The Effects of Changes in Social Security's Delayed Retirement Credit: Evidence from Administrative Data

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Abstract: The delayed retirement credit (DRC) increases monthly OASI (Old Age and Survivors Insurance) benefits for primary beneficiaries who claim after their full retirement age (FRA). For many years, the DRC was set at 3.0 percent per year (0.25 percent monthly). The 1983 amendments to Social Security more than doubled this actuarial adjustment to 8.0 percent per year. These changes were phased in gradually, so that those born in 1924 or earlier retained a 3.0 percent DRC while those born in 1943 or later had an 8.0 percent DRC. In this paper, we use administrative data from the Social Security Administration (SSA) to estimate the effect of this policy change on individual claiming behavior. We focus on the first half of the DRC increase (from 3.0 to 5.5 percent) given changes in other SSA policies that coincided with the later increases. Our findings demonstrate that the increase in the DRC led to a significant increase in delayed claiming of social security benefits and strongly suggest that the effects were larger for those with higher lifetime incomes, who would have a greater financial incentive to delay given their longer life expectancies.

Keywords: Social security, claiming behavior, retirement

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I. INTRODUCTION

Old Age, Survivors, and Disability Insurance (OASDI) – also known as Social Security - is the largest social insurance program in the United States. This program paid an average monthly benefit of $1,382 to 64.1 million Americans in 2019, with total OASDI expenditures in 2019 amounting to $1.06 trillion.[1] By far the largest group of OASDI recipients is retired workers, accounting for 45.1 million program beneficiaries with an additional 3.1 million dependents of this group in this same year. An individual with at least 10 years of earnings[2] can claim retired worker benefits as early as the month in which the individual turns 62, with an actuarial adjustment for each month of delaying claiming beyond that point.

Social Security represents a substantial source of income for retired workers and is the primary source of income for most individuals aged 65 and up. However, the trust fund is predicted to deplete its reserves by 2035 (Annual Trustees Report, 2020).[3] Understanding how individuals choose to claim benefits and respond to program rules is essential for implementing policies to address significant concerns surrounding the long-term solvency of this program. However, despite the substantial body of literature examining different aspects of Social Security, little attention has been paid to a key component of retirees' benefit calculation: the Delayed Retirement Credit (DRC). The DRC directly increases the financial incentive to claim retired worker benefits later through adjusting the monthly benefits upwards for those who delay claiming.

In this paper, we present new evidence on how changes in the Delayed Retirement Credit affect the timing of decisions to claim Social Security retirement benefits. The age that a person

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[1] 2020 Social Security Administration Annual Statistical Supplement, 4.A OASDI: Trust Funds, Table 4.A1 and Table 4.A2; 5.A OASDI Current-Pay Benefits: Summary, Table 5.A1

[2] An individual needs to earn 40 quarters of credits, where credits are received for work in covered earnings. See https://www.ssa.gov/benefits/retirement/planner/credits.html for more information.

[3] See https://www.ssa.gov/oact/tr/2020/tr2020.pdf for additional information. The Congressional Budget Office projects a depletion date 3 years earlier in 2032

claims Social Security, together with the average indexed monthly earnings (AIME), determines the benefit amount that will be received. Those who claim after their full retirement age (FRA) have their monthly benefits scaled up by the DRC, while benefits are reduced for those who claim earlier than FRA. Under the 1983 amendments, the DRC rose from just 3.0 percent per year for those born prior to 1925 to 8.0 percent per year for those born in 1943 and later. This DRC increase was phased in gradually in 0.5 percentage point increments every two years, so that those born in 1925 and 1926 had a 3.5 percent annual DRC, those born in 1927 and 1928 had a 4.0 percent annual DRC, and so forth through the 1943 birth cohort. For this cohort and all subsequent ones, the DRC was equal to 8.0 percent.

We study the effects of these increases on individual claiming behavior through using Social Security administrative micro data, with individual level records of claiming for eleven birth years spanning 1923 through 1933. We focus on the first five changes in the DRC since the latter five changes coincided with other changes to retired worker benefits including the increase in the full retirement age (FRA) and the elimination of the earnings test for earnings received after FRA. We leverage the highly detailed nature of this data set to empirically estimate the effect of DRC increases on the probability that individuals delay claiming beyond FRA.

We first present simulations of the expected present value of claiming at different ages and highlight two key relationships: increases in the DRC increase the financial incentives to claim later, and there is substantial heterogeneity in the propensity to claim later. In particular, we predict that those with low mortality rate profiles (who tend to have higher lifetime incomes) are more affected by DRC increases, given that they are expected to receive a larger monthly benefit for a longer period of time. This may constitute a form of adverse selection. To the extent that those with the longest life expectancies respond to the policy change by delaying claiming, the present value of outlays by the Social Security program may be significantly larger as a result of the policy change. Furthermore, the effects of increases in the DRC on claiming decisions might be nonlinear in nature. More specifically, earlier DRC increases may have occurred at such low levels (e.g. from 3.0 percent annual to 3.5 percent annual) that these did not change behavior on the margin while subsequent increases may have. We utilize the staggered rollout of the DRC increases that we analyze to empirically test this hypothesis.

We use Social Security administrative data from a 10 percent random sample of all beneficiaries and focus on individuals born between 1923 through 1933 to implement a differences-in-differences identification strategy to estimate the causal effects of DRC increases. The DRC is uniquely determined by the year of birth, which creates several cutoffs between those with birth months of December of year t and January of year t+1. We compare claiming decisions of individuals born around the end of the calendar year, to those of individuals born around the beginning of the next calendar year. The individuals in the two groups are likely similar in nature except that those born right after January of a new calendar year experience a higher DRC rate compared to those born right before January.

To account for the possibility that individuals born early in the calendar year claim differently than those born later in the previous calendar year for other reasons, we utilize an equal number of placebo birth windows. For example, there was no change in the DRC from year-of-birth 1925 to 1926 or from 1927 to 1928. However, there was an increase from 1926 to 1927. We are careful to construct our sample to account for simultaneous shifts in other Social Security program rules – namely the increases in the full retirement age (FRA, which began with the 1938 year-of-birth) and the earnings test (ET, which was passed in April 2000) – to isolate the effect of the DRC. Our main analysis focuses on the effects for men, due to large changes in the working and claiming behavior of women (and corresponding changes in the composition of women claiming retired worker benefits) during our period of interest.

We find consistent empirical evidence that men respond, albeit moderately, to increases in the DRC among the full sample of men born between 1923 through 1933. More specifically, a 0.5 percentage point increase in the DRC increases claiming at or beyond age 66 (12 months after FRA) by 0.23 percentage points. Furthermore, we find suggestive evidence of adverse selection, as the estimated effect for those in the top earnings decile is much larger in magnitudes than the effect for those with earnings below the median. These differences are however only suggestive and are not statistically significant.

We also uncover evidence of nonlinearity in the effects of the DRC. Given the staggered rollout of the DRC increases, we are able to separately estimate the effects of the earlier increases (e.g. 3.0 to 3.5 percentage points annually) compared to later ones (e.g. 5.0 to 5.5 percentage points). The effects of the DRC are concentrated among the later increases, where a shift from 3.0 percent to 4.5 percent did not produce a significant effect on later claiming behavior, while each 0.5 percentage point increase in the DRC during the shift from 4.5 to 5.5 percent increased the probability of claiming at 66 or after by 0.37 percentage points. Consistent with our results for the full sample, the effect is most pronounced among the highest earners. Taken together, we estimate that 69 percent of the observed rise in the fraction of male retired workers claiming later during our sample period (from the 1923 to 1933 cohorts) is explained by increases in the DRC.[4]

A substantial body of literature has explored retired worker claiming behavior, generally finding that individuals claim benefits sooner than what might be predicted through economic theory as optimal. While researchers have noted a variety of factors that contribute to early

[4] The fraction of male retired workers born in 1923 who claim at or beyond age 66 is 4.56% versus 5.64% for their counterparts born in 1933.

claiming behavior, there has been surprisingly little exploration of a whether a policy specifically aimed at promoting later claiming is effective. Furthermore, we examine an additional important component to claiming behavior: individuals have private information, such as health status, that affects their propensity to claim benefits. To the extent that those who are longer-lived also claim later and receive larger benefits for a longer amount of time, this form of adverse selection has important implications for the consequences of this policy and the solvency of the OASI program.

The outline of the paper is as follows. In Section 2, we provide details on Social Security's retired worker claiming rules along with background on the 1983 Social Security Amendments. We also briefly summarize the related previous literature. Section 3 describes our sources of administrative data from the Social Security Administration along with the construction of our analysis sample. We then provide an overview in Section 4 of OASI claiming patterns along with a simulation of the effect of the DRC increases on the present value of retired worker benefits. Sections 5 and 6 summarize our identification strategy and our empirical results, respectively, before we conclude in Section 7.

II. BACKGROUND

In this section, we provide details on how an individual's monthly Social Security retired worker (as opposed to disabled worker, spousal, etc.) benefits are determined and how these benefits vary by claiming age. We highlight several factors that might influence an individual's claiming decisions, including the Delayed Retirement Credit (DRC), the Full Retirement Age (FRA), and the earnings test (ET). In the next section we document how we isolate the effect of the DRC, taking into account simultaneous changes in these other factors. Lastly, we discuss this paper in the context of the existing literature.

