the actual expansion factors, as observed at the moment of retirement (not as in the first year of observation), are used so that the results can be interpreted for population totals.

The three panels in Table 4 report the results for the aggregated social costs and its two main components for the entire 13 year period in Euro prices of the year 2001 – first on the basis of ex post predictions, second on the basis of counterfactual health states, and third on the basis of perfect health for all (first-best counterfactual). While the first column reports the values for the entire sample population, the five subsequent columns

[11] The direction and size of the relationship between age and VSL is still debated. Jones-Lee (1989) finds an inverted U shaped relationship between VSL and age. Johannesson and Johansson (1997) find that the value of a statistical life increases with age whereas several more recent studies, such as Krupnick et al. (2002), find that the VSL falls with age especially after the age of 70.

[12] Some retirees in our dataset, such as home makers or unemployed workers, have no labour income in the year prior to retirement. Since their income is assumed constant for the remainder of the potential working life, we implicitly assume that these retirees do not incur any social costs.

report those for workers in different quintile positions relative to the health gradient. In each panel, the entries in the first row reports discounted labour income losses, the second row the gains from mortality reductions and the third row the social costs, calculated as the difference between the first two. Our ex post predictions suggest a total of about 1040 billion Euros of discounted labour income losses, gains of about 271 billion Euros from reduced mortality risks social costs amounting to about 769 billion Euros, which is on average 50 to 60 billion Euros per year. The social costs attributable to the lower quintiles (4 and 5) appear to be much larger than those in the higher quintiles (1 and 2); but these differences shrink in the two counterfactual scenarios.

The value of labour income losses in the counterfactual scenarios is influenced by two opposite effects: On the one hand, workers in better health tend to retire later and labour income losses will be lower. On the other hand, healthier workers also live longer which raises the probability that labour income is lost in the later years. In both counterfactuals, the first effect seems to dominate. By contrast, the value of health improvements from early retirement is lower the better a worker’s health state. Our counterfactual assumption that all workers enjoyed the mean health of the highest income quintile is responsible for the lower aggregate health benefits reported in the middle row of the second panel of Table 4. In the first-best counterfactual, the aggregate health benefits are zero because all workers are assumed to have the highest health state.

Of course, workers directly value the higher levels of health that they enjoy in the counterfactual scenarios, and the final row of the second and third panels in Table 4 reports the estimated aggregate willingness-to-pay for the implied mortality reductions. These estimates suggest that the welfare gains that may be within reach from better medical technology and more equitable and effective health care are likely to be several orders of magnitude greater than the estimated health benefits from early retirement in the current set-up.

Figure 5 shows the development of the social costs and its two main components over time. Substantial fluctuations are evident and the first trough coincides with the recession year of 1993, but after that they do not seem to be strongly correlated with the business cycle, suggesting that labour demand factors do not influence our results much. The peak of labour income losses lies in the year 2001, at about 100 billion Euros almost twice the level observed in 2004, when labour income losses were at their lowest. Over the whole 13 year period, the social costs of health-related early retirement seem to be in decline, and the relatively low costs observed since the pension law reforms in 2001 are mainly responsible for this trend.

## 6 Conclusion

A number of recent empirical studies have addressed the problem of the social costs of early retirement in Germany, yet identifying the part that is health-related has remained an unresolved task. It requires first to identify the workers who retire on health grounds, even as eligibility criteria and the relative attractiveness of different pathways to retirement, described in Viebrok (2003), change over time, and second to value the full costs and benefits in monetary terms, including costs and benefits of changes in health. In this paper, we have attempted to fill this gap and estimate social costs that amount to almost 60 billion Euros in every year of the sample period, 1992 to 2005. Assuming constant productivity, discounted aggregate losses of labour income streams amount to about 80 billion per year. The social costs are lower than the labour income losses because early retirement can slow down and partly reverse the work-related decline in health. We observe health improvements within two to three years after early retirement that cannot be observed for workers retiring at the "normal" age. These health benefits and the implied mortality risk reduction seem to be especially important for workers in the lowest quintile relative to their agegroup-specific health gradient, often workers in manual jobs that wear down their health more rapidly.

