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Which Households Benefit from Delayed Claiming?

Abstract

Although Social Security benefits are gender neutral, survivor benefits may lead to opposite claiming-age incentives for many husbands and wives. The typical wife's survivor benefit rises with the age at which her husband claims his retired-worker benefit. Yet, it does not depend on her retired-worker claiming age, which should lead her to claim her own benefit earlier. We find that spousal claiming age decisions tend to increase lifetime benefits of households. Married men claim later than single men, controlling for lifetime earnings; and married men with younger wives claim even later, though few delay as long as they should to maximize expected lifetime benefits. Married women, meanwhile, claim earlier than single women. Next, we find that married men have substantially lower mortality than single men, controlling for lifetime earnings and claiming age, further increasing the gain from delaying claiming. Incorporating these differences into household benefit calculations, we find that the return to delayed claiming of the husband's retired-worker benefit is substantially more than actuarially fair, but for different reasons by household type. For disadvantaged households, the return to delay by the husband arises more from gains to the survivor benefit than to his retired-worker benefit. For advantaged households, the return to delay arises largely from the gains to husbands' retired-worker benefits. Thus, adverse selection in claiming ages by high-earning men raises costs to the Old-Age and Survivors Insurance Trust Fund. In contrast, lower-earning men, who claim relatively early, on average, forgo an important gain from delay in the form of higher survivor benefits for their wives.

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1. Introduction

Social Security benefit payments to widows and widowers cost $112 billion a year and amount to 10.3% of benefit payments.[1] Survivor benefits are payable to a deceased worker's spouse if that amount exceeds the surviving spouse's own retired-worker benefit. Moreover, survivor benefits depend on the age at which the deceased spouse claimed retired-worker benefits, *but not* the age at which the survivor initially claimed as a retired-worker. These features create complicated and largely unresearched interactions of claiming-age incentives for both spouses, many years before the death of either spouse may be expected. The incentives of either spouse to claim early or late depend on both spouses' relative ages and life expectancy. This stands in contrast to the simple incentive faced when ignoring survivor benefits: The higher one's life expectancy, the later one should claim.

Consider the interesting, and common, case in which spouses within a couple face opposite claiming-age incentives. This arises when a lower-earning wife claims as a retired worker but expects to receive survivor benefits later in life, both because she is likely to outlive her husband, which is typical because most wives are younger than their husbands and have longer life expectancy, and she has lower, but not drastically lower, lifetime earnings than he does.[2] For 39.5% of married couples in the merged survey-

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[1] OASDI Trustees Report Table III.A5, Calendar year 2022 (U.S. Social Security Administration 2023).

[2] Although survivor benefits are gender neutral, for ease of exposition we will refer to the claimant of survivor benefits using female designations and her spouse using male designations. Moreover, we will impose additional heteronormativity and gender bias by

administrative data set that we analyze (who were born or whose spouses were born between 1930 to 1950), lifetime earnings are lower for the wife than the husband, but not so much lower that wives would immediately claim benefits as a dependent spouse rather than a retired worker. Moreover, the share of couples in this category has risen as earnings differences between spouses have declined for recent cohorts.[3] For such couples, because the amount of the wife's survivor benefit depends on her husband's claiming age it *lengthens* the effective time horizon for his retired-worker claiming decision, which should lead to later claiming by husbands if the goal is to increase expected household benefits. And, because her survivor benefit does not depend on her claiming age for retired-worker benefits, it *shortens* the effective time horizon for her own claiming-age decision, which should lead to earlier claiming by wives. These divergent incentives within the household arise even though Social Security benefits are gender neutral.

We analyze claiming ages of spouses and the impact on lifetime benefit payouts *at the household level*. We use administrative data from the U.S. Social Security Administration (SSA) merged with the Current Population Survey (CPS).[4] We estimate

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describing the higher-earning spouse as the husband and the lower-earning spouse as the wife, reflecting typical earnings histories in the birth cohorts we study.

[3] In other words, these are households in which the wife's average lifetime earnings are between 50% to 100% of the husband's, so she would claim benefits as a retired worker rather than a dependent spouse while married and as a survivor should she outlive her husband. The statistic reported here is for couples in which at least one spouse was born between 1930 to 1950; whose initial claim was not for Disability Insurance, spousal, or survivor benefits; and who worked at least 40 quarters.

[4] The CPS allows us to observe marital status for cohorts nearing retirement, while the administrative data allows us to observe their benefit eligibility, subsequent claiming-age

whether husbands with the greatest incentive to claim late and wives with the greatest incentive to claim early do so, which indicates whether spouses may act to increase the expected present value of household benefits.[5] We then incorporate couples' behavior into our previous analysis (Dushi et al. 2021), in which we demonstrated that the return to delayed claiming is more than actuarially fair for men who actually delay. Here, we consider whether adverse selection similarly arises among married couples, and the impact on the Old-Age and Survivors Insurance (OASI) Trust Fund and on inequity in benefit payouts. Extending the analysis to couples is important for understanding distributional consequences of Social Security, both because marriage is unequally distributed in the population (with high earners more likely to have long marriages that generate survivor benefits) and because married couples live longer on average than singles.

We find that men and women claim in ways that increase but do not maximize the expected present value of lifetime benefits, raising Social Security outlays but not to the extreme. Married men claim later than single men, controlling for lifetime earnings; and married men with younger wives, who will likely spend more years as a surviving spouse, claim even later. Married women claim earlier than single women, particularly if their PIA is small relative to that of their husbands. For plausible wealth levels and

decisions, and for older cohorts, mortality. We discuss further strengths and weaknesses of available data later.

[5] For risk-averse households with uncertain life spans, delaying claiming not only increases the value of annual benefits but effectively results in additional purchase of longevity insurance. Thus, the higher earner in expected utility maximizing households should delay claiming even beyond the age at which the expected present value of lifetime benefits is maximized (Sun and Webb 2011).

preference parameters (Sun and Webb 2011), expected household utility is maximized if lower-earning married women claim at age 62, which is what we observe. However, while higher-earning married men should delay until 68, few are observed to do this.

Next, we find that married men have significantly and substantially lower mortality than single men, even after controlling for claiming age and lifetime-earnings quartile. The effect of their lower mortality is to further increase the return on delayed claiming of retired-worker benefits by married men, and to redistribute Social Security wealth away from single individuals to married couples, already a more financially secure group on average. Holding female mortality constant, the lower mortality of married men will decrease the value of survivor benefits because wives will spend fewer years as a surviving spouse. But, we similarly find that married women have lower mortality than single women, partially offsetting this effect. We find, further, that men with higher lifetime earnings have substantially lower mortality, including within the group of married men.

When we incorporate these differences in claiming and mortality at the household level, we find that the return to delayed claiming of the husband's retired-worker benefit is substantially more than actuarially fair, but for different reasons for different types of households. Consider households disadvantaged by a combination of socioeconomic conditions and high male mortality, where the husband is in the lowest quartile of lifetime earnings among men and both spouses claim at age 62. For such households, the return to delay by the husband arises more from the gains accruing to the survivor benefit than to the husband's retired-worker benefit. In contrast, for households in which the husband has the low average mortality of those who are in the highest quartile of lifetime earnings and who claim at age 66, the return to delay arises largely from the gains to the husband's retired-worker benefit. For this latter group, the gain accruing to the survivor benefit from the husband's delay is quite small in present value, given his much lower mortality. We calculate these results for the 1939 to 1940 birth cohort, as later cohorts do not have a sufficiently long mortality history to make definitive projections. Nevertheless, it is likely that the sources of the return to delay will continue to diverge for subsequent birth cohorts as socioeconomic mortality differentials continue to widen (Sanzenbacher et al. 2019).[6]

These patterns deepen the insights that we gained in our earlier analysis in Dushi et al. (2021), which treated all men as single. There, we found evidence of adverse selection which particularly benefits high earners: As mortality and claiming-age patterns have diverged, the return to delayed claiming has become more than actuarially fair by a great deal for late claimers in high lifetime earnings quartiles but by only a little for early claimers in low lifetime earnings quartiles. Here, when we incorporate the availability of the survivor benefit in married households, it is apparent that the results are more nuanced. Men in the highest quartile of lifetime earnings, who claim later on average, raise costs to the OASI Trust Fund through higher retired-worker benefit payouts, while men in the lowest quartile of lifetime earnings, who claim relatively early on average, forgo an important gain to delay in the form of higher survivor benefits that would accrue to their wives.

[6] Moreover, various Social Security rules, including the return on delayed claiming at particular ages, have changed for more recent cohorts. Nevertheless, such rule changes do not alter the nature of the marriage-related, claiming-age incentives that we evaluate.

Our contributions extend the literature on claiming-age patterns and benefit equity. Coile et al. (2002) find that married men claim earlier, contrary to our finding; however, they focus on 1980 to 1981 claimants, when most claiming occurred at age 62. Our analysis is motivated by the observation that claiming ages have become much more dispersed for the recent cohorts that we consider. We also draw attention to another factor that affects the coordination of retirement timing among spouses, in addition to employer pension incentives (Coile 2004), health insurance (Blau and Gilleskie 2006), and complementary leisure (Gustman and Steinmeier 2000). Lastly, we extend recent research on the positive correlation between claim age and socioeconomic status (Coile et al. 2002; Hurd et al. 2004; Sass et al. 2013; Armour and Knapp 2021; Dushi et al. 2021), illustrating another demographic factor — marital status — that exacerbates these sources of inequality. And, similar to Liebman (2002), we illustrate another way in which features of Social Security benefit formulas intended to protect those facing vulnerabilities (in this case, widows) may undermine other redistribution goals within Social Security.

2. Social Security claiming incentives

The U.S. Social Security system provides three main types of retirement benefits based on covered employment:

1. a retired-worker benefit, based on an individual's lifetime earnings,
2. a spousal benefit payable to spouses of retired workers if that benefit exceeds the spouse's own retired-worker benefit, and
3. a survivor benefit payable to surviving spouses of retired workers if that benefit exceeds the surviving spouse's own retired-worker benefit.

The interactions among benefit types and claiming-age decisions for spouses who expect to receive survivor benefits in the future alter household lifetime benefits and claiming-age incentives for both spouses.[7]

2.1 Survivor benefits

Most women in the recently retiring cohorts that we study earned less than their husbands over their lifetime, are younger than their husbands, and have longer life expectancy. Therefore, they can expect to 1) outlive their husband, 2) receive a retired-worker benefit based on their own earnings for as long as their husband is alive, and 3) receive a survivor benefit after the death of their husband. The increasing prevalence of "dual entitlement" reflects age, longevity, and earnings differences between spouses in recent cohorts, along with program rules.

