Do Not Go Gentle into that Good Night:

The Effect of Retirement on Subsequent Mortality of U.S. Supreme Court Justices, 1801-2006*

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ABSTRACT

Mortality hazard and length of time until death are widely used as health outcome measures, and are themselves of fundamental demographic interest. Considerable research asks if labor force retirement reduces subsequent health and its mortality measures. Previous studies report positive, negative and null effects of retirement on subsequent longevity and mortality hazard, but inconsistent findings are difficult to resolve because 1) nearly all data confound retirement with unemployment of older workers, and, often, 2) endogeneity bias is rarely addressed analytically. To avoid these problems, albeit at loss of generalizability to the entire labor force, I examine data from an exceptional subgroup, of interest in its own right: US Supreme Court justices, 1801 - 2006. Using discrete time event history methods, I estimate retirement effects on mortality hazard and years-left-alive. Some substantive and methodological considerations suggest models that specify endogenous effects estimated by instrumental variables (***IV***) probit, ***IV*** Tobit and ***IV*** regression methods. Other considerations suggest estimation by endogenous switching (***ES)*** probit and ***ES*** regression. Estimates by both methods are consistent with the hypothesis that on average retirement decreases health, as indicated by elevated mortality hazard and diminished years left alive. These findings may apply to other occupational groups characterized by high levels of work autonomy, job satisfaction, and financial security.

[For reviewer convenience, this document includes a supplementary 13-page appendix with full details of all analyses; in the paper itself, those analyses are summarized in Table 4 on one page.]

[Appendix A1 reports a simulation study that addresses a reader's questions concerning the conditions under which instrumental variables estimates of a variable's effect are stronger than reduced form estimates of that same effect.]

I. INTRODUCTION

The length of life remaining until death (or its probabilistic determinant, mortality hazard) is of interest in its own right, and widely used as a measure of "objective" health (Bound 1991), health outcome and vitality (see e.g. the review of 27 studies by Idler and Benyamini 1997).[1] Considerable research investigates the effects of labor force status in general, and retirement in particular, on longevity, mortality and the health constructs they measure. Some studies report that retirement tends to shorten remaining life (Snyder and Evans 2006; Waldron 2001, 2002; Morris Cook and Shaper 1994), while others find the opposite (Munch and Svarer 2005; Handwerker 2007), or no effect at all (Tsai 2005; Litwin 2007; Mein Martikainen Hemingway Stansfeld and Marmot 2003). However, data limitations and consequent measurement and modeling problems appear to reduce the certainty of these findings, and make it difficult to resolve inconsistencies among them. Here, I reconsider existing data and methods pertaining to this topic. I attempt to avoid measurement and modeling problems by using 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 or manageable. That subgroup is justices of the U.S. Supreme Court from 1801 through 2006.

The Data Problem. In data from nearly all population segments, retirement is conflated with involuntary unemployment of older workers. Simply stated, involuntarily unemployed older workers, pension recipients whose employment was involuntarily terminated, and those who were "encouraged" by employers to retire tend to report that they are *voluntarily retired* (Gustman, Mitchell and Steinmeier 1995: s63; Stolzenberg 1989). Empirically, this misreporting would attribute to retirement the

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[1] Our use of the word "objective" follows Bound (1991) and the definition of the *The American Heritage Dictionary of the English Language*, Third Edition, 1996: a synonym for "observable" or perceived "by someone other than the person affected." In predicting health care expenditures, Shang and Goldman (2008) compare their predicted life expectancy measure ("predicted life expectancy with some noise" [p. 409]) to several health condition and health risk factor measures.

empirically-verified, pernicious effects of unemployment (e.g. Linn, Sandifer, and Stein 1985; Gerdtham and Johannesson 2003, Morris Cook and Shaper 1994; Voss, Nylen, Floderus, Diderichsen, Terry 2004). Further, misreporting of retirement status confuses key concepts. Although everyday language uses many demographic and labor force concepts imprecisely, 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 or physical intensity of paid work (see reviews in Moen, Kim and Hofmeister 2001; Lumsdaine 1995).[2] Thus the retirement decision is a rational voluntary action by the individual concerned, whether it is made with pleasure or regret, whether made to permit the retiree to care for an ailing relative, pursue leisure interests, escape from distasteful working conditions, or for any other reason. In this way, retirement differs fundamentally from involuntary changes in labor force status, including 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 by judicial regulations from workplace pressure to resign. That is, justices' retirement benefits and pay are fixed by law, their working conditions are free of employer manipulation, and regulations prevent them from receiving gifts, payments or other material

[2] 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 full and accurate information were available (Polivka and Rothgeb 1993:24). This social science definition of retirement differs from actuarial and accounting definitions, which usually include only recipients of money payments from pension funds (Society of Actuaries 1992).

inducements to resign or remain in office (United States Constitution. Article II. 1789; 5 U.S.C. App. §§ 501-505).[3] Justices are famously vocal about their intentions to remain on the court as long as they wish (e.g. Williams 1990),[4] and their behavior is generally consistent with these stated intentions, even in the presence of physical decay and "mental decrepitude" (Garrow 2000). 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 full pay as working justices. Historically, 49.5 percent of all former justices died in office, without retiring. In short, the law leaves retirement decisions to justices, evidence suggests that they make those decisions as they alone choose, and so Supreme Court data appear to avoid confounding retirement with involuntary unemployment.

The Endogeneity Problem. In addition to measurement problems, analyses of retirement effects are well-known for susceptibility to unrecognized endogeneity and consequent identification and estimation problems (Snyder and Evans 2006; Handwerker 2007). Endogeneity arises because voluntary retirement is, in the language of causal effects, a self-selected "treatment." Health and vitality, as indicated by mortality hazard and remaining length of life, may affect the decision to select this treatment, even as the treatment may affect various measures of health, vitality, mortality hazard and remaining life. For example, increases in mortality hazard and decreases in years of remaining life are substantially correlated with self-assessed subjective health (Idler and Benyamini 1997), and reduced subjective health is often mentioned as a reason for retirement, even by retirees who are able to work (Bound 1991, Reno 1971, Schwab 1974, Sherman 1985, Sickles and Taubman 1986). If these reciprocal effects exist empirically but are unrecognized analytically, they produce endogeneity bias, even in efforts to estimate only retirement effects on longevity, rather than longevity and retirement effects on each other.

[3] Justices can be, but none have been, terminated from office for treason, bribery, or serious crimes.

[4] 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).

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). That analysis uses an actuarial definition of retirement (i.e. receipt of pension benefits) well suited to pension fund financial analysis, but not as appropriate for the present purpose of understanding effects of retirement decisions on the individuals who make those decisions. Below, I note that unusual features of Supreme Court justice pension policies, and other peculiarities of those pensions, permit the use of *pension qualification* (rather than pension *receipt*) as an instrumental variable to identify retirement effects on health and longevity.

Finally, if mortality risk is determined according to one causal regime before retirement, but according to another regime 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). Below, I explain why Supreme Court data appears to permit instrumental estimation of retirement effects on subsequent longevity, as well as distinguishing between retirement and involuntary unemployment.

Retirement Effects in Special Populations. 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 characterized 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). Our analyses are relevant to ongoing policy debates about retirement in general (Gokhale 2004; Ashenfelter and Card 2002), and term limits for Supreme Court justices (Calabresi and Lindgren 2006). Even in the unlikely case that Supreme Court retirement and mortality patterns are dissimilar to those patterns in any other social group, and unrelated to retirement, health 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 (Garrow 2000; 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.[5] Separately, and apparently unaware 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 mortality following retirement. Finally, analyses presented here address a key question about nonpecuniary effects of work and employment: Is it economically irrational for justices to work after they become eligible to receive retirement pensions equal to their pay?[6] If continued life has sufficient value, and if work prolongs life, then the value received for unpaid work would be apparent.

