Retirement, Death, and the Distribution of Retirement Leisure

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ABSTRACT. Using longitudinal data from the Health and Retirement Study and an ordered SUR Tobit model, we jointly analyze retirement age and death age to see who consumes retirement leisure. We find that retirement age and death age are indeed correlated. Hispanic and non-white individuals, and those who suffer from poorer self-assessed health, heart or mobility problems all successfully compensate for differences in mortality risk: they have a statistically significantly different retirement or death age, but statistically indistinguishable total retirement leisure. Attempts to fix Social Security or pension funding by increasing pension eligibility ages would disproportionately affect some of these groups (C34, J11, J26).

I. Introduction

Ongoing discussion of the long-term solvency of the Social Security system and funding shortfalls in some notable defined benefit pension plans keep returning the institution of retirement in the United States to a prominent role in the public discourse. Critics of traditional pension plans contend that at current rates of funding these plans are unsustainable because of the increased amount of time that the average American spends in retirement. This trend is evident when examining the difference between retirement age and life expectancy presented in Figure 1 for men and in Figure 2 for women. For both men and women median retirement ages have consistently declined from 1950 to 2005 with a steady climb in life expectancy over the same time period. Both of these trends have contributed to the increased amount of retirement leisure consumed by the average man and woman. As rising life expectancies are generally considered to be a desirable trend for society, an often-cited solution to the financial problems is to delay pension eligibility ages, bringing total retirement leisure in line with the expectations of plan designers. This was indeed the tactic taken by the 1983 Greenspan Commission who, spurred by the observed trend in retirement leisure and the future retirement of the Baby Boomers,

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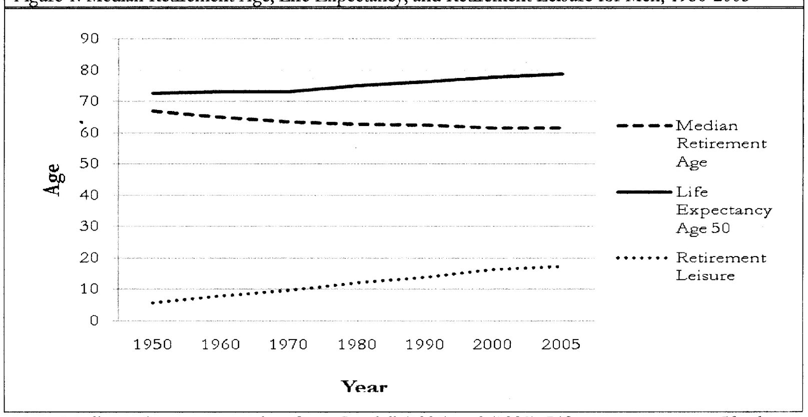
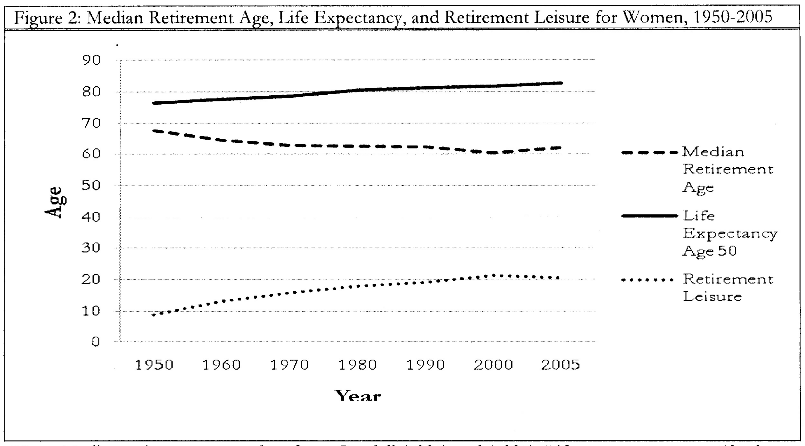
recommended a phased increase in Social Security eligibility ages to their current levels, a move which temporarily alleviated concerns over Social Security finances.

Figure 1: Median Retirement Age, Life Expectancy, and Retirement Leisure for Men, 1950-2005



Note: Median retirement ages taken from Gendell (2001) and (2008). Life expectancy at age 50 taken from Arias (2002) and (2007). Life expectancy data for 2005 actually 2004. Retirement leisure is the difference between life expectancy at age 50 and median retirement age.

Figure 2: Median Retirement Age, Life Expectancy, and Retirement Leisure for Women, 1950-2005



Note: Median retirement ages taken from Gendell (2001) and (2008). Life expectancy at age 50 taken from Arias (2002) and (2007). Life expectancy data for 2005 actually 2004. Retirement leisure is the difference between life expectancy at age 50 and median retirement age.

While delaying eligibility ages for retirement programs may solve the financial problems of Social Security and other defined benefit types of plans, the proposals may be based on a faulty belief that retirement ages can be used as a proxy for retirement leisure consumption and thus pension liability. Any common determinant of early retirement and early mortality, such as poor health, would invalidate this claim by showing that those who retire earlier may not be consuming more retirement leisure. Individuals presumably have more information about their mortality risk than administrators of retirement plans, whether private or governmental. The life-cycle model of intertemporal optimization suggests that individuals would use information about their own life expectancy to determine their own optimal retirement age. This leads to the compensation hypothesis: all else equal, individuals who expect to die earlier will retire earlier than individuals who expect to live longer (Ghilarducci and Neuman, 2004). If this is true, it suggests that eligibility age changes may have distributional effects which are being ignored, or at least improperly identified, in the policy analysis.