To quantify the potential welfare gains from targeted investments in health and medical technology, we calibrate two counterfactual scenarios in which the health gradient is absent – first by assigning the mean health observed in the highest quintile and second by assigning the highest level of self assessed health to all candidates for early retirement. Our estimate of the social costs in 2003 is about twice as high as the liabilities in Germany's social budget that Rehfeld (2006) attributes to health-related retirement in 2003.

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# Appendix

The German Socio-economic Panel (GSOEP) is a longitudinal microeconomic dataset on a wide variety of topics that is designed to be representative for the non-institutionalized resident population in Germany. The GSOEP dataset was provided by the Deutsches Institut für Wirtschaftsforschung (DIW). It started in 1984 and has been repeated annually since then. In 1990 it was expanded to include the East German population. In order to ensure a large sample size, additional samples were drawn over the course of time to include immigrants and to offset ongoing panel attrition. In addition the GSOEP has been adjusted and modified in a number of ways to ensure international comparability, for example through the rephrasing of questions and the introduction of new topics (DIW, 2007). Foreign residents, as well as East German residents and high income households are over-sampled in the GSOEP, while elderly people above 70 are under- represented. To adjust for disproportionate sampling of subgroups, non- response and attrition in the course of time the GSOEP provides pre-calculated cross sectional and longitudinal weighting factors. In each cross section the weighting factors are normalized according to population marginals in the micro census and add up to the non-institutionalized resident population in Germany. Longitudinal weighting factors are provided in the form of staying probabilities to account for panel attrition (Haisken-DeNew and Frick, 2005). The average life tables for the years 2003/05 were provided by the Statistisches Bundesamt. All regression analyses were performed using the statistical software package Stata 8.0. Some of the preparative calculations for the social costs formula were done in Excel version 2003.

### Variable Description

**Health:** *Self-rated health status* is measured on an ordinal 5 point scale [1 = very good; 5= very bad]. *Self-rated health* is measured on a continuous scale using the empirical normal transformation ranging from -2.196 very bad to 1.577 very good. **Education:** *No of years of formal education*, starting with 7 years = 0; more = 1; 2; 3… **Economic status:** *Employment status* is a dummy with working = 0; no work = 1. *Public service employee* is a dummy with no = 0; yes = 1. **Income:** *Post-government household income* is HH labour income + HH asset income + Private transfers + Public transfers + Social Security pensions + HH Private retirement income – Total HH Taxes. *Real household income* is post-government household income indexed for inflation in Euro prices of 1995. *Log household income* is the Log of Real household income. *Log equivalent household income* is the Log of (Real household income / No of equivalent persons in household). *Real individual labour income* is individual labour income indexed for inflation in Euro prices of 2000. **Wealth:** Wealth is the Log of (Imputed owner - occupied house rental value in Euros / Consumer price*100). **Health gradient:** *Relative position on the age group-specific health gradient* is a continuous measure. *Quintile position* is a discrete variable from 1 best to 5 worst. **Consumer Price Index** is from the OECD: 1992 = 83.7; 1993 = 87; 1994 = 90.1; 1995 =92.6; 1996 = 94.1; 1997 = 95.3; 1998 = 97.1; 1999 = 97.9; 2000 = 98.6; 2001 = 100; 2002 = 102 ; 2003 = 103.4; 2004 = 104.5; 2005 = 106.2.