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

II. HYPOTHESES

A. The Simple Model: Can the observed statistical relationship of retirement with subsequent health (as indicated by subsequent longevity and mortality hazard) be explained by a simple model (hereafter, the "Simple Model") that lacks retirement effects on subsequent health, mortality and longevity.[7] In the Simple Model, each justice's true health is exogenously determined, unobservable to

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[5] Preston (1977:171) writes, "Mortality levels obviously have a major influence on the structure of other elderly leadership groups such as union leaders, Supreme Court justices, and Communist Party officials." Thus, determinants of these anomalous mortality levels are of interest for the identical reason.

[6] This question was asked informally, by Gary S. Becker of U.S. Federal Judge Richard Posner (Personal communication, May 8, 2007).

[7] Thanks to Robert Willis for suggesting this approach.

the justice, tends to decline monotonically over time, and is indicated by mortality hazard and, consequently, time left to live. True health stochastically determines justices' subjective assessments of their own health. Justices tend to retire when they believe that their true health is poor. In brief, the Simple Model asserts that true health is a cause of both retirement and mortality. If correct, the Simple Model would explain the positive correlation of mortality and retirement, without any effect of retirement on health or its indicator, mortality.

However, the Simple Model is inconsistent with statistical data and historical narrative. In particular, although the Simple Model predicts that justices retire when health fades and death approaches, Garrow (2000) describes a regular historical pattern in which justices delay or refuse retirement, despite obvious, even lurid, "mental decrepitude" and physical decline. Consistent with Garrow, 49.5 percent (as of 2006) of all former justices never retire, but die in office. Some of those who die in office may perish while believing themselves to be in good health, but it seems doubtful, if not absurd to assert, that half of all previous justices died in office without prior awareness that they were at death's door. Further, although the Simple Model asserts that failing health is the primary signal for justices to resign, Stolzenberg and Lindgren (2009: Table 3) find that political circumstances, pension eligibility and other factors account for more than nine times as much of the variation in the hazard of retirement as years left alive. In short, the Simple Model would be a Procrustean bed for Supreme Court retirement and mortality data. Next, I consider three competing hypotheses about the effects of retirement on post-retirement longevity.

B. The Null Effects Hypothesis**:** *There is no effect of retirement on subsequent mortality hazard and subsequent longevity, on average and other things equal.* This hypothesis asserts that any apparent association between retirement timing and subsequent mortality risk is spurious. Although empirical research methods are poorly suited to testing hypotheses of "no effect," this hypothesis has considerable precedent. For example, Mein, Martikainen, Hemingway, Stansfeld and Marmot (2003) report that early retirement at age 60 has no effect on physical health. Tsai (2005) concludes that, after adding control variables to his analysis of retirement and mortality of Shell Oil employees, "early retirement at 55 or 60 is not associated with increased survival." After detailed examination of confounding variables and measurement issues, Litwin (2007) concludes, "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.'"

C. The Increased Mortality Hypothesis**:** *Retirement increases subsequent mortality hazard (and reduces subsequent 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 finding in several large U.S. national data sets that mortality hazard declines as retirement age rises, controlling for current age. Theoretically, the Increased Mortality Hypothesis arises from the observation that, compared to nonparticipation in the labor force, employment tends intensify 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; Bortz 1984 calls these effects "disuse syndrome"). Wagner et al. (1992) 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 attenuated mortality among retired workers who return to work after their Social Security benefits are reduced by a government policy change; 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).

D. The Reduced Mortality Hypothesis. *Retirement reduces subsequent mortality hazard (and increases subsequent 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." A competing risks analysis of Danish data reports that "early retirement prolongs survival for men" (Munch and Svarer 2005:17). 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 suggests that employment exposes many workers to life-shortening health risks (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 and perceived effort-reward imbalance (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 health impediments and their mortality indicators.

III. ANALYSIS STRATEGY, ESTIMATION AND DATA

Our analysis strategy exploits key features Supreme Court justices' employment, including the temporal organization of Court work, and the Federal judicial pension system.

A. Discrete Time Event History Models. In the language of causal inference, I seek to measure the effect of a time-related treatment (retirement) on time-related outcomes (mortality hazard and years-left-alive), for those who select the treatment. Accordingly, I apply event history data and methods with accommodations for self-selection (see below). I 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. Date rounding, co-occurrence of events, and time-varying independent variables are more easily accommodated by discrete time event history methods than by continuous time methods (Yamaguchi 1991). Discrete time methods also accommodate right censoring (Allison 1995), which occurs for justices who die in office without resigning from the Court. So, I 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, death, remaining years of life and other events and 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 during that year. *Years-left-alive* for each justice-year is the number of additional years after the current year until the relevant justice dies. These forward-looking measures differ from lifetime average annual mortality probability or age at death.

To assure that estimates of mortality hazard are in the interval [0,1] for which probabilities are defined, I use maximum likelihood probit analysis to measure the effect of retirement on *mortality hazard*.[8] To assure that estimates of *years-left-alive* are non-negative, I sometimes use a probit-transformation (described below) of *years-left-alive* or a Tobit analysis of *years-left-alive* (Amemiya 1985). These methods are well-known, but not commonly used together.

B. Endogeneity by Mediation. Above, I observe that endogenous retirement can be represented as endogenous mediation (shown in Figure 1) or endogenous switching (discussed below and shown in Figure 2). In endogenous mediation, an endogenous variable, here *retired* (an indicator for retirement from active service on the Court), mediates some of the effect of exogenous variables on the endogenous

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[8] I use probit rather than logit or gompit methods for consistency with our use of Tobit, instrumental variables probit, and endogenous switching methods.

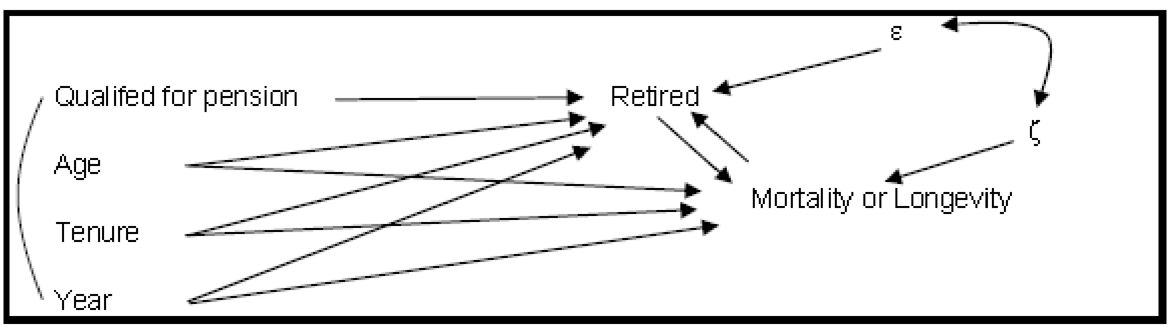
outcome variable, here *mortality hazard* or longevity. *Endogenous mediation* is the problem for which Instrumental Variables (***IV***) estimation is the standard solution (Amemiya 1985).