Delaying eligibility ages for retirement benefits may have minimal impact on the long-lived if they are already retiring later than the minimum age of eligibility. However, changes would impose a disproportionate burden on the short-lived, whose retirement leisure would be dramatically affected by such policies. To the extent that funding shortfalls in retirement programs are caused by over consumption of retirement leisure by the long-lived, fixing the shortfall at the expense of the short-lived appears to be an inequitable solution.

In this paper we use longitudinal data and an ordered Seemingly Unrelated Regression (SUR) Tobit model to measure the distribution of retirement leisure and jointly estimate the extent to which individuals successfully modify their retirement behavior to compensate for mortality risks of which they are aware. To the extent that this happens, adjustments to retirement eligibility ages would have a marginal impact that falls almost entirely on the short-lived and not the long-lived.

Despite the obvious applications to public policy, particularly related to Social Security, little research has been conducted examining the determinants of individual retirement leisure. Hurd et al. (2004) stress the important policy implications of the relationship between mortality risk and retirement age. They examine the effect of subjective survival probabilities on the decision to retire and the decision to claim Social Security benefits. Their study uses self-reported subjective survival probabilities as a proxy for the mortality risk they acknowledge to be the key policy implication. The paper presents a strong case for the importance of jointly examining retirement and mortality, however, if subjective life expectancy does not directly translate to actual mortality their approach may lead to misleading results.

Ghilarducci and Neuman (2004) explicitly examine the determinants of early retirement leisure, or leisure before the age of 65, using the Health and Retirement Study. Because the authors used observed retirement leisure as the dependent variable, data censoring issues caused by non-random death age observation limited their analysis to early retirement leisure. The limitation to early retirement leisure means that the significant factors in their study may be very different from factors influencing total retirement leisure. As the authors identified, retirement behavior correlated with mortality may not mean significantly different retirement leisure. However, in their analysis of early retirement leisure mortality differences between individuals likely did not have enough time to show up in the leisure measure. This made retirement age fluctuations the primary driver of early retirement leisure changes, which may not be the case with total retirement leisure.

Our study addresses the problems encountered by Ghilarducci and Neuman, and extends the analysis of Hurd et al., to measure total retirement leisure directly, incorporating findings from other studies on the individual determinants of retirement age and death age. While we apply the technique to retirement leisure it could be applied to any situation with endpoints that may only be observed in a particular order such as unemployment spells, disability leave, recessions, international treaties and accords, etc.

II. Econometric Strategy

A. RETIREMENT AGE AND DEATH AGE: CORRELATED EVENTS

A relatively simple first test of the compensation hypothesis would be to test whether retirement age and death age are correlated. If this is not the case, or if the correlation is negative, it would cast doubt on whether individuals are actually attempting to compensate for earlier death with earlier retirement. Because of the incomplete information in the data about retirement and death ages (some individuals were still living and/or working at the time of last observation), this correlation coefficient can be extracted by estimating a Tobit model without explanatory variables for retirement and death ages, then testing the resulting *r* parameter from the bivariate normal stochastic term against the null hypothesis that it is equal to zero. We reject this null hypothesis, finding that retirement and death ages are correlated with a coefficient of 0.2216 (95% confidence interval between 0.1308 and 0.3040). This is consistent with the life cycle model's prediction of compensation – those who die earlier tend to retire earlier also. But this does not explain the particulars of compensation: is compensation for earlier mortality more successful for some conditions as opposed to others? We examined further to investigate whether some determinants of retirement age and death age contributed to the correlation between the two events.

B. THE ORDERED SUR TOBIT MODEL

Time spent in retirement, the subject of our study, is the time that elapses between two events: retirement and death. If all the members of our population of interest were already dead, the simple variable constructed by taking the difference between the dates of these two events would be sufficient. This, however, is not the case; some members of our population are still working, and even more are still alive. Only sampling from the dead members of the population would both bias the results and fail to take advantage of information that we do know about individuals who are still working or still alive at a particular age. The situation calls for a two-equation model to examine the determinants of the time elapsed from a well-defined starting point to each event.

Even in a two-equation model using retirement age and death age to jointly define retirement leisure, a variety of problems arise. Data censoring leads to bias in ordinary least squares estimation of the two variables separately. However, it is also likely particularly in the context of an individual's retirement and death that the two events may be correlated in their error term. Any unobserved determinant of both events, for example an individual's discount rate which influences retirement as well as mortality through health investment and behaviors, would bias the individual equation results. The correlation between the two equations thus precludes using standard independent Tobit models to estimate the time elapsed until the two events separately.

A standard SUR bivariate Tobit model (Huang, 1999) might seem appropriate, but it too has a flaw in this context: in assuming that retirement age and death age are bivariate normally distributed, one assumes that there is some non-zero probability that retirement age is greater than death age. By our definition of retirement leisure, an individual cannot have a negative amount of it – that is, we assume that individuals stop working when they die. Thus, retirement age and death age cannot be jointly normally distributed.

Our approach is to use a modified SUR bivariate Tobit model that we call the ordered SUR Tobit. We develop a maximum likelihood model to estimate the determinants of the age of retirement and the age of death, allowing for correlation between the error terms, and where observation of both events is censored if it has not taken place by the time of observation and observation of retirement is censored if death takes place first. The ordered SUR Tobit model we develop takes advantage of all information available at the time of observation. A detailed derivation of the model can be found in the Appendix.