A. Social Security Claiming Rules

To qualify for retired worker benefits, an individual must have a sufficient work history, which includes 40 or more credits (at least 10 years) of covered employment or earnings that are subject to Social Security taxes and must also be at least 62 years old. In 2021, for a credit to "count", an individual must have earned at least $1,470, and a person who earned $5,880 during the year would earn 4 credits of coverage[5]. For those who are qualified, the monthly OASDI retired worker benefit depends primarily on three factors. Firstly, the worker's average indexed monthly earnings (AIME), which is an average of the 35 highest years of indexed covered earnings[6], determines the monthly primary insurance amount (PIA). The AIME is converted into the primary insurance amount (PIA) using progressive replacement rates of 90, 32, and 15 percent (SSA, 2019). For example, in 2020, the first $960 of AIME was replaced at a 90 percent rate while the next $4,825 of AIME was replaced at a 32 percent rate, with any remaining monthly earnings in the AIME (beyond $5,785) replaced at a 15 percent rate up to the taxable maximum.

A second important factor that affects the monthly retired worker benefit is the age at which the worker claims benefits. An individual who claims benefits at the full retirement age (FRA) receives 100 percent of the PIA while one claiming before that age receives benefits scaled down by monthly early retirement reduction factors to account for the longer period of benefit receipt. In contrast, each month that a worker delays claiming beyond the FRA results in benefits being scaled up by the monthly delayed retirement credit (DRC).

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[5] The amounts are of course lower in previous years: see https://www.ssa.gov/oact/cola/QC.html

[6] Only earnings that are subject to the OASDI payroll tax in each year count for the purposes of the AIME calculation. For example, in 2019, annual earnings above $132,900 were not subject to this tax and thus maximum monthly earnings for the purposes of the AIME calculation would be $11,075. Earnings in previous years are scaled up to account for increases in average economy-wide earnings over time. See: https://www.ssa.gov/oact/cola/rtea.html

Finally, a retired worker's benefits may be reduced due to Social Security's earnings test. The earnings test has been present since the creation of the Social Security program in 1935, though it was eliminated for those receiving benefits at or beyond the FRA in April of 2000. For an individual claiming retired worker benefits before FRA, the earnings test leads to a 50% phase out of OASDI benefits as annual earnings increase beyond a threshold ($18,960 in 2021).[7]

There are three ways the benefit amount can change after claiming: a cost of living adjustment (COLA), additional work, or an adjustment at full retirement age if one had received reduced benefits and exceeded the earnings limit. COLA affects monthly benefits by adjusting benefits following the Department of Labor's Consumer Price Index. In 2021, the announced adjustment was 1.3 percent, and affected more than 70 million beneficiaries. Additional work after receiving benefits could potentially increase the benefit amount if the additional earnings is higher than the highest thirty-five years of earnings before claiming. Lastly, a person who decided to take benefits early at a reduced rate while continuing to work may experience increases in benefits once they reach the full retirement age if they exceeded the allowable earnings limit and had some of the benefits withheld.

B. The 1983 Social Security Amendments

In 1981, President Ronald Reagan created the National Commission on Social Security Reform, a bipartisan group created due to the inability of Congress and the President to agree on reforms aimed at addressing the financial problems of two trust funds: The Old-Age and Survivors Insurance and The Disability Insurance. At that time, the program had run annual deficits for several consecutive years and the program's trust fund had shrunk to just one-sixth of annual

[7] Benefits that are withheld due to the earnings test are returned to the individual in the form of increased benefits after an individual reaches the Full Retirement Age.

expenditures (versus approximately 2.7 times today). These deficits were occurring despite a substantial increase during this period in the maximum amount of each worker's earnings subject to OASDI taxes along with a corresponding increase in the program's tax rate.

The National Commission was tasked with identifying sources threatening the financial insolvency of the program and "analyze potential solutions to such problems that will assure the financial integrity of the Social Security System … and provide appropriate recommendations" (Executive Order 12335). The recommendations of the National Commission made their way into the 1983 Social Security Amendments, including the increase in the FRA and in the DRC. Prior to this amendment, the DRC had stood at just 3.0% per year but this was gradually increased to 8.0% for those born in 1943 and later.

This legislation also increased the full retirement age, though these changes were phased in even more gradually. The full retirement age was left unchanged for workers born in 1937 and earlier (thus anyone aged 46 and up) while it was set at 67 for those born in 1960 and later. This FRA increase from 65 to 67 was phased in using two-month increments from the 1937 through 1943 birth cohorts and then again with two-month increments for the 1954 through 1960 birth cohorts. Interestingly, only half of this change from 38 years ago has so far been phased in, since workers reaching age 66 in 2020 still had an FRA of 66. Along with these changes, the 1983 amendments gradually reduced the amount of benefits at the earliest possible age of claiming of 62 at each age through the FRA. Once the changes are fully phased in, the fraction of an individual's PIA that the worker could receive when claiming at age 62 will be just 70 percent, versus 80 percent for those born in 1937 or earlier (and 75 percent for those born from 1943 through 1954).

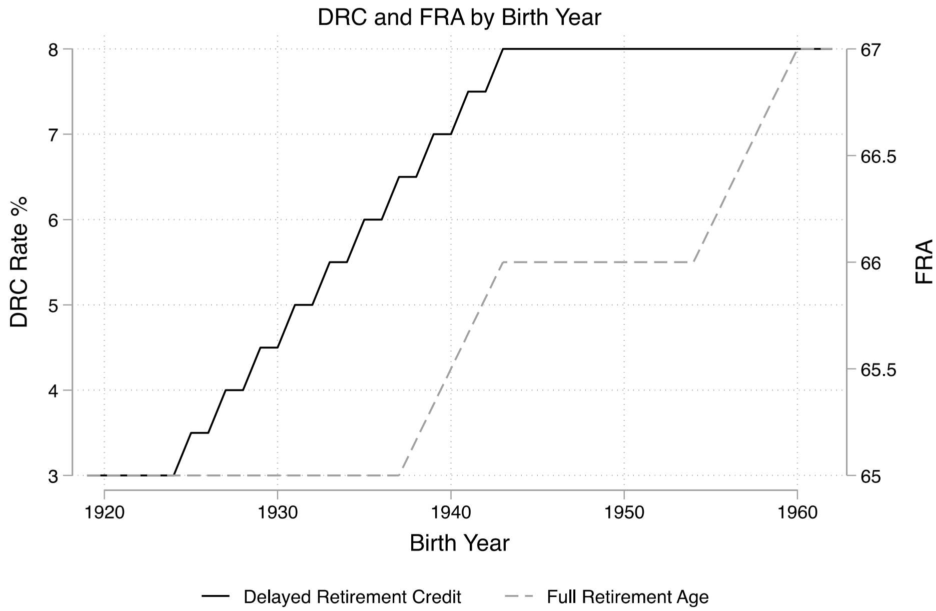

Figure 1: Delayed Retirement Credit and Full Retirement Age by Birth Year

Notes: Illustrates the Delayed Retirement Credit (DRC) and Full Retirement Age by birth year.

Figure 1 illustrates the evolution of the DRC and FRA for birth cohorts from 1919 through 1962. As the figure shows, the DRC starts at 3%, and steadily rises by half a percentage point starting with the 1925 birth cohort every two years. By the 1943 birth cohort, the DRC has risen to 8%, and remains steady at 8% for all subsequent birth cohorts. However, the Full Retirement Age is also changing as well, starting at age 65 up through the 1937 birth cohort and rising in two-month increments to age 66 for the 1943 through 1954 birth cohorts. This increases again to age 67 in two-month increments so that it reaches age 67 by the 1960 birth cohort. Given that both the DRC and FRA are dependent on birth year, simultaneous shifts in the FRA might confound our estimates for the effect of DRC changes, especially those occurring for the 1938 and later birth cohorts. We therefore discuss this in much greater detail in our sample selection and empirical analysis sections below.

We also summarize the effects of the FRA and DRC changes for the 1919 through 1954 birth cohorts in Table 1, illustrating how the FRA and DRC affects the amount of benefits received. In the last two columns, the values are the fraction of the PIA the individual will obtain when claiming at age 62 or age 70. For example, for the 1919 birth cohort, claiming at age 70 would lead to receiving 115 percent of the PIA in monthly benefits, while for the 1937 birth cohort, claiming at that same age would produces 132.5 percent of the PIA in monthly benefits.

Table 1. Social Security Benefit Schedule

Birth Year	FRA	DRC (%)	% PIA Claiming at Age 62	% PIA Claiming at Age 70
1919-24	65	3.00	80.00	115.00
1925-26	65	3.50	80.00	117.50
1927-28	65	4.00	80.00	120.00
1929-30	65	4.50	80.00	122.50
1931-32	65	5.00	80.00	125.00
1933-34	65	5.50	80.00	127.50
1935-36	65	6.00	80.00	130.00
1937	65	6.50	80.00	132.50
1938	65, 2 mo.	6.50	79.17	131.42
1939	65, 4 mo.	7.00	78.33	132.67
1940	65, 6 mo.	7.00	77.50	131.50
1941	65, 8 mo.	7.50	76.67	132.50
1942	65,10 mo.	7.50	75.83	131.25
1943-54	66	8.00	75.00	132.00
1955	66, 2 mo.	8.00	74.17	130.67
1956	66, 4 mo.	8.00	73.33	129.33
1957	66, 6 mo.	8.00	72.50	128.00
1958	66, 8 mo.	8.00	71.67	126.67
1959	66,10 mo.	8.00	70.83	125.33
1960+	67	8.00	70.00	124.00

Notes: Full retirement age (FRA), annual Delayed Retirement Credit (DRC), and PIA benefit adjustment rate for selected claiming ages by birth year.

C. Related Literature

Our work builds on a body of previous literature that has investigated the determinants of claiming Social Security benefits. Previous studies have found that in terms of strictly financial calculations, there is a substantial benefit to delaying claiming (Coile et al. 2002; Sun and Webb 2009; Shoven and Slavov 2014b). For example, Shoven and Slavov (2014a) finds that at relatively low interest rates, similar to those that prevail today, primary earners with average life expectancy should delay benefits to age 70 to maximize the expected present value of benefits. Coile et al (2002) finds that while incentives to delay vary according to life expectancy or the marital status of the claimant, in most cases the gains from delaying are substantial. Yet, very few individuals claim after FRA, even though people on average are claiming later than they did 30 years ago. (Purcell, 2020). Most of this previous research suggests that relatively little is known regarding the mechanisms through which people decide to claim earlier than what is predicted through economic theory. An investigation of the impact of the policy-induced increase in the delayed retirement credit therefore remains an important question to be answered.