In ***IV*** endogenous mediation models, identification of the effect of an endogenous mediating variable on an endogenous outcome variable requires at least one instrument. The instrument must have direct effects on the endogenous mediator, but is restricted to have only indirect effects on the endogenous outcome. Because those requirements must be satisfied logically and substantively, I now discuss reasons to believe that *pension-eligible* – a variable indicating whether or not a particular justice in a specific year would be eligible for retirement with a Federal judicial pension if retired – has a direct effect on retirement, but only indirect effects on subsequent mortality measures.

Pension-eligible appears to have the necessary direct effect on *retired* because pension eligibility removes from retirement the tremendous financial disincentive of salary loss. The Federal judicial pensions system was enacted in 1869 in specific response to the stated unwillingness of senile, sick and feeble justices to lose income by retiring from the Court (Yoon, 2006). Indeed, justices are especially likely to lack savings and other income sources because they are legally forbidden to practice law, or to receive income related to legal practice, fiduciary duties, writing articles, endorsements and other business and professional activities while in active service on the Court (5 U.S.C. App. §§ 501-505; http://www.uscourts.gov/ library/ conduct_outsideemployment.html, accessed 28 June 2009). Stolzenberg and Lindgren (2010) find that pension eligibility increases the annual odds of retirement from the Supreme Court by a multiple greater than 8, on average, net of the retirement effects of the justice's age, justice's time served on the Court, the calendar year, and indicators of political climate.

Three conditions justify the identifying restriction that pension eligibility has no direct effect on mortality or years-left-alive: First, mere *eligibility* for a pension does nothing to health or its mortality indicators – one must actually receive the pension to spend it in ways that affect mortality and longevity. Second, *pension-eligible* is behaviorally distinct from pension receipt: from 1801 to 2006, 23.3 percent of justice-years served on the Supreme Court were served by justices who were already *pension-eligible*.

Finally, pension eligibility is to some extent determined by government policies completely unrelated to the health, mortality hazard and longevity of particular justices. In particular, there were no pensions for justices until 1869, when retirement pay equal to full time pay was instituted for former justices over the age of 70 who had 10 or more years of Federal judicial service (i.e. service on any Federal court). Thus, justices who serve on lower courts before elevation to the Supreme Court might become pension-eligible at a younger age than those without previous Federal judicial service. Starting in 1954, pensions were awarded to former justices older than 65 years if they had at least 15 years of Federal judicial service, and to those over 70 with at least 10 years of service. In 1984, pension eligibility was extended to former justices over 65 years old with at least 10 years of Federal judicial service, for whom the sum of years of age and years of service exceeded 80 ("Rule of 80") (see Yoon 2006). For those justices who ever qualify for the pension, the first quartile of the distribution of age at time of first eligibility is 66 years; the median and third quartile are 70 years, and the maximum is 77 years.



Endogenous Mediation Model of Mortality Hazard

(***IV***1R) $\Pr[\text{retired}_{jt}] = f(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{qualified-for-pension}_{jt}, \varepsilon_{jt})$ [reduced form equation]

(***IV***1S) $\Pr[\text{retired}_{jt}] = f(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{qualified-for-pension}_{jt}, \Pr[\text{death}_{jt}], \varepsilon sp_{jt})$ [structural equation]

(***IV***2a) $\Pr[\text{death}_{jt}] = g(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{retired}_{jt}, \zeta_{jt})$ [structural equation]

Endogenous Mediation Model of Years-left-alive

(***IV***2R) $\Pr[\text{retired}_{jt}] = f(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{qualified-for-pension}_{jt}, \varepsilon_{jt})$ [reduced form]

(***IV***2S) $\Pr[\text{retired}_{jt}] = f(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{qualified-for-pension}_{jt}, \text{years-left-alive}_{jt}, \varepsilon sr_{jt})$ [structural equation]

(***IV***2b) $\text{years-left-alive}_{jt} = h(\text{age}_{jt}, \text{year}_{jt}, \text{tenure}_{jt}, \text{retired}_{jt}, \varpi_{jt})$ [structural equation]

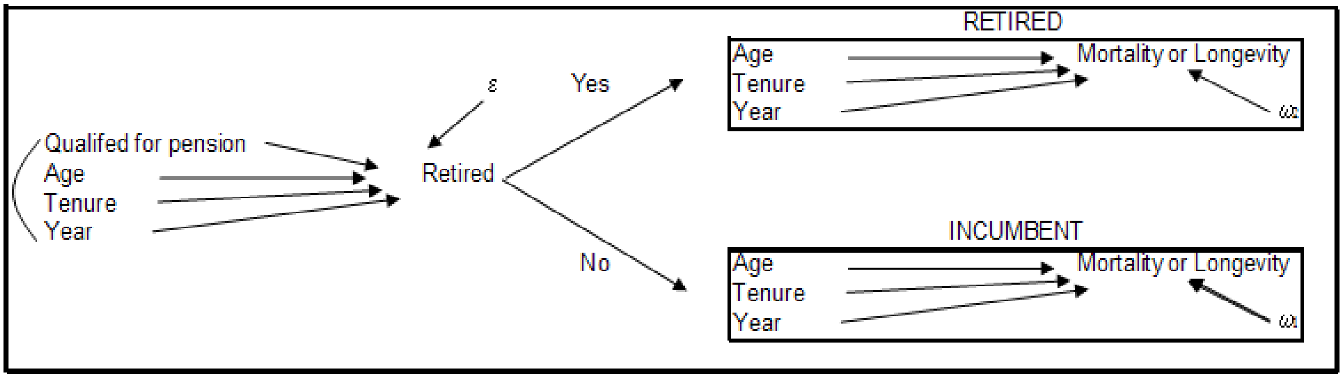
Notes: Equations (***IV***1S) and (***IV***2S) are neither identified nor estimated.
Subscript j refers to the j^{th} justice of the Court; subscript t refers to the t^{th} calendar year ;
f, *g*, and *h* are functions and can include transformations of independent and dependent variables;
ε_{jt}, εs_{jt}, 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.

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

For additional consideration of the suitability of *qualified-for-pension* as an instrument for *retired* in Figure 1, I also estimate ***IV*** analyses on the subset of 57 justices who become *qualified-for-pension* at some point in their Court tenure. In this subset, *pension-eligible* varies over the tenure of each justice, so *qualified-for-pension* can affect the probability of retirement in any justice-year. But, in this subset only, every retiree receives a pension, so qualified-for-pension cannot possibly serve as a proxy for pension receipt in retirement, which would indicate financial resources available to promote the health after retirement.

In Figure 1, equations (***IV***1), (***IV***2a) and (***IV***2b) summarize the endogenous mediation model of retirement and mortality. Subscript j refers to the jth justice of the Court. Subscript t refers to the tth time period (year). $Retired_{jt}$ equals 1 if the jth judge is retired at the start of the tth 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 eponymous characteristics or events, measured in whole years, pertaining to the jth individual during the tth time period.

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 is achieved *via* instrumental variables or nonlinearities.



Retirement Selection Model (Probit)

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

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

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

Conditional Model of Mortality Hazard (Probit)

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

($\boldsymbol{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)

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

($\boldsymbol{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.

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

D. Estimation and Tests. To constrain estimated hazards to the [0,1] interval for which they are defined, I 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, I 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 $(Y=50[\Phi(\Psi)]-0.5)$. 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 on mortality or labor force behavior. 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). Because many mathematical functions virtually duplicate the same values over a fixed domain, it is neither necessary nor possible to distinguish various functions that might produce the same nonlinear effects 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).