Wives are likely to outlive their husbands because of both age differences of spouses and longevity differences between men and women. The average age difference of spouses in our sample is about three years, while the average years of widowhood is almost five years for wives in the most advantaged group that we consider in our sample (those with husbands in the highest earnings quartile who claim at age 66) and reaches almost eight years for wives in the least advantaged group that we consider (those with husbands in the lowest earnings quartile who claim at age 62).

Wives are likely to qualify for a retired-worker benefit, rather than a spousal benefit, when they first claim, assuming they are not yet widows. This is because their

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[7] Because we seek to provide a relatively simple explanation of benefit determination and resulting incentives, we relegate some details and special cases to the footnotes. We incorporate all these program rules in our modeling with some minor, noted exceptions.

average lifetime earnings are lower, but not much lower, than their husbands.' The spousal benefit is at most (depending on spousal claim age) 50% of the higher-earner's retired-worker benefit that is payable were he to claim at his full retirement age (FRA). It only exceeds the lower earner's retired-worker benefit if lifetime earnings are quite disparate. This most commonly occurs when the lower earner has a shorter work history. Because retired-worker benefits are a function of *average* lifetime earnings, several years of nonemployment will lower the average, since it is computed over the highest 35 earning years. Yet, even if average lifetime earnings of the lower earner are only, say, 50% of the higher earner's, the lower earner will still get higher benefits claiming as a retired-worker than claiming as a spouse. This is because a progressive formula transforms average lifetime earnings into the annual retired-worker benefit, so the resulting retired-worker benefit for the spouse exceeds 50% of the higher earner's benefit. The lower-earning spouse must earn substantially less than 50% of average lifetime earnings of the higher-earning spouse to get spousal benefits.

Yet it is also likely the case, because their lifetime earnings are likely to be lower, that wives shift from a retired-worker to a survivor benefit if their husband dies first. The survivor benefit equals her husband's retired-worker benefit (subject to adjustments based on survivor benefit claiming age), which exceeds her own in most cases.[8]

As we show below, within our sample, for 80% of couples who are married at ages 50 to 61 in cohorts born between 1930 and 1950, the husband has higher average

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[8] For widows claiming after their FRA, the survivor benefit is subject to a minimum of 82.5% of the deceased spouse's primary insurance amount (PIA), placing a floor under the survivor benefit of widows whose spouses claimed early (Weaver 2001).

lifetime earnings than the wife. Therefore her retired-worker benefit is less than her husband's and, if he predeceases her, she will normally receive survivor benefits.[9] Yet, among those, in 49% — almost half — the wife has a retired-worker benefit that exceeds 50% of her husband's, so she initially claims as a retired worker. The fraction of married women receiving retired-worker benefits has continued to increase in recent cohorts (Karamcheva et al. 2015).

2.2 Claiming-age timing

Retired-worker benefits can be claimed as early as age 62 or as late as age 70, subject to actuarial adjustments in the monthly benefit. Benefits are reduced for early claiming and increased for late claiming (Jivan 2004).[10] However, if a wife shifts to receiving a survivor benefit after her age of full entitlement (FRA), it is adjusted only for *his* early or delayed claiming.[11] This is why delayed claiming by the husband increases

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[9] This is true for the case in which both claim at the same age relative to their FRA and widowhood occurs after the widow has attained her FRA. If, as is often the case, the wife claimed earlier than her husband relative to their respective FRAs, the survivor benefit may exceed the widow's retired-worker benefit even if her earnings exceeded those of her husband.

[10] The same adjustment is made when one claims a spousal benefit before the FRA, and adjustments based on the husband's retired worker benefit claiming age and the wife's survivor benefit claiming age are made when the wife claims a survivor benefit before the FRA (though neither case is the focus of our analysis). A widow is eligible for a survivor benefit at age 60 (earlier in certain circumstances). If the survivor benefit is claimed prior to the widow's FRA, it is subject to a reduction based on the number of months before the FRA that the benefit was claimed. The monthly amount of the reduction varies by birth cohort along with the FRA.

[11] If both husband and wife claim retired-worker benefits early, the widow receives the lesser of the husband's retired worker benefit, subject to the 82.5% floor, and the husband's PIA multiplied by the reduction for early claiming of widow's benefit.

not only his retired-worker benefit but also his wife's potential survivor benefit; while in contrast, delayed claiming by the lower-earning wife of a retired-worker benefit does not increase benefits over her full remaining expected life span, since her time horizon is shortened to her husband's life expectancy if she expects to outlive him.[12]

The actuarial adjustments (as we term both the Actuarial Reduction Factor which applies before the FRA and the Delayed Retirement Credit which applies after) now range between 7% to 8% per year of delay, while earlier cohorts in our sample were subject to both a smaller benefit reduction for early claiming and smaller return to delay for late claiming.[13]

2.3 Claiming-age hypotheses

In sum, claiming-age incentives differ in two ways for couples in which the wife claims retired-worker benefits initially and expects to claim survivor benefits later, compared to other couples. Because it increases his survivor's benefits as well as his own, delayed claiming by the husband has a greater return at the household level. And,

[12] For individuals born after January 1, 1962, the Bipartisan Budget Act of 2015 eliminated the option for a husband who reaches the FRA to elect to claim and immediately suspend benefits, allowing his wife to receive a spousal benefit based on his earnings record. We do not incorporate the "claim-and-suspend" strategy in our analysis, as it was seldom used in the cohorts we consider.

[13] For workers with a FRA of 65 (born before 1943), the actuarial reduction factor (ARF) reduces benefits by 20% for those claiming at age 62; for workers with a FRA of 66 (those born in 1943 to 1954), the ARF reduces benefits by a 25%. Each month of delay from age 62 to the FRA raises benefits fractionally. https://www.ssa.gov/oact/quickcalc/earlyretire.html. The Delayed Retirement Credit (DRC) is the increase in benefits for those who delayed claiming past their FRA. For workers born prior to 1917, the DRC was small, only 1% per year. The DRC increased across succeeding birth cohorts, reaching 8% per year for those born 1943 or later. https://www.ssa.gov/OP_Home/cfr20/404/404-0313.htm.

because it only increases retired-worker benefits for the duration of her husband's expected lifetime, delayed claiming by the wife will have a smaller return.

These effects arise for households in which the wife begins claiming as a retired-worker and would switch to claiming as a survivor, so we focus in particular on households in which the wife's retired-worker entitlement (her primary insurance amount — PIA) is between 50% to 100% of her husband's PIA. Formally, we test whether the wife claims earlier and the husband later when her PIA is between 50% to 100% of his, compared to other men and women who are either unmarried or have relative PIAs in other ranges. The effects we are interested in also arise for households in which the wife expects to outlive the husband, so we test whether the wife claims earlier and the husband later when the husband is some years older than the wife (we focus on an age difference of at least three years), compared to men and women who are either unmarried or have a smaller age difference. In fact, higher-earning women face the same claiming incentives as higher-earning men, conditioning on age — which we also explore.

Increases in the Delayed Retirement Credit have raised the return to delay for the cohorts we study, thus increasing the incentive to delay for married men whose wives expect a survivor's benefit, since their wives can benefit even more, while increasing it less for single men and for married men whose wives do not expect a survivor's benefit. Meanwhile, the same increases have had a greater effect on incentives to delay claiming retired-worker benefits for women who do not expect a survivor's benefit than for women who do, since the time horizon is shorter for the latter group.

Lastly, later cohorts' increases in life expectancy blunt the incentive to claim early for married women who expect survivor benefits, while inducing complicated incentives for married men, depending on the correlation in life expectancy within couples. As positive assortative matching by education and earnings has increased, we hypothesize that this correlation is positive (which we test) and possibly increasing (which we lack power to test). To the extent that both spouses' life spans have increased but husbands' by more (since longevity increases have slowed faster among women than men), this reduces the length of widowhood, blunting later-claiming incentives for husbands whose wives expect survivor benefits. It is, further, possible that, as longevity has increased, uncertainty about the date of death has decreased, which would reduce the expected length of widowhood, also blunting husbands' incentives to claim later.[14]

2.4 Previous research

Our contributions extend the literature on claiming-age patterns and benefits equity. Earlier studies highlighted the relationship between claiming age and mortality (Wolfe 1983; Waldron 2001, 2002, 2004; Duggan and Soares 2002) and between claiming age and socioeconomic status (Coile et al. 2002; Hurd et al. 2004; Sass et al. 2013; Waldron 2007, 2013, 2020b; Armour and Knapp 2021), with Jaimes (2019) and Dushi et al. (2021) considering both. In light of increasing dispersion in claiming ages

[14] When the standard deviation of longevity falls, then the expected length of widowhood falls as well. If widowhood were unbounded, then reducing the standard deviation of each spouse's longevity would reduce the likelihood of both long negative and long positive spells; but since negative spells of widowhood are not possible (as this corresponds to widowerhood, which is not of major interest in our current research since relatively few husbands receive survivor benefits), then the reduction in the frequency of positive spells reduces the expected length of widowhood.

for recent cohorts (Munnell and Chen 2015), Dushi et al. (2021) demonstrate that the return to delayed claiming has become sharply positive for men who actually delay claiming — both because late claimers have been living longer and have higher lifetime earnings. This conclusion counters earlier work that treated everyone as having average life expectancy (Heiland and Yin 2014). Yet, calculations in those and other earlier papers (including Shoven and Slavov 2014 and Dushi et al. 2021) treated everyone as unmarried, ignoring the ambiguity of claiming-age incentives arising within couples.

Several papers have analyzed claiming ages of women (Waldron 2020a) and the coordination of retirement timing between spouses. We draw attention to the role of Social Security benefit determination, in addition to other factors including employer pension incentives (Coile 2004), health insurance (Blau and Gilleskie 2006), and complementary utility gained from joint retirement (Gustman and Steinmeier 2000). The analysis of claiming-age delays by Coile et al. (2002) considers the role of marital status. While they find a dissimilar result — that married men claim earlier than single men — than ours, they focus on a much earlier cohort, mostly born in 1918 to 1919 and claiming in 1980 to 1981. As we show in Dushi et al. (2021), at that time, a large majority of individuals claimed as early as possible, at age 62. The increasing penalties to early claiming since then, combined with increases in life expectancy, strengthen incentives to time benefit claims more carefully.