# Tables

## Table 1: Results of Stratified Parametric Duration Models for Early Retirement

| Coefficients | | 17<br>Weibull | 17a<br>Cox | 19<br>Weibull | 19b<br>Cox |
|---|---|---|---|---|---|
| | | Household income | | Per Equivalent HH member | |
| **Self-rated health** | | **- 0.24** | **-0.27** | **- 0.23** | **-0.27** |
| | P>\|z\| | 0.000 | 0.000 | 0.000 | 0.000 |
| **Household income** | | **- 0.04** | **0.02** | **-0.01** | **0.05** |
| | P>\|z\| | 0.542 | 0.795 | 0.938 | 0.578 |
| **2nd quintile** | | **- 7.44** | | **- 10.09** | |
| | P>\|z\| | 0.004 | | 0.000 | |
| **3rd quintile** | | **- 15.39** | | **- 15.90** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| **4th quintile** | | **- 12.83** | | **-18.09** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| **5th quintile** | | **- 17.05** | | **- 17.18** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| **Cons** | | **81.81** | | **79.12** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| | **ln_p** | | | | |
| **2nd quintile** | | **0.31** | | **0.41** | |
| | P>\|z\| | 0.003 | | 0.000 | |
| **3rd quintile** | | **0.56** | | **0.58** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| **4th quintile** | | **0.48** | | **0.64** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| **5th quintile** | | **0.60** | | **0.61** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| **Cons** | | **0.82** | | **0.79** | |
| | P>\|z\| | 0.000 | | 0.000 | |
| No. of subjects | | 14280 | 53328723 | 14280 | 53328723 |
| No. of failures | | 2867 | 12955664 | 2867 | 12955664 |
| Time at risk | | 26122296 | 1.06415e+11 | 26122296 | 1.06415e+11 |
| Log likelihood | | - 2976149.9 | -17680.8 | - 2883200.2 | -17665.3 |
| No. of obs | | 74604 | 74604 | 74604 | 74604 |
| Wald $chi^2$ | | 452.72 | 384.26 | 439.45 | 387.96 |
| Prob > $chi^2$ | | 0.000 | 0.000 | 0.000 | 0.000 |
| AIC | | **5952341.8** | | **5766442.4** | |

*Source: Own Calculations using Stata version 8.0*

*Note*: The coefficients of our additional control variables – dummies for housing wealth male, no employment of males and females, respectively, public sector employment, sex, married females, and Eastern residence as well as the year of birth and years of education, – are available from the authors on request.

**Table 2: Five State Fractile Markov Estimates for Transition Probabilities between Quintiles on the Health Gradient for Population 40+**

| Observations | Transition end state Quintile 1 | Quintile 2 | Quintile 3 | Quintile 4 | Quintile 5 |
|---|---|---|---|---|---|
| | First order, time stationary estimates of the one-year transition probabilities | | | | |
| 16306 | **0.682** | 0.229 | 0.06 | 0.021 | 0.008 |
| 19440 | 0.189 | **0.494** | 0.24 | 0.064 | 0.012 |
| 18906 | 0.052 | 0.234 | **0.439** | 0.232 | 0.043 |
| 18729 | 0.021 | 0.065 | 0.221 | **0.503** | 0.19 |
| 19349 | 0.008 | 0.014 | 0.039 | 0.171 | **0.768** |
| Stationary Distribution | 0.174 | 0.204 | 0.201 | 0.202 | 0.219 |
| Sample Distr. | 0.176 | 0.21 | 0.204 | 0.202 | 0.209 |
| | First order, time stationary estimates of the five-year transition probabilities | | | | |
| 7354 | **0.525** | 0.296 | 0.122 | 0.042 | 0.015 |
| 9285 | 0.204 | **0.371** | 0.277 | 0.122 | 0.027 |
| 8506 | 0.103 | 0.231 | **0.319** | 0.266 | 0.081 |
| 7886 | 0.049 | 0.106 | 0.224 | **0.378** | 0.243 |
| 6393 | 0.024 | 0.037 | 0.077 | 0.216 | **0.645** |
| Stationary Distribution | 0.164 | 0.202 | 0.206 | 0.213 | 0.215 |
| Sample Distr. | 0.187 | 0.236 | 0.216 | 0.2 | 0.162 |
| | One-year transitions iterated 5 times | | | | |
| | **0.303** | 0.273 | 0.201 | 0.139 | 0.084 |
| | 0.229 | **0.252** | 0.219 | 0.177 | 0.123 |
| | 0.17 | 0.219 | **0.22** | 0.208 | 0.182 |
| | 0.12 | 0.175 | 0.204 | **0.233** | 0.268 |
| | 0.072 | 0.117 | 0.163 | 0.243 | **0.405** |