I analyze data on the universe of Supreme Court justices of the United States from 1801 through 2006. Rather than join a debate over the appropriateness of sampling-based significance tests for population data, I report standard errors and significance tests for all equation parameters estimated here, but take no position on their appropriateness. Because data contain multiple observations per justice, each justice constitutes an observational cluster, and I 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, and the predictions themselves are based on nonlinear functions of model estimates, ordinary standard errors are not readily available, so I use bootstrapping to calculate them. Although McCullagh (2000) criticizes bootstrapping with clustered data, Feng Mclerran and Grizzle (1998) and Field and Welsh (2007), find that bootstraps perform well, particularly when the number of clusters is 50 or more. All analyses reported here have more than 50 clusters.

E. Data. I examine data on all justices of the United States Supreme Court from 1801 through 2006.[9] Data are an annual event-history data set consisting of one observation for each year in which

[9] I 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

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 lives after leaving the Court. Table 2 provides descriptive statistics for justice-years used in analyses here, as well as historical statistics on all justices of the Court. The solid, stair-step line in Figure 5b provides a Kaplan-Meier plot of raw survival data for justices. From 1801 through the end of 2006, 95 justices served on the Court and subsequently died; one (O'Connor) served, retired in 2006 and lives as this paper is written; and 9 have neither resigned nor died. Collectively, justices served 1825 justice-years on the court, and lived 427 justice-years in retirement.[10] Two women have served as justices (O'Connor and Ginsburg); at the end of 2006, both live and only O'Connor has retired. Obviously, gender controls are not possible in these analyses, nor are inferences about gender differences. A reader speculated that results would differ if women were excluded from statistical analyses. However, when analyses were re-done with Justices O'Connor and Ginsburg omitted from the data, findings were unchanged or virtually identical.

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.[11] Analyses of mortality hazard also include justice-years for justices who have not died as of 2006, 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 or resigned during that year, or before starting service the next year.

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. I added post-retirement data for justices who did not die in office.

[10] 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.

[11] Years-left-alive is unobservable for justices still alive as this research is done.

2. ***Death***. A dummy (0,1) variable equal to 0 for a justice-year unless the justice died that year.

3. ***Year, Year1788**, **ln(Year1788).*** *Year* is the calendar year. *Year1788* is *Year* - 1788.[12] *ln(Year1788)* is the natural logarithm of *Year1788*; the logarithmic transformation improves the fit of some models. I 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, I 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 for fitting nonlinear tenure effects.

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

7. ***Years-left-alive***. In each justice-year, *Years-left-alive* indicates future longevity, or remaining years of life. *Years-left-alive* for a justice-year is the difference between the calendar year of the justice-year and the calendar year in which the justice ultimately dies.

IV. RESULTS

A. Retirement Analyses. Because all of our ***ES*** and ***IV*** analyses of mortality and longevity require a regression or probit analyses of *retired,* I report those 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. Although

[12] 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.

probit analysis does (and 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 models for estimating the effect of *retired* on *years-left-alive*, and the upper panel of Table 4 presents empirical estimates of these estimates. See Appendix 2 for details of analyses.

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, Models 1a-1d all indicate negative effects of retirement on future longevity, all are statistically significant (α≤.01, 1-tailed robust cluster-corrected test). **All analyses hold constant functions 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 (α≤.01, 1-tailed robust cluster-corrected test).
- **Model 1b** applies a probit transformation to *years-left-alive*, yielding a coefficient of -.2847 (α≤.01, 1-tailed robust cluster-corrected test). To express that coefficient in intuitively meaningful terms, I evaluate its effect in years at 11 *years-left-alive* (the median of *years-left-alive*), where the *retirement effect is 3.74 years less remaining life, on average*.
- **In Model 1c**, I 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 6.60 years less life for the retired than for incumbents (significant, α≤.01, 1-tailed test, 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 ($\alpha \leq .01$, 1-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.

<u>*B2.* ***IV*** *Regression and* ***IV*** *Tobit Analyses*</u>. 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 -13.3562 for retired (significant, $\alpha \leq .05$, 1-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.[13]
- **Model 3b** applies a probit transformation to years-left-alive, as well as ***IV*** estimation, yielding a coefficient of -1.0366 (significant, $\alpha \leq .05$, 1-tailed robust cluster-corrected test). The curved, unbroken line in Figure 3 shows model 3b estimates of *years-left-alive* if incumbent or if retired. Other things equal, Model 3b estimates that an incumbent justice with 11 *years-left-alive* (the median of *years-left-alive*) would survive 9.24 fewer years if retired.
- **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. I use Model 3c to predict the *years-left-alive* if incumbent and the *years-left-alive* if retired for each justice-year. The mean difference between these predictions is 13.58 fewer years-left-alive for the retired. Figure 3 plots these predictions in the patterned line. ***IV*** probit and ***IV*** Tobit predictions shown in Figure 3 are similar.

[13] A skeptical reader proposes that the estimated effect of *retired* on mortality must be weaker when estimated with IV analyses than when estimated in analyses that ignore endogeneity. Appendix A1 uses simulation to examine that concern.

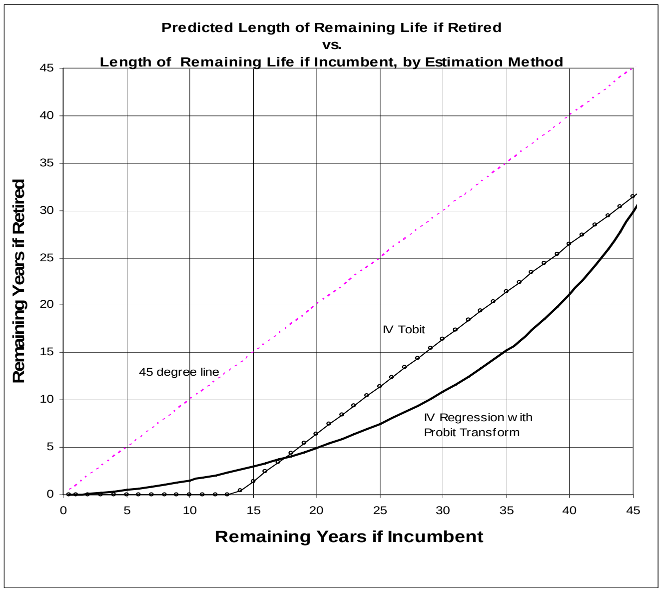


Figure 3 – Predicted Length of Remaining Life If Retired vs. Predicted Length of Remaining Life if Incumbent, for IV Tobit and IV Regression with Probit Transformation

<u>*B3. Endogenous Switching Analyses*</u>. Models 5a and 5b use endogenous switching regression to model the endogeneity of *retired* effects on *years-left-alive.*

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

Figure 4 is the scatter-plot of the ratio of predicted years-left-alive-if-incumbent to predicted years-left-alive-if-retired, by age. Curves are fitted by fractional polynomial regression for indicated half-century periods. In all periods shown in Figure 4, the ratio is largest at the youngest ages (and apparently would be larger yet below age 55), declines as age increases, and then rises again. As indicated by the graph, for the period 1951-2006, justices who are incumbent at age 65 have twice the years-left-alive as those who are retired, other things equal.