The previous empirical research on claiming patterns, however, consistently finds that people typically claim significantly earlier than what economic theory suggests would be optimal (Goda et al. 2018, Shoven and Slavov 2014a, Shoven et al. 2017, Rohwedder and van Soest 2006, Mastrobuoni 2009). Numerous factors, such as financial literacy, mistrust in the program, liquidity constraints, mortality risk, undervaluing of annuitized income, social norms, and framing are studied. There is not a consensus among these and related studies regarding the influence of these factors, with the exception of interest rates and mortality risk. However, considerably less work has focused on examining how policy levers can be used to influence claiming behavior and whether individuals' claiming margins are responsive to program incentives. The closest study to our paper, Pingle (2006), utilizes the changes in the DRC to explore the effects on men's employment, using the Survey on Income and Program Participation (SIPP). This paper finds substantial DRC-induced labor supply increases for men in the 65 to 69 age group, but it does not explore the effect of the DRC changes on Social Security claiming decisions. Additionally, Benitez-Silva and Yin (2009) find suggestive evidence that increases in the DRC did not lead to substantial changes in claiming behavior.

The existing literature on the determinants of individual claiming decisions and the effects of the DRC has two main shortcomings. First, existing studies are often limited to using publicly available data sources. As a consequence, it is difficult to precisely measure Social Security claiming behavior as well as other key factors, such as lifetime earnings, benefit amounts, and exact year of birth (given that year of birth determines benefit rules). Another challenge for many of the previous studies is the absence of a plausibly exogenous source of variation with which to reliably estimate causal effects. Since there is an exact relationship between the age at which people claim Social Security, the calendar year in which they claim, and the delayed retirement credit level (which is determined by their birth year), identification of the true effect of the change in DRC that takes into account both the effect of age and of calendar year is difficult.

The DRC is determined uniquely by birth year and claiming year, so disentangling calendar year and age effects is a challenge. Studies note this difficulty and treat policy changes that vary at the cohort level as plausibly exogenous shocks (Mastrobuoni 2006, Pingle 2006, Duggan et al 2007, Engelhardt et al 2020). While researchers take care to test the robustness of the ensuing results, it is difficult to rule out the possibility that people from different birth years are different in unobserved ways. For example, if researchers do not account for the time trend effect where people claim later due to increased life expectancy in general or due to changes in interest rates over time, they might erroneously conclude that an increase in DRC raises the likelihood of claiming beyond the full retirement age.

Lastly, solely determining whether the DRC produces an effect provides an incomplete picture on the broader consequences of this policy change. Finkelstein and Poterba (2004) test for the presence of adverse selection in the private annuity market in the U.K., finding that individuals self-selected into contracts based off of their private information on their mortality risk. This suggests that health may be an important determinant of claiming behavior and that there may be important heterogeneity to uncover.

In this study, we leverage the rich administrative micro data from the Social Security Administration to track an individual's lifetime earnings, date of claiming, Social Security claiming type, and demographics, including birth date. In addressing the inherent identification concerns, we utilize a differences-in-difference design to estimate the causal effect of the substantial increase in Social Security's delayed retirement credit (from 3.0 to 8.0 percent annually). Using our data on claimant characteristics, we can investigate whether the reaction to the greater financial incentive to delay claiming is statistically meaningful, and whether the effects vary with claimant characteristics. Given that those with higher incomes on average have much lower mortality rates (and thus more to gain by delaying claiming), we can test for the presence of adverse selection in the claiming of an annuity (Social Security) in our setting as well.

III. DATA & ANALYSIS SAMPLE

We use several sources of administrative data from the Social Security Administration to create the primary dataset for our empirical analyses. We begin with the 10 percent sample of the Master Beneficiary Record (MBR) extract, containing the history of OASDI receipt for the primary beneficiary and other beneficiaries (e.g. spouse or child dependents), and merge it with the corresponding Master Earnings File (MEF), which provides the annual taxable earnings data from 1951 through 2016, and with the Numident extract. The Numident is used as a source of demographic and death data for all individuals, including those who are unobserved in the MBR.

Table 2. Sample Restrictions

	Total	Male	Female
Sample is primary beneficiaries with matched records in NUMIDENT, MEF, and MBR files[1]	14,848,075	8,268,923	6,579,152
Remaining sample after exclusions:			
Keep those born in 1923-1933, inclusive	2,035,642	1,149,995	885,647
Exclude DI beneficiaries	1,692,399	917,362	775,037
Keep only those Fully insured at age 61 for OASI[2]	1,548,571	881,904	666,667
Keep only those Alive at age 62	1,536,911	873,213	663,698
Keep only OASI retirement beneficiaries, i.e. exclude spouse/widow or all other beneficiaries	1,509,049	864,810	644,239
Exclude those with OASI claim prior to age 62	1,508,310	864,272	644,038
Exclude those with OASI claim after age 70	1,498,364	862,144	636,220
Exclude if race is missing	1,495,243	860,451	634,792
Restrict the sample to birth months 1-3 and 10-12[3]	673,610	387,265	286,345
Final Sample	673,610	387,265	286,345

[1] We start with this sample, for which we could match the records in NUMIDENT file, in the MEF (those with at least a year of positive earnings), and in MBR file (primary beneficiaries).

[2] To be fully insured for OASI, workers born in 1929+ must have at least 40 quarters of coverage or 10 years of work.

[3] This is our regression analysis sample. Hence, for the 1923 birth cohort we also exclude those born in January-March, and for the 1933 cohort we exclude those born in October-December.

To create our main analysis sample, we make a series of sample restrictions to limit to individuals who are the primary claimants (beneficiaries) of Social Security retirement benefits (i.e., they claimed benefits on their own earning records) as documented in Table 2. We restrict to individuals who do not claim Social Security Disability Insurance (SSDI) benefits at any time, those who are insured for OASI retired worker benefits (i.e., have at least 40 quarters of coverage or 10 year of work), and who are still alive at age 62 (the first possible age for claiming retired worker benefits). We also limit to those who claim between the ages of 62 and 70. Very few individuals claim retired worker benefits before age 62 or after age 70; we remove those who claim prior to age 62 since no individual is permitted to claim OASI before age 62 and therefore must be an error in the data. Those who claim after the age of 70 do not receive further actuarial adjustments so any delay beyond age 70 is not an optimal choice and will lead to lifetime benefit losses; given our focus in this paper on the effect of DRC, we exclude them from our analysis.

The main outcome variable of interest in our analysis is whether an individual claims Social Security retirement benefits for the first time on or after one year beyond their full retirement age. An individual will start receiving DRC adjusted benefits when claiming at any month after the FRA, so those who claim between FRA and one year after FRA will also receive a monthly benefit that exceeds the PIA due to the delayed retirement credit. We focus on a full year (12 or more months) after the FRA to better ensure that we are capturing the effects of the DRC-induced later claiming and to limit measurement error that might arise for those who accidentally claim just a few months after FRA. For example, we might observe an individual who claims at 2 months after FRA, but this individual actually intended to claim at the FRA and it took some time to file for benefits. Furthermore, it might not be widely known that DRC benefits are applied at the monthly level. Claiming at ages one year after FRA provides a natural starting point for plausibly attributing the claiming that we observe in the data as intentional later claiming.

Our main analysis focuses on cohorts born between 1923 and 1933, which includes several birth cohorts with an annual DRC of 3.0 percent (1923 – 1924), up to cohort with a DRC of 5.5 percent (1933), and 8 birth cohorts with a DRC between 3.5 percent and 5.0 percent (1925 – 1932).

These birth cohort restrictions do not fully encompass all DRC increases, since the DRC rises to 8.0 percent for the 1943 and subsequent birth cohorts. However, we restrict up to the 1933 birth cohort to produce a sample of individuals who do not also face concurrent changes in other Social Security claiming factors so as to more reliably isolate the effect of the DRC. Firstly, the Full Retirement Age (FRA) starts to increase beginning with the 1938 birth cohort. Behaghel and Blau (2012) have shown that individuals are highly responsive to the changes in the FRA, with claiming patterns for many moving in lockstep with FRA increases. Therefore, we remove birth cohorts born on or after 1938, and thus only consider cohorts that faced an unchanging FRA of exactly age 65 in our main analysis.

Secondly, we exclude birth years whose claiming is most affected by the removal of the earnings test in 2000 for those beyond FRA. On April 7th of 2000, Congress enacted the Senior Citizens' Freedom to Work Act of 2000, which removed the earnings test for those at or older than FRA (see Song and Manchester (2007) for a full discussion of the legislation and the predicted effects on labor supply). While an individual need not claim simultaneously with retirement, this change in the earnings test rules likely affected working decisions for these groups (Gelber et al. shows that the earnings test affected employment) and thus claiming behavior. Even if a person's work decisions were not affected, the individual may decide to claim earlier due to the post-FRA earnings test elimination. For example, an individual who previously worked well beyond age 65 may have decided to delay claiming social security benefits due to the earnings test. With this eliminated, the individual might now decide to claim.

Figure 2. Fraction of Individuals Claiming 1 Year After FRA by Birth Year

Notes: The figure plots the share of individuals who claim Social Security after one year past their FRA for male. The numbers are authors' calculations from the 10% MBR sample.

