Does Labor Force Retirement Accelerate Mortality? General Considerations and Evidence from U.S. Supreme Court Justices, 1801-2006

r-stolzenberg@uchicago.edu
Department of Sociology, University of Chicago

ABSTRACT: Does labor force retirement accelerate death? Previous studies report positive, negative and null effects of retirement on mortality hazard. It is difficult to resolve these inconsistencies because 1) nearly all data confounds unemployment with retirement, and, often, 2) endogeneity bias is uncorrected. We analyze data that avoids these problems, from an exceptional subgroup, of interest in its own right: US Supreme Court justices, 1801 - 2006. Using discrete time event history methods, we examine retirement effects on mortality hazard and years-left-alive. Substantive and methodological considerations suggest several models. Some models specify endogenous effects estimated by instrumental variables (***IV***) probit, ***IV*** Tobit and ***IV*** regression methods. Others specify effects estimated by endogenous switching (***ES)*** probit and ***ES*** regression. All results are consistent with the hypothesis that retirement increases mortality hazard and reduces years left alive. Retirement effects on mortality hazard compare to the effects of smoking 2 packs of cigarettes daily.

I. INTRODUCTION

Does labor force retirement affect subsequent longevity and mortality risk? Some studies report that retirement tends to reduce the length of remaining life (Snyder and Evans 2006; Waldron 2001 2002), while others find the opposite (Munch and Svarer 2005; Handwerker 2007), or no retirement effect on mortality at all (Tsai 2005; Litwin 2007; Mein Martikainen Hemingway Stansfeld and Marmot 2003). This paper reconsiders existing research data and methods, and concludes that data limitations and consequent measurement and modeling problems reduce the certainty of earlier empirical results and perhaps produce inconsistencies among them. To avoid these measurement and modeling problems, we apply the venerable demographic strategy of analyzing unusual data from an exceptional population subgroup, of interest in its own right, in which these data problems and their methodological consequences are absent. That subgroup is justices of the U.S. Supreme Court from 1801 through 2006.

In data from nearly all other population segments, retirement is conflated with involuntary unemployment of older workers. Simply stated, involuntarily unemployed older workers and pension recipients who were forced to retire by employers tend to report that they are *voluntarily retired* (Gustman, Mitchell and Steinmeier 1995: s63; Stolzenberg 1989). This misreporting attributes to retirement the powerful, empirically-verified, pernicious effects of unemployment (e.g. Linn, Sandifer, and Stein 1985; Gerdtham and Johannesson 2003). More fundamentally, misreporting of retirement status confuses key concepts. Since at least 1990, retirement from the civilian labor force has been defined in social science research as a *worker's decision* to withdraw from the labor force, or to substantially reduce the hours, intellectual demands and/or physical intensity of paid work (see reviews in Moen, Kim and Hofmeister 2001; Lumsdaine 1995).[1] Whether made with pleasure or regret, the retirement decision is a rational voluntary action by the individual concerned. Retirement is distinguished fundamentally from involuntary labor force withdrawal, labor force exit due to disability, and job loss by firing or layoff.

Usual retirement misreporting problems are obviated in Supreme Court data because justices are Constitutionally protected from firing, and sheltered from workplace pressure to resign. Justices' retirement benefits and pay are fixed by law, their working conditions are not subject to employer manipulation, and regulations prevent them from receiving gifts, payments or other inducements to vacate the Court.[2] Justices are famously vocal and enthusiastic about their intentions to remain on the court as long as they are alive and wish to do so (e.g. Williams 1990).[3] After 1800, 23.3 percent of all years served on the Court have been served by justices already eligible to retire with pension benefits equal to their pay as working justices.

In addition to measurement problems, analyses of retirement effects are well-known for susceptibility to unrecognized or ignored endogeneity and consequent identification and estimation problems (see Snyder and Evans 2006; Handwerker 2007). Endogeneity concerns arise because voluntary retirement is, in the language of causal effects, a self-selected "treatment." Moreover, retirement and mortality both tend to occur late in life and share common causal antecedents such as age

[1] This current social science research usage is consistent with the common language definition of retirement as "*withdrawal* from one's occupation, business, or office" [emphasis added] (American Heritage 1996), although it differs from some other definitions. For example, the U.S. Current Population Survey (CPS) accepts, solely as an expedient, jobless respondents' description of their labor force status as "retired," if they are at least 50 years old; thus, CPS respondents who are coded as "retired" include persons who would be classified as disabled, unemployed, or otherwise if accurate information were available (Polivka and Rothgeb 1993:24). The social science research definition of retired persons differs sharply from actuarial and accounting definitions, which usually include only recipients of money payments from pension funds (Society of Actuaries 1992).
[2] Justices can be, but none have been, terminated from office for treason, bribery, or serious crimes.
[3] Justice Thurgood Marshall is reported to have stated for publication, "I have a lifetime appointment and I intend to serve it. I expect to die at 110, shot by a jealous husband" (Williams 1990).

and health. Further, it has been hypothesized that mortality and retirement propensities partially cause each other. If reciprocal effects exist empirically but are ignored analytically, they would produce endogeneity bias. Finally, if mortality risk is determined according to one causal regime or specification before retirement, but according to another regime or specification after retirement, then that situation would be described as endogenous switching, and it too would represent a form of endogeneity (Quandt 1972; Mare and Winship 1988).

A recent analysis seeks to overcome endogeneity problems by using Social Security policy change as an instrumental variable to identify retirement effects (Snyder and Evans 2006; however see Handwerker 2007). But that analysis uses an actuarial definition of retirement (*receipt of pension benefits*) well suited to pension fund financial analysis, but not suitable for the present purpose of understanding retirement decisions of individuals. Below, we show that Supreme Court data permits appropriate instrumentation of retirement effects on subsequent longevity, as well as distinguishing between retirement and involuntary unemployment.

Substantively, our focus on Supreme Court justices builds upon findings of occupational differences in mortality and retirement patterns (Guralnik 1962; Kitagawa and Hauser 1973; Fletcher 1983, 1988; Hayward and Hardy 1985; Hayward Grady Hardy and Sommers 1989; Norman, Sorlie and Backlund 1999). Further, our concentration on Supreme Court justices extends a body of mortality research and labor force exit studies of very small social groups that are defined by their members' high achievement, influence and power (e.g., Abel and Kruger 2005; Redelmeier and Singh 2001a, b; Gavrilov and Gavrilova 2001; Waterbor, Cole, Delzell and Jelkovich 1988; Treas 1977; McCann 1972; Quint and Cody 1970). From the policy perspective, our analyses contribute new information to ongoing debates about retirement in general (Gokhale 2004; Ashenfelter and Card 2002), and mandatory retirement and term limits for Supreme Court justices (Calabresi and Lindgren 2006).

Even in the unlikely event that Supreme Court retirement and mortality patterns are unique and unrelated to retirement and mortality in the general population, Supreme Court demography itself is a topic of perennial popular interest, periodic political significance, longstanding legal importance and general governmental consequence, both in the United States and elsewhere that American government exerts power (see e.g. Garrow 1998; Woodward and Armstrong 1979; Toobin 2007; *New York Times* 2007; *USA Today* 2007). Preston (1977) suggests the demographic importance of Supreme Court justice mortality patterns, but the topic has escaped previous demographic study. Separately, and without apparent awareness of relevant demographic research, a long, contentious, self-critical literature in law and political science examines pre-retirement deaths of Supreme Court justices (see the review and critique by Stolzenberg and Lindgren, 2009), but that research does not consider life after retirement. Finally, analyses presented here address a question asked informally, but seriously, of a U.S. Federal judge by Gary S. Becker: Is it economically irrational for justices to keep working after they become eligible to retire with pensions equal to their pay?[1] If continued life has sufficient value, and retirement accelerates death, then the rationality of unpaid work, even at little or no marginal financial gain, is apparent.

The next section reviews relevant previous findings and theory, and presents hypotheses. Section III considers methodological issues and data. Section IV presents results. And Section V discusses findings.

II. HYPOTHESES

This paper considers three competing hypotheses about the effects of retirement decisions on subsequent longevity:

A. The Increased Mortality Hypothesis. *Retirement increases subsequent mortality hazard (and therefore reduces ensuing longevity), on average and other things equal.*

In an early empirical result, McMahan Folger and Fotis (1956) find that military personnel live about two years in retirement for every three years served on active duty, suggesting that delayed retirement prolongs life after retirement. Waldron (2001, 2002) reports findings in several large U.S. national data sets that mortality hazard declines as retirement age rises, controlling for current age. More theoretically, the hypothesis of negative effects of retirement on longevity arises from the observation that, compared to nonparticipation in the labor force, most employment tends to raise the intensity of social, physical and mental activity. Increased physical activity reduces the incidence of "depression, fractures, coronary heart disease and mortality" (Wagner, LaCroix, Buchner and Larson 1992: 452; see Bortz 1984, who calls these effects "disuse syndrome"). Wagner et al. speculate that "Although most of the evidence available pertains to physical activity, inactivity in other aspects of life – intellectual, social, interpersonal" reduces physical health, mental health and longevity. Snyder and Evans (2006) find reduced mortality hazard among retired workers who return to work after their Social Security benefits are reduced; Snyder and Evans speculate that work at older ages prolongs life by reducing social isolation, and they cite evidence that social contact reduces mortality risk (Berkman and Syme, 1979; Blazer, 1982; House, Landis, and Umberson, 1988; Berkman, 1995, 2000; Cohen et al., 1997; Colantonio et al., 1993; Zuckerman, Kash, and Ostfeld, 1984; Putnam, 2000; Seeman et al. 1987). Others report that any activity, including work, is an antidote to the "powerful adverse effects on physical health and functional status" of depression (Wagner et al 1992:458; see also Camacho et al 1991 and Farmer et al 1988).

B. The Reduced Mortality Hypothesis. *Retirement reduces subsequent mortality hazard (and increases longevity)*, on average *and other things equal*.

Tsai (2005) writes, "There is a widespread perception that early retirement is associated with longer life expectancy and later retirement is associated with early death." In a competing risks model of Danish death rates, Munch and Svarer (2005:17) find that "early retirement prolongs survival for men." Mein Martikainen Hemingway Stansfeld and Marmot (2003) report that early retirement at age 60 was associated with an improvement in mental health, particularly among high socioeconomic status groups. Voluminous evidence that employment tends to expose workers to life-shortening health risks: The effect of work on mortality is so strong that occupations are characterized by the mortality patterns of their incumbents (Guralnik 1962; Kitagawa and Hauser 1973; Fletcher 1983, 1988; Norman, Sorlie and

Backlund 1999). For office workers like lawyers and judges, the most apparent work-related health and mortality risks are stress – a physiological reaction to the belief that resources are insufficient to accomplish tasks that are perceived to be necessary – and effort-reward imbalance – a belief that returns received from work are insufficient to justify the effort required to produce them - (House, Landis, and Umberson 1988; House, Kessler, et al. 1990; Marmot and Theorell 1988; Cohen and Syme 1985; Marmot and Wilkinson 1999; Peter Siegrist Hallqvist Reuterwall and Theorell 2002; Siegrist Peter Cremer and Seidel 1997; Peter et al., 1998; Siegrist et al., 1990). The Reduced Mortality Hypothesis reasons that retirement reduces or eliminates work-related exposure to these and other mortality accelerants.