The maximum likelihood estimation technique produces results for the covariates in the individual retirement and death age equations. While the individual equation results are interesting in their own right, the results do not tell us the joint effect of a covariate on retirement leisure. A simple SUR Tobit model would recover the effect of a variable on retirement leisure by subtracting its effect on retirement age from its effect on death age. If a chi-squared test of the equality of a covariate's coefficients across the retirement and death age equations rejects the equality of the coefficients, we conclude that coefficient has a statistically significant effect on retirement leisure. In this case, the zero net effect may be due to a covariate failing to influence either retirement or death age, or it may be due to correlations between the two events causing the individual effects to offset each other.

Our approach is slightly more complex. This difference in coefficients from the maximum likelihood estimation can only affect retirement leisure as predicted by the SUR Tobit if the resulting difference between the predicted death age and retirement age is non-negative. If it changes the difference from one negative number to another, there is zero effect on actual retirement leisure. If the change in a variable changed the predicted difference from positive to negative, the marginal effect will be closer to zero than predicted by the SUR Tobit model. This model thus predicts asymmetrical marginal effects – the effect of increasing a variable on retirement leisure might not be the same as the corresponding decrease in the variable. In any event, the actual marginal effect of a variable will be the same as or be closer to zero than the effect predicted by the SUR Tobit model.

The marginal retirement leisure effects reported in Table 2 are the average of the upward and downward marginal effects. While technically the marginal effects are asymmetrical, we present the average of the effects due to the fact that the differences between the upward and downward effects are small in practice. Because of the complexity of this estimator of effects on retirement leisure, we bootstrapped the estimator to predict statistical significance, sampling with replacement from our original sample to generate 1000 re-samplings. We repeated the experiment for each of these 1000 re-samplings and ordered the predicted marginal effects of each variable. When all but the bottom 0.5% and top 0.5% of the samples yielded marginal effects with the same sign, we designated the results significant at the 99% level; similar calculations were used to determine the other significance levels.

III. The Health and Retirement Study Data Sample

The Health and Retirement Study (HRS) is a biennial longitudinal data set which began in 1992 interviewing individuals between the ages of 51 and 61 on a variety of topics. A detailed description of the HRS can be found in National Institutes of Health (2007). For our estimation we utilize a subset of the HRS born between the years of 1931 and 1935 who were interviewed in the initial wave of the survey. We also limited our sample to those individuals who had worked at least ten years in their lifetimes so that the concept of retirement would be applicable to the group. We then followed individuals from 1992 to 2004 in order to observe retirement and death. After removing 239 observations due to missing values we arrived at a final sample of 3,410 individuals.

As the joint effect on retirement leisure is our primary concern in the present analysis we chose to use the same covariates in the retirement and death age equations of the ordered SUR Tobit model. These covariates are broken down into three main categories: demographics, health, and financial variables. All time varying information except for retirement and death ages is taken from the 1992 wave of the survey, and all financial information is in 1992 dollars.

We present the variable definitions and summary statistics in Table 1.

TABLE 1–Summary statistics of HRS data sample (N=3,410)[a]

Variable	Variable Definition	Mean	Std. Dev.
Retirement Age N=2,845	First self-reported retirement age	59.36	5.7963
Death Age N=598	Observed death age	66.10	3.4420
Demographics			
Coupled	1=Yes, 0=No	0.75	0.4345
Female	1=Yes, 0=Male	0.48	0.4996
Non-white	1=Yes, 0=White	0.20	0.4007
Hispanic Ethnicity	1=Yes, 0=Non-Hispanic	0.07	0.2597
School years	Number of years spent in school	12.12	3.1887
Health Insurance	1=Yes, 0=No	0.87	0.3359
Lose Employer Health Ins. At Ret.	1=Yes, 0=No	0.28	0.4485
Blue-collar High Skilled[b]	1=Yes, Base=low-skill blue-collar	0.27	0.4435
White-collar High Skilled[b]	1=Yes, Base=low-skill blue-collar	0.30	0.4565
White-collar Other[b]	1=Yes, Base=low-skill blue-collar	0.24	0.4269
Goods producing Industry	1=Yes, 0=No	0.28	0.4497
Midwest	1=Yes, Base=Northeast	0.25	0.4328
South	1=Yes, Base=Northeast	0.41	0.4921
West	1=Yes, Base=Northeast	0.15	0.3607
Health Status and Health Behavior			
Good Health	1=Good or better, 0=Fair or poor	0.78	0.4160
Ever Smoked	1=Yes 0=No	0.65	0.4768
Smoke Now	1=Yes, 0=No	0.25	0.4328
Arthritis	1=Had condition, 0=Never had	0.37	0.4841
Cancer	1=Had condition, 0=Never had	0.06	0.2300
Diabetes	1=Had condition, 0=Never had	0.10	0.3062
High Blood Pressure	1=Had condition, 0=Never had	0.38	0.4845
Heart Problem	1=Had condition, 0=Never had	0.13	0.3317
Lung Disease	1=Had condition, 0=Never had	0.06	0.2426
Psychiatric Problem	1=Had condition, 0=Never had	0.06	0.2377
Stroke	1=Had condition, 0=Never had	0.02	0.1541
Sum of ADL Measures	Responses of ‘some difficulty’	0.08	0.3949
Sum of Large Muscle Measures	Responses of ‘some difficulty’	0.63	1.0533
Sum of Mobility Measures	Responses of ‘some difficulty’	0.56	1.0846
Relative Probability of Living to Age 75	Self-report relative to life table prob.	0.93	0.4298
Wealth and Pension			
Value of IRA	Value of IRA/Keogh in $1992	22,793	66,392
Net Financial Wealth	Financial assets less debt in $1992	53,338	162,945
Net Housing Value	Home value less mortgages in $1992	65,162	75,108
Defined Contribution Plan Value	Present value in $1992	16,941	61,563
Defined Benefit Plan at Last Job	1=Yes, 0=No	0.28	0.4494
Defined Benefit Plan at Previous Jobs	1=Yes, 0=No	0.19	0.3886