This partially explains why in Figure 2, we see that the fraction of men claiming retired worker benefits one year or more after FRA fell very sharply from 5.6% for the 1933 cohort to 1.5% for the 1935 cohort after steadily increasing for previous birth cohorts. This dramatic fall explains the phenomenon where workers who previously delayed claiming due to the presence of earnings test no longer saw the need to delay once it was removed in 2000. Table 3 shows the birth cohorts directly affected by the earnings test. Those born in 1930 turn 70 in 2000 when the change in earnings test is introduced, so the first row in Table 3 shows that their claiming behaviors are unaffected (denoted with a 0) from age 62 to 70. In contrast, those born in 1938 turn 62 in 2000, so the second to the last row in Table 3 shows that their claiming behaviors are affected (denoted with 1) by the policy change for the entire duration between age 62 and age 70. Those born between

1931 and 1937, however, are affected for a subset of the duration, which makes it difficult to separate the effect of increase in DRC on claiming ages from change in behavior induced by introduction of the change in earnings test.

Table 3. Claiming Behavior Affected by Earnings Test

Birth Year	62	63	64	65	66	67	68	69
1930	0	0	0	0	0	0	0	0
1931	0	0	0	0	0	0	0	1
1932	0	0	0	0	0	0	1	1
1933	0	0	0	0	0	1	1	1
1934	0	0	0	0	1	1	1	1
1935	0	0	0	1	1	1	1	1
1936	0	0	1	1	1	1	1	1
1937	0	1	1	1	1	1	1	1
1938	1	1	1	1	1	1	1	1
1939	1	1	1	1	1	1	1	1

Note: This table indicates whether the claiming behavior at each age for each birth year could have been affected by the removal of the earnings test in 2000. For the claiming age and birth year affected by the removal of the earnings test, we assign 1. We assign 0 if otherwise. We exclude birth years greater than or equal to 1934 in the main sample due to the removal of earnings test.

To identify the share of those claiming after or at age 66, who are unaffected by the earnings test, we restrict the sample of birth years to those born on or before 1933. We note that while those born before April of 1934 are also arguably unaffected by the earnings test for this outcome variable, we exclude the birth year 1934 entirely since those born later in the year would be affected. The earnings test likely induced a change in behavior even for those born before April 1934, as seen by the steep decline initiating with the birth month of January 1934 in Figure 2.

Finally, we further restrict the sample to men, due to large changes in the working and claiming behavior of women over the study period (similar to Coile et al. (2002), Pingle (2006)).

We include results for women separately in the appendix. There are several reasons why we analyze results for women separately from those of men in this period. First, there are large compositional changes in the share of women who claim benefits as the primary claimant, compared to claiming as a spouse. Given that we are concerned with the behavior of the primary claimant, for women, we cannot rule out that changes in the DRC over time happen concurrently with the changes in the composition of those who decide to claim as the primary or spouse. In fact, the fraction of aged women who claim as the primary beneficiary decreased from 59.5 percent in the 1923 cohort to 51.2% in the 1933 cohort, offset by increases in those who claim as the spouse.[8] Meanwhile, consistently around 97 percent of men claim as the primary beneficiary. Second, women in these birth years tend to have significantly lower lifetime earnings compared with men, and thus their decisions to claim might look very different from those of men's. For example, in the 1933 birth cohort, the median PIA was $626 for men, while it was just $329 for women. The corresponding difference in AIMEs are even greater. Given the age difference between spouses, along with the fact that couples tend to align the timing of their retirement, the DRC is likely to affect the claiming decisions for women differently. Table 2 illustrates changes in the number of observations as we apply all of our above-mentioned selection criteria, leaving us with a sample of 387,265 men born from 1923 through 1933 who were alive on their 62nd birthday and not previously on SSDI.[9]

[8] Based of authors' calculations. We tabulate share who claim as the primary for OASI benefits, out of the total number who claim as the primary for OASI, aged spouse, and aged widower benefits. This variable captures claim type by the type of initial entitlement, and are thus mutually exclusive categories. Some of the initial benefits could come from dual entitlement.

[9] Note for our final regression sample, we additionally restrict to individuals born in months of October-December and January-March. Our regression methodology (discussed in Section V) relies on comparing individuals born right around the end of the calendar year, since DRC is determined by birth year. Restricting our sample in this manner allows us to compare those who are plausibly similar in nature with the exception of the DRC.

IV. TRENDS IN CLAIMING PATTERNS

A. Historical Claiming Patterns

By far the most popular choices among retired workers are to claim benefits immediately at age 62, the earliest possible claiming age, or to claim at the FRA (age 65 for all individuals in our sample). In fact, across these 1923 through 1933 birth cohorts, 55.7 percent claim at age 62. Meanwhile, on average 15.5 percent claim at age 65. Figure A1 displays the fraction of male individuals in our sample who claim retired worker benefits at each age (in years and months), with one graph for each year-of-birth between 1924 and 1933, inclusive. The share claiming at each age sums to one for each birth year. It is evident from this figure that across all 10 birth cohorts displayed, claiming immediately at age 62 and at the FRA are by far the most common. In fact, in the earlier birth groups, more than 40 percent claimed benefits at exactly age 62 and 0 months. Despite this fraction slightly waning over time, claiming at 62 remains by far the most popular choice for all 10 cohorts, with this group receiving 80 percent of the PIA.

As Figure A1 shows, claiming after the FRA is overall uncommon compared to claiming at the earliest eligibility age or at the full retirement age. Figure 2 illustrates the total fraction who claim on or after one year (i.e. 12 or more months) after their FRA by birth year. This share does trend gradually upward from 4.6 percent to 5.6 percent for the 1923 to 1933 birth cohorts. However, this fraction plummets to just 1.5 percent for the 1935 birth cohort, presumably as a result of the 2000 elimination of the earnings test for Social Security recipients at or older than the full retirement age. Some older workers presumably delayed claiming their retired worker benefits prior to this policy change to avoid having their benefits decreased due to the earnings test. But the elimination of that earnings test encouraged more workers to claim earlier. Therefore, this sharp change in claiming behavior further reiterates the need to make sample restrictions due to the introduction of the earnings test.

B. How Does DRC Affect Claiming?

We begin with simulations of the net present value of claiming Social Security retired worker benefits at different ages, to highlight two notions: (1) increases in the DRC makes claiming benefits after the FRA financially more attractive and (2) there is heterogeneity in claiming incentives, where those with longer life expectancies and lower discount rates have a stronger incentive to claim later. These simulations are similar in spirit to Coile et. al. (2002), Sun and Webb (2009), Shoven and Slavov (2014a), and Alleva (2016).

The stream of Social Security benefits can be modeled as an annuity, where an individual will receive constant interval installments for the rest of life. The decision to start benefits at age 62 (the earliest qualifying age) or at a later age is influenced by the fact that monthly benefits are adjusted by a rate schedule. Therefore, claiming at age 62 means that the retiree will start receiving benefits immediately, but the amount obtained each month will be less than, say, claiming at age 64, 66, or 68. However, waiting to age 68 presents other challenges: there are forgone benefits in the six years from age 62 through age 68, which implies the receipt of fewer years of benefits[10].

[10] One need not simultaneously claim Social Security benefits and retire from working, but for many, these decisions occur at the same time. This simulation abstracts from this and assumes those who claim later do not work for longer. Adding the tradeoff of working longer would make the simulation richer in nature. However, for the purposes of making the model as simple and transparent as possible to illustrate the relationship between claiming age and DRC, we believe including the working longer margin would only introduce complexity and not allow us to highlight these key channels. Working longer likely affects the worker's AIME/PIA and to the extent the earnings are subject to the earnings test, might in turn affect the DRC and actuarial adjustment factor, creating substantial complexity.

In the simplest possible formulation to highlight the impact of the DRC and other key parameters, we assume that individuals behave as if they seek to maximize the following expected present value of their Social Security benefit streams of claiming at different months:

$$\mathrm{EPV_{m,b}} = \left(\mathrm{PIA} \cdot \delta_{\mathrm{m,b}}\right) \sum_{\mathrm{t=m}}^{\mathrm{A}} [S_t (1 + \mathrm{r})^{-\mathrm{t}}]$$

In this equation, $m = 0,1, \ldots ,96$ is the number of months since age 62 that an individual can choose to claim benefits. PIA is the individual's primary insurance amount, and it is adjusted by the adjustment factor, $\delta_{m,b}$, which is dependent on the month of claiming and the birth year, b. This adjustment factor is either the early retirement factor for those who claim before the FRA, or the delayed retirement credit for claiming after the FRA. We abstract from the effects of the cost-of-living adjustment since it will be offset by inflation and thus the PIA (multiplied by the adjustment factor) remains constant for all of the remaining years. Then, we sum over the discounted cash flow of receiving the benefit at all future ages, at real discount rate r and up to the maximum age, A. However, we must also account for the probability of an individual being alive at each age, which is captured by the inner term, S_t. S_t. is the probability of an individual surviving to month *t* conditional on being alive at age 62 and 0 months ($S_0 = 1$). For example, for the age 62 and 0 month perspective, the Social Security benefit the individual will receive at age 70 and 0 months is the monthly benefit, discounted by the probability of surviving to age 70, and the real interest rate over the 8 years (captured by the $(1 + r)^{-t}$ term). The survival rates are obtained from the SSA historical and projected death probabilities for men, derived from the 2018 Trustees Report. We allow survival rates to differ by year of birth. The PIA adjustment factor is detailed in table 2.A20 of the 2019 *Annual Statistical Supplement*.

Our main simulation results can be summarized in Figure 3, where the expected present value of claiming at different ages, in age 62 dollars, is graphed for each claiming age and a select group of birth years. For simplicity the PIA is set for both birth years to $1000 with a 3% real discount rate (Coile et al 2002). We select two birth cohorts, 1923 and 1933, with DRC 3.0% and 5.5%, respectively. The vertical dashed lines delineate the full retirement age. As shown in the left panel, the drastic drop off in EPV after the FRA for the 1923 birth cohort that flattens out for the 1933 group suggests that the increases in the DRC substantially raise the EPV for a person with average mortality rates of claiming at later ages and thus helps make claiming later more actuarially fair. We also display the same information in Table 4, which shows the difference in the EPV at each age compared to claiming at age 65. As the first two columns illustrate, the loss in EPV for the 1933 birth cohort for claiming after 65 is smaller compared to the 1923 cohort.