C. The Null Effects Hypothesis. *There is no effect of retirement age on mortality hazard, on average and other things equal.*

This hypothesis asserts that any apparent association between retirement timing and subsequent mortality risk is spurious or the result of measurement or analysis problems. Although empirical research methods are poorly suited to testing hypotheses of no effect, this hypothesis has considerable precedent: After adding appropriate control variables to his analysis of retirement and mortality of Shell Oil employees, Tsai (2005) concludes that “early retirement at 55 or 60 is not associated with increased survival,” and “Employees who retired at 60 had similar survival to those who retired at 65.” Mein Martikainen Hemingway Stansfeld and Marmot (2003) report that early retirement at age 60 was not associated with any effect on physical health. After a careful discussion and examination of confounding variables and measurement issues, Litwin (2007) concludes that “respondents who had prematurely exited the [Israeli] labour force did not benefit from disproportionately longer lives when compared with the respondents who retired ‘on time.’”

In short, previous empirical studies and logic offer some support for hypotheses of negative, positive or null effects of retirement on mortality risk and longevity. However, confidence in previous findings is undermined by measurement problems, conceptual issues, or unrecognized endogeneity. The next section describes a strategy for circumventing those problems.

III. ANALYSIS STRATEGY, ESTIMATION AND DATA

A. Discrete Time Event History Models. In the language of causal inference, we seek to measure the effect of a time-related treatment (the timing of retirement) on time-related outcomes (mortality hazard and years-left-alive), for those who select the treatment. Accordingly, we use event history data and methods, with accommodations for self-selection, to test hypotheses and to construct models of retirement effects. We use discrete time methods with a one-year time period because Court terms and data are organized on an annual basis: Justices customarily resign at the end of the Court's annual term; the Court structures its activities into annual sessions; and Court pension-eligibility rules are based on completed years of service and whole years of age. Consequently, dates and times for Supreme Court careers tend to be rounded to whole years; multiple resignations in the same year tend to occur simultaneously; and relevant time-varying political circumstances tend to exist for whole years, rather than for shorter intervals. Date rounding, co-occurrence of events, and time-varying independent variables are easily accommodated by discrete time event history methods, but not as easily by continuous time methods (Yamaguchi 1991). Discrete time methods also accommodate right censoring, which occurs for the 44.5 percent of all justices who died in office without first resigning from the Court (Allison 1995). So, we test hypotheses with discrete time event history models in which the time unit is one year, the unit of analysis is the *justice-year*, and variables indicate retirement status, mortality, remaining years of life and other characteristics of a particular justice in a specific year.

Our analyses measure retirement effects on two outcome measures: annual mortality hazard and years-left-alive. *Annual mortality hazard* is the probability that a specific justice who is alive at the start of a particular year dies before the end of that year. *Years-left-alive* for each justice-year is the number of additional years after the current year until the relevant justice dies. To assure that estimates of mortality hazard are in the interval [0,1] for which probabilities are defined, we use maximum likelihood probit analysis to measure the effect of retirement on mortality hazard. To assure that estimates of years-left-alive are non-negative, some of our analyses use a probit-transformation of years-left-alive (described below) and other analyses use a Tobit analysis of years-left-alive (Amemiya 1985). All of these methods are well-known, but not commonly used together.

B. Endogeneity by Mediation. Above, we observe that the endogeneity of retirement can be represented in two different ways. In the first representation, shown in Figure 1, an endogenous variable, *retired*, mediates some of the effect of exogenous variables on the endogenous outcome variable, *mortality hazard* (or longevity). We call this representation *endogenous mediation*; it is the problem for which Instrumental Variables (***IV***) estimation is the standard solution (Amemiya 1985).

In endogenous mediation models, identification of the effect of an endogenous mediating variable on an endogenous outcome variable requires an instrumental variable, an exogenous variable that has direct effects on the endogenous mediator, but is restricted to have only indirect effects on the endogenous outcome. In Figure 1, we reason that *pension-eligibility* has the necessary direct effect on *retirement* because Supreme Court pensions are equal to the full pay of Supreme Court justices; pension eligibility removes the need to work for pay. To make the necessary identifying restriction in Figure 1, we observe that mere *eligibility* for a pension could have no mortality or longevity effect – one must actually receive the pension to spend it in ways that affect mortality and longevity. In passing, *pension-eligible* is empirically distinct from pension receipt: from 1801 to 2006, 23.3 percent of the justice-years served on the U.S. Supreme Court were served by justices who were eligible to retire from the bench with a Federal pension.

For additional consideration of the suitability of *pension-eligible* as an instrument for *retired* in Figure 1, we also estimate ***IV*** analyses on the subset of 57 justices who become *qualified-for-pension* while still incumbent. As for the full population of justices, we hypothesize (and find) that pension eligibility explains some variation in retirement decisions. But because all 57 of these pension-eligible justices would receive pensions upon retirement, there can be no variation in their eligibility for pension benefits at the time that those benefits could affect their mortality hazard and longevity.

Equations (***IV***1), (***IV***2a) and (***IV***2b) summarize the endogenous mediation model of retirement and mortality. Subscript j refers to the j^{th} justice of the Court. Subscript t refers to the t^{th} time period. $Retired_{jt}$ equals 1 if the j^{th} judge is retired at the start of the t^{th} time period; otherwise, $Retired_{jt}$ equals 0. f, g, and h are functions that can involve nonlinear and nonadditive transformations of variables. ε, ϖ and ζ are random disturbances. Variables age_{jt}, $calendar\ year_{jt}$, $pension\text{-}eligibility_{jt}$ and $death_{jt}$ and $tenure_{jt}$ are measures of the eponymous characteristics or events, measured in whole years, pertaining to the j^{th} individual during the t^{th} time period.

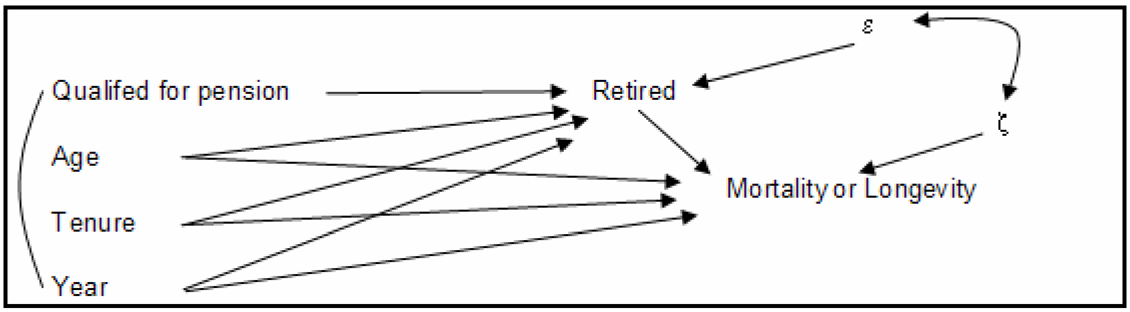

Figure 1 –Endogenous Mediation by Retirement of Mortality or Longevity
(Not a linear additive path model)

Endogenous Mediation Model of Mortality Hazard

(***IV***1) $\Pr[retired_{jt}] = f(age_{jt}, year_{jt}, tenure_{jt}, qualified\text{-}for\text{-}pension_{jt}, \varepsilon_{jt})$

(***IV***2a) $\Pr[death_{jt}] = g(age_{jt}, year_{jt}, tenure_{jt}, retired_{jt}, \zeta_{jt})$

Endogenous Mediation Model of Years-left-alive

(***IV***1) $\Pr[retired_{jt}] = f(age_{jt}, year_{jt}, tenure_{jt}, qualified\text{-}for\text{-}pension_{jt}, \varepsilon_{jt})$

(***IV***2b) $years\text{-}left\text{-}alive_{jt} = h(age_{jt}, year_{jt}, tenure_{jt}, retired_{jt}, \varpi_{jt})$

Notes: Subscript j refers to the j^{th} justice of the Court; subscript t refers to the t^{th} calendar year ; *f*, *g*, and *h* are functions and can include transformations of independent and dependent variables; ε_{jt}, ζ_{jt} and ϖ_{jt} are errors. Equations ***IV***1 and ***IV***2a are estimated by maximum likelihood instrumental variables probit analysis. Equation ***IV***2b is estimated by instrumental variables regression.

C. Endogeneity by Switching. The second representation of endogeneity discussed above is Endogenous Switching (***ES***). In ***ES*** models here, exogenous variables are the same as in the endogenous mediation model; retirement is endogenous; but retired and incumbent justices can experience separate causal regimes (parameter values) for mortality hazard or longevity. If a justice is retired, then mortality (or longevity) is unobserved in the equation for incumbents; if a justice is incumbent, then mortality (or longevity) is unobserved in the equation for retirees. Identification can be achieved via instrumental variables or, if none, nonlinearities.

Finally, for comparison to research that disregards endogeneity, we add analyses that do so also.

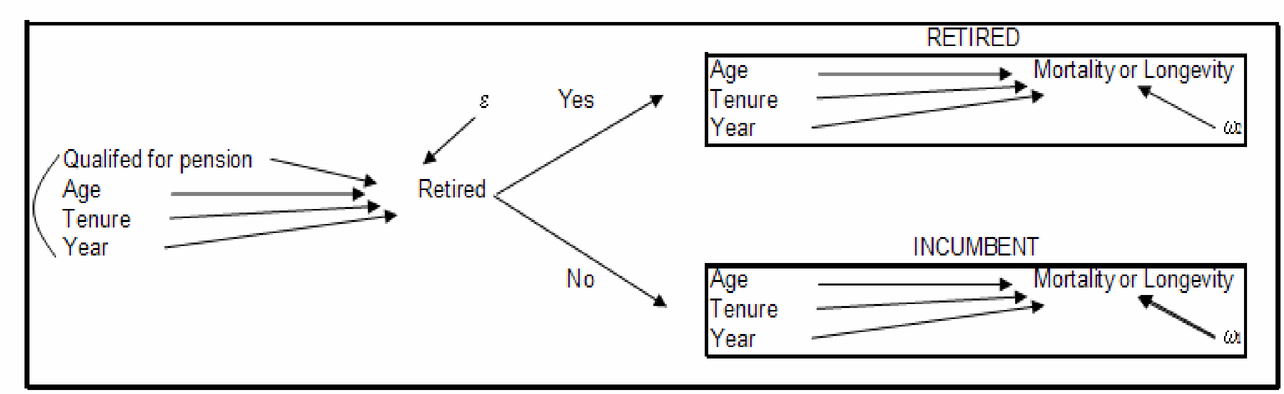

Figure 2 –Endogenous Switching by Retirement of Mortality or Longevity Processes
(Not a linear additive path model)