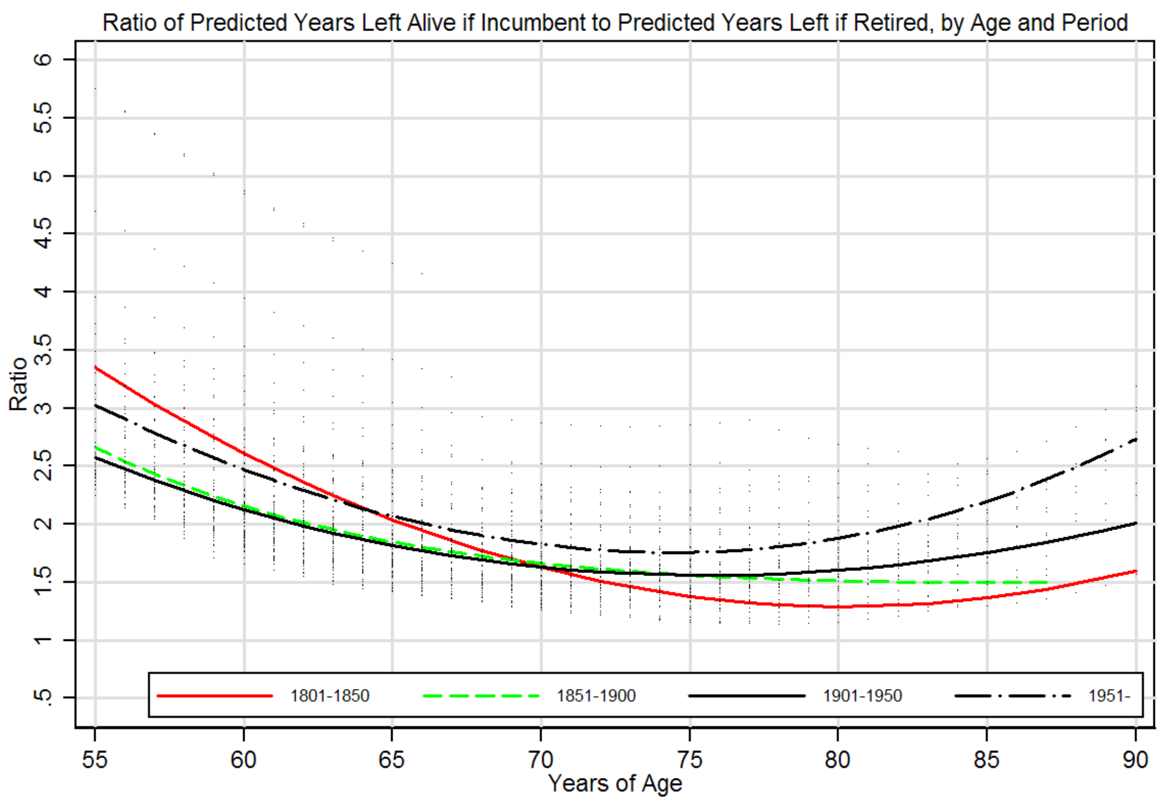


Figure 4 -- Ratio of expected-years-left-alive-if-incumbent to expected-years-left-alive-if-retired vs Calendar Year, by Age

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* 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 5a graphs the effect of this coefficient on mortality hazards. The distance from the solid line to the lower "equal values" line is the estimated effect of retirement on mortality hazard: In the metric of probabilities, 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. **IV** Analysis*. Model 4 is an ***IV*** probit analysis of retirement effects on mortality hazard. The coefficient of *retired* in Model 4 is .7538361 – significantly different from zero ($\alpha \leq .01$, 1-tailed robust cluster-corrected test), and larger than the Model 2 estimate. Based on the Model 4 coefficient, the upper broken line in Figure 5a shows the ***IV*** probit (Model 4) estimate of the retirement effect on mortality hazard. Other things equal, an incumbent with an annual mortality hazard of 5 percent would face a hazard of 18.6 percent if retired, according to Model 4.

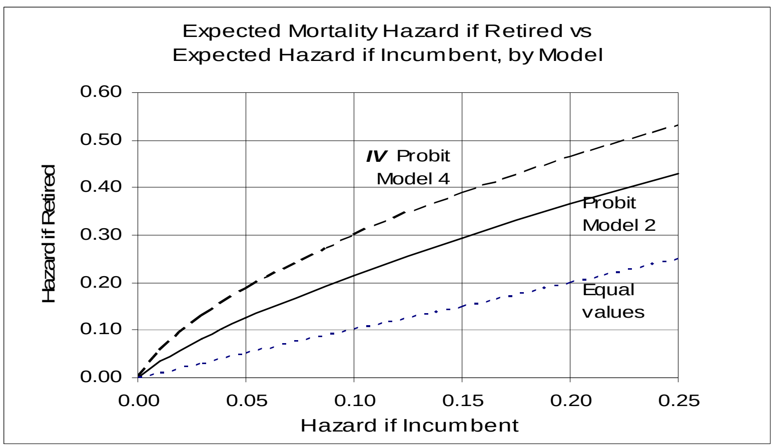


Figure 5a -- Expected Mortality Hazard after Retirement vs Expected Hazard before Retirement, by Model, with Plotted Equal Values Line

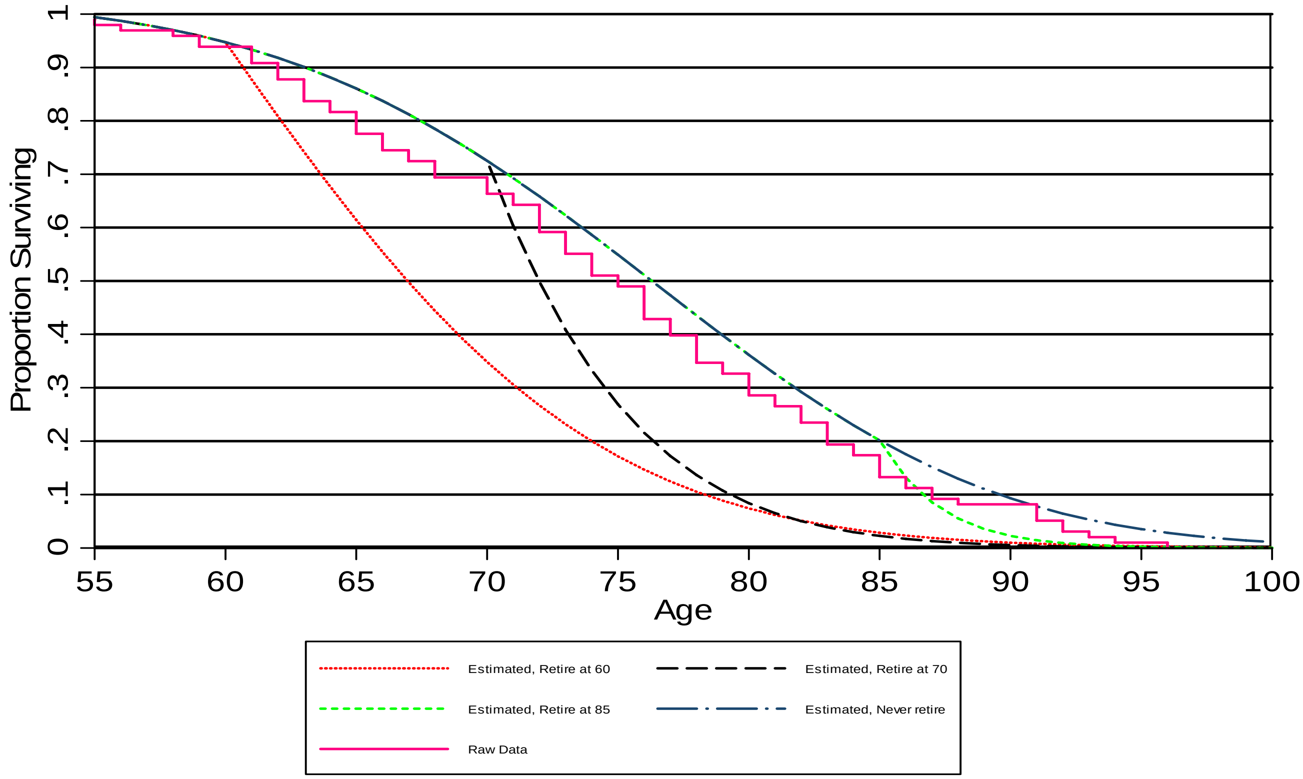


Figure 5b – Raw and *IV* Probit Estimated Survival Functions for Justices who are Alive at Age 55, by age and Retirement Age, with Controls for Calendar Year and Tenure on the Court

Notes: Tenure increases annually until retirement, then is fixed. Estimates assume approximate mean calendar year for the data, 1911, and age of 50 years upon taking oath of office (5 years tenure at age 55). Age is age on first day of year.