| Agegroup | Boundary 1 | Range 2 | Range 3 | Range 4 |
|---|---|---|---|---|
| 43 | -0.385 | 0.272 | 0.230 | 0.294 |
| 48 | -0.419 | 0.293 | 0.254 | 0.317 |
| 53 | -0.423 | 0.303 | 0.249 | 0.314 |
| 58 | **-0.430** | **0.304** | **0.250** | **0.318** |
| 63 | -0.408 | 0.282 | 0.250 | 0.335 |
| 68 | -0.359 | 0.246 | 0.222 | 0.305 |
| 73 | -0.332 | 0.217 | 0.196 | 0.304 |
| 78 | -0.337 | 0.233 | 0.218 | 0.355 |

*Source: Own Calculations using Stata Version 8.0*

**Table 3: Lost Working Years due to the Health Gradient**

| | Covariate values at age 40 assumed constant thereafter | Covariate values at retirement assumed constant before |
|---|---|---|
| Mean time to retirement in days / Mean age at retirement | | |
| Ex Post Predictions | 6947 / 59.02 | 6791 / 58.59 |
| Counterfactual mean health of upper income quintile | 7070 / 59.35 | 7029 / 59.24 |
| Counterfactual best health | 7953 / 61.79 | 7953 / 61.79 |
| Lost working years due to the health gradient | | |
| total | 4,020,458 | 7,816,393 |
| per retiree | 0.322 | 0.6264 |
| months per retiree | 3.86 | 7.51 |
| Lost working years due to imperfect health | | |
| total | 32,900,000 | 38,000,000 |
| per retiree | 2.634 | 3.043 |
| months per retiree | 31.61 | 36.51 |

**Table 4: The Social Costs of Health-related Early Retirement, in billion Euros**

| | Total | Health Gradient | | | | |
|---|---|---|---|---|---|---|
| | | Quintile 1 | Quintile 2 | Quintile 3 | Quintile 4 | Quintile 5 |
| *Ex post predictions* | | | | | | |
| Aggregated labour income losses | 1040 | 122 | 192 | 194 | 219 | 313 |
| Health benefits from early retirement | 271 | 66.4 | 53 | 39 | 43 | 70 |
| Social costs | 769 | 55.6 | 139 | 155 | 176 | 243 |
| *Counterfactual health* | | | | | | |
| Aggregated labour income losses | 995 | 98.6 | 171 | 184 | 207 | 334 |
| Health benefits from early retirement | 230 | 21 | 25 | 38 | 48 | 98 |
| Social costs | 765 | 77.6 | 146 | 146 | 159 | 236 |
| Value of direct mortality reductions from counterfactual health improvement | 22500 | 8980 | 5970 | 4580 | 2180 | 808 |
| *First-best counterfactual* | | | | | | |
| Aggregated labour income losses | 600 | 74.1 | 111 | 110 | 110 | 195 |
| Health benefits from early retirement | 0 | 0 | 0 | 0 | 0 | 0 |
| Social costs | 600 | 74.1 | 111 | 110 | 110 | 195 |
| Value of direct mortality reductions from counterfactual health improvement | 58300 | 16300 | 13000 | 11800 | 9070 | 8380 |

*Own Calculations using Stata version 8.0*

# Figures

Figure 1: Health Distribution Across Self Assessed Health Categories Pooled for all Waves 1992 to 2005