Retirement Selection Model (Probit)

(***ES***1) Pr[retired_{jt}] = $f(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{qualified-for-pension}_{jt}, \varepsilon_{jt})$

${}_1\delta_{jt} = \mu(E[\Pr[\text{Retired}_{jt}]])$

${}_2\delta_{jt} = \mu(E[\Pr[\text{Not Retired}_{jt}]]) = \mu(1 - E[\Pr[\text{Retired}_{jt}]])$

Conditional Model of Mortality Hazard (Probit)

(***ES***2a1) $\Pr[\text{death}_{jt} \mid \text{incumbent}] = \pi(\text{age}_{jt}, \text{year}_{jt}, {}_1\delta_{jt}, \zeta_{jt})$

(***ES***2a2) $\Pr[\text{death}_{jt} \mid \text{retired}] = \theta(\text{age}_{jt}, \text{year}_{jt}, {}_2\delta_{jt}, \zeta_{jt})$

Conditional Model of Years-left-alive (Regression)

(***ES***2b1) $\text{years-left-alive}_{jt} \mid \text{incumbent} = \eta(\text{age}_{jt}, \text{calendar year}_{jt}, {}_1\delta_{jt}, {}_1\varpi_{jt})$

(***ES***2b2) $\text{years-left-alive}_{jt} \mid \text{retired} = \gamma(\text{age}_{jt}, \text{calendar year}_{jt}, {}_2\delta_{jt}, {}_2\varpi_{jt})$

Notes: f, μ, η and γ are functions and can include transformations of independent and dependent variables; ε_{jt} and ϖ_{jt} are errors; E is the expectation operator. Equation ***ES***1, ***ES***2a1 and ***ES***2a2 are estimated by maximum likelihood selection-corrected probit analysis. Equation ***ES***2b1 and ***ES***2b2 are estimated by maximum likelihood. The ES estimator is also called the selection-corrected regression estimator.

D. Estimation and Tests. To constrain estimated hazards to the [0,1] interval for which they are defined, we use probit, ***IV*** probit, and selection corrected probit methods to estimate mortality hazard models. To constrain estimated years-left-alive to the non-negative values for which it is defined, we use Tobit, ***IV*** Tobit, regression with probit transformation of years-left-alive, and ***IV*** regression with a probit transformation of years-left-alive (Stolzenberg 2006: 56). The probit transformation is as follows: Where Y is years-left-alive, Ψ is the transformed value of Y, Φ is the Normal cumulative distribution function, and Φ^{-1} is the inverse Normal cumulative distribution function, $\Psi = \Phi^{-1}((Y+0.5)/50)$.

Transformation back to years is computed from the inverse ($\widehat{Y} = 50[\Phi(\widehat{\Psi})] - 0.5$), much as probability estimates in probit regression are computed from estimated probits. Table 1 summarizes this combination of estimation methods and mortality measures.

Although they are not the subject of this paper and serve here only as control variables, age, tenure and calendar year are well known to have nonlinear effects. These nonlinearities are variously described as compression of morbidity (Fries 2005), historical change, decreasing (or increasing) marginal effects, and, in failure-time analysis, the whimsically-named, ∪-shaped "bathtub distribution" (Hjorth 1980). Although well-known, none of these nonlinear patterns have been linked to a specific mathematical function. Because many mathematical functions virtually duplicate the same values over a fixed range, it is neither necessary nor possible to distinguish among the various functions that might produce the nonlinear effects observed in a specific data set. Rather, it is sufficient to use log-fractional polynomial transformations of these variables to parsimoniously *permit but not require* time variables to have nonlinear effects. Log-fractional polynomial transformations are a simple, mathematically well-behaved, and rich generalization of polynomial regression (Royston and Altman 1994; Gilmour and Trinca 2005).

We analyze data on the universe – not a sample – of Supreme Court justices of the United States from 1801 through 2006. Although there is strong reason to believe that sampling-based significance tests are inappropriate for population data, presentation of sample tests on population data are commonplace and many readers expect them. So we report standard errors, and significance tests for all equation parameters estimated here. Because data contain multiple observations per justice, each justice constitutes a cluster, and we calculate robust standard errors with first-order Taylor series linearization correction for clustering (Huber-White "sandwich" estimators; Binder 1983). For some analyses in which statistics of interest are population means of analysis forecasts or predictions, themselves based on nonlinear functions of model estimates, ordinary standard errors are not available, so we use bootstrapping methods to calculate them. Although McCullagh (2000) criticizes bootstrapping with clustered data, Feng Mclerran and Grizzle (1998) and Field and Welsh (2007), find that the method performs well, particularly when the number of clusters is 50 or more; data examined here have 110 clusters.

E. Data. We examine data on all justices of the United States Supreme Court from 1801 through 2006.[4] Data are an annual event-history data set consisting of one observation for each year in which each justice of the Court was alive, starting in the year in which the justice takes office on the Court, and ending in the year in which the justice dies. These are the data used in Stolzenberg and Lindgren (2008), with one additional justice-year observation for each year that each justice is alive after retiring from the Court. Table 2 provides descriptive statistics for these data. From establishment of the Supreme Court in 1789 until the end of 2006, 110 justices served a total of 1895 justice-years on the court, and lived 457 post-resignation justice-years.

In 2006, except for Justice Sandra Day O'Connor, all justices who had previously resigned from the court had died. Our statistical analyses of longevity are estimated over all 1971 justice-years after the year 1800, for 91 justices who died before the year 2007 and who either died in office, or who resigned from the court at the age of at least 55 years. (Years-left-alive is unobservable for justices still alive as this research is done; for Supreme Court justices, resignation before age 55 is presumed to indicate a change of jobs rather than a departure from the labor force altogether.) Our analyses of mortality hazard also include justice-years for justices who have not died as this research is done, for a total of 2132 justice-years. Variables are as follows:

1. ***Retired***. A dummy (0,1) variable equal to 0 for a justice-year unless the corresponding justice retired, resigned or accepted "senior status" during that year, or before starting service the next year.
2. ***Death***. A dummy (0,1) variable equal to 0 for a justice-year unless the justice died that year.

[4] We started with database kindly supplied to Professor James Lindgren by Professor Albert Yoon (see Yoon 2006), based on information obtained from the Administrative Office of the U.S. Courts (Federal Judicial Center 2006). Lindgren checked some of those data against various sources including the *Congressional Record*, corrected errors and added more data from the Federal Judicial Center (2006), and the U.S. Supreme Court (2006) for 1789-1868 and 2003-2006. We added post-retirement data for justices who did not die in office.

3. ***Year, Year1788*, *ln(Year1788)*.** *Year* is the calendar year. *Year1788* is *Year* - 1788.[5] *ln(Year1788)* is the natural logarithm of *Year1788*; the logarithmic transformation improves the fit of some models. We include calendar year to hold constant mortality and retirement trends.

4. ***Age, Age^2, Age^3***. *Age* is the age of the justice in years at the start of the justice-year. Probabilities of death and retirement increase with age. In some analyses, we add Age^2 and Age^3 to the analysis, to fit nonlinear age effects.

5. ***Tenure, $Tenure^3$, $Tenure^3$ x ln(Tenure)***. *Tenure* is years of service on the Court. The annual probability of job quitting in the working population is known to first decline as tenure increases, and then increase with additional tenure (Stolzenberg 1989). $Tenure^3$ and $Tenure^3$ *x ln(Tenure)* prove useful transformations of tenure with fractional polynomial models.

6. ***Qualified-for-pension.*** A dummy variable equal to 0 unless the justice is eligible for a Federal judicial pension.

7. ***Years-left-alive, $Years\text{-}left\text{-}alive^2$*** . In each justice-year, *Years-left-alive* indicates future longevity, or remaining years of life. *Years-left-alive* for each justice-year is the difference between the calendar year of the justice-year and the calendar year in which the justice ultimately dies. $Years\text{-}left\text{-}alive^2$ proves to be a useful transformation of future longevity.

IV. RESULTS

A. Retirement Analyses. Because all of our ***ES*** and ***IV*** analyses of mortality and longevity require a regression or probit analyses of retirement timing, we report those retirement analyses first, in Table 3. Independent variables in these retirement models are *qualified-for-pension*, and, to fit expected nonlinear temporal effects, polynomials of *age*, *year1788*, and *tenure*. Because qualified-for-pension serves as an identifying instrument for retirement, a key result in Table 3 is the expected positive, statistically significant ($\alpha \leq .05$, 1-tailed robust cluster-corrected test) coefficient of qualified-for-pension.

---

[5] Subtracting 1788 from calendar year preserves all information and avoids rounding problems that occurred in initial analyses with STATA version 8 that used calendar year.

Although the probit analysis does (and the regression does not) constrain probability estimates to [0,1], probability estimates from these two models are similar, with a Pearson correlation of 0.8411.

B. Years-left alive-analyses. Table 1 defines nine different models for estimating the effect of retirement on years-left-alive, and the upper panel of Table 4 presents empirical estimates of these effects. See Appendix 2 for details.

B1. Analyses that Ignore Endogeneity. Models 1a-1d ignore endogeneity, but are presented for comparison to IV and ES analyses. Consistent with the Increased Mortality Hypothesis, retirement effects in Models 1a-1d all indicate negative effects of retirement on future longevity, all are statistically significant ($\alpha \leq .05$, 1-tailed robust cluster-corrected test), and all are based on analyses that hold constant the effects of age, tenure and calendar year.

- **In Model 1a**, ordinary regression estimates an average of *3.6 years less remaining life* (the coefficient of *retired*) for those who are retired than for those who are not retired ($\alpha \leq .01$, one-tailed robust cluster-corrected test).
- **Model 1b** applies a probit transformation to years-left-alive, yielding a coefficient of -.2847 ($\alpha \leq .01$, one-tailed robust cluster-corrected test). To express that coefficient in intuitively meaningful terms, we evaluate its effect (in years) at 11 years-left-alive (the median of years-left-alive). At 11 years left-alive, the *retirement effect is 3.74 years less remaining life, on average*.
- **In Model 1c**, we estimate separate models of probit-transformed years-left-alive for retired and incumbent justices. For each regression, the regression prediction of probit-transformed years-left-alive is computed for each justice-year, the predictions from each equation are re-transformed into years, and the predicted years left alive if incumbent is subtracted from the predicted years left alive if retired. The mean difference between years left alive if retired and years left alive if incumbent is *6.596452 years less life for the retired* than for incumbents (significant, $\alpha \leq .01$, 1-tailed test, based on clustered bootstrap standard error, with 1391 replications).

- **Model 1d** is the Tobit regression of *years-left-alive* on *age, age^2, age^3, year1788, $year1788^2$, tenure, $tenure^2$* and *retired*. Model 1d resembles Model 1b, but uses Tobit analysis rather than probit transformation to assure that predicted longevity is never negative. The significant ($\alpha \leq .01$, one-tailed robust cluster-corrected test) coefficient of 3.3338 for *retired* in Model 1d indicates an average of *three-and-a-third fewer years-left-alive for the retired* than for incumbents.