C.3. Analyses of Pension Qualifiers Only. Here, for additional robustness of identifying assumptions, I limit analyses to justices who become *pension-eligible* before departure from the Court. In Table 5, I re-estimate the coefficients of *retired* in that limited sample for analyses 3b, 3c and 4. These coefficients are approximately equal to, and statistically significant at lower α levels than the corresponding coefficients in Table 4. Further, 95 percent confidence intervals around each coefficient in Table 5 overlap the point estimates for the same quantities in Table 4. In short, findings in Table 4 and Table 5 lead to the same conclusions and do not differ meaningfully.

C.4. Analyses that Censor Early Retirement Years. In the "simple model" described above, justices simply delay retirement until they perceive that that their health is so poor that death is imminent, at which point they resign. If the "simple model" were correct, then estimates of retirement effects on mortality would weaken or disappear if analyses ignored deaths in the first year or two after retirement. So, for additional robustness, I estimate two additional ordinary probit and IV probit analyses of mortality hazard. The first additional analysis censors data from the first year after retirement. The second analysis censors data from the first two years after retirement. Results are shown in the text table below. In brief, censoring the first, or the first two years of retirement would lead to conclusions that are identical to those drawn from analyses that do not censor the first one or two years of retirement.

Comparison of Effect of Censoring the First and Second Years of Retirement on Coefficient of Retired in Ordinary Probit and IV Probit

	All Years		Retired Year 1 deleted		Retired Years 1 & 2 deleted	
	Probit	IV Probit	Probit	IV Probit	Probit	IV Probit
Retired	.4963	.7538	.5453	.6916	.5728	.6859
(Z)	(3.49)	(3.48)	(3.58)	(3.28)	(3.50)	(3.30)
n	2132		2092		2055	

D. Endogenous Switching Analysis. In Model 6, retirement effects on mortality hazard are measured with separate selection-corrected probit analyses for retired and incumbent justices. For each justice-year in Model 6, I use observed values of independent variables and estimated parameters from the "retired" equation to calculate the expected mortality hazard if the relevant justice were retired in the corresponding justice-year. Separately, I 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 .0433. *If retirement prevailed in all justice-years, then mean expected annual hazard would be larger by about one-third, .0567, all else equal.*[14] For parsimony, I 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. Lines in Figure 6 are obtained by fractional polynomial regression of this ratio on *age,* fitted separately for four half-century

[14] *1/3≈31 percent=(.0567353*-.0433389*)/* .0433389. Estimates 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, and the mean if incumbent is .0380390.

historical periods. In all periods, the ratio declines with increasing age until about age 70, and then increases. 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.

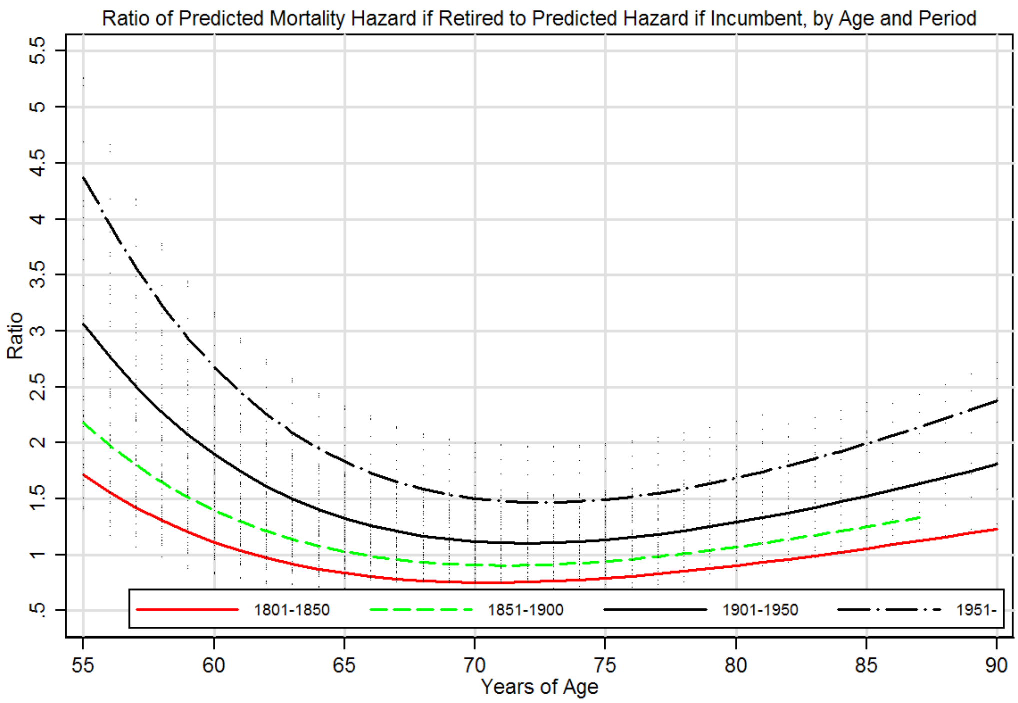


Figure 6 – Ratio of Predicted Mortality Hazard if Retired to Predicted Hazard if Incumbent, by Age and Period

V. DISCUSSION

This paper considers the hypothesis that labor force retirement diminishes mortality-based measures of the health of U.S. Supreme Court justices. Because Supreme Court justices have Constitutionally guaranteed freedom to keep their positions as long as they choose, Supreme Court data are unusually well-protected against commonplace confounding of voluntary retirement with unemployment. In addition, since 1869, Supreme Court pensions equal Supreme Court salaries, obviating purely financial explanations of retirement effects. Analyses here use ***IV*** and ***ES*** estimation to accommodate the endogeneity of retirement, and various probit and Tobit methods to deal with logical constraints on estimates of times and probabilities. Permutation of these models, methods and dependent variables provides 12 different tests of the hypothesis that labor force retirement accelerates death. To investigate various possible alternative explanations of findings, additional analyses examine retirement effects after deletion of a) justices who never qualify for pensions, b) female justices, or c) justice-years pertaining to the first one or two years of retirement. Although standard errors are large, as is usual in IV and ES estimation, all tests are statistically significant ($\alpha \leq .05$, 1-tailed, robust standard error corrected for clustering) and *inconsistent with the hypothesis that retirement prolongs life*. In particular:

1. 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, the current remaining life expectancy of 65-year-old Americans is 18.7 years (U.S. Census Bureau 2008: Table 101).