Our analysis also fits in with papers that focus on increasing differences in marriage based on socioeconomic status. For example, Chiappori et al. (2017) highlight increasingly positive assortative matching in marriage. Liebman (2002) demonstrated the impact of multiple sources of differences — including in employment, marriage, and mortality — in undermining redistribution within Social Security from high to low earners. Our analysis builds on his by emphasizing continued trends in in these factors as well as the role of claiming-age responses.

3. Research methodology

Our analysis of claiming-age incentives and mortality differences by marital status consists of three steps. First, we estimate the relationship between claiming age and marital status, controlling for lifetime earnings quartile, in order to investigate the hypotheses that we laid out earlier. Second, we estimate a mortality model in order to determine whether adverse selection in claiming ages arises among married couples. Third, we compute the expected present value of household level benefits by marital status and claiming age and the resulting return to delayed claiming. To undertake these steps, we merge administrative data from SSA, in which we observe claiming age and lifetime earnings, with survey data, which reveals marital status for a cross-section of individuals.

3.1 Estimation of claiming-age patterns

To investigate our hypotheses about claiming ages, we estimate ordered probit models of claiming age separately for men and women, as follows:

$$
\begin{aligned}
P(claim\ age_i = j) \\
&= \Phi(\beta_1 marital\ status_i + \beta_2 age\ diff_i + \beta_3 PIA\ diff_i + \beta_4 educ_i \\
&\quad + \beta_5 lifetime\ earnings_i + \beta_6 cohort_i)
\end{aligned}
$$

We specify the left-hand side variable $claim\ age_{ij}$ for individual *i* as taking one of the values $j = \{62,\ 63, 64, 65, 66+\},$.[15] As we noted earlier, the critical right-hand side variables are $marital\ status$, $age\ diff$, $PIA\ diff$, and $cohort$, all of which are related to the individual- and couple-specific return to delay.[16] However, several issues may make it difficult to ascribe a causal interpretation to these variables. Notably, the individual-specific delay depends on variables (such as subjective mortality expectations) that are likely to be correlated with variables that we do observe and seek to control for — such as marital status and age, as well as employment and earnings choices, which determine PIA. For example, a husband who expects to live longer should work longer and claim later, both of which increase lifetime resources. He may be more likely to be married and may have a relatively high-earning wife (given positive assortative matching). It is possible that our nonlinear transformations of key variables (such as dummies for particular values of age and PIA differences, rather than the absolute age and PIA or even age difference and PIA difference) isolate variation arising from Social Security incentives that affect claiming-age decisions directly while reducing the influence of correlations with unobservable variables that also affect

[15] Relatively few individuals claim after age 66, so we do not model later claiming ages separately. In addition to the coefficients on the right-hand side variable in the regression equation, the ordered probit estimates threshold values corresponding to the ordered left-hand side variable. With the left-hand side claiming-age variable taking one of five values, the estimated threshold values are those at which the right-hand side index $X\hat{\beta}$ (where $X\beta$ refers to the right-hand side of the regression equation) passes the threshold value determining the probability of claiming at each age *j* relative to the base claiming age of 62.

[16] Claiming age behavior may also be related to other observable demographic differences, for example in race and ethnicity. In future work, we will undertake a comprehensive examination of how minority status affects claiming ages and marriage-induced redistribution from Social Security.

claiming ages. Lastly, while causal estimates of individual responses to claiming-age incentives are of interest, they are not necessary for our other purpose of determining how current claiming-age patterns affect Social Security payouts. For those calculations, it is helpful to include covariates in our estimation that are correlated with omitted variables of interest.

Because claiming-age patterns, marital status, and age and earnings differences between spouses (the relationships in which we are interested) are likely to vary with socioeconomic status (SES), we add several controls that are available in the CPS or administrative data. We allow claiming age to vary with education (less than high school, high school or some college, and college) and lifetime earnings quartile (specifically the Social Security Average Indexed Monthly Earnings, with the quartiles calculated for men and women separately). Because the return to delayed claiming has changed over time, we control for birth cohort.

We are interested in the impact of marital status and age and earnings differences between spouses on claiming ages. The variable *marital status* is a dummy taking the value of 1 if the individual reports being married or separated in the CPS, and 0 otherwise. The variable *age diff* is a dummy for whether the husband is more than three years older than the wife, since in such couples the husband is relatively more likely to die first. The variables *earnings diff* control for whether the full retired-worker benefit (as measured by the PIA) of the wife is less than 50%, between 50% and 100%, or more than 100% of the husband's full retired-worker benefit, since this determines whether the wife is likely to claim initially as a spouse, initially as a retired worker and then as a survivor if her husband dies, or only as a retired worker.

We estimate the model with differing samples and combinations of controls, separately by gender. For example, we estimate it for the full sample of married and single individuals and omit $PIA\ diff$, and thus compare single individuals with individuals whose wives are of a similar age and with individuals whose spouses are somewhat younger. We then limit the sample to married individuals and explore the impact of PIA differences along with age differences. The goal is to determine what differences in claim ages are robust as we consider these different specifications.

3.2 Estimation of mortality patterns

We are interested broadly in the impact of the structure of survivor benefits on Social Security outlays to married couples and specifically on whether adverse selection in claiming arises for married couples. Although Social Security is not sold in a market, the decision to delay is effectively the purchase of an annuity, and pricing for that purchase is uniform. Risk-based selection occurs when information about mortality is private or is public but not used in pricing.[17] In the current setting, we seek to determine whether men who claim later are risky from the standpoint of Social Security — either because they are long-lived or have a spouse who will receive survivor benefits for a long while (whether because they are relatively young or long-lived).

To do this we need to compute annual survival probabilities for our sample of retiring cohorts, even though not all cohort members have died. Therefore, as in Dushi et al. (2021), we estimate a Gompertz survival model, a parametric demography model

[17] Adverse selection does not require that individuals possess or act on information about their mortality. It can also occur if claiming decisions are influenced by socioeconomic characteristics that are correlated with mortality.

commonly used to analyze mortality. Dushi et al. (2021) showed both that baseline mortality at age 62 varied with birth cohort, claiming age, and lifetime earnings quartile, and that the annual age-related increase in mortality varied with birth cohort and lifetime earnings quartile. That study relied solely on administrative data and was therefore unable to investigate whether mortality also varied with marital status and age difference between husband and wife, characteristics that may also be correlated with claiming age and lifetime earnings quartile. Our linking of administrative with CPS data permits us to estimate a Gompertz model with a richer set of covariates, including both marital status and age difference between husband and wife.

The Gompertz model is a two-parameter distribution with a hazard that takes the form:

$$h(t) = e^{\lambda} e^{\gamma t} \tag{1}$$

The parameter λ incorporates individual characteristics X_{1i} that affect baseline mortality at age 62:

$$\lambda_i = e^{X_{1i}\beta_1} \tag{2}$$

and similarly the parameter γ incorporates individual characteristics X_{2i} that affect the exponential growth rate of mortality at subsequent ages t:

$$\gamma_i = e^{X_{2i}\beta_2}\, t \tag{3}$$

Reflecting the mortality estimation results in Dushi et al. (2021), we estimate a flexible model in which baseline mortality, the annual age-related increase in mortality, and the effects of claim age, lifetime earnings quartile, marital status, age difference between husband and wife, and educational attainment vary by birth cohort. We estimate models separately for men and women, ignoring possible interdependencies between spouses' mortality (Carriere 2000). We find that even with our large sample, some of our coefficients are imprecisely estimated, and we therefore also report pooled results including birth cohort dummies, imposing the constraint that the above coefficients are equal for all birth cohorts. We also estimate models separately on married and single men and married and single women, pooled across birth cohorts, allowing the relationship between mortality and explanatory variables, including the annual age-related increase in mortality, to vary by marital status.

The earliest cohorts we study have largely complete mortality histories. For example, 81.83% of men in the 1929 to 1930 birth cohort who lived to age 62 had died by 2020, the date to which we have mortality data.[18] We do not analyze mortality of the most recent retirees, since few members of their cohort have died. For incomplete cohorts in between those extremes, the model relies on the empirical regularity first reported by Gompertz (1825) and which we also observe in more complete cohorts, that the annual percentage increase in mortality varies little with age, up to quite advanced ages.[19] In effect, the Gompertz model treats a worker who delays claiming or who has other characteristics associated with low mortality as having, at all ages, a biological age that is x years younger than their chronological age. We acknowledge that if there is an upper limit to life expectancy common to both individuals with high and low mortality in their 60s, we might expect a higher age-related increase in mortality among those who have low mortality in their 60s. This restriction may lead us to overestimate

[18] Authors' calculations using SSA cohort life tables.

[19] We curtail our analysis at the 1945 birth cohort and age 72 because mortality at older ages for more recent birth cohorts has yet to be observed.

the life-expectancy of low mortality individuals, especially for incomplete cohorts for which estimates are based solely on mortality at younger ages.[20] Given that benefits payable at very advanced ages are subject to substantial time and mortality discounting, we consider our assumption of exponentially increasing mortality a reasonable approximation.

3.3 Impact of marriage and claiming age on lifetime benefits

Once we estimate how both claiming age and mortality differ by marriage, we use the estimated coefficients to compute expected lifetime benefits by marital status and claiming age, as well as birth cohort and lifetime earnings. We distinguish the impacts on expected lifetime benefits of the four separate components of the household's benefits — the husband's and wife's retired worker benefits, the wife's spousal benefit, and the wife's survivor benefit. Relative to a baseline in which both husband and wife have population average mortality based on Social Security Administration cohort life tables, we then examine the impact of 1) lower mortality of retired-worker benefit claimants due to our exclusion of higher mortality Social Security Disability Insurance (SSDI) recipients, 2) the extent to which lower mortality of high earners and late claimants results in higher lifetime benefits, and finally 3) the further impact of the lower-than-population-average mortality of married men and women.

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[20] We tabulated mortality for participants with selected characteristics and compared the exponential increases in mortality. We observe substantial fluctuations in the rate of increase even after applying a five-year moving average, reflecting both our sample size and natural variations in mortality.

We focus our analysis on two households types: a high-mortality household, with mortality rates of age-62 claimants married to each other and in the lowest AIME quartile, and a low-mortality household, with mortality rates of age-66 plus claimants married to each other and in the highest AIME quartile.[21] The difference between these two groups indicates the extent to which mortality differentials contribute to inequality in lifetime benefits.