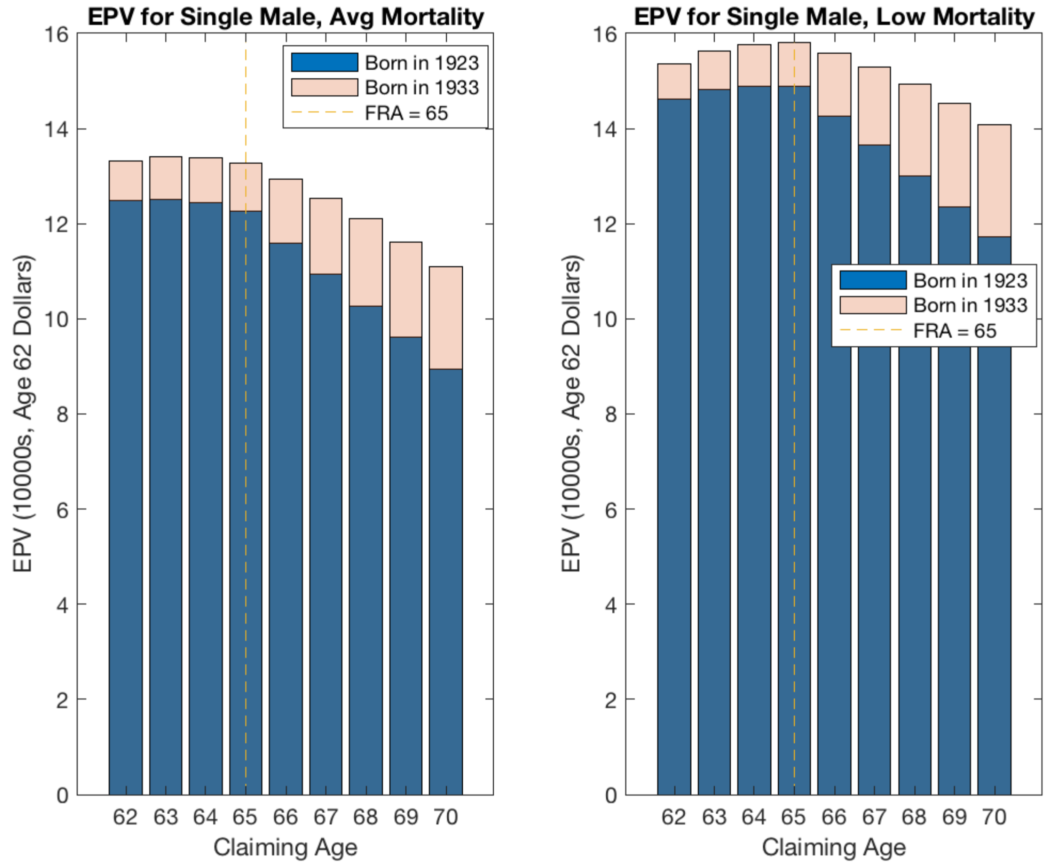

Figure 3. Expected Present Value of Claiming

Notes: The expected present value is the sum of Social Security benefit streams for claiming at different months. Average mortality profiles for each birth cohort were taken from the 2018 Trustees Report, while low mortality refers to a mortality profile with 64% of the death probabilities at each age for the average profile. Benefits are discounted by the real discount rate and survival probability at each age, calendar year, and sex. The PIA is fixed at $1000 and the real discount rate is 3%, annually. The DRC of the 1923 and 1933 birth cohorts are 3% and 5.5% respectively. Vertical lines delineate the full retirement age of 65.

Figure 3 also shows the EPV profiles for those with low mortality rate profiles and lower mortality is also associated with claiming later. Intuitively, for those with longer life expectancies, claiming later implies a larger monthly benefit for a longer period of time. Here, we simulate low mortality as an individual in each of the birth cohorts as having 64 percent of the average probability of death at each age.[11] These same EPV are also shown in columns (3) and (4) of Table 4. For a fixed birth cohort, increased life expectancy is associated with optimal claiming closer to age 65. In this table, we also vary the real discount rate, in columns (5) – (8). Low discount rates operate in the same manner as low mortality profiles, indicating the individual has stronger preferences for the future. In these simulations, the combination of low mortality and low discount rates can mean such that claiming after FRA is optimal, as shown in the last column.

Table 4: EPV Relative to Age 65 Claiming by Birth Year, Mortality, and Discount Rate

	Average Mortality		Low Mortality					
Discount Rate	3%		3%		1%		0.50%	
Birth Year	1923	1933	1923	1933	1923	1933	1923	1933
Age								
62	2230	330	-2710	-4470	-11350	-14080	-14150	-17200
63	2470	1240	-730	-1880	-6260	-8050	-8070	-10080
64	1690	1100	140	-420	-2500	-3380	-3370	-4360
65	0	0	0	0	0	0	0	0
66	-6630	-3410	-6150	-2260	-5480	-190	-5230	530
67	-13300	-7360	-12420	-5180	-11320	-1300	-10880	50
68	-19950	-11800	-18790	-8690	-17480	-3300	-16920	-1390
69	-26573	-16650	-25230	-12730	-23940	-6130	-23320	-3760
70	-33124	-21850	-31700	-17250	-30670	-9760	-30050	-7010

Note: Displays the difference in EPV at a particular age and at age 65 for each birth year, mortality profile, and real discount rate. The expected present value is the sum of Social Security benefit streams for claiming at different months. Average mortality profiles for each birth cohort were taken from the 2018 Trustees Report, while low mortality refers to a mortality profile with 64% of the death probabilities at each age for the average profile. Benefits are discounted by the real discount rate and survival probability at each age, calendar year, and sex. The PIA is fixed at \$1000 and the real discount rates are annual. The DRC of the 1923 and 1933 birth cohorts are 3% and 5.5% respectively.

Next in Figure 4 we illustrate how optimal EPV claiming varies by the individual's AIME. Income affects the EPV of Social Security benefits through two channels: (1) higher income translates into higher AIME and PIA, and (2) higher income individuals experience lower

[11] We chose the value 64% based off the mortality profiles of Bosley et. al. (2018), which documents mortality by AIME. While this paper does not produce mortality profiles by birth cohort and is limited to birth cohorts after 1930, we approximate the average mortality for the top quintile AIME for the 1933 birth cohort. Averaging across the age groups produce 64%. See Appendix A for more detailed information.

mortality rates. In this simulation, we fix the birth cohort at 1933 (where the FRA is 65 and the DRC is 5.5%) and the real discount rate at 3%, annually. Using our sample, we find the AIME quintile cutoffs by birth year for men, to illustrate the EPV of claiming at different ages by AIME quintile. Bosley et. al. (2018) produce mortality ratios for AIME quintiles, sex, year, and age, where the mortality ratio is the death rate (by single year of age) of a particular quintile divided by the death rate of the overall group. The overall pattern is that higher AIMEs have significantly lower death probabilities, but this difference decreases for older ages. Using these estimates, for each quintile, we obtain adjusted death probabilities, where we multiply the mortality ratio by the probability of death for a given calendar year and age[12]. The resulting simulation is shown in Figure 4, where the AIME of each quintile is the average of the cutoffs and the EPV is scaled by the PIA of the quintile. In comparing the lowest and highest quintiles, we can see that claiming later is more attractive for those with higher AIMEs, driven by the differential mortality ratios. We also note the fact that those with higher incomes are more likely to claim later. If this group also responds more to DRC increases, this may reflect a form of adverse selection - to the extent that those with higher incomes claim later further depleting Social Security trust funds, such an empirical finding might actually have negative consequences for the program's solvency.

[12] See Appendix A for detailed information on construction of the mortality profiles by AIME.