2. ***ES*** analyses estimate an average annual mortality hazard of 4.3 percent for justices if incumbent, and about one-third higher (5.7 percent) if retired. The ***ES*** hazard analysis implies that if mortality hazards were constant, then justices would to live 5.5 years longer if incumbent than if retired, on average.[15] This difference is roughly the same as the difference found in the ***ES*** analyses of years-left-alive.

3. Figure 6 indicates that after 1955, ***ES*** analyses estimate that retirement would have increased mortality hazard for every justice in every year. From 1901 until 1955, retirement would have increased mortality hazard, on average, but retirement would have reduced mortality hazard for some justices in some years. From 1851 to 1900, retirement would have reduced average annual mortality hazard for justices between the ages of 67 and 76. And from 1801 to 1850, retirement would have reduced average annual mortality hazards for justices between the ages of 63 and 82.

4. In endogenous mediation models *without correction for endogeneity bias*, I estimate that if incumbent justices had mortality hazard of 5 percent in a particular year, they would have an average

[15] Based on the geometric distribution. If mortality 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

hazard of more than 12 percent if they were retired in that year, other things equal. In models *with **IV** correction for endogeneity bias*, if justices had a 5 percent mortality hazard if incumbent, then retirement would raise that average hazard to more than 18 percent.[16]

In short, results are consistent with the hypothesis that, on average, voluntary retirement substantially accelerates death of Supreme Court justices. If justices do indeed lengthen their lives by working well into old age, then perhaps they are quite rational to eschew retirement, even if it brings a generous pension.

VI. A CAUTIOUS CONCLUSION

Have these analyses neglected some exogenous variable that both causes retirement and accelerates mortality? I hope not, but neglected variables are always possible. For example, readers have suggested Presidential political party as possible confounding omitted variables. Stolzenberg and Lindgren (forthcoming) do use similar data and methods to consider the effect of Presidential political party on the timing of retirement from the Supreme Court, but, regardless of the effect of Presidential

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[16] For comparison, 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. That smoking effect is about midway between our instrumented and uninstrumented probit estimates of the effect of retirement on one-year mortality hazard. So, even the smallest of the ***IV*** point estimates of hazard effects can be characterized as comparable to the one-year mortality hazard effects of heavy smoking. Of course, length of exposure matters too: smoking typically starts much earlier than retirement, so lifetime effects of smoking would be much greater than lifetime effects of retirement, even if the annual hazard rate effects of smoking and retirement were identical. 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.

party on resignations, I am aware of no suggestion anywhere that the political party of the U.S. President could directly affect the mortality hazard of individual Supreme Court justices.[17] Family caregiving responsibilities is also cited as a possible omitted variable, because other research reports that workers sometimes retire to care for sick or disabled relatives (Raymo and Sweeney 2006) and care-giving is found to reduce the health of care-givers (Reinhard and Horwitz 1995). But the care-givers described in that research are mostly middle aged women with modest financial resources, demographically and economically dissimilar to the mostly-male, highly educated, and well-paid individuals who are the past and present justices of the Supreme Court.[18] So it seems improbable, at best, that for Supreme Court

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[17] A reader asks for this paper "to convince readers why voluntary retirement is a rational decision at all." However, this paper is an analysis of an *effect of retirement on mortality*, not an examination of the causes of retirement. Using similar data, Stolzenberg and Lindgren (2010) examine determinants of retirement and death in office by Supreme Court justices.

[18] Available information indicates that only one Supreme Court justice has cited care-giving as a reason for retirement (Sandra Day O'Connor). But Justice O'Connor has placed her husband in a nursing facility, where he is attended by professional caregivers (Zernike 2007). Further, it is impossible to know the correspondence between O'Connor's actual and stated reasons for retirement, and it is difficult to know what effect, if any, her caregiving responsibilities might have on her mortality, as she remains alive at this writing. In addition, Supreme Court justices are well-paid, so there is reason to believe that they could buy caregiver services in place of their own labor, as Justice O'Connor has done. And, further yet, evidence suggests that Supreme Court justices differ markedly from most contemporary caregivers: According to a 1997 report (National Alliance for Caregiving and AARP 1997: p. 10), caregivers are disproportionately female (74 percent), less than 50 years of age (64 percent), have low household income (median $35000), and lacking professional or graduate education (91 percent). In short, almost everything that is known about caregiving effects on caregivers applies to a population segment that is very dissimilar to Supreme Court justices. Finally, the paper now reports that analyses

justices, the need to perform unpaid family health care work explains the observed association between justices' mortality hazard and retirement. In brief, political circumstances and family health care responsibilities do not seem to be omitted variables that rob our analyses of internal validity.

And there are the usual questions of external validity. Supreme Court data appear to solve important measurement problems that afflict most other retirement and mortality data, but small numbers always require caution. Further, Supreme Court justices are not average labor force participants. Nor is work at the Court comparable to work at construction sites, high schools, coal mines, or grocery stores, to name just a few places where people work. These and other limitations should be considered seriously. At best, our findings suggest general patterns in certain other population segments. For example, Supreme Court justices may resemble others characterized by very high achievement, who hold jobs with high employment security, high job autonomy, pleasant working conditions, low work-related physical demands, and high levels of work satisfaction. Although unusual, such persons are a socially and economically important segment of the working population. Studies suggest that such workers tend to retire later from the labor force than those who are not so characterized (see Raymo, Warren, Sweeney, Hauser, and Ho 2008; Raymo and Sweeney 2006; Hayward, Grady, Hardy and Sommers 1989). It seems reasonable to hypothesize that these talented workers who like to work at their very good jobs may well react to retirement in much the same way as Supreme Court justices. Obviously, more data would be needed before generalizing to larger population segments.

As we await that data, our results are added to analyses of other, sometimes larger, portions of the population that find negative effects of retirement on subsequent health and longevity. Much of what is known about work and employment effects on health and longevity has been discovered or tested on seemingly unusual population subgroups, including civil servants in England (Stansfeld et al 1995), residents of Alameda County, California (Camacho et al 1991), and the Wisconsin high school

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were repeated after omitting Justice O'Connor and Justice Ginsburg, with no consequent change in findings and virtually no change in estimates.

graduating class of 1957 (Marks and Shinberg 1997). Finally, Preston (1977:171) aptly notes that low mortality among "elderly leadership groups such as union leaders, Supreme Court justices, and Communist Party officials" contributes to their grip on power, thereby making their longevity more consequential than their small numbers might suggest. How interesting, then, that analyses reported here suggest that, at least for U.S Supreme Court justices, the tenacious grip on power seems to contribute to longevity, even as longevity prolongs their hold on high office.