*B2. **IV** Regression and **IV** Tobit Analyses*. In Models 3a-3c, instrumental variables estimation is used to accommodate the endogeneity of retirement. Again, all of these analyses hold constant the effects of age, tenure and calendar year, and all indicate a negative impact of retirement on future longevity, consistent with the Increased Mortality Hypothesis.

- **Model 3a** estimates a coefficient of -13.35622 for retired (significant, $\alpha \leq .05$, one-tailed robust cluster-corrected test), indicating an average of *13.4 years less remaining life* for those who are retired than for those who are not retired. Although this coefficient is almost 4 times the size of the coefficient obtained in the corresponding un-instrumented analysis 1a, a 95 percent confidence band around the model 3a estimate overlaps the model 1a estimate, indicating that the IV and OLS model estimates cannot be distinguished from each other.
- **Model 3b** applies a probit transformation to years-left-alive, as well as ***IV*** estimation, yielding a *coefficient* of -1.0366 (significant, $\alpha \leq .05$, one-tailed robust cluster-corrected test). This coefficient is 3.6 times the size found in the corresponding OLS Model 1b, but a 95 percent confidence around the model 3b estimate overlaps the Model 1b estimate, indicating that the IV and OLS model estimates cannot be distinguished from each other. If evaluated at 11 *years-left-alive* (the median of *years-left-alive*), the *retirement effect in Model 3b is 9.24 fewer years of remaining life*. The curved, unbroken line Figure 3 plots other values of those retirement effects in years.[2]
- **Model 3c** combines ***IV*** estimation to accommodate the endogeneity of retirement with Tobit analysis to accommodate the restriction of years-left-alive to nonnegative values. Model 3c finds a

mean retirement effect of 13.58 *fewer years-left-alive for the retired* than for incumbents. Figure 3 plots those retirement effects in the patterned line, as estimated by the Tobit model.

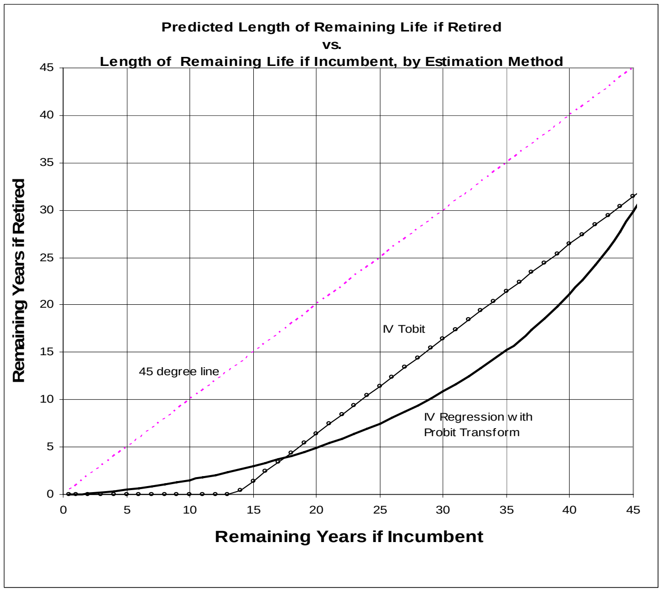

Figure 3 – Predicted Length of Remaining Life If Retired vs. Predicted Length of Remaining Life if Incumbent, for IV Tobit and IV Regression with Probit Transformation

<u>*B3. Endogenous Switching Analyses*</u>. Models 5a and 5b use endogenous switching regression to model the endogeneity of retirement effects on *years-left-alive.* Both analyses find statistically significant negative effects of retirement on years-left-alive (all are statistically significant ($\alpha \leq .05$, one-tailed robust cluster-corrected test).

- In **Model 5a**, years-left-alive is measured in its natural metric, and separate equations , corrected for endogenous selection bias, are estimated for the effects of *age, year1788* and *tenure* on *years-left-alive.* Each equation is used to predict the years-left-alive for each justice in each justice-year if retired and, separately, if incumbent. The mean difference between these estimates is *5.7903 fewer years of remaining life for the retired* than for incumbents. Significance testing is accomplished by clustered bootstrapping, with 1391 replications; this result is statistically significant ($\alpha \leq .01$, 1-tailed test).
- **Model 5b** follows the same procedure as Model 5a, except that the probit transformation is applied to *years-left-alive* before the analysis, and switching regression estimates are transformed back from

the probit transformation metric to years before calculating the difference in remaining life for each justice. The mean of that difference is *6.8810 fewer years of remaining life for the retired* than for the incumbent, after holding constant *age*, *tenure*, and *year1788* (significant, $\alpha \leq .01$, 1-tailed test, with clustered bootstrap standard error, with 1391 replications). In each and every justice-year, predicted years-left-alive-if-incumbent exceeds predicted years-left-alive-if- retired.

Figure 4 shows the ratio of predicted years-left-alive-if-incumbent to predicted years-left-alive-if-retired, by half-century time period and age. Points in Figure 4 are ratio values for each justice-year. Curves are fitted by fractional polynomial regression for the indicated period. In all periods shown in Figure 4, the ratio is largest at age 55, declines and then rises again. That pattern is most pronounced in the period 1951-2006, in which the ratio reaches a minimum of 1.76 at 74years of age and then rises to more than 2 at 85 years of age. In short, the longevity advantage of incumbency is greatest at the youngest ages, although it rises again at very old ages.

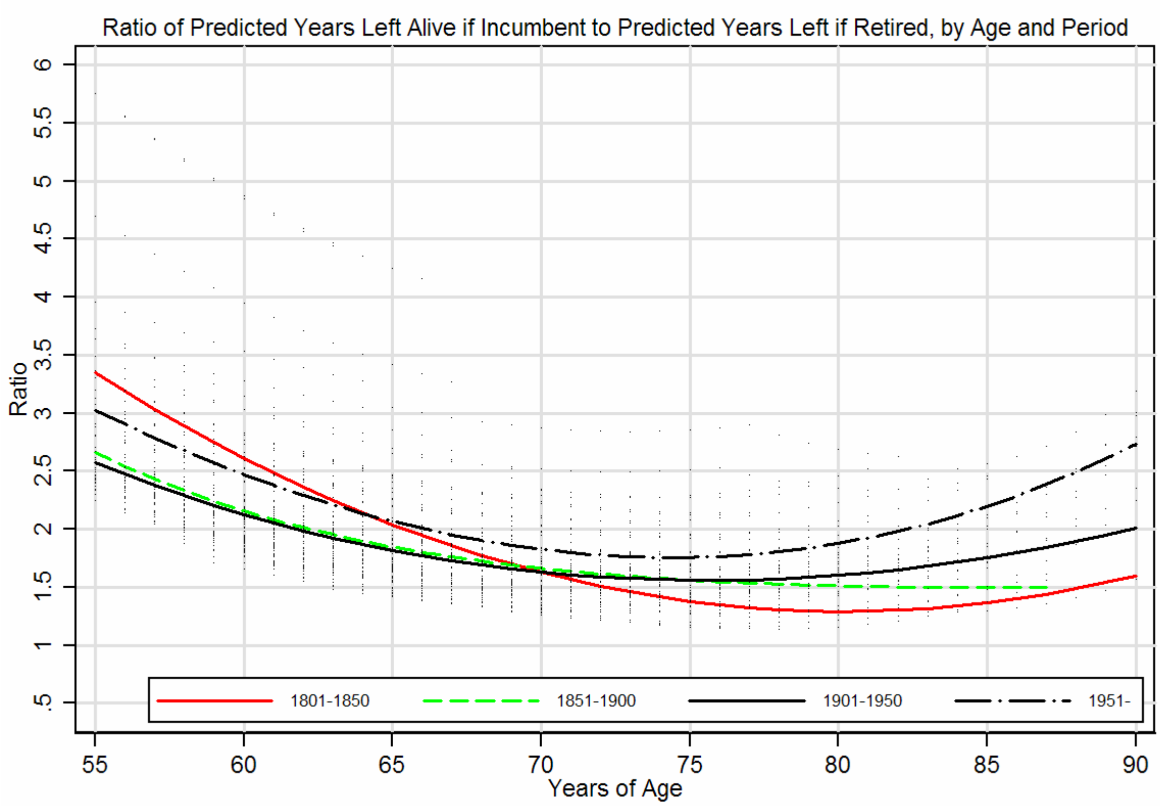

Figure 4 -- Ratio of expected-years-left-alive-if-incumbent to expected-years-left-alive-if-retired vs Calendar Year, by Age

C. Mortality Hazard Analyses. Table 1 defines three probit models for estimating the effect of retirement on annual mortality hazard. Empirical estimates of those effects are shown in the lower panel of Table 4. See Appendix 2 for analysis details. Results of all analyses are consistent with the Increased Mortality Hypothesis.

C1. Analysis that Ignores Endogeneity. Model 2 is the ordinary probit regression of mortality hazard on *year1788, age,* age^2*, tenure* and *retired*. The coefficient of *retired* is .4962814 (significant, $\alpha \leq .01$, 1-tailed robust cluster-corrected test). The solid line in Figure 5 graphs the effect of this coefficient on mortality hazards. The distance from the solid line to the lower dashed "equal values" line shows the estimated effect of retirement on mortality hazard: In the metric of probabilities (but not in the metric of probits), the average retirement effect increases as the mortality hazard for incumbent justices increases. According to Model 2, on average, an incumbent justice with an annual mortality hazard of 5 percent would face a hazard *2.5 times higher*, or 12.5 percent, if retired..

C2. Instrumental Variables Analysis. Model 4 is an IV probit analysis of retirement effects on mortality hazard. The coefficient of retirement in Model 4 is .7538361 – significantly different from zero ($\alpha \leq .05$, one-tailed robust cluster-corrected test), larger by half than the coefficient in Model 2, but not significantly different from the Model 2 estimate ($\alpha \leq .05$, two-tailed robust cluster-corrected tests). Based on the Model 4 coefficient, the upper broken line in Figure 5 shows the ***IV*** probit (Model 4) estimate of the retirement effect on mortality hazard. On average, an incumbent justice with an annual mortality hazard of 5 percent would face a hazard of 18.6 percent if retired, according to Model 4.