[a]All information except for retirement age and death age is taken from the initial 1992 survey wave. Retirement age and death age are updated from successive waves of the survey.
[b]The occupation categories follow the classification of Quinn (1996).

In the top two rows of the table we present information about our two dependent variables. Of our entire sample we have observations on retirement age for 2,845 individuals or 83% of the sample, with an average retirement age of just over 59. To construct our retirement age variable we use information from the initial interview in 1992 and add information from each successive wave of the survey, filling in missing values if the individual reported that they had retired since the previous wave. We observe a death age for fewer individuals as could be expected given the relatively young ages of the sample, but still have information for 598 individuals, or about 18% of the sample. Combining the dependent variables we have complete retirement and mortality information for 460, or 13.5% of the sample. Limiting our analysis to just these individuals, even if doing so did not introduce sample bias issues, would have severely limited our study.

Along with other control variables a particularly interesting variable we include is a measure of the individual's expected probability of living until age 75 relative to the implied life table probability of living until age 75 conditioning on age and sex. Studies of subjective life expectancy in the HRS have shown that the measures predict actual mortality even when controlling for health status (Hurd and McGarry, 2002; Hurd et al., 1999). We include the life expectancy information under the hypothesis that individuals may alter their retirement behavior to compensate for below average expectations of retirement leisure consumption due to earlier expected deaths. This idea, termed the "compensation hypothesis" by Ghilarducci and Neuman (2004) would be one of the clearest examples of how retirement and death age may be related. A few other studies have examined the effect of subjective life expectancy on retirement behavior alone, finding conflicting results on whether an effect exists (Bloom et al., 2006; Hurd et al., 2004). We also include a quadratic term for relative life expectancy to check for non-linear effects.

The three groups of covariates cover a wide range of factors that prior studies examining retirement and mortality independently have found to be significant predictors of one, or both, of the events. Presumably individuals use all of this available information when making the optimal retirement decision, and thus, the variables can be possible dimensions upon which compensatory behavior can be based. Even if individuals do not actively modify their retirement behavior based on a piece of information, for instance race, finding that non-whites retire and die significantly earlier with no significant difference in retirement leisure still shows that the group is compensated for their earlier death, no matter how passive was their participation in the compensation. By showing who consumes less retirement leisure, and who retires earlier to compensate for earlier death, our study will help identify who will be adversely affected by increases in pension eligibility ages.

IV. Who Consumes Retirement Leisure?

A. RESULTS OF THE ORDERED SUR TOBIT MODEL

The results including all of the covariates are presented in Table 2, with the results for our primary model specification presented in the first panel. The first column presents the coefficients for the various variables' effects on retirement age while the second column presents the coefficients for the variables' effects on death age. These results show the effects of the covariates on the retirement leisure endpoints individually and can be compared to previous literature on retirement and mortality. The third column presents the net effect on retirement leisure.

Looking at the retirement age and death age equations we find a number of significant factors with virtually all matching our expectations from prior research. These results are interesting by themselves and demonstrate that our sample displays similar results to existing research. However, the more intriguing results for the joint effect on retirement leisure are in the final columns of the first results panel of Table 2. The net effect on retirement leisure is approximately equal to the negative of the retirement age coefficient (earlier retirement implies more retirement leisure) plus the death age coefficient. The coefficient is not precisely equal to the difference between the coefficients for reasons discussed in section II.

We find that women consume 4.716 more years of retirement leisure than men due to significantly earlier retirement, 1.433 years, and significantly later death, 3.283 years. While the longevity results are well documented, the fact that women retire earlier than men further increases their relative leisure consumption. Coupled individuals experience a similar effect where the effect of retirement and death work in the same direction, extending retirement leisure for members of the group. In these cases using retirement age as a proxy for retirement leisure would give the correctly signed effect, but it would severely underestimate the magnitude.