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

Variable	N Justice-Years	Mean	Standard Deviation	Minimum	Maximum
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 2b- Historical Descriptive statistics: Characteristics of Former Supreme Court Justices in their Final Year on the Court, by Vitality at End of Service, 1789-2006

			Former Justices			Current
Variable Name and Brief Description	**Statistic**	**All Justices**	**All Former**	**Died in Office**	**Retired**	**Justices 2006**
n except Future Longevity	n	110	101	49	52	9
Year	Mean	1909.5	1900.9	1880.2	1920.4	2006
(Calendar Year)	S.D.	62.82	58.21	48.18	60.51	0
	Min	1793	1793	1798	1793	2006
	Max	2006	2006	2005	2006	2006
Age	Mean	69.58	69.85	68.82	70.83	66.56
(Justice's Age in Years)	S.D.	9.691	9.631	8.381	10.67	10.44
	Min	48	48	48	48	51
	Max	91	91	87	91	86
Tenure	Mean	16.15	16.34	16.67	16.02	14
(Justice's years of service	S.D.	9.84	9.896	9.831	10.04	9.46
on the Supreme Court)	Min	0	0	2	0	0
	Max	36	36	34	36	31
Pension qualified	Mean	0.5182	0.5050	0.3061	0.6923	0.6667
(dummy)	S.D.	0.502	0.5025	0.4657	0.466	0.5000
Future Longevity	n	100	100	49	51	n/a
(Years between current	Mean	4.35	4.35	0	8.529	n/a
year and year of	S.D.	7.258	7.258	0	8.242	n/a
Justice's death; n/a for	Min	0	0	0	0	n/a
justices now alive)	Max	34	34	0	34	n/a

Notes: At this time (June, 2006), except for Sandra Day O'Connor, all former justices of the Supreme Court are deceased. Two individuals were appointed to the Supreme Court two times: John Rutledge retired in 1791 and 1795. Charles Evans Hughes retired in 1916 and 1941.

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 (Alternate Metric Effect Measure)**	**Effect Estimate (Alternate Metric Effect)**	**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 mean difference between the predicted value if incumbent and the predicted value if retired, of *years-left-alive* or *mortality hazard*, . See Appendix 2 for complete results for all 12 analyses. 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 failed 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. **na** not available; computations failed to converge. * Statistically significant, 1-tailed test, 5%. ** Statistically significant, 1-tailed test, 2%. ***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.

Table 5 – Coefficient of *Retired* from *IV* Models Estimated from only Justices who became *Pension-Eligible*

Model	Analysis Method	N justices	N justice-years	Coefficient of *Retired*	Cluster Corrected Robust Standard Error
3b	Instrumental variables regression *with probit transform*	50	1269	-1.089993	.378642***
3c	Instrumental variables Tobit	50	1269	-10.4401	4.832908**
4	Instrumental variables probit	57	1411	.9069707	.2824573***

Notes: This table reproduces the analyses shown in corresponding lines of Table 4, after excluding justices who resigned or died in office before qualifying for pension benefits. See notes for Table 4. Standard errors are robust and cluster-corrected ** Statistically significant, 1-tailed test, 2%. ***Statistically significant, 1-tailed test, 1%.

Appendix A1 – Simulation Analyses to Address the Question, Are IV Estimates of the Effect of an Endogenous Regressor Necessarily Weaker than Reduced Form Estimates?

A skeptical reader proposes that the estimated effect of *retired* on mortality must be weaker when estimated with IV analyses than when estimated in analyses that ignore endogeneity. This appendix uses simulation to examine that concern by (see also general treatments that permit an entirely mathematical approach, e.g. Amemiya 1985 and Wooldridge 2001). In a related matter, Lochner and Moretti (2004) show mathematically that statistical significance levels for IV estimates can be higher than significance levels for reduced form estimates.

The model represented in figure A1 is used to construct the simulation data. L is the observed dependent variable of interest, analogous to longevity. R is an observed endogenous regressor, analogous to retirement. Z is an observed exogenous variable that affects R directly, analogous to *qualified for pension* in our Supreme Court analyses. Z affects L only indirectly, through its effect on R, thus making Z suitable for use as an instrument for R. εRR and ε LL are exogenous disturbances. S is an exogenous variable that is unobservable by the hypothetical data analyst. Thus, S is relegated to treatment as a component of disturbances εR and εL, and thereby generates a correlation between them. The analysis depicted in Figure A2 ignores the endogeneity of R. The effect of R on L in Figure A2 is estimated by ordinary least squares (OLS); Figure A2 is underidentified, and OLS estimates of its parameters are biased. The analysis in Figure A3 uses Z as an instrument for R. Figure A3 is just identified and effects are estimated by IV regression. The model is specified mathematically as follows:

(A1) $\varepsilon RR \sim N_{0,1}$
(A2) $\varepsilon LL \sim N_{0,1}$
(A3) $S \sim N_{0,1}$
(A4) $Z \sim N_{0,1}$
(A5) $\Sigma = I_4$ where I_4 is the identity matrix of order 4 and Σ is the variance-covariance matrix of εRR, εLL, S, Z.
(A6) $\varepsilon R = \varepsilon RR + 0.5\Delta S$ where Δ is a coefficient equal to 1 or -1.
(A7) $\varepsilon L = \varepsilon LL + \Omega S$ where Ω is a coefficient equal to 1 or -1.
(A8) $R = Z + \varepsilon R$

(A9) $L = \Phi R + \varepsilon L$ where Φ is a coefficient equal to 1 or -1.

Equation (A10) is used to estimate by OLS the model of L shown in Figure A2.

(A10) $\hat{L}_{ols} = \beta_0 + \beta_1 R$

Equations (A11) and (A12) are used to estimate by IV regression the model shown in Figure A3.

(A11) $\hat{R} = \pi_0 + \pi_1 Z$

(A12) $\hat{L}_{IV} = \alpha_0 + \alpha_2 \hat{R}$

I estimate simulations under eight conditions corresponding to values of -1 or +1 for each coefficient Δ, Ω, and Φ. Using Stata 9.2 (with a random number seed X075bcd151f123bb5159a55e5002286574 6ad), I generate 100,000 replicates of εRR , ε LL , S, and Z, and then use OLS and IV (two stage least squares) regression to fit the models indicated in Figures A1 and A2. Table A1 shows results. Although the skeptical reader proposed that the IV estimates would be weaker than the OLS estimates of Φ, results show that in half the simulations, the absolute size of the OLS estimate is smaller than the absolute size of the IV estimator.

Table A1 – Simulation Results, N=100,000 Each

Design Number	Design Parameter			Estimate of Φ		Larger Absolute Estimate
	Δ	Ω	Φ	OLS	IV	
1	-1	-1	1	1.23	1.01	OLS
2	-1	-1	-1	-0.77	-0.99	IV
3	-1	1	1	0.78	1.00	IV
4	-1	1	-1	-1.22	-1.00	OLS
5	1	1	1	1.23	1.00	OLS
6	1	1	-1	-0.77	-1.00	IV
7	1	-1	1	0.78	1.01	IV
8	1	-1	-1	-1.22	-0.99	OLS

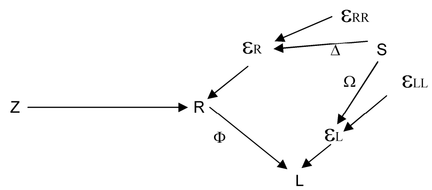

Figure A1 -- True Structure of Simulated Data
(R and L are observed variables of substantive interest; Z is observed but not of substantive interest; ε_R, ε_{RR}, ε_L, ε_{LL} and S are disturbances; Δ, Ω, and Φ are coefficients.)

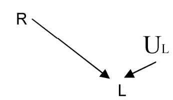

Figure A2 -- Mis-Specified Model of Simulated Data
(U_L is estimation error; S is unobserved, Z is ignored; model is not identified; estimated by OLS)

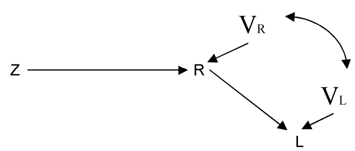

Figure A3 -- "Observed" Simulated Data
(V_L and V_R are estimation errors; S is unobserved; Z is ignored; model is identified; estimated by IV Regression)

APPENDIX 2

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 – 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. 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(like-lihood	-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. 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.