In addition to the expected present values of lifetime benefits, we also report the realized gain to delaying claiming based on mortality differences by claiming age, calculated from differences in expected lifetime benefits. We can calculate this by, for example, comparing lifetime benefits for age-66 claimants if they had claimed at age 62 instead. This latter analysis reveals the extent to which returns on delay by both husband and wife differ by household type and whether households that actually delay face more favorable returns on delay than those who opt to claim early.

We assume both mortality and benefit rules for the 1939 to 1940 birth cohort (which turned 65 in (2004/2005), a cohort that is relatively recent within our sample but old enough to have sufficient mortality experience. For both outcomes (expected lifetime benefits and the realized return to delaying claiming), we use as a metric the expected present value (EPV) of lifetime benefits per dollar of annualized PIA (the benefit payable at the FRA), payable at the claiming age and discounted back to age 62. This metric tells us the amount of benefits someone expects over their remaining lifetime at age 62, by virtue of their claiming age, life expectancy, and marital status.

[21] We undertake these calculations on a pooled sample of single and married individuals, because we are imposing a constant γ (mortality growth rate) for each gender in our estimates.

For example, a worker born 1939 to 1940 who was entitled to $1 a year of benefits at their age 65 FRA, but who claimed at age 70, would by virtue of the actuarial adjustments receive $1.25 a year. For that worker, we report the present value at age 62 (conditional on being alive at that age) of an annual benefit of $1.25 starting at age 70, multiplied by annual survival probabilities. For a couple, we assume that the husband and wife are the same age, and that the wife's PIA is one-half of that of her husband, so she is eligible for the retired-worker benefit, not the spousal benefit, upon initially claiming, while she will receive a survivor benefit of at least twice her benefit if she outlives her husband. Therefore, the lower-earning wife receives 50 cents a year of benefits per dollar of his PIA, adjusted for her early or late claiming, when both are alive; the husband receives $1 a year of benefits, adjusted for his early or late claiming, while he is alive; and the surviving widow also receives $1 a year of benefits, adjusted for early or late claiming by the husband of his retired worker benefit and for the wife's age when she was widowed. We assume a 3% real interest rate. Some individuals who delay claiming do not survive to collect benefits, and our expected present values for those who delay is averaged over those who survive to delay and those who we estimate, based on our mortality model, planned to delay but died before claiming.

4. Data

4.1 Current Population Survey

To undertake this analysis, we merge administrative and survey data.[22] We form our sample using a large survey, the Current Population Survey Annual Social and Economic Statistical Supplement (the CPS or CPS-ASEC, also known as the March CPS), in order to observe married couples who have been matched to their Social Security earnings, claiming-age, and mortality information from SSA administrative records. The disadvantage of the CPS is that it is not a panel, so we only observe marital status at a point in time, and marital status may have changed by the time they are making claiming decisions. Therefore, we consider a fairly narrow age range of married couples, who are observed near their claiming age in the CPS and subsequently in the administrative data. Consequently, in order to increase the sample size for our analysis, we use available years of the matched CPS survey from 1991, 1994, and 1996 to 2010.

In each survey year we restrict the analysis sample to respondents born in 1930 through 1950, who were in the 50 to 61 age range at the time of the survey, and who had reached at least age 70 by year 2020 — the latest year that administrative records were available at the time of this analysis. In this way we can observe the full claiming-

[22] Administrative data alone reports incomplete information on marital status, only including this information when one member of the couple has made a spouse or survivor claim. This results in a selected sample; for example, we do not know future eligibility for survivor benefit claims among those who have claimed initially as retired workers. Survey data alone (even in the case of the Health and Retirement Study (HRS)) does not report enough information to compute Social Security benefit entitlements with any accuracy, nor are benefit claiming ages reported accurately. Moreover, the HRS sample size is small for our purposes.

age distribution for these birth cohorts. Among couples, we restrict the sample to those in which at least one spouse was born within this period. We further restrict the sample to those couples for whom we had the spouse's survey information and also matched administrative records. We keep only those in which the respondent's initial claim was for a retired-worker benefit, dropping those who initially claim SSDI, spousal, or survivor benefits.[23] Next, we exclude respondents who were not fully insured (i.e., had not worked at least 40 quarters) for a worker or survivor benefit and those who died before reaching age 62. Among couples, we drop those where both spouses died before reaching age 62. Finally, we restrict the sample to those who claimed retired worker benefits between the ages of 62 to 70 (since earlier claiming is not allowed and later claiming does not raise benefits). Our final sample consists of 142,150 observations, of which 108,269 are husbands and wives in coupled households. Note that for those couples where one of the spouses died before reaching age 62, we do not drop them from the sample but include them in the sample of singles for analysis purposes. While in our discussion we focus mainly on couples, we report estimates for singles as well, and separately for men (76,547 observations) and women (65,603 observations).

All the demographic information for this analysis is drawn from what CPS respondents report at the time of the survey, since the administrative data does not report educational level or reliable information on race (Martin 2016). Since respondents' report of age at the time of survey is subject to reporting error, we use information in the administrative records to determine their age (and their spouse's) at

[23] Workers who claim disability benefit are automatically converted to a retired worker benefit upon attaining their FRA, so they have no claiming-age decision.

the time of the survey. We assign each spouse's information (from the survey or administrative data) to the other spouse.

4.2 Administrative data

Because the CPS is not a panel survey, we cannot observe if or when respondents apply for Social Security benefits (whether retired worker, SSDI, spousal or survival). Hence, we match survey respondents with their Social Security administrative records and track their claiming age, benefit type and amounts, lifetime earnings, and mortality. We use information from the Master Beneficiary Record (MBR) file, the Summary Earnings Record (SER), and NUMIDENT. The matched data files contain information on Social Security taxable wages from 1951 through 2020, Social Security benefits, date of birth, and date of death. Our analysis sample comprises those with a matched record in the three files noted above.

Our measure of lifetime earnings is Average Indexed Monthly Earnings (AIME) computed at age 62, so it is independent of claim age and later work decisions. The OASI claiming age is derived from the MBR file, and mortality data from the MBR and NUMIDENT files.

5. Claiming-age estimates

We investigate claiming-age patterns by estimating an ordered probit model, where the left-hand side variable takes values 0 to 4 for claiming at age 62, 63, 64, 65, or else 66 and older. We estimate models separately for men and women and control for education and lifetime earnings quartile. We are interested in determining whether, conditional on education and lifetime earnings quartile, married men who are three or more years older than their wives claim later; and whether married men whose wives' PIA is between 50% to 100% of theirs (and thus whose wives are most likely to claim initially as retired-workers and later as survivors) claim later than other men. We look similarly for whether wives in the same situation claim earlier than other women.

5.1 Claiming-age patterns of husbands

Specifications 1 to 3 in Table 1 report results for all, single, and married men respectively, with specifications 1 and 3 including controls for whether, conditional on marriage, the husband is older or of similar age (defined as at least or less than three years older) as his wife. The rest of the specifications focus on married men. Specifications 4 and 5 control in different ways for relative PIA of spouses, specifications 6 and 7 report results separately for married men who are older or of similar age to their wife, and specifications 8 to 10 report results separately for married men with wives whose PIA is less than 50%, 50% to 100%, or more than 100% of that of her husband.

In all our specifications, education and lifetime earnings quartiles are almost always statistically significant and vary little in impact across specifications. Men with higher education and higher lifetime earnings claim significantly later than men with low education and low lifetime earnings.[24] The effects are significantly different from each other across each category, while they are quite large in magnitude for the highest education and lifetime earnings group. For example, those in the highest lifetime earnings quartile are 14.5 percentage points (pp) less likely to claim at 62 and 11.1 pp

[24] In Dushi et al. (2021), we established this result for lifetime earnings quartile. Here, we demonstrate that it persists even controlling for education.

more likely to claim at 66 or older, relative to those in the lowest earnings quartile; the second and third earnings quartiles have statistically significant differences that are considerably smaller.[25] These patterns will matter for our analysis later because any differences in mortality by earnings will magnify the impact of differences in claiming age by earnings, as high earners will get their higher benefits over a longer lifetime.

Our key hypotheses involve differences in claiming incentives of married couples, and especially whether, for couples in which the wife is likely to claim a survivor benefit, husbands claim later than other men and wives claim earlier than other women. Such couples are ones in which the wife is younger than the husband (we focus on an age difference of three or more years) and the wife's PIA is within 50% to 100% of the husband's. A consistent finding across all specifications is that married men with a wife at least three years younger claim significantly later than men with a wife of a similar age. In specification 1, for the full sample, they are 4.2 percentage points less likely to claim at age 62, and 3.1 percentage points more likely to claim at 66 or older, than are unmarried men. When the sample is restricted to married men, they are 3.5 percentage points less likely to claim at 62 and 2.6 percentage points more likely to claim at 66 or older than are men with wives who are less than three years younger; notably, the latter

[25] These values (and others reported in the text) are ordered probit marginal effects. They report how the probability of claiming at a given age changes when the variable in question shifts from taking a value of 0 to a value of 1 (in the case of dummy variables like these). These probabilities are computed at the mean values of the other right-hand side variables. The ordered probit marginal effects are available from the authors upon request but are not reported in the paper because there are so many; there is a marginal effect for each right-hand side variable as it applies to each of the five ordered outcomes.

specification controls for PIA differences.[26] Men whose wife is of a similar age claim somewhat later than single men, but the magnitude of the effect is about one-third that for the group above. These patterns will increase Social Security outlays, but not by a great deal as the claim-age differences are not that substantial. Moreover, Sun and Webb (2011) demonstrate that household welfare is maximized if husbands delay claiming until age 68 (given the implicit value of longevity insurance that this offers).[27]

The direction of the effects, if not their magnitude, are consistent with the greater incentives to delay claiming faced by married men and the even greater incentives faced by married men with younger wives. The finding that men with younger wives claim later than men with similar-age wives is consistent with Coile et al. (2002). But that study found, counterintuitively, that married men claimed earlier than single men. The difference in findings may result from differences across birth cohort or in control variables. The earlier study focused on men who claimed in 1980 to 1981, mostly the 1918 to 1919 birth cohort, which predated the decline and more (and substantial) recent increase in claim ages that we documented in Friedberg et al. (2021). They also

[26] The difference between the younger wife and similar-age wife coefficients is 0.0863, and in specifications 3 to 5 where the sample is only married men and the base case is a similar-age wife, the younger-wife coefficients are 0.0915, 0.0953, and 0.0967 respectively. For specifications 6 and 7, we need to compare the cut points of 0.564 and 0.467, yielding a difference of 0.097.