TABLE 2–Results from the order SUR Tobit model

	Using first reported retirement age N=3,410			Using last reported retirement age N=3,410			Using only individuals not retired at first survey N=2,388		
	Retirement Age	Death Age	Retirement Leisure	Retirement Age	Death Age	Retirement	Retirement Age Leisure	Death Age	Retirement
Demographics									
Coupled	-1.242***	1.599***	2.841***	-1.062***	1.587***	2.648***	-0.723***	2.058***	2.776***
Female	-1.433***	3.283***	4.716***	-1.485***	3.290***	4.773***	-1.243***	3.640***	4.868***
Non-white	-0.502	-1.441***	-0.940	-0.816**	-1.440***	-0.624	-0.013	-1.315**	-1.300*
Hispanic	2.126***	0.654	-1.472	1.648***	0.663	-0.985	1.026**	0.507	-0.519
School Years	-0.065	-0.009	0.056	-0.017	-0.008	0.010	-0.020	0.044	0.064
Health Ins.	-0.955**	0.467	1.422*	-1.165***	0.476	1.640**	0.225	0.864	0.639
Lose Emp HI at Ret.	1.441***	-0.047	-1.488***	1.386***	-0.051	-1.437***	0.939***	-0.268	-1.206
Blue Col. High Skill	-0.446	0.526	0.972	-0.379	0.525	0.904	-0.381	0.736	1.116
White Col. High Skill	0.661	0.688	0.027	0.527	0.679	0.152	0.263	0.445	0.182
White Collar Other	0.606	0.441	-0.165	0.530	0.420	-0.110	0.136	-0.603	-0.738
Goods Prod. Industry	0.229	-0.027	-0.257	0.168	-0.016	-0.185	-0.370	-0.152	0.218
Midwest	0.026	-0.208	-0.234	0.009	-0.202	-0.212	-0.200	-0.023	0.177
South	-0.023	-0.061	-0.038	-0.005	-0.056	-0.051	-0.214	0.026	0.239
West	-0.890**	-0.041	0.849	-0.795*	-0.045	0.750	-0.763**	0.477	1.239
Health Status & Behavior									
Good Health	1.626***	1.664***	0.038	1.939***	1.665***	-0.273	0.433	0.956	0.523
Ever Smoked	-0.477	-1.548***	-1.070**	-0.320	-1.566***	-1.245**	-0.383	-1.919***	-1.534**
Smoke Now	-0.047	-3.053***	-3.006***	-0.313	-3.050***	-2.736***	0.045	-3.444***	-3.482***
Arthritis	-0.627**	0.290	0.918*	-0.537**	0.288	0.825*	-0.034	0.000	0.033
Cancer	-0.509	-3.741***	-3.232***	-0.404	-3.750***	-3.345***	0.634	-3.792***	-4.415***
Diabetes	-0.706*	-3.280***	-2.574***	-0.754*	-3.283***	-2.529***	-0.568	-3.176***	-2.604**
High Blood Pressure	-0.080	-1.069**	-0.989**	-0.109	-1.072**	-0.963**	0.145	-1.147**	-1.290**
Heart Problem	-1.072***	-2.206***	-1.134	-0.978**	-2.208***	-1.229*	-0.642*	-2.598***	-1.954**
Lung Disease	0.360	-3.037***	-3.396***	-0.095	-3.041***	-2.945***	0.112	-3.234***	-3.339**
Psych. Problem	-0.056	-0.952	-0.896	-0.045	-0.977	-0.932	0.706	-2.089*	-2.790**
Stroke	-0.915	-3.236***	-2.322*	-0.346	-3.221***	-2.874**	1.937*	-2.585	-4.509**
ADL Measure	-0.689*	-0.690	-0.001	-0.665*	-0.683	-0.018	0.696	-1.703*	-2.395
Large Muscle Measure	-0.524***	0.232	0.757***	-0.495***	0.235	0.730***	-0.143	0.234	0.377
Mobility Measure	-0.356**	-0.387*	-0.031	-0.325**	-0.392*	-0.067	-0.262	-0.472	-0.209
Rel Prob Live to 75	1.731	1.381	-0.350	0.866	1.378	0.512	-0.105	3.306	3.136
(Rel Prob Live to 75)^2	-0.992	-0.600	0.392	-0.482	-0.604	-0.122	0.084	-1.682	-1.764
Wealth and Pension									
Value of IRA ($10k)	-0.021	0.014	0.035	-0.031***	0.014	0.045	-0.028	0.110	0.138***
Net Fin Wealth ($10k)	-0.026***	0.019	0.045***	-0.027***	0.020	0.046***	-0.005	0.036	0.041
Net Housing Val ($10k)	-0.017	-0.008	0.009	-0.017	-0.008	0.009	-0.006	-0.043	-0.037
DC Plan Value ($10k)	0.621**	0.398	-0.223	0.244	0.406	0.162	-0.611**	0.665	1.275*
DB Plan Last Job	2.469***	0.162	-2.307***	1.942***	0.168	-1.774***	-0.966***	-0.070	0.896
DB Plan Previous Jobs	-3.420***	0.566	3.986***	-3.153***	0.549	3.702***	-0.753**	0.635	1.386

***Statistically significant at the 1% level **at the 5% level *at the 10% level.

Using retirement age or death age as leisure proxies by themselves would be relatively accurate in some situations. In the case of smoking, cancer, high blood pressure, lung disease, and strokes, the covariate seems to affect death age but has no significant effect on retirement age; in these cases retirement leisure is reduced by decreased life expectancy uncompensated by movement of the retirement age variable. Similar results occur for health insurance, arthritis, large muscle problems, net financial wealth, and the Defined Benefit pension indicators except that it is retirement age which is significantly altered with no corresponding change in death age.