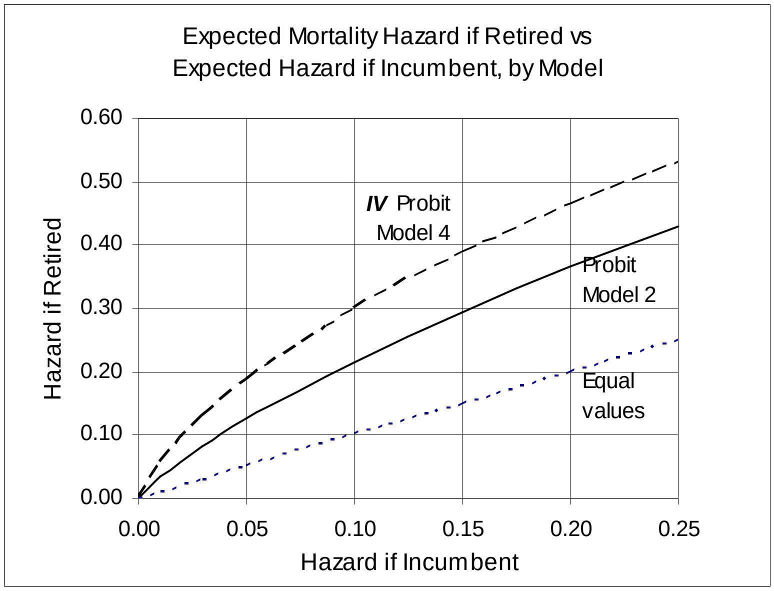

Figure 5 -- Expected Mortality Hazard after Retirement vs Expected Hazard before Retirement, by Model, with Plotted Equal Values Line

C3. Endogenous Switching Analysis. In Model 6, retirement effects on mortality hazard are measured with separate selection-corrected probit analyses for retired and incumbent justices. Unlike our IV analyses, switching analyses estimate all independent variable coefficients separately for retired and incumbent justices. For each justice-year in Model 6, we use observed values of independent variables and estimated parameters from the "retired" equation to calculate the expected mortality hazard if the relevant justice was retired in the corresponding justice-year. Separately, we use the same observed values of independent variables with estimated parameters from the "incumbent" equation to calculate the expected mortality hazard if the relevant justice was incumbent on the Court in the corresponding justice-year. If incumbency occurred in all justice-years, then the mean expected annual hazard would be .0433389. *If retirement prevailed in all justice-years, then the mean expected annual hazard would be about one-third (1/3≈31 percent=(.0567353*-.0433389*)/* .0433389*) higher, .0567353, all else equal.*[6] For parsimony, we calculate the ratio of *Expected-mortality-hazard-if-retired* to *Expected-mortality-hazard-if-incumbent*. In Figure 6, that ratio is scatter plotted vs. age, and lines show

[6] Based on justice-years for which the justice's age is at least 55 years. If all ages are included, then the mean hazard if retired is .0516245, the mean if incumbent is .0380390, a slightly larger proportional increase than for those who are at least 55 years old.

fractional polynomials of age fitted to the data, separately for four half-century historical periods.[7] In all periods, the ratio declines with increasing age until about age 70, and then increases. Up until about 1950, fitted lines indicate average ratios of less than one for justices in their 60's and 70's. 40.4 percent of the plotted points in Figure 6 indicate a ratio below 1. *However, after 1900, the fitted line is always above 1.0. And, in a result not visible from Figure 6, after 1955 there are no individual justice-years whatsoever for which estimated mortality hazard is lower if retired than if incumbent.*

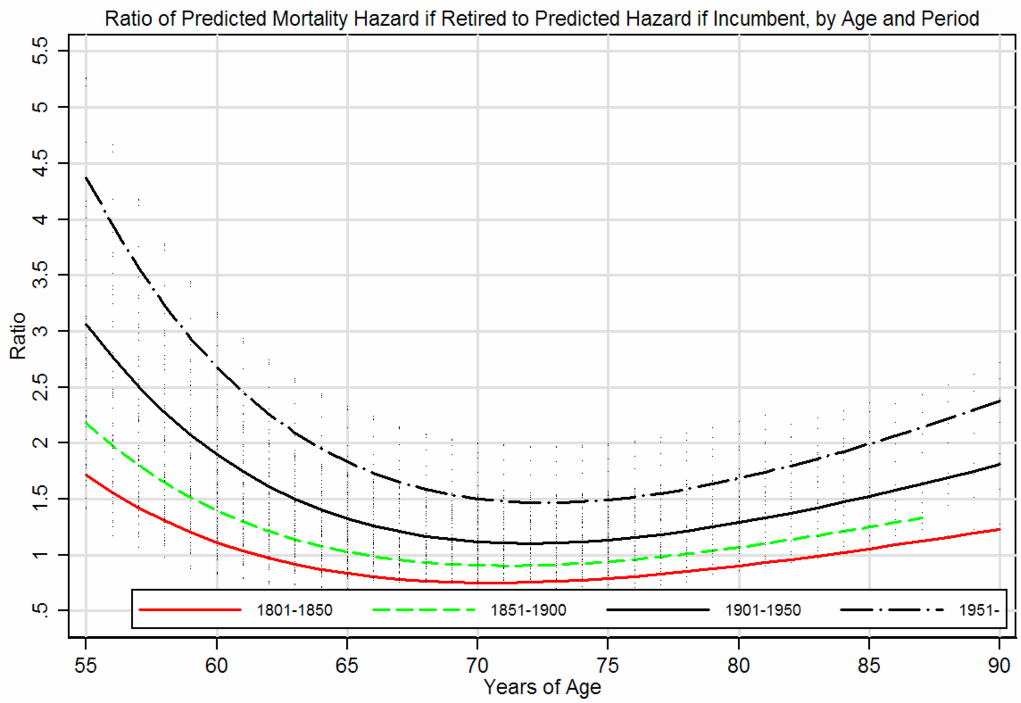

Figure 6 – Ratio of Predicted Mortality Hazard if Retired to Predicted Hazard if Incumbent, by Age and Period

V. DISCUSSION AND CONCLUSIONS

This paper considers the hypothesis that labor force retirement accelerates death, against the alternative hypothesis that retirement retards mortality, or does not affect it at all, other things equal. Following the venerable strategy of using data on unusual subgroups to avoid methodological problems in general population data, we use U.S. Supreme Court justice data to avoid confounding voluntary retirement with involuntary unemployment, job loss, employer-forced retirement or similar: The U.S.

---

[7] Results of the regression of Ratio of Predicted-Mortality Hazard if Retired to Predicted Hazard if Incumbent is as follows, with robust, cluster-corrected t-statistics below coefficients:

Ratio =	14.43587	-.3823013 age	+.0026407 age^2	+.0000193 $year1788^2$	-3.409846 (1/year1788)	R^2=0.8822
	(27.90)	(-27.09)	(28.05)	(15.80)	(-2.94)	N = 1711

Constitution gives Supreme Court justices unmatched freedom to keep their jobs as long as they like, for as long as they live. And justices do exercise that freedom: 23.3 percent of all justice years served after 1800 have been served by justices who had already qualified to retire with full pension benefits. In addition, Supreme Court data provide natural separation of retirement effects from the consequences of personal finances; Supreme Court pensions equal Supreme Court salaries, providing retiring justices with a seamless financial transition from employment to retirement.

We consider two forms of endogeneity of retirement effects: endogenous mediation, and endogenous switching. We use ***IV*** estimation to accommodate endogenous mediation, and, separately, ***ES*** estimation to accommodate switching endogeneity. For comparison, we also estimate models that ignore endogeneity. Because probabilities are defined only for the interval [0,1] and future longevity is defined only for non-negative lengths of time, we estimate mortality hazards with probit regression methods, and we estimate future longevity with probit transformations of years-left-alive, and, separately, with Tobit regressions. Permutation of these models, methods and dependent variables provides 12 different tests of the hypothesis that labor force retirement accelerates death. Findings from all 12 tests are statistically significant ($\alpha \leq .05$, one-tailed, robust standard error corrected for clustering) and consistent with the hypothesis that retirement accelerates death. In particular,

- The smallest point estimate of the average effect of retirement on longevity is an average loss of 3.3 years of life; the largest point estimate is an average loss of 13.6 years. For comparison, those estimates range from about one-sixth to three-quarters of the current remaining life expectancy of 65-year-old Americans, 18.7 years (U.S. Census Bureau 2008: Table 101).
- Endogenous switching analyses estimate an average mortality hazard of 4.33 percent for justices if incumbent, and about one-third higher (5.67 percent) if retired. To compare these hazards to estimated years-left-alive, we note that the ***ES*** hazard analysis implies that if mortality hazards were

constant, then justices would to live 5.46 years longer if incumbent than if retired, on average.[8] This difference is roughly the same as the difference found in the ***ES*** analyses of years-left-alive.

- In endogenous mediation models *without correction for endogeneity bias*, we find that if an incumbent justice had a mortality hazard of 5 percent in a particular year, that justice would have an estimated hazard of more than 12 percent if instead he or she were retired in that year, other things equal. In models *with **IV** correction for endogeneity bias*, if a justice had a 5 percent mortality hazard if incumbents, then retirement would raise that hazard to more than 18 percent. For comparison, we consider mortality effects of cigarette smoking. A recent study finds that smoking two or more packs of cigarettes a day (compared to never smoking) would raise nonsmokers with a 5 percent mortality hazard to a 15.8 percent hazard.[9] That smoking effect is about midway between our instrumented and uninstrumented probit estimates of the effect of retirement on mortality hazard. So, even the smallest of the ***IV*** point estimates of hazard effects can be characterized as comparable to the mortality hazard effects of heavy smoking. Effects of this magnitude seem to need no further interpretation.

In short, whatever model, estimation method or mortality measure we use, results are consistent with the hypothesis that voluntary retirement substantially accelerates death. Justices of the Court do not constitute a representative sample of the U.S. labor force as a whole. But Supreme Court data eliminates insoluble and endemic measurement problems in general population data, and Court data permits ***IV*** estimation of endogenous mediation models. Further, we note that much of what is known about the health and longevity effects of work and employment has been discovered or tested on population subgroups no less unusual than members of the U.S. Supreme Court. These groups include civil servants in England (Stansfeld et al 1995), residents of Alameda County, California (Camacho et al 1991), and

---

[8] Based on the geometric distribution. If x is geometrically distributed with annual mortality probability of p, then the expected years until death is 1/p. 5.458 = 1/0.0433 - 1/0.0567

[9] Computed from Rogers Hummer Krueger and Pampel (2005: 272), who report that the largest estimated logistic regression coefficient for a dummy variable for smoking two or more packages of cigarettes a day, compared to never having smoked, is 1.274.

the Wisconsin high school graduating class of 1957 (Marks and Shinberg 1997). And, finally, as Preston (1977) aptly notes, the combination of great power vested in Supreme Court justices, and the antiquity of so many of the justices themselves, makes their demography in general and their mortality in particular subjects of considerable interest in their own right.

REFERENCES

Amemiya, T. 1985. Advanced econometrics. Cambridge, Mass.: Harvard University Press.

American Heritage. 1996. The American Heritage Dictionary of the English Language, Third Edition 1996 by Houghton Mifflin Company

Anderson, Kathryn H. 1985 The effect of mandatory retirement on mortality Journal of Economics and Business 37: 81-88

Ashenfelter, Orley and David Card 2002 Did the Elimination of Mandatory Retirement Affect Faculty Retirement? The American Economic Review, Vol. 92, No. 4 (Sep., 2002), pp. 957-980

Berkman, Lisa F. 2000 "Social Integration, Social Networks, Social Support, and Health," in Lisa F. Berkman and Ichiro Kawachi (Eds.), *Social Epidemiology* (New York: Oxford University Press, 2000).