[27] The calculations in Sun and Webb (2011) are, importantly, intended to demonstrate optimal behavior in the absence of liquidity constraints that might drive early claiming. Recent trends, including lower interest rates and mortality, would, if anything, lead to later optimal claiming than their calculations suggest.

included controls for subjective survival beliefs, not available in our CPS data, that predict claiming and are likely correlated with marital status.[28]

We next discuss the role of PIA differences. The PIA difference should have an approximately monotonic effect on the husband's claiming age, with the greater the difference, the higher his claim age. If the wife's PIA is less than 50% of the husband's, then she will claim spousal benefits initially, and both her spousal and later survivor benefits are reduced when he claims early. If the wife's PIA is between 50% and 100% of the husband's, then she will claim retired-worker benefits initially, which only depend on her claiming age; but if she outlives her husband, she will switch to receiving survivor benefits, which are reduced when he claims early. And if the husband's PIA is smaller than the wife's, his claiming age does not affect her benefits at all.

The estimated effects are consistent with these hypotheses, though they are rather small when we test them in the restricted sample of married men. The coefficients on the dummies for the wife's PIA being between 50% to 100% of her husband's or more than 100% of her husband's, are statistically significantly different than zero and from each other. They are both negative, so they are associated with earlier claiming by the husband, and the latter is more negative than the former. However, they are both small in magnitude. If the wife's PIA is between 50% to 100% of the husband's, then he is 0.8 percentage points more likely to claim at age 62 and 0.66 percentage points less likely to claim at 66 and older, than if the wife's PIA is less

[28] In contrast to Coile et al. (2002), we include controls for AIME quartile and educational attainment, both of which are correlated with both marital status and claiming ages and might be expected to reduce any positive correlation between marital status and claiming age.

than 50% of the husband's. If the wife's PIA is more than 100% of her husband's, the values are 2.1 percentage points more likely and 1.5 percentage points less likely.

5.2 Claiming-age patterns of wives

Table 2 reports similar results for women. As with men, higher levels of educational attainment and higher lifetime earnings are associated with delayed claiming, and if anything the magnitudes are a little greater. For example, women in the highest lifetime earnings quartile (among women of the same birth cohort) are 19.7 percentage points less likely to claim at 62 and 13.1 percentage points more likely to claim at 66 or older, relative to those in the lowest earnings quartile. To the extent that high earners have greater longevity, the claiming-age responses will magnify differences in lifetime benefits. Moreover, to the extent that high-earning women are married to high-earning men, it will increase differences across households.

The key finding is that, as predicted by a model in which couples claim strategically to maximize lifetime benefits, married women claim significantly and substantially earlier than single women. Within the sample of married women, though, differences are smaller (though still significant) than for men based on age difference between spouses; women who are at least three years younger than their husband claim somewhat earlier than women whose husbands are similar in age. The PIA differences matter more for married women than for married men, on the other hand. Women whose PIA is less than half of their husband's claim earliest, while women whose PIA exceeds that of their husband delay claiming.

Although statistically significant, some of these effects are small in magnitude. The exceptions are the difference between the likelihood of married and single women claiming early (with married women around 13 to 14 percentage points less likely to claim at 62 and around 10 percentage points more likely to claim at 66 or older, depending on the spousal age difference, compared to single women), and the difference based on the ratio of the husband's and wife's PIA (with married women whose PIA is more than 100% of their husband's 11.2 percentage points less likely to claim at 62 and 7.2 percentage points more likely to claim at 66 or older than women whose PIA is less than 50% of their husband's).

5.3 Mortality estimates

In order to investigate life expectancy by claiming age for incomplete cohorts (whose members have not all died), we build on the Gompertz mortality models estimated by Dushi et al. (2021). Those models showed that baseline mortality at age 62 differs significantly with birth cohort, claiming age, and lifetime earnings quartile and that the annual age-related increase in mortality also differed significantly with birth cohort and lifetime earnings quartile. We extend the analysis by allowing baseline mortality to also vary, in one specification, with marital status and spousal age and PIA difference, parameterized as before.[29] These models allow us to test formally whether the mortality of married individuals differs from that of single individuals, after controlling

[29] Our sample size, and more particularly, the number of deaths in more recent cohorts, does not permit us to estimate a model in which mortality varies with both the age difference between husband and wife and the relative PIA of husband and wife. If, as some demographers believe, there is a natural limit to the human life span, it is plausible that the age-related increase in mortality will be higher among groups with lower age-62 mortality. But our sample size is insufficient to estimate a model in which gamma, the age-related increase in mortality, is allowed to vary with our covariates.

for mortality differences associated with claim age and lifetime earnings; whether men married to younger women have lower mortality than men married to women of similar age; and whether mortality varies with the relative PIA of husband and wife. In addition to estimated coefficients, we report remaining life expectancy at age 62 based on those coefficients.[30] As in Dushi et al., we estimate the model separately for each two-year birth cohort, in effect allowing all explanatory variables to vary by birth cohort. Both claiming-age patterns and the annual age-related increase in mortality have changed considerably across the 1929/1930 to 1943/1944 birth cohorts, and we reject the hypothesis of a constant age-related increase across cohorts.

Throughout, we compare our results to SSA cohort life tables, partly to investigate the difference between population-average life expectancy and life expectancy for our sample, which is relatively healthy because we exclude beneficiaries with a previous DI claim, and partly to gauge whether our cross-cohort changes in life-expectancy seem reasonable. We do this by calculating average mortality for each cohort in our sample, weighted by claiming age, lifetime earnings quartile, marital status, and age difference between spouses. Overall, we find that our model yields reasonable estimates of baseline mortality, age-related mortality increases, and consequently life expectancy, up to the 1941 to 1942 birth cohort.

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[30] The raw coefficient estimates show how much log mortality at each age changes as a result of the covariate in question. To illustrate, in the specification for men born in 1931 to 1932, λ, γ, and the age-66 claiming-age coefficient are estimated to be -4.298, 0.083, and -0.240 respectively. Age-70 log mortality of an age-66 claimant equals -4.298 + 0.083*(70-62)-0.240, or -3.874, which when exponentiated equals 2.08%.

Even with our large sample, when we estimate our model separately for each two-year birth cohort, some of the coefficients are imprecisely estimated while others vary considerably from year to year. We report results for each cohort and for a pooled sample with cohort dummies, for which the coefficients are more precisely estimated. In the following discussion, we often highlight results for the 1939 to 1940 birth cohort, 36.4% of whom have died, yielding a reasonable fit.

Table 3 reports results for a pooled sample of married and single men. The pooled sample implicitly imposes the constraint that the age-related increase in mortality is the same for all men within a cohort, albeit from different baseline age-62 mortality rates. The base case is a single man with less than a high school education in the lowest AIME quartile who claims at age 62. When exponentiated, the constant reports age 62 mortality of the base case — thus for men born in 1939-40, the exponent of the -4.097 constant yields age 62 mortality of 1.66%.

Consistent with Dushi et al. (2021), we find that men in higher AIME quartiles have significantly and substantially lower mortality. To illustrate, when exponentiated, the -0.419 coefficient for men in the highest earnings quartile who were born 1939 to 1940 equals 0.658, so that they have a mortality rate at all ages that is 34.2% (1-0.658) lower than that of men in the lowest AIME quartile. The coefficients for birth cohort dummies for more recent birth cohorts are increasingly negative, reflecting declining mortality rates across birth cohorts. A comparison of the 1943 to 1944 and the 1929 to

1930 constants reveals that baseline mortality declined by 38.2% over the 14 year period (1-exp(3.973-4.455)), close to the population level decline over the period.[31]

Our key variables of interest are those capturing the relationships between mortality and marital status. We include two dummy variables, one for men married to women who are at least three years younger and the other for men married to women who are closer in age. The coefficients are -0.438 and -0.406 respectively, which when exponentiated equate to mortality reductions of 35.5% and 33.2%. The two coefficients are not significantly different from each other. The implication is that the return to delayed claiming of the retired worker benefit is greater for married than for single men. Assuming married women have population average mortality, the lower mortality of married men reduces the expected present value of their wife's survivor benefit because it will start at a later date and be payable for a shorter period. In unreported results, we find that the relationship between male mortality and relative PIA of husband and wife is small and lacking in statistical significance.

Table 4 reports similar results for women. As with men, married women have lower mortality than single women. The 1939 to 1940 coefficients for married women married to a man the same age and to a man at least three years older are -0.374 and -0.298 respectively, which equate to mortality reductions of 31.2% and 25.8%, respectively. The lower mortality of married women has only a small effect on the expected present value of their retired worker benefit because, regardless of mortality, they stand a very good chance of living to transition to survivor benefit. But it increases

[31] We do not expect it to be identical because the characteristics of the sample have changed over the period — for example, levels of educational attainment have increased — and the sample excludes SSDI recipients, a high mortality group whose size has fluctuated.

the expected present value of their survivor benefit. We return to this question when we discuss returns to delay.

Other coefficients are similar but not identical to those for men. Differences between mortality rates of women in the first three AIME quartiles rarely attain statistical significance in the models estimated on two-year birth cohorts, but are statistically significant in the pooled sample. For the 1935 to 1936 birth cohort onward, women in the highest AIME quartile have significantly lower mortality than those in the lowest quartile, albeit less of a reduction than for men. The AIME quartile four coefficient for 1939-40 is -0.179, which equates to a mortality reduction of 16.4%. The claim age coefficients are similar for both men and women, although as previously mentioned women who delay claiming are a smaller and likely more select group.

We then estimate models in which the annual age-related increase in mortality and the reductions in mortality associated with delayed claiming are allowed to vary between single and married men and between single and married women and models for married men and women that also allow mortality to vary with the relative PIA of husband and wife. We detected no statistically significant correlations between mortality and relative PIA, and therefore report results for models that exclude relative PIA.

Among men, the reductions in mortality associated with delayed claiming or membership of a higher AIME quartile are greater for married than for single men. Among women, we detect few significant differences between married and single women in the relationship between claiming age and mortality and between AIME quartile and mortality.