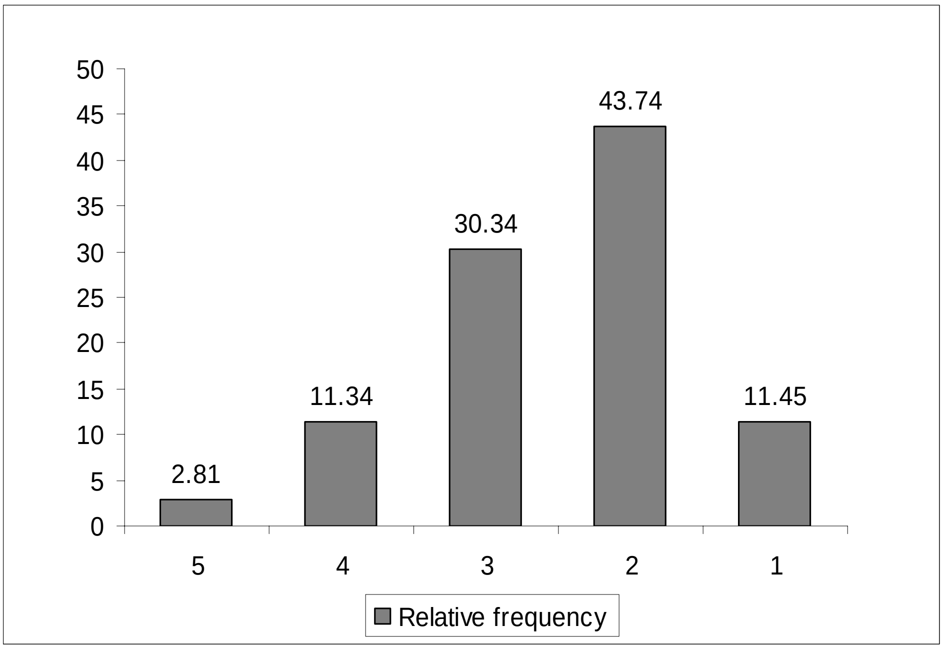



Figure 2: Mean Health at Retirement by Age