Berkman, Lisa F., and S. Leonard Syme, "Social Networks, Host Resistance, and Mortality: A Nine-Year Follow-up Study of Alameda County Residents," *American Journal of Epidemiology* 109 (June 1979), 186–204.

Binder, David A. 1983. On the Variances of Asymptotically Normal Estimators from Complex Surveys. International Statistical Review 51: 279-292.

Blazer, Dan G. "Social Support and Mortality in an Elderly Community Population," *American Journal of Epidemiology* 115 (May 1982), 684–694

Bortz, W. M. II. 1984. The Disuse Syndrome. *West. J. Med.* 141(5):691-94

Buchner DM, Beresford SAA, Larson EB, LaCroix AZ, Wagner EH. Effects of physical activity on health status in older adults II: Intervention Studies. *Annual Review of Public Health.* 1992. 13:469-488.

Calabresi, Steven G., and Lindgren, James. 2006a. "Term Limits for the Supreme Court: Life Tenure Reconsidered." *Harvard Journal of Law and Public Policy* 29: 770-877.

Camacho, T. C., Roberts, R. E., Lazarus, N. B., Kaplan, G. A., Cohen, R. D. 1991. Physical activity and depression: Evidence from the Alameda County Study. *Am. J. Epidemiol.* 134(2):220-31

Cohen, Sheldon, W. J. Doyle, D. P. Skoner, B. S. Rabin, and J. M. Gwaltney, "Social Ties and Susceptibility to the Common Cold." *Journal of the American Medical Association* 277 (June 1997),1940–1944.

Colantonio, Angela, Stanislav V. Kasl, Adrian M. Ostfeld, and Lisa F. Berkman, "Psychosocial Predictors of Stroke Outcome in an Elderly Population," *Journal of Gerontology* 48:5 (1993), 261–268.

Crimmins, Eileen M., Mark D. Hayward, and Yasuhiko Saito, "Changing mortality and morbidity rates and the health status and life expectancy of the older population." Demography 31 (1994) pp. 119-175.

Farmer, M. E., Locke, B. Z., Moscicki, E. K., Dannenberg, A. L., Larson, D. B., et al. 1988. Physical activity and depressive symptoms: The NHANES I epidemiologic follow-up study. *Am. J. Epidemiol.* 128:1340--51

Feng, Ziding, Dale Mclerran and James Grizzle 1998 "A Comparison of Statistical Methods for Clustered Data Analysis with Gaussian Error" *Statistics In Medicine* Volume 15 Issue 16, Pages 1793 – 1806 Published Online: 4 Dec 1998

Field, C.A. and A. H. Welsh 2007 "Bootstrapping clustered data" *Journal of the Royal Statistical Society: Series B (Statistical Methodology)* Volume 69 Issue 3, Pages 369 – 390 Published Online: 22 May 2007

Fries, James F. 2005 "The Compression of Morbidity" *The Milbank Quarterly* 83 (4) , 801–823

Gerdtham UG, M Johannesson 2003 "A note on the effect of unemployment on mortality" J Health Econ. 2003 May;22(3):505-18

Gokhale, Jagadeesh 2004 "Mandatory Retirement Age Rules: Is It Time To Re-evaluate?" Testimony before the U.S. Senate Special Committee on Aging Washington D.C. 20510 September 9, 2004

Gustman, Alan L. and Thomas L. Steinmeier 2001. “Retirement Outcomes in the Health and Retirement Study” National Bureau for Economic Research. Research Paper No. 7588. Revised, September, 2001.

Handwerker, Elizabeth W. 2007. "Nixing the Notch: The Effects of the Social Security Notch on Overall Incomes, Labor Supply, and Mortality in Retirement." Working paper. Washington, D.C.: U.S. Bureau of Labor Statistics.

Hjorth, Urban 1980 “A Reliability Distribution with Increasing, Decreasing, Constant and Bathtub-Shaped Failure Rates” Technometrics, 22: 99-107

House, James S., Karl R. Landis, and Debra Umberson, “Social Relationships and Health,” *Science* 241 (July 1988), 540–545.

Korpelainen, Helena. 2000 “Fitness, Reproduction and Longevity among European Aristocratic and Rural Finnish Families in the 1700s and 1800s.” Proceedings of the Royal Society: Biological Sciences. Vol. 267, No. 1454, pp. 1765-1770

Linn, M., R. Sandifer, and S. Stein 1985 “Effects of unemployment on mental and physical health”. Am J Public Health. 75(5): 502–506

Litwin, Howard 2007 “Does early retirement lead to longer life ?”Ageing & Society 27, 2007, 739–754.

Manton, K. G, et. Al., “Time varying covariates of human mortality and aging: Multidimensional generalizations of the Gompertz.” Journal of Gerontology: Biological Sciences 49:B169-B190.

Manton, K. G. and Stallard, E., “Heterogeneity and its Effect on Mortality Measurement” Chapter 12 in Methodologies for the Collection and Analysis of Mortality Data. International Union for the Scientific Study of Population, Vallin J, Pollard JH, Heligman L, Eds. 1984. Ordina Editions http://www.cds.duke.edu/publications/DocLib/cds_1080.pdf

Mare, Robert D. and Christopher Winship. 1988. "Endogenous Switching Regression Models for the Causes and Effects of Discrete Variables." Pp. 132-160 in Common Problems/Proper Solutions: Avoiding Error in Quantitative Research, edited by J. S. Long. Newbury Park, CA: Sage.

Marks N.F. and D.S. Shinberg 1997 Socioeconomic differences in hysterectomy: the Wisconsin Longitudinal Study. American Journal of Public Health. September; 87(9): 1507–1514.

Marmot, Michael and Richard G. Wilkinson (eds.) 1999. *Social Determinants of Health*. Oxford: Oxford University Press.

McCullagh, Peter 2000 “Resampling and Exchangeable Arrays” Bernoulli, Vol. 6, No. 2 (Apr., 2000), pp. 285-301

McMahan, C. A.; John K. Folger; Stephen W. Fotis 1956 Retirement and Length of Life *Social Forces*, 34: 234-238.

Mein G, P Martikainen, H Hemingway, S Stansfeld, M Marmot 2003 “Is retirement good or bad for mental and physical health functioning? Whitehall II longitudinal study of civil servants” J Epidemiol Community Health 2003;57:46–49

Munch, J. R. and Svarer, M. 2005. Mortality and socio-economic differences in Denmark: a competing risks proportional hazard model. Economics and Human Biology, 3, 1, 17–32.

Olshansky, S. Jay, Bruce A. Carnes, Jacob Brody 2002 “A Biodemographic Interpretation of Life Span” Population and Development Review, Vol. 28, No. 3, (Sep., 2002), pp. 501-513

Peter R ; Siegrist J ; Hallqvist J ; Reuterwall C ; Theorell T .2002 . Psychosocial work environment and myocardial infarction: improving risk estimation by combining two complementary job stress models in the SHEEP Study. J Epidemiol Community Health (Journal of epidemiology and community health.) 2002 Apr; 56(4): 294-300

Polivka, Anna and Jennifer Rothgeb 1993 “Redesigning the CPS Questionnaire” Monthly Labor Review 116: 10-28.

Putnam, Robert D., *Bowling Alone: The Collapse and Revival of American Community* (New York: Simon & Schuster, 2000).

Quandt R. 1972. “A new approach to estimating switching regression” . Journal of the American Statistical Association 1972; 67: 306-310.

Rogers, Richard G., Robert A. Hummer, Patrick M. Krueger and Fred C. Pampel 2005 "Mortality Attributable to Cigarette Smoking in the United States" *Population and Development Review* 31, No. 2 (Jun., 2005), pp. 259-292

Rones, Philip 1985 "Using the CPS to track retirement trends among older men." Monthly Labor Review 108: 46-49.

Seeman, Teresa E., George Kaplan, Lisa Knudson, Richard Cohen, and Jack Guralnik, "Social Network Ties and Mortality among the Elderly in the Alameda County Study," *American Journal of Epidemiology* 126 (1987), 714–723.

Siegrist, J. R. Peter, P. Cremer, D. Seidel. 1997. "Chronic work stress is associated with atherogenic lipids and elevated fibrinogen in middle-aged men." Journal of Internal Medicine. 242: 149-56.

Society of Actuaries. 1992. "Transactions of Society of Actuaries: 1991-92 Reports: Report of the Retirement Plans Experience Committee: Mortality among members of uninsured pension systems"

Sorokin, Pitirim 1925 "Monarchs and Rulers: A Comparative Statistical Study" *Social Forces* Vol. 4, No. 1 (Sep., 1925), pp. 22-35

Stansfeld, Stephen, Amanda Feeney, Jenny Head, Robert Canner, Fiona North, Michael Marmot 1995 "Sickness Absence for Psychiatric Illness: The Whitehall II" Social Science and Medicine 40 (2, Jan), 189-197

Stolzenberg Ross M. 2006 Multiple Regression Analysis. Chicago: The University of Chicago

Tsai, Shan, Judy Wendt, Robin Donnelly, Geert de Jong, and Farah Ahmed. 2005. "Age at retirement and long term survival of an industrial population: prospective cohort study" British Medical Journal October 29; 331(7523): 995

U.S. Census Bureau, *Statistical Abstract of the United States: 2008* (127th Edition) Washington, DC, 2007

Wagner EH, LaCroix AZ, Buchner DM, Larson EB. Effects of physical activity on health status in older adults I: observational studies. *Annual Review of Public health.* 1992:(13):145-68.

Waldron, Hilary 2002 "Do Early Retirees Die Early? Evidence from Three Independent Data Sets" ORES Working Paper Series, No. 97, Social Security Administration, Office of Policy, Office of Research, Evaluation, and Statistics.

Waldron, Hilary. 2001. *Links Between Early Retirement and Mortality*. ORES Working Paper Series, No. 93, Social Security Administration, Office of Policy, Office of Research, Evaluation, and Statistics. August.

Williams, Juan 1990 "The Thurgood Marshall Nobody Knows" *Ebony* May 1, 1990.

Zuckerman, Diana M., Stanislav V. Kasl, and Adrian M. Ostfeld, "Psychosocial Predictors of Mortality among the Elderly Poor: The Role of Religion, Well-Being and Social Contacts," *American Journal of Epidemiology* 119 (March 1984), 410–123.