5.4 Relationship between marital status and mortality

Table 5 reports life expectancy at age 62 for men and women in the 1933 to 1934 birth cohort by claim age (62, 63, 64, 65, or 66 plus), by AIME quartile, and by marital status and spousal age difference (single, married with wife at least three years younger and married with wife less than three years younger).[32] Consistent with Dushi et al. (2021), life expectancy is substantially greater for those who delay claiming and for those in higher lifetime earnings quartiles. For example, among male claimants with a wife at least three years younger who are in the lowest AIME quartile, age-62 life expectancy is 20.4 years among those who claim at age 62, versus 24.7 years among those who claim at age 66. In comparison, age-62 life expectancy for the same type of men but in the highest AIME quartile is 23.8 years for those who claim at age 62, versus 28.3 years for those who claim at age 66. Moreover, the table shows that, longevity is considerably greater for married men than for singles, and slightly higher for married men married to younger women (which would slightly reduce expected survivor benefit payouts to such households) compared to men married to women of a similar or greater age. As is well known, women live longer than men, while the differences by lifetime earnings and marital status are similar to those for men. However, the relationship between claiming age and longevity among women is weaker — perhaps because of the incentive for married women who expect a survivor benefit to claim early.

We can further calculate the number of years that a married woman can expect to spend as a surviving spouse. Greater male longevity reduces the expected duration

[32] Age-62 mortality of those who claim after age 62 includes those who died without claiming, based on our assumption of exponentially increasing mortality.

of survivor benefits and greater female longevity increases the expected duration of those benefits, and it is an empirical question whether the net effect is an increase or decrease. For couples in the 1939 to 1940 birth cohort who are the same age, our results show that the expected length of widowhood is substantially greater for households with low socioeconomic status. For the most disadvantaged households we consider in terms of male mortality (households with men in the lowest AIME quartile who have the mortality of those who typically claim at age 62), the average length of widowhood is 7.78 years. In contrast, for the most advantaged households we consider (households with men in the highest AIME quartile who have the mortality of those who typically claim at age 66), the average length of widowhood is 4.90 years. These calculations suggest that survivor benefits may be more important for low SES households, compared to high SES households.

5.5 Impact on Social Security benefits

Now, we investigate the combined impact of mortality and claiming patterns by marriage on expected lifetime benefits. Table 6 reports how the expected present value (EPV) of lifetime benefits for spouses and for households as a whole depend on claim ages of husbands and wives, comparing 1) couples with the mortality of the lowest AIME quartile, age-62 claimants born in 1939 to 1940; and 2) couples with the mortality of the highest earnings quartile, age-66 claimants born in 1939 to 1940. The first row shows the EPV of the total household benefit, which consists of the EPV of the husband's retired-worker benefit (in the second row), the EPV of his wife's retired-worker benefit (in the third row), and the EPV of his widow's survivor benefit (in the fourth row); the household in question has a husband who claims the retired-worker benefit with a spouse who is three years younger. The averages account for expected mortality of the husband and wife, as estimated in our mortality models. In each case, we assume the husband's benefit payable at his FRA is $1 a year (PIA is $1/12) and the wife's benefit payable at her FRA is 50 cents a year. The assumed interest rate is a real 3%.

At a baseline claiming age of 62 for the husband, the high-mortality household (with average mortality of age-62 claimants in the lowest AIME quartile) has an EPV of lifetime benefits of $21.10 per dollar of PIA. This consists of $12.30 resulting from the husband's retired-worker benefit (received in the event that he remains alive), $5.27 from the wife's retired-worker benefits (received in the event that both spouses remain alive), and $3.53 from the wife's survivor benefit (received in the event that only she remains alive). For the same household, if the husband claims at 66, the EPV of their household lifetime benefit is $22.16 per dollar of PIA — yielding a return to his delay (comparing the gain in the EPV of the husband's benefit and the wife's survivor benefit) of 6.73% — that is, the household gets 6.73% higher lifetime benefits if the husband in a low-mortality household delays claiming from age 62 to 66. This is substantially higher than the gain to delay for a similar high-mortality man if he were assumed to be single, as we assume in Dushi et al. (2021). The difference can be explained largely because much of the gain from delaying accrues to the widow benefit, not to the husband's retired-worker benefit.

The return to delay for the low-mortality household that we consider (with average mortality of age-66 claimants in the highest AIME quartile) is somewhat higher. In this case, the gain to delaying claiming from age 62 to age 66 for a low-mortality man is 8.75%. Notably, the widow benefit is lower in expectation for the low-mortality household, by virtue of the husband's considerably longer life expectancy, and the full gain in lifetime benefits is achieved through the higher EPV of his retired-worker benefit alone.

In fact, the survivor benefit narrows outcomes across households by SES, compared to our analysis in Dushi et al. (2021). While men with high lifetime earnings who delay claiming gain considerably from delay through their greater longevity, men with low lifetime earnings would benefit their wives by delaying claiming. Yet, we do not see these delays in claiming occurring in many low SES households. And moreover, with marriage rates declining more among low-earning than high-earning households, the protective role of the survivor benefit is becoming less available for lower-earning households.

6. Conclusions

The structure of Social Security survivor benefits creates complicated and largely unresearched interactions of claiming-age incentives for couples in which the lower-earning spouse within a couple claims benefits as a retired-worker but expects to receive survivor benefits later in life. This is because survivor benefits depend on the age at which the deceased spouse claimed retired-worker benefits, *but not* the age at which the survivor initially claimed as a retired-worker. For example, in the common case in which a lower-earning wife claims as a retired worker but would receive survivor benefits if she outlives her husband (which occurs for 39.5% of married couples in the sample that we analyze), husbands have an incentive to claim retired-worker benefits later, not just to increase his own retired-worker benefit but also to increase his spouse's possible survivor benefit. In contrast, wives have an incentive to claim retired-worker benefits early, since the time horizon for this decision is not her own expected longevity but her husband's. This stands in contrast to the simple incentive faced when ignoring survivor benefits: The higher one's life expectancy, the later one should claim.

We find that men and women claim in ways that increase but do not maximize the expected present value of lifetime benefits, raising Social Security outlays but not to the extreme. We also find that systematic mortality differences by marriage further increase the returns to delay for married men.

Incorporating these differences into household benefit calculations, we find that the return to delayed claiming of the husband's retired-worker benefit is substantially more than actuarially fair, but for different reasons by household type. For households facing disadvantages, the return to delay by the husband arises more from gains to the survivor benefit than to his retired-worker benefit. For households with advantages, the return to delay arises largely from the gains to husbands' retired-worker benefit.

These patterns deepen the insights that we gained in our earlier analysis in Dushi et al. (2021), which treated all men as single and found evidence of rising adverse selection that particularly benefits high earners. Here, when we incorporate the availability of the survivor benefit in married households, it is apparent that the results are more nuanced. Men with higher lifetime earnings, who claim later on average, raise costs of the OASI Trust Fund through higher retired-worker benefit payouts, while men with lower lifetime earnings, who claim relatively early on average, forgo an important gain to delay in the form of higher survivor benefits that would accrue to their wives.

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Tables

Table 1: Claiming-age coefficient estimates, different ordered probit specifications and samples, men

Dependent variable: claim age	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Sample:	All men	Single men	Married men	Married men	Married men	Married men, wife 3+ years younger	Married men, wife <3 years younger	Married men, wife's PIA> husband's	Married men, wife's PIA<50% husband's	Married men, wife's PIA ∈50-100% husband's
Independent variables:										
Single (omitted)										
Married, wife 3+ years younger	0.117***		0.0915***	0.0953***	0.0967***					
	(0.0117)		(0.00942)	(0.00948)	(0.00950)					
Married, wife < 3 years younger	0.0307***									
	(0.0118)									
Education < high school (omitted)										
High school graduate or some college	0.105***	0.121***	0.103***	0.108***	0.109***	0.101***	0.107***	0.139***	0.0880***	0.118***
	(0.0140)	(0.0306)	(0.0158)	(0.0158)	(0.0159)	(0.0205)	(0.0248)	(0.0363)	(0.0243)	(0.0258)
Completed college	0.482***	0.487***	0.481***	0.488***	0.490***	0.505***	0.459***	0.487***	0.513***	0.472***
	(0.0145)	(0.0320)	(0.0163)	(0.0164)	(0.0164)	(0.0213)	(0.0253)	(0.0364)	(0.0252)	(0.0270)
AIME quartiles (1 omitted)										
AIME 2	0.0927***	0.1000***	0.0867***	0.0754***	0.0762***	0.0594***	0.122***	0.0705***	0.0889***	0.0545**
	(0.0118)	(0.0243)	(0.0135)	(0.0138)	(0.0138)	(0.0177)	(0.0207)	(0.0252)	(0.0255)	(0.0223)
AIME 3	0.0816***	0.122***	0.0684***	0.0528***	0.0519***	0.0568***	0.0863***	0.0701**	0.0720***	0.00561
	(0.0119)	(0.0266)	(0.0134)	(0.0140)	(0.0141)	(0.0180)	(0.0202)	(0.0299)	(0.0242)	(0.0227)

AIME 4	0.396***	0.458***	0.383***	0.365***	0.361***	0.354***	0.417***	0.492***	0.337***	0.335***
	(0.0123)	(0.0303)	(0.0136)	(0.0144)	(0.0145)	(0.0184)	(0.0202)	(0.0346)	(0.0238)	(0.0240)
wife's PIA> husband's				-0.0465***						
				(0.0123)						
wife's PIA < 50% of husband's PIA (omitted)										
wife's PIA 50-100% of husband's PIA					- 0.0208**					
					(0.0104)					
wife's PIA >100% of husband's PIA					-0.0579***					
					(0.0135)					
1930 cohort (omitted)										
	-0.131***	0.0265	-0.144***	-0.144***	- 0.144***	-0.135**	-0.180*	-0.152	-0.110	-0.187**
1933-1934	(0.0468)	(0.176)	(0.0485)	(0.0485)	(0.0485)	(0.0543)	(0.108)	(0.105)	(0.0740)	(0.0811)
	-0.0645	0.148	-0.0834*	-0.0833*	-0.0832*	-0.0417	-0.211**	-0.0649	-0.0624	-0.140*
1935-1936	(0.0449)	(0.167)	(0.0467)	(0.0467)	(0.0467)	(0.0525)	(0.102)	(0.100)	(0.0716)	(0.0778)
	-0.100**	0.160	-0.129***	-0.129***	- 0.129***	-0.122**	-0.154	-0.105	-0.0897	-0.224***
1937-1938	(0.0434)	(0.160)	(0.0452)	(0.0452)	(0.0452)	(0.0512)	(0.0974)	(0.0974)	(0.0692)	(0.0754)
	-0.104**	0.0432	-0.121***	-0.122***	- 0.122***	-0.111**	-0.150	-0.0638	-0.119*	-0.194***
1939-1940	(0.0424)	(0.157)	(0.0441)	(0.0441)	(0.0441)	(0.0501)	(0.0953)	(0.0957)	(0.0674)	(0.0737)
	0.0106	0.142	-0.00560	-0.00715	- 0.00708	0.00277	-0.0259	-0.00938	0.00954	-0.0708
1941-1942	(0.0419)	(0.154)	(0.0436)	(0.0436)	(0.0436)	(0.0497)	(0.0942)	(0.0953)	(0.0666)	(0.0726)
	0.0807*	0.195	0.0672	0.0660	0.0663	0.0726	0.0520	0.0975	0.0704	-0.00749
1943-1944	(0.0414)	(0.153)	(0.0431)	(0.0431)	(0.0431)	(0.0491)	(0.0933)	(0.0935)	(0.0659)	(0.0717)
	0.395***	0.436***	0.396***	0.396***	0.397***	0.409***	0.375***	0.384***	0.431***	0.301***
1945+	(0.0411)	(0.152)	(0.0428)	(0.0428)	(0.0428)	(0.0488)	(0.0928)	(0.0925)	(0.0657)	(0.0710)
	0.580***	0.567***	0.599***	0.598***	0.600***	0.615***	0.569***	0.522***	0.640***	0.505***