Perhaps the most interesting results arise when examining the compensation hypothesis of Ghilarducci and Neuman (2004). The authors hypothesized that workers who expect to die earlier would retire earlier to compensate for expected lost leisure. If this behavior is taking place, using retirement age as a proxy would completely misidentify the true effect on retirement leisure. We do not see this happening with the relative probability of living to age 75 as we expected, but we may see this happening in the case of subjective personal health assessment, heart disease, and mobility. Those in worse health, with heart disease, and with more mobility problems retire and die significantly earlier, but the significant effects on both of the endpoints offset each other such that the net effect on retirement leisure is insignificant. Given the subjective nature of the health assessment and knowledge of these two conditions, it is quite possible that these variables are capturing compensatory behavior. We find a similar result for diabetes in that individuals with diabetes both retire and die significantly earlier. What is particularly interesting though is that they still consume significantly less retirement leisure as their earlier deaths overwhelm their earlier retirements. These individuals may have been trying to compensate, but were not able to do so successfully. We may also see some evidence of compensatory behavior with race, Hispanic ethnicity, and Defined Contribution (such as 401(k)s) value. Non-Hispanics and those with less value in a Defined Contribution plan retire significantly earlier, while non-whites die significantly earlier, with no significant effect on retirement leisure.

Even if the above results are not deliberate attempts to alter retirement to compensate for expected lost leisure, from a purely measurement perspective the results provide evidence that joint estimation is important in order to accurately analyze retirement leisure effects and related issues. In terms of subjective health, heart disease, and mobility using retirement age as a proxy would suggest over-consumption of retirement leisure and death age would suggest under-consumption, neither of which is correct. In the less extreme cases of Hispanic ethnicity, Defined Contribution plan value, and race where only one endpoint was significant, using the information as a proxy would still be inaccurate.

B. IMPLICATIONS OF RESULTS FOR INCREASES IN ELIGIBILITY AGES

For policies recommending raising pension eligibility ages, particularly Social Security, our results have some obvious implications. We have shown that some people do not seem to engage in compensatory behavior and thus show predictable and easily identifiable results regarding retirement leisure, but others do seem to compensate and therefore have less predictable and identifiable results when using single equation models. From a policy perspective, whether the compensation is active or passive is not all that important as the results still show which groups do have a correlation between retirement age and death age. This information should be used when evaluating the costs of any proposed policy regarding eligibility ages.

Before evaluating the policy effects it is important to emphasize that actual retirement behavior may not change just because eligibility ages increase. Using the current Social Security ages of 62 for receipt of partial benefits and 67 for full benefits for concreteness, if individuals are already optimally retiring well after these eligibility ages, increases may not alter their behavior. In our sample that does not seem to be the case. Using actual retirement ages when observed and predicted retirement ages (at variable means) when not observed, the predicted average retirement age in our sample is 62.6 years old (ranging between 50.6 and 69.4). While clearly not everyone retires at the average age, the majority of workers and particularly the lower tail of the distribution are retiring in age ranges where they would presumably be affected by Social Security eligibility age increases. One might argue that those retiring prior to the original eligibility ages would also be unaffected by increases, but this ignores the fact that people likely take expected future payments into consideration when choosing to retire. Even if not currently receiving checks, delaying these payments will likely alter decision making for some people. Assuming then that most in our sample will be affected by the eligibility age increases we can use our results to show which groups are most affected and what distributional effects may have been missed by not jointly estimating retirement and death.

Presumably, eligibility age increases are aimed at those individuals who retire earlier and actually do consume more retirement leisure. This group over-consumes retirement leisure, thus contributing more to pension costs, and at least from a financial perspective their behavior should be adjusted accordingly. In theory, early benefits are reduced so that the lifetime stream of benefits is actuarially fair. However, if individuals do not have average life expectancies then this may not be true. This is of particular importance from a funding perspective if individuals live longer than expected. From our analysis we find that this group includes women, coupled workers, workers with health insurance they keep when they retire, workers with arthritis, workers with large muscle problems, wealthier workers, and workers with defined benefit plans from previous jobs. Women and couples are particularly noteworthy in this group as they not only retire earlier but live longer, making them over-consumers of leisure due to both ends of retirement. While these groups would be adversely affected by Social Security eligibility age increases these are precisely the groups which the policy is designed to affect.

A second group would also be particularly affected by the policies. These workers do not retire significantly earlier, but do consume significantly less retirement leisure due to significantly earlier deaths. Our analysis finds that this group consists of smokers (ever and now), and those with cancer, high blood pressure, lung disease, and who have had a stroke. Presuming their retirements are now delayed by the eligibility age increases these already disadvantaged groups would be further harmed by the policies even though they are ironically helping the funding problems through their earlier deaths. These types of workers are always a point of contention regarding eligibility age policies, but given that they are relatively easy to identify with single equation models it has generally been recognized that those in worse health will be disproportionately harmed by these polices if they cannot exit the labor force through another avenue such as Social Security Disability.

From the perspective of our study the most important group of workers that will be affected by the policies are those who engage in compensatory behavior. The reason these workers are so interesting is that they are erroneously attributed with excessive retirement leisure consumption due to faulty measurement. This group would include anyone who retired earlier but did not actually consume more retirement leisure. In these cases the earlier retirement simply kept the individuals with the same retirement leisure. As described above our study finds that this group would include non-Hispanic workers, non-white workers, workers who do not consider themselves to be in good health, and workers suffering from heart problems, diabetes, and mobility problems. If the policies alter retirement ages, these individuals would experience a disproportionate segment of the costs of these policies, as they would be the target of reforms but would not actually be the guilty parties consuming greater retirement leisure. While perhaps not as troubling from a policy perspective as the second group of workers given that they are not starting at a disadvantage, the effects on these workers would generally not be identified properly using single equation models and the costs would not be correctly applied to the policy analysis. Only an analysis jointly estimating retirement and death like our study would capture the effects on these groups.