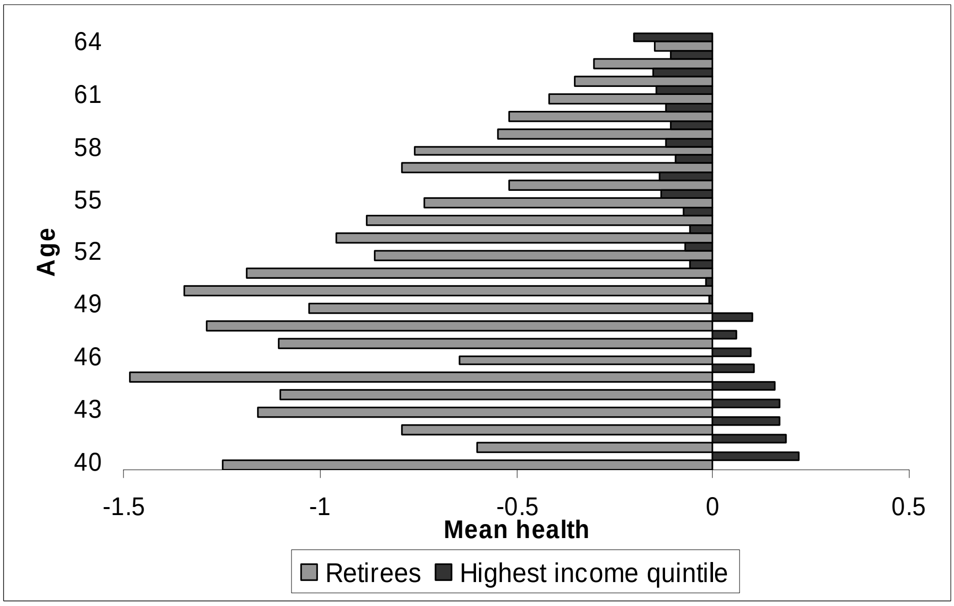
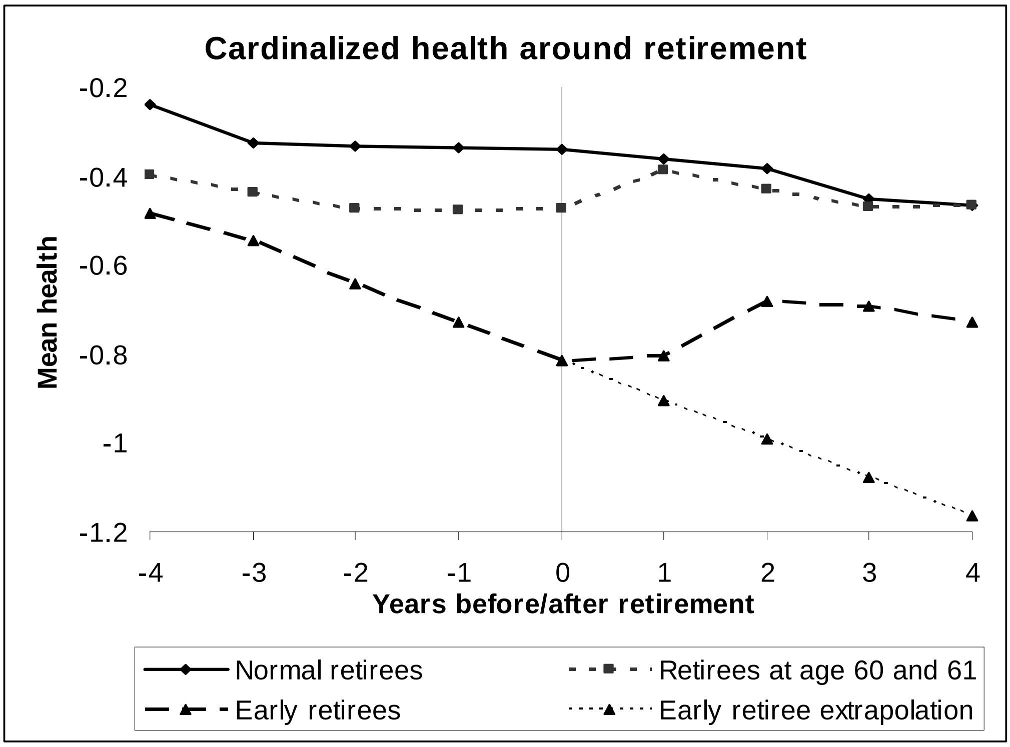
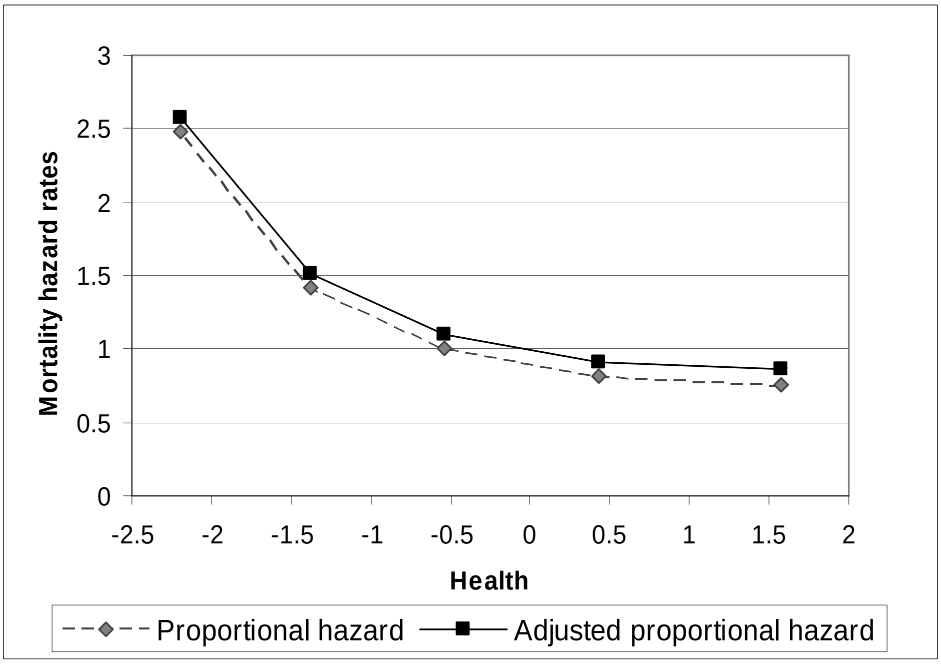