Age-63 threshold	0.633***	0.737***	0.588***	0.573***	0.564***	0.467***	0.613***	0.571***	0.530***	0.429***
	(0.0421)	(0.152)	(0.0436)	(0.0438)	(0.0440)	(0.0488)	(0.0933)	(0.0924)	(0.0665)	(0.0716)
Age-64 threshold	0.824***	0.932***	0.778***	0.763***	0.755***	0.655***	0.806***	0.759***	0.724***	0.616***
	(0.0421)	(0.152)	(0.0437)	(0.0438)	(0.0440)	(0.0488)	(0.0933)	(0.0925)	(0.0665)	(0.0716)
Age-65 threshold	1.082***	1.157***	1.044***	1.029***	1.020***	0.953***	1.035***	1.018***	1.004***	0.870***
	(0.0421)	(0.152)	(0.0437)	(0.0439)	(0.0441)	(0.0489)	(0.0934)	(0.0926)	(0.0666)	(0.0717)
Age-66+ threshold	1.645***	1.630***	1.630***	1.615***	1.607***	1.612***	1.549***	1.572***	1.617***	1.446***
	(0.0423)	(0.152)	(0.0440)	(0.0441)	(0.0443)	(0.0493)	(0.0936)	(0.0932)	(0.0670)	(0.0720)
Observations	76,547	14,417	62,130	62,130	62,130	33,206	28,924	12,523	25,065	24,542

Note: Claiming age takes five possible values: 62, 63, 64, 65, 66+. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 2: Claiming-age coefficient estimates, different ordered probit specifications and samples, women

Dependent variable: claim age	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Sample:	All women	Single women	Married women	Married women	Married women	Married women, wife 3+ years younger	Married women, wife <3 years younger	Married women, wife's PIA> husband's	Married women, wife's PIA<50% husband's	Married women, wife's PIA ∈50-100% husband's
Independent variables:										
Single (omitted)										
Married, wife 3+ years younger	-0.392***		-0.0354***	-0.0516***	-0.0638***					
	(0.0118)		(0.0111)	(0.0112)	(0.0113)					
Married, wife < 3 years younger	-0.356***									
	(0.0109)									
Education < high school (omitted)										
High school graduate or some college	0.0949***	0.156***	0.0772***	0.0871***	0.103***	0.0601*	0.0921***	-0.0953*	0.133***	0.153***
	(0.0169)	(0.0276)	(0.0215)	(0.0216)	(0.0216)	(0.0311)	(0.0299)	(0.0496)	(0.0364)	(0.0323)
Completed college	0.370***	0.346***	0.388***	0.396***	0.417***	0.334***	0.426***	0.149***	0.642***	0.395***
	(0.0178)	(0.0293)	(0.0225)	(0.0225)	(0.0226)	(0.0330)	(0.0309)	(0.0507)	(0.0382)	(0.0339)
AIME quartiles (1 omitted)										
AIME 2	0.273***	0.298***	0.260***	0.252***	0.192***	0.391***	0.171***	0.177***	0.178***	0.198***
	(0.0132)	(0.0258)	(0.0154)	(0.0155)	(0.0165)	(0.0242)	(0.0201)	(0.0507)	(0.0219)	(0.0297)
AIME 3	0.515***	0.538***	0.496***	0.471***	0.355***	0.642***	0.396***	0.365***	0.380***	0.365***
	(0.0131)	(0.0251)	(0.0155)	(0.0157)	(0.0193)	(0.0244)	(0.0201)	(0.0468)	(0.0560)	(0.0287)
AIME 4	0.525***	0.421***	0.579***	0.506***	0.390***	0.747***	0.465***	0.480***	-	0.390***
	(0.0136)	(0.0250)	(0.0163)	(0.0175)	(0.0208)	(0.0258)	(0.0211)	(0.0446)		(0.0312)

wife's PIA> husband's				0.165***						
				(0.0144)						
wife's PIA < 50% of husband's PIA (omitted)										
wife's PIA 50-100% of husband's PIA					0.171** (0.0165)					
wife's PIA >100% of husband's PIA					0.306***					
					(0.0198)					
1930 cohort (omitted)										
	0.0195	0.0182	0.0124	0.0173	0.0118	0.0630	-0.0162	-0.232	-0.0389	0.213
1933-1934	(0.0795)	(0.147)	(0.0949)	(0.0950)	(0.0950)	(0.145)	(0.126)	(0.212)	(0.152)	(0.150)
	-0.00481	0.0487	-0.0422	-0.0296	-0.0345	-0.0544	-0.0360	-0.308	-0.110	0.181
1935-1936	(0.0729)	(0.134)	(0.0874)	(0.0874)	(0.0875)	(0.131)	(0.117)	(0.200)	(0.139)	(0.139)
	0.00351	0.0266	-0.0219	-0.0163	-0.0245	-0.0753	0.0171	-0.243	-0.0488	0.153
1937-1938	(0.0701)	(0.128)	(0.0841)	(0.0842)	(0.0842)	(0.126)	(0.113)	(0.186)	(0.133)	(0.135)
	0.0358	0.0600	0.00304	0.00662	-0.00457	-0.0292	0.0253	-0.206	-0.0380	0.175
1939-1940	(0.0684)	(0.125)	(0.0822)	(0.0822)	(0.0823)	(0.123)	(0.111)	(0.180)	(0.130)	(0.132)
	0.125*	0.119	0.108	0.108	0.0946	0.0748	0.131	-0.0517	-0.0259	0.307**
1941-1942	(0.0675)	(0.123)	(0.0810)	(0.0811)	(0.0811)	(0.121)	(0.109)	(0.177)	(0.128)	(0.131)
	0.224***	0.160	0.236***	0.234***	0.221***	0.238**	0.231**	0.103	0.108	0.422***
1943-1944	(0.0669)	(0.122)	(0.0802)	(0.0803)	(0.0803)	(0.119)	(0.108)	(0.175)	(0.126)	(0.130)
	0.447***	0.397***	0.452***	0.453***	0.441***	0.463***	0.446***	0.348**	0.305**	0.652***
1945+	(0.0666)	(0.122)	(0.0797)	(0.0798)	(0.0798)	(0.119)	(0.108)	(0.174)	(0.125)	(0.129)
	0.562***	0.421***	0.607***	0.601***	0.579***	0.569***	0.633***	0.545***	0.414***	0.775***
Age-63 threshold	0.433***	0.407***	0.928***	0.939***	0.984***	0.947***	0.967***	0.330	0.952***	1.013***
	(0.0777)	(0.132)	(0.0953)	(0.0955)	(0.0958)	(0.127)	(0.140)	(0.216)	(0.152)	(0.156)
Age-64 threshold	0.618***	0.573***	1.123***	1.134***	1.180***	1.116***	1.184***	0.498**	1.178***	1.202***
	(0.0777)	(0.132)	(0.0955)	(0.0956)	(0.0960)	(0.128)	(0.140)	(0.216)	(0.152)	(0.156)
Age-65 threshold	0.980***	0.929***	1.489***	1.501***	1.548***	1.445***	1.583***	0.819***	1.623***	1.554***
	(0.0779)	(0.132)	(0.0959)	(0.0960)	(0.0964)	(0.128)	(0.141)	(0.216)	(0.153)	(0.157)
Age-66+ threshold	1.711***	1.656***	2.228***	2.240***	2.291***	2.175***	2.333***	1.631***	2.302***	2.296***
	(0.0789)	(0.133)	(0.0977)	(0.0979)	(0.0983)	(0.131)	(0.143)	(0.219)	(0.158)	(0.159)

Observations	65,603	19,464	46,139	46,139	46,139	19,462	26,677	9,849	14,921	21,369

Note: Claiming age takes five possible values: 62, 63, 64, 65, 66+. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 3: Mortality model estimates, by birth cohort and for pooled sample, men