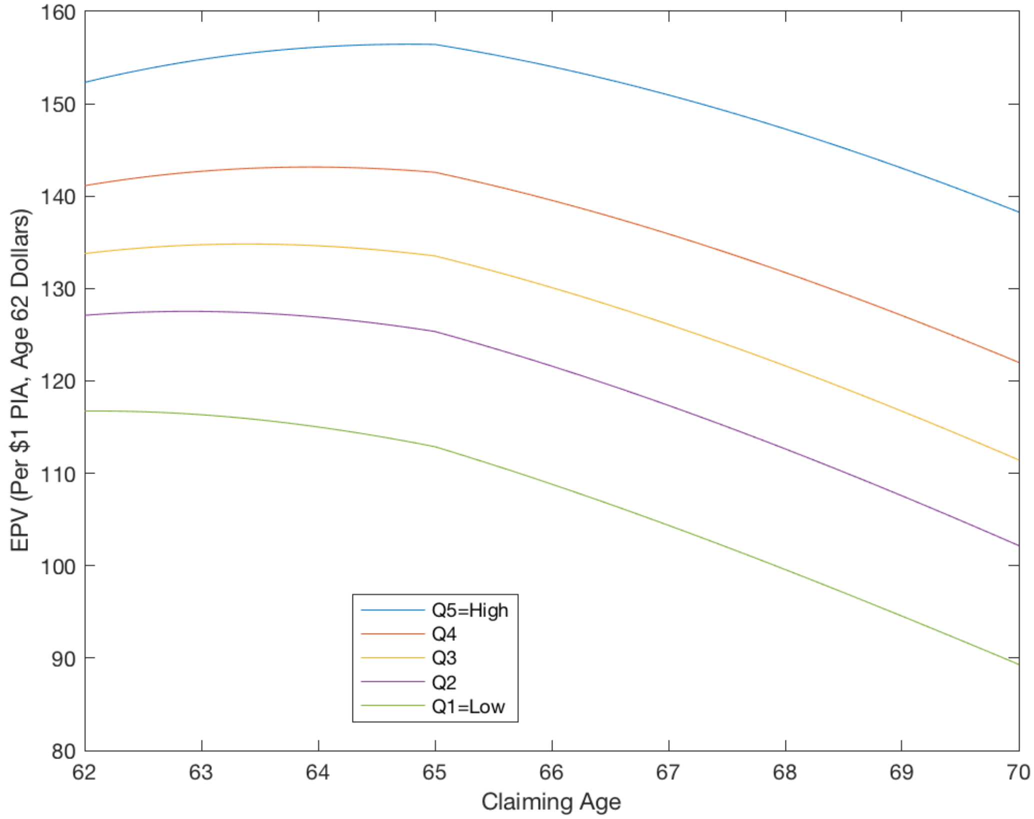

Figure 4. Expected Present Value (EPV) for AIME Quintiles

Notes: Simulations illustrate the expected present value by claiming age and AIME quintile. The EPV is the sum of Social Security benefit streams for claiming at different months. Average mortality profiles were taken from the 2018 Trustees Report, while mortality ratio by AIME quintile is constructed from Bosley et. al. (2018). Benefits are discounted by the real discount rate and survival probability at each age and calendar year. All simulations use birth year 1933, and real discount rate of 3%. PIA for each AIME quintile is derived from the average AIME of each AIME quintile cutoff.

In summary, our simple simulations reveal that the increases in the DRC makes claiming later more actuarially fair (or more than actuarially fair). Actuarially fair would imply that the EPV of claiming at different ages should be equivalent. In our simulations, we have shown that for later birth cohorts, who face higher DRC rates, their EPV profiles are "flatter" in nature, as compared with earlier birth cohorts who see sharp dropoffs in EPV if they were to claim after their full retirement age. In addition, those with longer life expectancies might even find it more than actuarially fair to claim after their FRA.

The financial incentive to claim later differs by individuals' expectations of future life expectancies and perceptions of discount rates. Therefore, on the margin, increases in the delayed retirement credit are likely to induce later claiming by some individuals. However, this setup assumes that individuals seek only to maximize the expected present value of their life-time Social Security benefits and that there are no frictions. No doubt a host of other factors also influence claiming decisions by OASI retired worker recipients that are currently not captured in the simulation. For example, individuals might face liquidity constraints or incomplete information on Social Security benefit rules. In Section VI, we include a discussion of some of these other factors when interpreting our empirical results.

V. EMPIRICAL STRATEGY

We estimate the causal effects of the increase in the delayed retirement credit by using a differences-in-differences identification strategy. In the ideal experiment, the DRC would be randomly assigned, so that we can attribute observed claiming behavior to differences in the DRC. In the absence of the random assignment, we can compare individuals who are plausibly very similar in nature, except for having different DRC rates. For those who are born around January 1st of a year, individuals born at the end of the previous calendar year are likely similar in nature to those born on or immediately after January 1st of the new calendar year. Since each individual's DRC is assigned by the birth year, we compare differences in claiming behavior between people born after January 1st of the year when DRC increased by 0.5 percentage points ("post" time period), and people born right before the January 1st cutoff who did not see a corresponding change in their DRC ("pre" time period). We assume that an individual's birth year is not manipulated for the purposes of obtaining a particular DRC rate 65 years in the future, which seems reasonable given that the legislation causing the DRC changes was not passed until 1983, several decades after all of the individuals in our sample were born.

To control for the possibility that claiming decisions are influenced by calendar month-of-birth, we add additional robustness to our empirical design by introducing placebo windows in which the DRC remained constant. Since DRC increased every two birth years, birth cohorts in the window that experienced DRC changes have two 'placebo control' windows with constant DRC before and after the change in DRC. To illustrate this, consider a birth window that includes those born in October, November, and December of 1926 and those born in January, February, and March of 1927, where the DRC increased from 3.5 percent for the 1926 birth cohort to 4.0 percent for 1927 birth cohort. A simple comparison of claiming patterns of those born in 1926 who had a 3.5 percent DRC with that of those born in 1927 who had a 4.0 percent DRC might be biased, given that there may be a month effect where people's likelihood of claiming in each month might differ across calendar months. To eliminate this month effect, we compare the claiming patterns of people in the window where DRC changes (1926-1927), with the corresponding claiming patterns of people born in the birth windows where the DRC does not change (1925-1926 and 1927-1928). Therefore, the 1926-1927 birth window is the "treated" unit, while the adjacent birth windows are the "control" units.

Each birth window where the DRC increases by 0.5 percentage points and the surrounding birth windows where the DRC does not change is an individual experiment and can be thought of as an individual difference-in-differences. We can pool together all 5 experiments that occur throughout our main sample. Our main sample includes birth years 1923-1933, where the DRC increases from 3.0 to 5.5 percent. We begin with focusing on a 3-month bandwidth around each January 1st cutoff, where each birth window runs from October-December, and January-March.

We estimate the average difference in claiming behavior between those who are born soon after January 1 with those born shortly before January 1 in those six-month windows when the DRC changes compared to windows where the DRC does not change with the following specification:

$$Y_i = \alpha + \delta_w + \gamma_m + \beta \cdot DRC\ Change_i \cdot PostJan1_i + X_i \theta + e_i$$

In this specification, Y_i is a claiming outcome of individual i, $DRC\ Change$ is a variable that is equal to 1 if an individual i was born in the window of six months around a January 1st where the DRC changes, $PostJan1$ takes on the value of 1 if individual i was born after the Jan 1st cutoff and zero otherwise. Therefore, β is our coefficient of interest, which captures the effect of experiencing a 0.5 percentage point higher DRC, compared to those born in comparable windows around January who did not experience changes in DRC. We also include fixed effects for each January window, δ_w, to account for the possibility that claiming patterns are changing over time due to macroeconomic conditions, social norms, or other factors. There are also a full set of six calendar month-of-birth fixed effects, γ_m. We also include a set of individual-level controls X_i, such as the individual's PIA decile and race. This specification pools together all 5 DRC changes in our sample, as well as pooling together all January windows where the DRC does not change as the effective control groups.

Our current specification treats all DRC changes symmetrically, regardless of the base DRC rate. As a result, increases from 3.0 to 3.5 percent are treated the same as increases from 5.0 to 5.5 percent. Given the EPV simulations described above, it seems plausible that the changes at lower DRCs may be less likely to influence behavior than the changes at higher DRCs. We further split our sample into DRC changes from 3.0% to 4.5% and 4.5% to 5.5% to understand whether our overall pooled results may be driven by later DRC increases or whether there are consistent effects throughout. In other words, this serves to test whether the effects of DRC increases are non-linear in nature.

Our main outcome variable of interest is an indicator variable for claiming at least one year after FRA (i.e. claiming on or after the month in which an individual reaches age 66). We test for the presence of adverse selection in claiming by extending our analysis to subsamples of individuals with different PIAs. As predicted in the simulations, those with higher PIAs (and thus, the lowest mortality rates), have stronger incentives to claim later. We explore heterogeneous DRC effects for those in the above median PIA, below median PIA, and top decile PIA.

Table 5. Summary Statistics - Men

	Mean	St. Dev	Min	Max	N
Claim at 66+	0.043	0.203	0	1	387265
Birth Year	1928	2.901	1923	1933	387265
DRC Changes in Window	0.499	0.500	0	1	387265
Birth Month: Jan, Feb, Mar	0.506	0.500	0	1	387265
Black	0.076	0.266	0	1	387265
Other Race	0.025	0.156	0	1	387265
PIA	486.48	177.57	5.79	1262.43	387265

Note: Summary statistics for our main regression sample, derived from the 10% Social Security administrative data. Includes men in birth cohorts 1923-1933 for those born in months Jan-March and Oct-Dec. A Birth Window is a 6-month window around January of a calendar year (Oct-March). The PIA is the Primary Insurance Amount, derived from the age 62 AIME.

In Table 5, we tabulate summary statistics of the main variables for the regression sample. We have a total of 387,265 men in our sample who were born between 1923 and 1933. Out of the 10 birth windows (e.g. 1923-24, 1924-25, etc.) between 1923 and 1933, where birth year changes from October of one year to March of the next year, there are five windows in which the DRC changes by 0.5 percentage points. In this regression sample, 4.3% of all people claim on or after the month in which they reach age 66. The average primary insurance amount (PIA) (adjusted to

2012 dollars using the GDP deflator) is \$486.48 for our sample. While this amount may seem relatively low compared with the average 2019 monthly retired worker benefit of about \$1,500, it is worth remembering that the maximum earnings subject to social security taxes was quite low during the 1950s and 1960s. For example, in 1954 average annual earnings were \$3,156 while the maximum taxable earnings were just \$3,600. The corresponding numbers in 1965 were \$4,659 and \$4,800, respectively. In contrast in 2019 the maximum taxable wage was 2.5 times higher than average annual earnings (\$132,900 versus \$54,100). The relatively low taxable maximum in earlier years depressed the AIME (and thus the PIA) for any workers with relatively high earnings below what it would have been had the average earnings to maximum taxable wage ratio been similar to that of recent years. In the analysis, we use PIA decile dummies as controls. PIA is calculated at age 62 for every individual in the sample, such that working past 62 does not impact the decile variable.