Figure 3: Development of Mean Health Around Retirement




*Source: Own calculations based on the GSOEP 1992 - 2005*

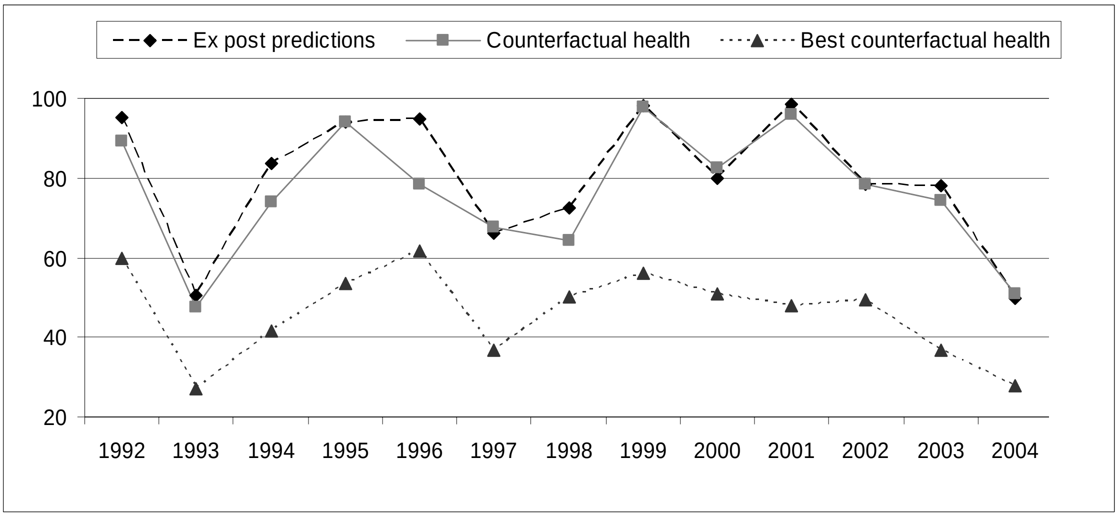
Figure 4: Mortality Hazard Rates Dependent on Health



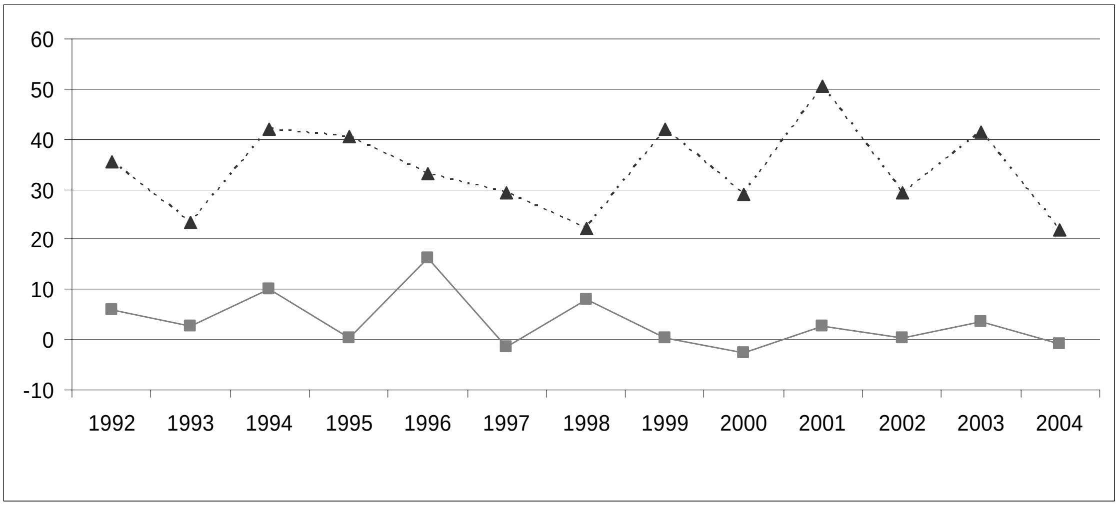

*Source: Own calculations based on the GSOEP 1992 - 2005*

Figure 5: Components of Social Costs in Early Retirement, in billion Euros

*Panel A: Annual Aggregate Labour Income Losses*




*Panel B: Annual Gains from Improved Health*




*Panel C: Annual Social Costs of Early Retirement*