Independent	Birth cohorts							Pooled	
variables	**1929-30**	**1931-32**	**1933-34**	**1935-36**	**1937-38**	**1939-40**	**1941-42**	**1929-50**	**1929-44**
Claim age:									
62 (omitted)									
63	-0.00507	-0.390***	-0.0274	0.0221	-0.0884	0.107	0.0709	-0.00824	-0.0168
	(0.145)	(0.105)	(0.0917)	(0.0845)	(0.0778)	(0.0790)	(0.0776)	(0.0267)	(0.0310)
64	-0.189*	-0.0695	-0.276***	-0.191***	-0.160***	-0.180**	-0.105	-0.158***	-0.168***
	(0.107)	(0.0733)	(0.0736)	(0.0641)	(0.0600)	(0.0706)	(0.0750)	(0.0235)	(0.0258)
65	-0.153	-0.0703	-0.233***	-0.248***	-0.260***	-0.289***	-0.261***	-0.204***	-0.221***
	(0.111)	(0.0767)	(0.0679)	(0.0613)	(0.0589)	(0.0540)	(0.0518)	(0.0203)	(0.0225)
66+	0.0624	-0.0508	-0.463***	-0.218	-0.133	-0.166	-0.253**	-0.449***	-0.319***
	(0.195)	(0.116)	(0.149)	(0.261)	(0.179)	(0.156)	(0.126)	(0.0294)	(0.0427)
AIME quartiles (1 omitted)									
AIME 2	0.117	0.00365	-0.113*	-0.0950	-0.0115	-0.0759	-0.220***	-0.121***	-0.0839***
	(0.110)	(0.0701)	(0.0661)	(0.0602)	(0.0555)	(0.0567)	(0.0543)	(0.0189)	(0.0219)
AIME 3	-0.0312	-0.156**	-0.257***	-0.172***	-0.225***	-0.209***	-0.465***	-0.302***	-0.258***
	(0.111)	(0.0710)	(0.0679)	(0.0620)	(0.0578)	(0.0587)	(0.0578)	(0.0198)	(0.0228)
AIME 4	0.0201	-0.261***	-0.398***	-0.413***	-0.382***	-0.419***	-0.684***	-0.498***	-0.444***
	(0.112)	(0.0736)	(0.0709)	(0.0663)	(0.0614)	(0.0634)	(0.0632)	(0.0216)	(0.0244)
Single (omitted)									
Married, wife	-0.265*	-0.164*	-0.391***	-0.359***	-0.333***	-0.438***	-0.281***	-0.385***	-0.367***
3+ yrs ygr	(0.145)	(0.0938)	(0.0811)	(0.0684)	(0.0603)	(0.0560)	(0.0552)	(0.0197)	(0.0236)
Married, wife	-0.0245	-0.00113	-0.302***	-0.292***	-0.273***	-0.406***	-0.303***	-0.356***	-0.316***
<3 yrs ygr	(0.162)	(0.106)	(0.0909)	(0.0755)	(0.0659)	(0.0608)	(0.0597)	(0.0208)	(0.0257)
1930 cohort (omitted)									
1931-1932								-0.0415	-0.0428
								(0.0463)	(0.0463)
1933-1934								-0.178***	-0.176***
								(0.0458)	(0.0458)
1935-1936								-0.256***	-0.248***
								(0.0448)	(0.0449)
1937-1938								-0.271***	-0.260***
								(0.0443)	(0.0444)
1939-1940								-0.374***	-0.359***
								(0.0448)	(0.0450)
1941-1942								-0.382***	-0.365***
								(0.0451)	(0.0452)
1943-1944								-0.428***	-0.428***

								(0.0462)	(0.0467)
1945+								-0.506***	
								(0.0443)	
Estimated γ	0.0992***	0.0953***	0.0913***	0.0860***	0.0844***	0.0800***	0.0727***	0.0819***	0.0850***
	(0.00534)	(0.00359)	(0.00366)	(0.00359)	(0.00362)	(0.00408)	(0.00449)	(0.00136)	(0.00143)
Estimated λ	-4.455***	-4.372***	-4.091***	-4.140***	-4.162***	-4.097***	-3.998***	-3.746***	-3.847***
	(0.176)	(0.114)	(0.0995)	(0.0860)	(0.0777)	(0.0736)	(0.0713)	(0.0488)	(0.0512)
Observations	843	2,162	2,841	3,981	5,325	6,427	7,976	76,484	38,958

Note: Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 4: Mortality model estimates, by birth cohort and for pooled sample, women

Independent variables	Birth cohorts							Pooled	
	1929-30	1931-32	1933-34	1935-36	1937-38	1939-40	1941-42	1929-50	1929-44
Claim age:									
62 (omitted)									
63	-0.186	-0.348	-0.101	-0.0117	0.0705	-0.0302	0.138	0.0106	-0.00801
	(0.291)	(0.214)	(0.160)	(0.141)	(0.115)	(0.114)	(0.104)	(0.0383)	(0.0473)
64	-0.298	-0.500***	-0.216*	-0.0381	-0.115	-0.464***	-0.0247	-0.167***	-0.168***
	(0.218)	(0.167)	(0.124)	(0.100)	(0.0925)	(0.104)	(0.0976)	(0.0345)	(0.0394)
65	-0.352	-0.311*	-0.280**	-0.0190	-0.351***	-0.421***	-0.312***	-0.286***	-0.305***
	(0.219)	(0.170)	(0.129)	(0.103)	(0.0991)	(0.0849)	(0.0779)	(0.0322)	(0.0371)
66+	-0.422	-0.106	-0.246	0.0547	0.130	-0.106	-0.119	-0.368***	-0.209***
	(0.382)	(0.197)	(0.174)	(0.160)	(0.134)	(0.140)	(0.123)	(0.0376)	(0.0497)
AIME quartiles (1 omitted)									
AIME 2	0.348*	0.0689	-0.267**	-0.00909	0.182**	-0.0743	-0.0660	-0.0896***	-0.0302
	(0.191)	(0.137)	(0.109)	(0.0900)	(0.0832)	(0.0793)	(0.0747)	(0.0270)	(0.0325)
AIME 3	0.232	0.150	-0.0392	-0.108	0.0427	-0.102	-0.270***	-0.179***	-0.0953***
	(0.194)	(0.137)	(0.105)	(0.0929)	(0.0854)	(0.0814)	(0.0802)	(0.0281)	(0.0335)
AIME 4	0.0423	-0.0888	-0.104	-0.236**	-0.0444	-0.179**	-0.339***	-0.293***	-0.234***
	(0.210)	(0.144)	(0.107)	(0.0960)	(0.0876)	(0.0829)	(0.0818)	(0.0290)	(0.0347)
Single (omitted)									
Married, wife	-0.299	-0.315**	-0.493***	-0.217**	-0.307***	-0.298***	-0.415***	-0.371***	-0.349***
3+ yrs ygr	(0.188)	(0.135)	(0.103)	(0.0882)	(0.0760)	(0.0757)	(0.0716)	(0.0258)	(0.0310)
Married, wife	-0.366**	-0.300**	-0.405***	-0.335***	-0.524***	-0.374***	-0.498***	-0.476***	-0.455***
<3 yrs ygr	(0.183)	(0.123)	(0.0945)	(0.0835)	(0.0745)	(0.0720)	(0.0679)	(0.0243)	(0.0293)
1930 cohort (omitted)									
1931-1932								0.0143	0.0131
								(0.0825)	(0.0825)
1933-1934								-0.0803	-0.0788
								(0.0770)	(0.0770)
1935-1936								-0.101	-0.0939
								(0.0747)	(0.0748)
1937-1938								-0.201***	-0.189**
								(0.0738)	(0.0739)
1939-1940								-0.270***	-0.255***
								(0.0738)	(0.0740)
1941-1942								-0.313***	-0.295***
								(0.0740)	(0.0743)
1943-1944								-0.393***	-0.390***

								(0.0750)	(0.0755)
1945+								-0.505***	
								(0.0730)	
Estimated γ	0.0971***	0.0960***	0.0915***	0.0946***	0.0879***	0.0892***	0.0918***	0.0884***	0.0910***
	(0.00914)	(0.00691)	(0.00573)	(0.00537)	(0.00525)	(0.00556)	(0.00608)	(0.00201)	(0.00216)
Estimated λ	-4.861***	-4.715***	-4.534***	-4.794***	-4.786***	-4.727***	-4.720***	-4.381***	-4.495***
	(0.271)	(0.182)	(0.138)	(0.122)	(0.106)	(0.103)	(0.0981)	(0.0798)	(0.0830)
Observations	348	727	1,440	2,264	3,458	4,715	6,374	65,577	27,496

Note: Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 5: Life expectancy at age 62, 1933-34 birth cohort

	Single				**Married, wife <3 years younger**				**Married, wife 3+ years younger**			
AIME quartile	**1**	**2**	**3**	**4**	**1**	**2**	**3**	**4**	**1**	**2**	**3**	**4**
	Men											
Men's claim age 62	16.56	19.16	18.04	19.77	19.63	20.43	21.87	23.04	20.39	21.20	22.65	23.84
63	17.55	20.22	19.07	20.85	19.75	20.55	21.99	23.16	20.51	21.32	22.78	23.97
64	17.52	20.19	19.04	20.82	22.04	22.87	24.36	25.58	22.83	23.67	25.18	26.41
65	19.13	21.92	20.72	22.56	21.54	22.36	23.85	25.05	22.32	23.16	24.66	25.88
66	18.47	21.21	20.03	21.85	23.84	24.69	26.22	27.46	24.65	25.51	27.06	28.30
	Women											
Women's claim age 62	20.64	24.02	21.01	22.15	24.26	26.49	24.70	25.10	25.06	27.31	25.50	25.91
63	19.21	22.49	19.57	20.67	26.53	28.81	26.98	27.39	27.35	29.64	27.81	28.22
64	22.26	25.73	22.65	23.82	26.36	28.64	26.81	27.22	27.18	29.47	27.63	28.05
65	22.92	26.42	23.31	24.49	26.84	29.13	27.30	27.71	27.67	29.96	28.12	28.54
66	23.04	26.55	23.44	24.62	25.80	28.07	26.25	26.66	26.62	28.90	27.07	27.48

Table 6: Expected present value of lifetime benefits per dollar of PIA at the full retirement age, 1939 to 1940 birth cohort

Claim age	**62**	**63**	**64**	**65**	**66**	**67**	**68**	**69**	**70**
	Based on mortality of men who are in AIME quartile 1 and typically claim at age 62								
Total household benefit	21.10	21.30	21.68	21.95	22.16	22.30	22.34	22.30	22.19
Husband's retired-worker benefits	12.30	12.42	12.52	12.54	12.51	12.40	12.22	11.98	11.67
Wife's retired-worker benefit	5.27	5.27	5.27	5.27	5.27	5.27	5.27	5.27	5.27
Survivor benefit	3.53	3.62	3.88	4.14	4.39	4.63	4.85	5.06	5.26
Husband's gain to delaying claiming		1.26%	3.65%	5.39%	6.73%	7.57%	7.85%	7.62%	6.93%
	Based on mortality of men who are in AIME quartile 4 and typically claim at age 66								
Total household benefit	22.59	22.93	23.39	23.75	24.05	24.27	24.40	24.45	24.42
Husband's retired-worker benefits	14.56	14.85	15.14	15.34	15.49	15.56	15.55	15.46	15.30
Wife's retired-worker benefit	5.86	5.86	5.86	5.86	5.86	5.86	5.86	5.86	5.86
Survivor benefit	2.17	2.22	2.38	2.54	2.70	2.85	2.99	3.13	3.26
Husband's gain to delaying claiming		2.05%	4.81%	6.96%	8.75%	10.07%	10.86%	11.13%	10.94%

Note: These calculations assume a 3% interest rate. Women are assumed to claim retired-worker benefits at age 62 and survivor benefits upon the death of the husband. The table computes the expected present value of benefits, depending on the age at which the husband claims retired-worker benefits.