While any public policy necessarily imposes costs on particular groups of society for the benefit of others, it is important to understand just who actually will be bearing the costs. Our analysis shows that in this context using retirement age as a proxy for retirement leisure severely distorts the view of the distribution of retirement leisure and likely leads to unintended consequences of eligibility age reforms particularly for those who may be engaging in compensatory behavior. Studies examining retirement decisions and pension financing may reach suspect conclusions if they do not examine both retirement and death together.

C. SPECIFICATIONS AND MODEL TESTS

Using our ordered SUR Tobit we find a variety of factors that influence the amount of retirement leisure consumed by individuals. However, there are a few potential questions regarding the results. The first concern is in regard to the retirement age reported by individuals. Although a self-reported retirement age seems to be a relatively straightforward variable, in reality individuals often report more than one retirement age, either because they answered the question incorrectly one time, or the individual returned to work.

The first specification on the table is based on our primary definition of retirement age. In this specification, the earliest date at which a worker self-identifies as "retired" is used to calculate retirement age. This first date would likely be the primary retirement decision in the traditional sense as this was most likely to be when they left their career job. To check for potential biases of using the first retirement age, we estimate a specification where we update our retirement age information if the individual reported a later retirement age after first being questioned. Death age is calculated the same in the two specifications.

Results using the updated retirement age are presented in the second panel of Table 2. There are two differences between the results from the first two specifications. First, using the second specification, workers who are non-white retire earlier than white workers, although in the first specification there is no significant difference between the retirement ages of the two populations. This suggests that white workers are more likely to return to work after reporting they have retired than workers who are not white. In terms of the predicted policy effects, this confirms what we had expected in our baseline estimates that non-whites were compensated for their earlier deaths through their retirement behavior. Secondly, the value of a worker's IRA has a significant effect on retirement age in the second specification, but not in the first. This suggests workers might be coming out of retirement for financial reasons.

A second concern arises from the fact that some of our individuals had already retired when they were first interviewed. While we were able to obtain information about their retirement age retrospectively, the problem is that some of the time varying information is obtained from the individual after they had retired. If retirement itself changes the individual's characteristics, what we are using to explain the individual's retirement may be much different than what actually caused the individual to retire. This is particularly problematic for information about net wealth, IRA value, housing value, and Defined Contribution plan value. Individuals who had already retired at the time of first interview likely have less-than-peak values for these variables because being retired they have already started spending down some or all of these assets. To check for issues relating to this problem we restrict our sample to only those individuals who had not yet retired by the time of first interview when we collected their time variant information. Thus all information could be considered to potentially cause retirement and not be altered by retirement itself. On the other hand, this specification is flawed because it systematically excludes individuals who retired earlier.

Results testing this concern regarding the timing of retirement in our sample are presented in the final section of Table 2. Initial analysis indeed shows differences between these results, but we cannot determine the extent to which these differences are due to the status at time of first observation (more likely for the financial variables; less likely for demographic and health variables) and the extent to which these differences are due to selection bias in the restricted sample.

A third explanation of the differences in these results is more benign. To the extent that certain variables have significant effects in specifications one and two but not in specification three, it bears noting that the sample size in specification three is roughly one-third smaller, potentially reducing the precision of our estimates and the number of significant coefficients. Closer examination of the results suggests that this might be the case as only stroke, Defined Contribution value, and Defined Benefit from last job have significant results with different signs than the first specification. Of these variables it is likely that the pension variables do suffer from the timing issue. A case can be made that the health insurance variable suffers from this same problem as well although it is not significant. However, the coefficients of the other retirement age covariates that were significant originally do have the same sign in the restricted sample, but simply lose their significance. This analysis suggests the timing of measurement may play a role in the results but that it is not an important issue for most of the covariates.

V. Conclusion

Although retirement leisure has indirectly become a topic of concern politically and socially due to Social Security funding issues, little is known about the actual distribution. Simple analysis of the topic using retirement age as a proxy for retirement leisure consumption may have distorted results because retirement age and death age are correlated. Using longitudinal data and a newly developed ordered SUR Tobit model we exploit all information available to estimate whether individuals modify their retirement behavior to compensate for increased expected mortality. Our model accounts for potential correlations between retirement and death ages and captures total retirement leisure, not just changes in the two end points. We then use this more precise information about who consumes retirement leisure to analyze the effects of increases in Social Security eligibility ages.

In terms of the effect on retirement and death age individually our results closely match those from prior research. When examining the joint effect of retirement and death on retirement leisure our estimation strategy finds a number of instances where using retirement age or death age as simple proxies for retirement leisure would be reasonably accurate. In these instances those who consume more retirement leisure due to earlier retirement, and those who consume less due to earlier death, would be adversely affected by increases in eligibility ages, assuming their retirement behavior is actually altered.

More importantly we find a variety of factors that significantly affect retirement and death age independently but do not influence retirement leisure jointly. For these factors, the Ghilarducci-Neuman compensation hypothesis does seem to hold: individuals who do not consider their overall health to be good, individuals with heart problems, and individuals with mobility problems expect to die earlier (correctly), and thus retire earlier to compensate, leaving no significant net effect on retirement leisure. Hispanic and non-white workers may also be engaging in this type of behavior. Analyses using only one determinant of retirement leisure would capture these results and conclude that these individuals consume significantly different retirement leisure even though they do not. These results show that eligibility age policies aimed at early retirees need to take total retirement leisure consumption into account as the policies may not be aimed at individuals who are actually consuming more retirement leisure and driving pension funding problems.