Table 1 – Twelve Analyses of Retirement Effects on Mortality Hazard and Years-left-alive

Dependent Variable	**Identification and Estimation Methods**		
	Endogeneity Ignored	*Instrumental Variables*	*Endogenous Switching*
Years-left-alive	**Model 1** **1a** Regression **1b** Regression* **1c** ANCOVA** **1d** Tobit	**Model 3** **3a** ***IV***Regression **3b** ***IV***Regression* **3c** ***IV***Tobit	**Model 5** ***5a ES*** Regression 5***b ES*** Regression*
Mortality Hazard	**Model 2** Probit	**Model 4** ***IV*** Probit	**Model 6** ***ES*** Probit

Note: *To avoid negative estimates of years-left-alive, regressions are estimated with the following probit transformation of years-left-alive: Where Y is years-left-alive, Ψ is the transformed value of Y, Φ is the Normal cumulative distribution function, and Φ^{-1} is the inverse Normal cumulative distribution function, $\Psi = \Phi^{-1}((Y+0.5)/50)$. For comparison purposes only, those regressions are also estimated with no transformation of remaining length of life. **Analysis of Covariance with probit transformation of years-left-alive; these are separate analyses for retired and incumbent justice years.

Table 2 Descriptive Statistics for Discrete Time Event History Data, 1801-2006

Variable	N Justice-Years	Mean	Standard Deviation	Mini-mum	Maxi-mum
Year	2252	1911.467	57.66898	1801	2006
Age	2252	65.2873	10.55461	33	96
Tenure	2252	12.26643	8.387604	1	37
Qualified-for-pension	2252	.2801954	.4491943	0	1
Qualified-for-pension if not retired	1825	.2334247	.4231261	0	1
Retired	2252	.1896092	.3920789	0	1
Years-left-alive	2091 [b]	12.97131	9.359188	0	42
Age resigned from Court, if resigned	1297	72.54625	10.50579	47	90

Notes: [a]Excludes 135 justice-years for justices serving on Court in 2006. [b]Excludes 161 justice-years for justices serving on court in 2006, and the one live former justice in 2006, Sandra Day O'Connor.

Table 3 – Robust, Cluster-Corrected Standard Errors and Test Statistics for First Stage Regression and Probit Analyses of *Retired* on Exogenous Variables

	Regression Analysis			***Probit Analysis***		
Independent variable	**Coefficient**	**Standard Error**	**t**	**Coefficient**	**Standard Error**	**Z**
Qualified-for-pension	.1528307	.0185742	8.23	.230499	.1259244	1.83
Age	-.0231956	.0055049	-4.21	-.0691558	.0729478	-0.95
Age^2	.0005687	.0000404	14.07	.0015023	.0005143	2.92
Age^3	-3.61×10^{-6}	1.36×10^{-7}	-26.57			
Year1788	-.002061	.0005303	-3.89	.005611	.0010305	5.44
$Year1788^2$	9.98×10^{-6}	2.23×10^{-6}	4.48			
Tenure	-.0098628	.0025242	-3.91	-.0967758	.0081613	-11.86
$Tenure^2$	.0002073	.0000733	2.83			
Constant	.306169	.1710694	1.79	-3.117741	2.582586	-1.21
N justice-years	1971			1971		
R^2 or Psueudo R^2	0.5603			0.4366		
F or Ln(likelihood)	$F_{(8,1962)}$ =	312.52		Ln(likelihood)=	-505.261	

Notes: The dependent variable in both of these analyses is the dummy variable *retired*. n for these analyses identical to n's for ***IV*** analyses in Table 4.

Table 4 – Results of 12 Analyses of Retirement Effects on Remaining Years of Life and Annual Mortality Hazard, 1801-2006 (**a**)

Dependent Variable, Model and Analysis Method		**Effect Measure**	**Effect Estimate (Alternate Estimate)**	**Effect Metric (Alternate Metric)**	**Effect Standard Error**
ANALYSES OF YEARS-LEFT-ALIVE					
Model 1a	Regression	Coefficient of dummy variable *retired*	-3.574253	years	1.39657***
Model 1b	Regression *with probit transform*	Coefficient of dummy variable *retired* *(Difference in expected years-left)*(**b**)	-.2846948 (-3.74) **(b)**	*probit transform* years (years) **(b)**	.1147357*** (*na)*
Model 1c	Separate regressions *with probit transform* for retired and incumbent justice-years	Difference between retired and not-retired mean predicted years-left-alive	-6.596452	years	2.590661***
Model 1d	Tobit	Coefficient of dummy variable *retired*	-3.333841	years	.6437869***
Model 3a	Instrumental variables regression	Coefficient of dummy variable *retired*	-13.35622	years	7.325909*
Model 3b	Instrumental variables regression *with probit transform*	Coefficient of dummy variable *retired* *(Difference in expected years-left)* (**c**)	-1.036574 (-9.24) **(c)**	*probit transform* years (years) **(c)**	.6031232* (*na)*
Model 3c	Instrumental variables Tobit	Coefficient of dummy variable *retired*	-13.58067	years	7.669498*
Model 5a	Endogenous switching regression	Difference between retired and not-retired mean predicted years-left-alive	-5.790303	years	1.698439***
Model 5b	Endogenous switching regression *with probit transform*	Difference between retired and not-retired mean predicted years-left-alive	-6.881007	years	1.494889***
ANALYSES OF MORTALITY HAZARD					
Model 2	Probit	Coefficient of dummy variable *retired*	.4962814 (.075366) **(d)**	*probit* (probability)	.142235*** (*na)*
Model 4	Instrumental variables probit	Coefficient of dummy variable *retired*	.7538361 (.13646) **(e)**	*probit* (probability)	.2168247*** (*na)*
Model 6	Endogenous switching probit	Difference between retired and not-retired mean predicted mortality hazard	.0133964	probability	**na**

Notes: **(a)** This table reports the coefficient of current retirement status (*Retired*) or the difference between the mean predicted value of *years-left-alive* or mortality hazard. See Appendix 2 for complete results. All instrumental variables analyses are based on two equations, one predicting the hazard of retirement, and one predicting years-left-alive or the hazard of mortality. Analyses are based on data for the years 1801 through 2006, for justices who died in office at any age, or resigned from the Supreme Court at the age of 55 years or older. The unit of analysis in all analyses is the justice-year; n = 1971 for analyses of years-left-alive; n = 2132 for analyses of mortality hazard Different n's occur because years-left-alive is not observable for the living (all incumbent justices in 2006, and the retired but living Sandra Day O'Connor). **(b)** For incumbent justices who otherwise have 11 years-left-alive, coefficient indicates a retirement effect of -3.74 years, on average. **(c)** For incumbent justices who otherwise have 11 years-left-alive, coefficient indicates a retirement effect of -9.24 years, on average. *na* indicates that repeated efforts to obtain a bootstrap estimate of the standard error of this effect. **(d)** For justices who would have an annual mortality hazard of.05 if incumbent, that probability would increase to .125366, an increase of .075366. **(e)** For justices who would have an annual mortality hazard of.05 if incumbent, that probability would increase to .18646 if retired, an increase of .13646. * Statistically significant, 1-tailed test, 5%. ** Statistically significant, 1-tailed test, 2.5%. ***Statistically significant, 1-tailed test, 1%. See text for discussion of significance tests. Standard errors are robust and cluster-corrected, except in Models 1c, 5a, 5b and 6, for which standard errors are obtained by cluster-sample bootstrapping with 1391 replications.

Results from IV Re-Estimation Excluding Justices who Never became Eligible to Receive Pension

Model 4; 57 clusters, 1411 justice-years, coefficient of *retired* = .9069707, cluster corrected robust standard error = .2824573, t = 3.21.
Model 3b, 50 clusters, 1269 justice-years, coefficient of *retired* = -1.089993, cluster corrected robust standard error =.378642, t= -2.88
Model 3c, 50 clusters, 1269 justice-years, coefficient of *retired* = -10.4401, cluster corrected robust standard error = 4.832908, t= -2.16

APPENDIX

COMPLETE ANALYSIS RESULTS FOR TABLE 4

>>>>INCLUDED FOR REVIEWERS, BUT NOT NECESSARILY FOR PUBLICATION<<<<

Model 1a -- OLS Regression of Years Left Alive, with Robust Standard Errors Corrected for Clustering

Independent Variable	Coefficient	Robust S.E.	t
retired	-3.574253	1.39657	-2.56
tenure	-.2186793	.0706771	-3.09
year1788	.1025659	.0657349	1.56
year1788^2	.000191	.0002771	0.69
age	-1.678824	.3520718	-4.77
age^2	.0109136	.0028048	3.89
age*year1788	-16.50542	10.7429	-1.54
constant	75.7191	10.2454	7.39
N	1971		
R^2	0.4503		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 1b -- OLS Regression of Probit-Transformed Years Left Alive, with Robust Standard Errors Corrected for Clustering

Independent Variable	Coefficient	Robust S.E.	t
Age	-.0772333	.0238676	-3.24
Age^2	.0003088	.000177	1.74
Age^3	6.13e-07	5.98e-07	1.02
year1788	.0010377	.004187	0.25
year1788^2	8.01e-06	.000016	0.50
tenure	-.0017634	.0133526	-0.13
tenure^2	-.0005205	.0003079	-1.69
retired	-.2846948	.1147357	-2.48
Constant	2.695483	.6648214	4.05
N	1971		
R^2	.4115		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 1c -- Separate regressions of Probit-Transformed Years Left Alive for retired and incumbent justice-years

Retired Justices Only

Independent Variable	Coefficient	Robust S.E.	t
tenure	-.0185406	.0075543	-2.45
year1788	.0151861	.0078526	1.93
$year1788^2$	-.0000121	.0000279	-0.44
age	.08479	.1267269	0.67
age^2	-.0005418	.0008467	-0.64
age*year1788	-1.421189	1.290685	-1.10
constant	-4.633766	4.483483	-1.03
N	334		
R^2	0.2467		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Incumbent Justices Only

Independent Variable	Coefficient	Robust S.E.	t
tenure	-.0162098	.0067212	-2.41
year1788	.0003398	.004854	0.07
$year1788^2$	.0000218	.0000212	1.03
age	-.0706071	.0259633	-2.72
age^2	.0003379	.0002459	1.37
age*year1788	-.3532584	.8616054	-0.41
constant	2.490507	.6571686	3.79
N	1637		
R^2	0.3696		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 1c -- Separate regressions of Years Left Alive for retired and incumbent justice-years

Retired Justices Only

Independent Variable	Coefficient	Robust S.E.	t
tenure	-.1767526	.0845531	-2.09
year1788	.2012675	.1025484	1.96
$year1788^2$	-.0003024	.000309	-0.98
age	.2642597	1.54243	0.17
age^2	-.0017258	.0104017	-0.17
age*year1788	-14.23091	15.50226	-0.92
Constant	-7.309449	54.63108	-0.13
N	334		
R^2	0.2632		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Incumbent Justices Only

Independent Variable	Coefficient	Robust S.E.	t
tenure	-.2358315	.0937814	-2.51
year1788	.0701248	.0756628	0.93
$year1788^2$	.0003229	.0003019	1.07
age	-1.610293	.3359931	-4.79
age^2	.0103302	.0029666	3.48
age*year1788	-15.82976	12.32282	-1.28
Constant	75.08222	8.914693	8.42
N	1637		
R^2	0.4199		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 1d -- Tobit Regressions of Years Left Alive