VI. EMPIRICAL RESULTS

A. Results

We present our main results from our difference-in-differences estimation strategy, shown in Table 6. Columns (1) – (4) reflect the results using the full 1923 - 1933 birth cohorts for men, where the DRC increased from 3.0 percent to 5.5 percent, but the FRA remained constant at age 65. Furthermore, the 2000 earnings test change would not have affected the decision to claim one or more years after FRA for any individuals in our sample. The outcome variable is an indicator variable that equals one if the individual claims retired worker benefits in the month that the individual reaches age 66 or later and is otherwise equal to zero. Each column adds control variables so that the fourth specification includes calendar month fixed effects, window fixed effects, and individual-level control variables including race and ten indicator variables for each cohort-specific PIA decile.

Table 6. The Effect of Increase in Delayed Retirement Credit on Claiming Behavior of Men

Outcome: Claim 1+ Year Post FRA

Sample	Full				Below Median PIA	Above Median PIA	Top PIA Decile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Change Window x Post Jan	.0025*	.0024*	.0025*	.0023*	.0021	.0025	.0065
	(.0013)	(.0013)	(.0013)	(.0013)	(.0017)	(.0020)	(.0062)
Change Window	-.0000	-.0000					
	(.0009)	(.0009)					
Post Jan	-.0018**						
	(.0009)						
Black				.0279***	.0248***	.0368***	.0602***
				(.0015)	(.0016)	(.0032)	(.0166)
Other Race				.0308***	.0271***	.0420***	-.0032
				(.0026)	(.0029)	(.0059)	(.0163)
Constant	.0435***	.0434***	.0478***	.0442***	.0467***	.0292***	.0995***
	(.0007)	(.0009)	(.0013)	(.0017)	(.0020)	(.0021)	(.0062)
Month FE		x	x	x	x	x	x
Window FE			x	x	x	x	x
PIA Decile FE				x	x	x	x
Mean Dep. Var	.0432	.0432	.0432	.0432	.0359	.0505	.1050
N	387,265	387,265	387,265	387,265	193,198	194,067	38,830

*** p<0.01, ** p<0.05, * p<0.1

Note: Data derived from the 10% Social Security Administration data, with restrictions outlined in Table 2. Columns (1) - (4) include the full sample of men in the 1919-1933 birth years and subsequent columns restrict to samples as indicated. Full Sample refers to the inclusion of all PIA deciles. All regressions utilize a 6-month window around January, thus including only those born in October- December and January-March. Those born in months without a full DRC window are also removed; for example, we remove those born in 1919 January - March, since there are no corresponding Oct-Dec observations. Mean Dep. Var is the average fraction claiming at 66+ for those in windows where DRC changes, born in Oct-Dec for the relevant samples. Robust standard errors in parenthesis.

The estimated effect of the DRC increase in all four specifications is positive, suggesting that – as expected – men are more likely to delay claiming when there is a greater financial incentive to do so. The coefficient of interest is on the interaction term of "Change Window" and

"Post Jan", which is the effect of experiencing a 0.5 percentage point increase in the DRC, compared to birth windows in which the DRC remains the same. We note that this coefficient remains stable even with the inclusion of additional controls. Taking column (4) as the most conservative specification, we estimate that a 0.5 percentage point increase in the DRC increases claiming at age 66+ by 0.23 percentage points, and this is statistically significant at the 10% level. Given that 4.3 percent of those born before January in windows where the DRC changes claim at age 66 or over, this represents a 5.32 percent increase from the mean. We additionally include regression specifications for women in the appendix. As discussed in the sample selection section, the claiming decisions of women in this sample are likely confounded with large composition effects of who decides to be the primary claimant, and that the claiming decisions of women are likely linked to that of their spouses, given that women tended to have lower lifetime earnings.

As predicted in our simulations, those with higher PIAs might be more affected by increases in the DRC, since claiming later is more attractive for these individuals who will obtain a larger monthly benefit for a longer period of time (since they face lower mortality rate profiles). We also test whether we observe this form of adverse selection in the data. Columns (5) – (7) of Table 6 displays results of regression specifications where we restrict to the below median PIA, above median PIA, and top decile PIA, respectively. We find that DRC did not have a significant effect on claiming behavior for any of the PIA groups. While this result may look puzzling, we attribute the loss of statistical significance in the aggregate effect to the reduction in number of observations in each specification. Since the magnitudes of the effects are relatively small and close to zero, one needs more statistical power, and thus a greater number of observations to avoid making a type II error. In terms purely of magnitudes, we find suggestive evidence of adverse selection, as our point estimates indicate that a 0.5 percentage point increase in DRC increased the share of men claiming later for those with PIA below median by 0.21 percentage points, while for those in the top PIA decile, the share claiming later was increased by 0.65 percentage points.

Table 7. Effect of Change of DRC on Claiming Behavior of Men

Outcome: Claim 1+ Year Post FRA

Change in DRC	3 → 4.5					4.5 → 5.5				
Birth years with DRC Change	1924-25, 1926-27, 1928-29					1930-31, 1932-33				
Control birth years	1923-24, 1925-26, 1927-28, 1929-30					1929-30. 1931-32, 1933				
Sample	Full		Below Median PIA	Above Median PIA	Top PIA Decile	Full		Below Median PIA	Above Median PIA	Top PIA Decile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Change Window x Post Jan	.0015	.0012	.0009	.0016	.0013	.0040*	.0037*	.0024	.0051	.0171
	(.0015)	(.0015)	(.0020)	(.0023)	(.0072)	(.0022)	(.0022)	(.0028)	(.0033)	(.0108)
Change Window	-.0090***	-.0089***	-.0109***	-.0069***	.0059	.0093***	.0086***	.0050**	.0122***	.0287***
	(.0017)	(.0016)	(.0022)	(.0025)	(.0074)	(.0019)	(.0019)	(.0024)	(.0029)	(.0090)
Post Jan	-.0004	.0007	.0002	.0012	.0143**	-.0026	-.0014	-.0009	-.0019	.0105
	(.0015)	(.0015)	(.0019)	(.0022)	(.0068)	(.0022)	(.0021)	(.0027)	(.0033)	(.0108)
Black		.0289***	.0274***	.0335***	.0494**		.0283***	.0229***	.0436***	.0808***
		(.0017)	(.0020)	(.0038)	(.0195)		(.0024)	(.0026)	(.0053)	(.0266)
Other Race		.0307***	.0271***	.0417***	-.0161		.0303***	.0271***	.0388***	.0143
		(.0032)	(.0035)	(.0073)	(.0183)		(.0039)	(.0043)	(.0088)	(.0265)
Constant	.0480***	.0428***	.0441***	.0306***	.0846***	.0441***	.0402***	.0442***	.0198***	.1027***
	(.0014)	(.0018)	(.0022)	(.0023)	(.0060)	(.0017)	(.0024)	(.0027)	(.0028)	(.0081)
Month FE	x	x	x	x	x	x	x	x	x	x
Window FE	x	x	x	x	x	x	x	x	x	x
PIA Decile FE	x	x	x	x	x	x	x	x	x	x
Mean Dep. Var	.0408	.0408	.0347	.0467	.0940	.0472	.0472	.0377	.0567	.1250
N	273,713	273,713	136,385	137,328	27,378	152,449	152,449	76,289	76,160	15,218

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Note: Data derived from the 10% Social Security Administration data, with restrictions outlined in Table 2. Columns (1) - (5) includes all birth cohorts for which DRC increased from 3 to 4.5% (1923-1930) and columns (6)- (10) includes all birth cohorts for which DRC increased from 4.5 to 5.5% (1929-1933). Full Sample refers to the inclusion of all PIA deciles. All regressions utilize a 6-month window around January, thus including only those born in October- December and January-March. Those born in months without a full DRC window are also removed; for example, we remove those born in 1919 January - March, since there are no corresponding Oct-Dec observations. Mean Dep. Var is the average fraction claiming at 66+ for those in windows where DRC changes, born in Oct-Dec for the relevant sample. Robust standard errors in parenthesis.

We also find suggestive evidence, as displayed in Table 7, for the presence of non-linearities in the effect of the DRC. Our simulations reveal that raising the DRC by 0.5 percentage points when it is already fairly low (such as at 3 percent) might not be enough to shift individuals into claiming later. This suggests that later increases in the DRC would potentially shift the behavior of more individuals who are closer to the margin of delaying claiming. In Table 7, we disaggregate the five total “mini-experiments” or DRC changes, into changes in the DRC from 3 to 4.5 percent (columns (1) – (5)) and changes in DRC from 4.5 to 5.5 percent (columns (6) – (10)). In addition to the birth windows where DRC increased, we have included the surrounding birth windows where the DRC remained constant as the controls. Using our most preferred conservative specification in column (2), we see that increases in DRC by 0.5 percentage points had no effect on the probability of claiming later for the 3.0 to 4.5 percent change group. On the other hand, increases in the DRC did have a significantly positive effect for the two later DRC changes. In column (7), we see that for the DRC increases from 4.5 to 5.5 percent, a 0.5 percentage point increase in the DRC led to a 0.37 percentage point increase in claiming at 66 or later and given a base of 4.72 percent claiming at 66 or later, this constituted a 7.9 percent increase. Furthermore, consistent with our findings for the full pooled sample, the effects of the DRC are concentrated among the high earners, albeit this is suggestive evidence due to the absence of statistical significance given the reduced number of observations. In terms purely of magnitudes, those with a PIA below the median see an increase in the share claiming after age 66 by 0.24 percentage points, while those in the top PIA decile see an increase by 1.71 percentage points.