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Appendix

Suppose the existence of two events of interest that occur to an individual in time: retirement and death. Let R^*_i be the time elapsed from this individual's birth until individual *i* retires or would have retired had individual *i* lived that long. For individuals who retire before they die, R_i^* is that individual's retirement age; for individuals who die without retiring, R_i^* is the (unobservable) age at which they *would have retired*. Let D^*_i be the time elapsed from birth until death takes place for the same individual. D^*_i is observable if it has taken place by the time of observation. R^*_i is observable if it has taken place by the time of observation and if death did not take place before retirement. Define A_i (the individual's age) to be the time elapsed from birth until the time of observation. Thus, retirement leisure can be defined as $D_i^* - R_i^*$ if $D_i^* > R_i^*$ and zero otherwise.

Assume that the determinants of R^*_i and D^*_i can be modeled as follows:

$$R_i^* = x_iB + \varepsilon_i^r \quad (1)$$
$$D_i^* = y_iG + \varepsilon_i^d \quad (2)$$

In Equations (1) and (2), ε_i^r and ε_i^d are bivariate normally distributed with means of zero, variance of σ^2_r and σ^2_d, and correlation of ρ. This bivariate observation of the error term for any individual *i* is independent of the error terms for other individuals $j \neq i$. Given this model, one can calculate the likelihood of any set of observations of R_i, D_i, and A_i. There exist four cases that could apply to the information observed (or not observed) regarding R^*_i and D^*_i:

Case 1: R^*_i and D^*_i observed as R_i and D_i.
Case 2: R^*_i observed as R_i, D^*_i not observed but greater than or equal to A_i.
Case 3: D^*_i observed as D_i; R^*_i not observed but greater than D_i.
Case 4: R^*_i and D^*_i both not observed, but greater than A_i.

Define dummy variables W_i and L_i such that L_i is equal to zero if D^*_i is actually observed (i.e. the individual stops living before the time of last observation) and one otherwise, while W_i is equal to zero if R^*_i is actually observed (i.e. the individual stops working before both event two and the time of last observation) and one otherwise. Then, the likelihood function corresponding to this model would be as follows:

$$
\begin{aligned}
L = & \prod_{(1-W_i)(1-L_i)=1} L(R_i = x_i B \wedge D_i = y_i G) \prod_{(1-W_i)(L_i)=1} L(R_i = x_i B \wedge y_i G \geq A_i) \\
& \prod_{(W_i)(1-L_i)=1} L(x_i B \geq D_i \wedge D_i = y_i G \prod_{(W_i)(L_i)=1} L(x_i B \geq A_i \wedge y_i G \geq A_i) \\
= & \prod_{(1-W_i)(1-L_i)=1} L(R_i = x_i B | D_i = y_i G) L(D_i = y_i G)
\end{aligned} \tag{3}
$$

$$
\begin{aligned}
& \prod_{(1-W_i)(L_i)=1} L(y_i G \geq A | R_i = x_i B) L(R_i = x_i B) \\
& \prod_{(W_i)(1-L_i)=1} L(x_i B \geq D_i | D_i = y_i G) L(D_i = y_i G) \\
& \prod_{(W_i)(L_i)=1} L(x_i B \geq A_i \wedge y_i G \geq A_i)
\end{aligned} \tag{4}
$$

$$
\begin{aligned}
& \ln L(B, G, \sigma_D, \sigma_R, \rho | \{D_i, R_i, A_i, x_i, y_i, L_i, W_i\}_{i=1\ldots n}) = \\
& \sum_{i=1}^{n} (1 - L_i)(1 - W_i) \ln \frac{1}{\sigma_R \sigma_D \sqrt{1-\rho^2}} \phi\left(\frac{D_i - y_i G}{\sigma_D}\right) \phi\left(\frac{R_i - x_i B - \rho \frac{\sigma_R}{\sigma_D}(D_i - y_i G)}{\sigma_R \sqrt{1-\rho^2}}\right) \\
& + L_i (1 - W_i) \ln \frac{1}{\sigma_R} \phi\left(\frac{R_i - x_i B}{\sigma_R}\right)\left(1 - \Phi\left(\frac{A_i - y_i G - \rho \frac{\sigma_D}{\sigma_R}(R_i - x_i B)}{\sigma_D \sqrt{1-\rho^2}}\right)\right) \\
& + (1 - L_i) W_i \ln \frac{1}{\sigma_D} \phi\left(\frac{D_i - y_i G}{\sigma_D}\right)\left(1 - \Phi\left(\frac{D_i - x_i B - \rho \frac{\sigma_R}{\sigma_D}(D_i - y_i G)}{\sigma_R \sqrt{1-\rho^2}}\right)\right) \\
& + L_i W_i \ln\left(1 - \Lambda\left(\frac{A_i - x_i B}{\sigma_R}, \frac{A_i - y_i G}{\sigma_D}, \rho\right)\right)
\end{aligned} \tag{5}
$$

In equation (5), Φ is the cumulative density function of the standard normal distributions, φ is the probability density function of the standard normal distribution, and Λ is the cumulative density function of the bivariate standard normal distribution with correlation ρ. This log likelihood function would then be maximized using a Newton-Rhapson or similar algorithm to obtain a consistent maximum likelihood estimate, with all the standard properties thereof.