Independent Variable	Coefficient	Robust S.E.	t
age	-1.519081	.156942	-9.68
age^2	.0073067	.001218	6.00
age^3	.0000106	4.40e-06	2.40
year1788	.0261405	.0152114	1.72
$year1788^2$	.000066	.0000645	1.02
tenure	-.1343282	.071263	-1.88
$tenure^2$	-.0040159	.0021061	-1.91
retired	-3.333841	.6437869	-5.18
Constant	76.02991	4.896246	15.53
Sigma	7.128618	.1172019	
N	1971		
LR $\chi^2_{(8)}$	1148.78		
Ln(likelihood)	-6454.7319		
Pseudo R^2	0.0817		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 2 Probit Analysis of Mortality Hazard

Independent Variable	Coefficient	Robust S.E.	t
year1788	-.0039512	.001188	-3.33
age	.063109	.0491088	1.29
age^2	-.0002155	.0003329	-0.65
tenure	.0149047	.0057127	2.61
retired	.4962814	.142235	3.49
Constant	-4.90723	1.752648	-2.80
N	2132		
Wald $\chi^2_{(5)}$	83.19		
Ln(pseudo-Likelihood)	-324.40042		
Pseudo R^2	0.1373		

Notes: Estimated over all justice-years 1801-2006 for justices who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 3a – IV Regression of Years Left Alive

First-stage regression of Retired

Independent Variable	Coefficient	Robust S.E.	t
age	-.0231956	.0055049	-4.21
age^2	.0005687	.0000404	14.07
age^3	-3.61e-06	1.36e-07	-26.57
year1788	-.002061	.0005303	-3.89
$year1788^2$	9.98e-06	2.23e-06	4.48
tenure	-.0098628	.0025242	-3.91
$tenure^2$	.0002073	.0000733	2.83
qual4pen	.1528307	.0185742	8.23
Constant	.306169	.1710694	1.79
N	1971		
R^2	0.5603		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Instrumental variables (2SLS) regression

Independent Variable	Coefficient	Robust S.E.	t
retired	-13.35622	7.325909	-1.82
age	-1.914292	.4524699	-4.23
age^2	.0144326	.0055055	2.62
age^3	-.000024	.0000243	-0.99
year1788	.0123105	.0627544	0.20
$year1788^2$	.0001476	.0002517	0.59
tenure	-.2095038	.1850819	-1.13
$tenure^2$	-.0018258	.003969	-0.46
Constant	82.83009	12.20146	6.79
N	1971		
R^2	0.3661		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 3b -- IV Regression of Probit Transformed Years Left Alive

First-stage Regression of Retired

Independent Variable	Coefficient	Robust S.E.	t
age	-.0231956	.0055049	-4.21
age^2	.0005687	.0000404	14.07
age^3	-3.61e-06	1.36e-07	-26.57
year1788	-.002061	.0005303	-3.89
$year1788^2$	9.98e-06	2.23e-06	4.48
tenure	-.0098628	.0025242	-3.91
$tenure^2$	.0002073	.0000733	2.83
qual4pen	.1528307	.0185742	8.23
Constant	.306169	.1710694	1.79
N	1971		
R^2	0.5603		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Instrumental variables (2SLS) regression

Independent Variable	Coefficient	Robust S.E.	t
retired	-1.036574	.6031232	-1.72
age	-.1018938	.0317218	-3.21
age^2	.0007896	.0004223	1.87
age^3	-1.89e-06	1.95e-06	-0.97
year1788	-.0001655	.0044944	-0.04
$year1788^2$	.0000153	.0000185	0.82
tenure	-.0059157	.0140778	-0.42
$tenure^2$	-.0004328	.0003397	-1.27
Constant	3.076704	.8184089	3.76
N	1971		
R^2	0.3354		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 3c Instrumental Variables Tobit Regressions of Years Left Alive

Instrumental variables Tobit regression of Years left Alive

Independent Variable	Coefficient	Robust S.E.	t
retired	-13.58067	7.669498	-1.77
age	-1.857586	.4564751	-4.07
age^2	.0138752	.0056318	2.46
age^3	-.0000235	.0000252	-0.93
year1788	.0096202	.0651035	0.15
$year1788^2$	.0001654	.0002628	0.63
tenure	-.1920307	.1921549	-1.00
$tenure^2$	-.0027865	.0042272	-0.66
Constant	81.30999	12.24022	6.64
N	1971		
Wald $\chi^2_{(8)}$	373.53		
Ln(pseudo-likelihood	-6504.8751		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

First-stage Probit Analysis of Retired

Independent Variable	Coefficient	Robust S.E.	t
age	-.0231956	.0183781	-1.26
age^2	.0005687	.000168	3.39
age^3	-3.61e-06	5.09e-07	-7.08
year1788	-.002061	.0012595	-1.64
$year1788^2$	9.98e-06	5.57e-06	1.79
tenure	-.0098628	.0047287	-2.09
$tenure^2$	.0002073	.0001231	1.68
qual4pen	.1528307	.0565794	2.70
Constant	.306169	.4996489	0.61
N	1971		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 4 Instrumental Variables Probit Analysis of Mortality Hazard

A. IV Probit Analysis of Mortality Hazard

Independent Variable	Coefficient	Robust S.E.	t
retired	.7538361	.2168247	3.48
year1788	-.0041319	.0011201	-3.69
age	.0862405	.0536082	1.61
age^2	-.0004365	.0003811	-1.15
tenure	.0202197	.0074704	2.71
Constant	-5.534755	1.884165	-2.94
N	2132		
Wald $\chi^2_{(5)}$	89.57		
Ln(pseudo-Likelihood)	-332.88078		

Notes: Estimated over all justice-years 1801-2006 for justices who did not retire before age 55. Standard errors are robust and corrected for clustering.

B. First Stage Probit Analysis of Retired

Independent Variable	Coefficient	Robust S.E.	t
year1788	-.0004743	.0010528	-0.45
$year1788^2$	2.01e-06	4.28e-06	0.47
age	-.0296732	.0171831	-1.73
age^2	.0006281	.0001588	3.95
age^3	-3.71e-06	5.22e-07	-7.11
tenure	-.0101764	.0046932	-2.17
$tenure^2$	.00024	.0001304	1.84
qual4pen	.1381422	.051353	2.69
Constant	.4392054	.4753172	0.92

Model 5a – Endogenous Switching Regression Analyses of Years Left Alive

First-stage Regression of Retired

Independent Variable	Coefficient	Robust S.E.	t
qual4pen	.230499	.1259244	1.83
age	-.0691558	.0729478	-0.95
age^2	.0015023	.0005143	2.92
year1788	.005611	.0010305	5.44
tenure	-.0967758	.0081613	-11.86
Constant	-3.117741	2.582586	-1.21
LR chi2(5)	783.18		
Pseudo R^2	0.4366		
Log likelihood	-505.261		
N	1971		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Endogenous Switching Regression of Years Left Alive, Retired Only

Independent Variable	Coefficient	Robust S.E.	t
age	.6692842	1.600425	0.42
age2	-.0063831	.0098246	-0.65
year1788	.0081152	.0167509	0.48
tenure	-.1543455	.0818264	-1.89
Constant	-4.431863	64.57724	-0.07
Wald $\chi^2_{(4)}$	10.98		
Ln(pseudo-likelihood	-1556.475		
N	1971		
Uncensored N	334		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Endogenous Switching Regression of Years Left Alive, Incumbent Only

Independent Variable	Coefficient	Robust S.E.	t
age	-1.89451	.4557748	-4.16
age2	.0117676	.0044845	2.62
year1788	.0466799	.017038	2.74
tenure	-.2745009	.1324152	-2.07
Constant	84.64221	12.10133	6.99
N	1971		
Wald $\chi^2_{(4)}$	343.86		
Uncensored N	1637		
Ln(pseudo-likelihood	-6030.015		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 5b – Endogenous Switching Regression Analyses of Probit-Transformed Years Left Alive

Endogenous Switching Regression of Probit Transformed Years Left Alive, Retired Only

Independent Variable	Coefficient	Robust S.E.	t
age	.0740601	.1462538	0.51
age^2	-.000717	.0008651	-0.83
year1788	.000472	.0017244	0.27
tenure	-.0115464	.0105761	-1.09
Constant	-2.386027	6.16838	-0.39
N	1971		
Uncensored N	334		
Wald $\chi^2_{(4)}$	16.84		
Ln(pseudo-likelihood	-777.8706		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Endogenous Switching Regression of Probit Transformed Years Left Alive, Incumbent Only

Independent Variable	Coefficient	Robust S.E.	t
age	-.0748656	.0203591	-3.68
age^2	.0003159	.0001846	1.71
year1788	.0029542	.0011001	2.69
tenure	-.0132261	.0070124	-1.89
Constant	2.607436	.5710812	4.57
N	1971		
Uncensored N	1637		
Wald $\chi^2_{(4)}$	322.77		
Ln(Pseudo-likelihood)	-1769.277		

Notes: Estimated over all justice-years 1801-2006 for justices deceased by 2006 who did not retire before age 55. Standard errors are robust and corrected for clustering.

Model 6 Endogenous Switching Probit Analysis of Mortality Hazard

A. First Stage Probit Analysis of Retired

Independent Variable	Coefficient	Robust S.E.	t
qual4pen	.2148681	.1238097	1.74
age	-.0462563	.0734785	-0.63
age^2	.0013982	.0005168	2.71
year1788	.0036605	.0009559	3.83
tenure	-.1017056	.0081515	-12.48
Constant	-3.918372	2.611989	-1.50
N	2132		
LR $\chi^2_{(5)}$	796.54		
Ln(likelihood	-527.21186		
Pseudo R^2	0.4303		

Notes: Estimated over all justice-years 1801-2006 for justices who did not retire before age 55. Standard errors are robust and corrected for clustering.

B. Probit Analysis of Mortality Hazard for Retired Only

Independent Variable	Coefficient	Robust S.E.	t
age	-.0608349	.1993165	-0.31
age^2	.0006625	.0011915	0.56
year1788	-.0016844	.0022062	-0.76
tenure	.013234	.019291	0.69
Constant	-.6299868	8.359182	-0.08
N	2132		
Censored N	1798		
Wald $\chi^2_{(4)}$	15.24		
Log likelihood	-641.3718		

Notes: Estimated over all justice-years 1801-2006 for justices who did not retire before age 55. Standard errors are robust and corrected for clustering.

C. Probit Analysis of Mortality Hazard for Incumbent Only

Independent Variable	Coefficient	Robust S.E.	t
age	.1172049	.1118509	1.05
age^2	-.0005329	.0008795	-0.61
year1788	-.0042365	.0014241	-2.97
tenure	.0058439	.0147427	0.40
Constant	-6.850955	3.654688	-1.87
N	2132		
Censored N	334		
Wald $\chi^2(4)$	39.43		
Log likelihood	-735.6234		

Notes: Estimated over all justice-years 1801-2006 for justices who did not retire before age 55. Standard errors are robust and corrected for clustering.

[1] Personal communication, May 8, 2007.

[2] Explain computations.