The 1983 Social Security amendments substantially increased the benefits of claiming later through changes in the Delayed Retirement Credit, and this study reveals that men are modestly responsive in their claiming decisions, with most pronounced effects among the highest earners and among the later DRC changes. Using Table 7, we can assume no effect of the DRC for the 3 to 4.5 changes and then assume an effect of 0.37 percentage points for each 0.5 percentage point increase in DRC from 4.5 up through 5.5 percent. This suggests the changes in DRC from 3 to 5.5

percent increased the probability of claiming at age 66 by 0.74 percentage points. Given that 4.56 percent claimed at 66 or over in the 1923 birth cohort and this increased to 5.64 percent by the 1933 birth cohort, the DRC explains 69 percent of the total increases in claiming later.[13] [14]

B. Discussion

We highlight that informational frictions could be dampening the effects of DRC increases. Liebman and Luttmer (2012) conducted a survey in 2008 to those between the ages of 50 and 70 to understand knowledge of Social Security program rules and to test interventions. While respondents largely knew that later claiming produces larger monthly benefits, they found that individuals were "unaware that benefit increases from delaying claiming are higher between the full-benefit age (generally age 66 in our sample) and age 70 than between age 62 and the full retirement age" and that individuals perceived a benefit increase in claiming between the ages of 70 and 74, which is not the case. Therefore, knowledge regarding the DRC might be quite limited or at least incomplete and thus reduce responsiveness to this change in financial incentives.

While we observe that changes in the DRC do encourage some men to claim retirement benefits later on the margin, we note that the overall rates of claiming after the Full Retirement Age are still quite low – as shown in Figures 2 and A1, the most popular age to claim by far is immediately at age 62, with another spike at the Full Retirement Age, and claiming later is overall rare. Therefore, while the DRC does mildly promote later claiming, we note the overall low rates of claiming later and highlight other factors at play.

[13] Our study focuses on the 1923-1933 birth cohorts where DRC increased from 3 to 5.5 percent, yet the DRC increased further to 8 percent by the 1943 birth cohorts. Ideally, we would like to present back of the envelope calculations of how much of the total change in claiming later for the 1923 through 1943 birth cohorts can be explained by the increase in DRC from 3 to 8 percent. However, the observed rates of claiming for the 1935 and later birth cohorts are affected by the change in the earnings test, which substantially decreased rates of claiming later. In recent years, rates of later claiming have risen substantially, perhaps partly due to rising life expectancies and lower liquidity constraints.

[14] The increase in probability of claiming later by 0.74 is obtained by doubling 0.37. 69% = 0.74/(5.64-4.56)

In the context of the simulations presented in this paper, claiming is a function of the real discount rate and individual's expectations of their own health and life expectancies. High discounting of the future and beliefs regarding poorer health could be especially important. For example, Goda et al (2018) found that many who claim earlier have sufficient liquidity to delay, but that those who claim earlier have worse self-reported health. In addition, individuals might be maximizing lifetime benefits subject to market frictions, which our simulations do not currently take into account. For example, liquidity and credit constraints might prevent claiming later. Consistent with this, Shoven et al (2017) fielded a survey and found that the most common reason for claiming before the FRA was that individuals needed the money.

VII. CONCLUSION

Social Security represents the most important source of income for most elderly Americans. A number of changes have been made to the program during the last several decades. Arguably the most important set of changes were included in the 1983 amendments to Social Security, which increased the payroll tax rate and the full retirement age while also increasing the DRC. While this last provision actually represented an *increase* in the amount of benefits by increasing the actuarial adjustment beyond FRA, policymakers hoped it would also improve Social Security's financial standing. The most plausible mechanism for such an improvement was to lead individuals to claim later than they otherwise would have and to work longer as a result. At the time of the Amendments, Social Security benefits would be decreased for those working beyond FRA who had already claimed benefits.

We show in this paper that men did change the timing of their claiming of retired worker benefits in response to the large increases in the DRC. We take care to disentangle the effect of the DRC increase from other policy changes that took effect during our study period, including the increase in the full retirement age and the elimination of the earnings test for those receiving Social Security at or beyond their FRA. However, those who were most responsive to the DRC appear to be those with the largest lifetime earnings and highest expected present value of Social Security benefits. This adverse selection in response to claiming incentives predicts that the DRC increases may actually have increased Social Security payments to workers and might even imply negative consequences to the OASI trust fund's finances.

It is worth noting that over the last several years, there has been a substantial increase in the fraction claiming benefits beyond the FRA. For example, from 2009 through 2019, the fraction of male retired workers claiming one or more months after their full retirement age (66 for essentially all post-FRA claimants during this period) increased from 4.1 percent to 16.2 percent after having remained relatively unchanged in the preceding several years.[15] However, the DRC reached 8 percent for the 1943 and subsequent birth cohorts. Therefore, the DRC increases would essentially have been fully phased in for those 66+ and thus the DRC is unlikely to be the primary driving force for the 2009 to 2019 increase. A variety of other factors could have contributed to this increase in claiming later. The economic recovery following the Great Recession could have encourage later claiming to the extent that individuals desired to work longer following lost income or take on jobs due to better employment prospects. In addition, life expectancy continued to rise. Lastly, it could also be the case that it takes time for individuals to learn about the DRC or adopt new norms in delaying claiming. Future research should explore the causes along with the consequences of this change given the importance of Social Security benefits to most elderly Americans.

[15] Estimates from 2020 Social Security Administration Annual Statistical Supplement, Table 6.B5 – Number, average age, and percentage distribution, by sex and age, selected years 1940-2019

REFERENCES

Alleva, B.J., 2016. Discount rate specification and the social security claiming decision. *Social Security Bulletin*, *76*, p.1.

Behaghel, L. and Blau, D.M., 2012. Framing social security reform: Behavioral responses to changes in the full retirement age. *American Economic Journal: Economic Policy*, *4*(4), pp.41-67.

Benitez-Silva, H. and Yin, N., 2009. An Empirical Study of the Effects of Social Security Reforms on Benefit Claiming Behavior and Receipt Using Public-Use Administrative Microdata. *Social Security Bulletin*, *69*(3).

Bosley, T., Morris, M. and Glenn, K., 2018. Mortality by Career-Average Earnings Level. *Actuarial Study*, *124*.

Coile, C., Diamond, P., Gruber, J. and Jousten, A., 2002. Delays in claiming social security benefits. *Journal of Public Economics*, *84*(3), pp.357-385.

Deshpande, M., Fadlon, N. and Gray, C., 2020. How sticky is retirement behavior in the U.S.? Responses to Changes in the Full Retirement Age. (No. w27190). National Bureau of Economic Research.

Duggan, J.E. and Soares, C.J., 2002. Actuarial nonequivalence in early and delayed social security benefit claims. *Public Finance Review*, *30*(3), pp.188-207.

Duggan, M., Singleton, P. and Song, J., 2007. Aching to retire? The rise in the full retirement age and its impact on the social security disability rolls. *Journal of Public Economics*, *91*(7-8), pp.1327-1350.

Engelhardt, G.V., Gruber, J. and Kumar, A., 2020. Early Social Security Claiming and Old-Age Poverty: Evidence from the Introduction of the Social Security Early Eligibility Age. *Journal of Human Resources*, pp.0119-9973R1.

Finkelstein, A. and Poterba, J., 2004. Adverse selection in insurance markets: Policyholder evidence from the UK annuity market. *Journal of Political Economy*, *112*(1), pp.183-208.

Gelber, A.M., Jones, D., Sacks, D.W. and Song, J., Forthcoming. The employment effects of the social security earnings test. *Journal of Human Resources*

Goda, G.S., Ramnath, S., Shoven, J.B. and Slavov, S.N., 2018. The financial feasibility of delaying Social Security: evidence from administrative tax data. *Journal of Pension Economics and Finance*, *17*(4), pp.419-436.

Heiland, F. and Yin, N., 2014. Have we finally achieved actuarial fairness of social security retirement benefits and will it last?. *Michigan Retirement Research Center Research Paper*, (2014-307).

Liebman, J. and Luttmer, E., 2012. The Perception of Social Security Incentives for Labor Supply and Retirement: The Median Voter Knows More Than You'd Think. *Tax Policy and the Economy*, *26*, pp.1-42.

Martin, P.P., 2016. Why Researchers Now Rely on Surveys for Race Data on OASDI and SSI Programs: A Comparison of Four Major Surveys. *Research and Statistics Note*, (2016-01).

Mastrobuoni, G., 2006. The Social Security Earnings Test Removal. Money Saved or Money Spent by the Trust Fund? (Vol. 25). Industrial Relations Section, Princeton University.

Mastrobuoni, G., 2009. Labor supply effects of the recent social security benefit cuts: Empirical estimates using cohort discontinuities. *Journal of Public Economics*, *93*(11-12), pp.1224-1233.

Mastrobuoni, G., 2011. The role of information for retirement behavior: Evidence based on the stepwise introduction of the Social Security Statement. *Journal of Public Economics*, *95*(7-8), pp.913-925.

National Commission on Social Security Reform, 1983. Report of the National Commission on Social Security Reform. *Social Security Bulletin*, *46*(2), pp.2-7.

Pingle, Jonathan, F., 2006. Social security's delayed retirement credit and the labor supply of older men. Finance and Economics Discussion Series 2006-37. Board of Governors of Federal Reserve System (U.S.).

Purcell, P.J., 2020. Employment at Older Ages and Social Security Benefit Claiming, 1980–2018.

Rohwedder, S. and van Soest, A., 2006. The impact of misperceptions about Social Security on saving and well-being. *Michigan Retirement Research Center Research Paper No. WP*, *118*.

Shoven, J.B. and Slavov, S.N., 2014a. Does it pay to delay social security?. *Journal of Pension Economics and Finance*, *13*(2), pp.121-144.

Shoven, J.B. and Slavov, S.N., 2014b. Recent changes in the gains from delaying social security. *Journal of Financial Planning* 27 (3), pp. 32–41

Shoven, J.B., Slavov, S.N. and Wise, D.A., 2017. Social Security claiming decisions: survey evidence (No. w23729). National Bureau of Economic Research.

Social Security Administration. 2018. *2018 Trustees Report*

Social Security Administration. 2019. *Annual Statistical Supplement*

Social Security Administration. 2020. *Annual Statistical Supplement*

Social Security Administration. 2020. *2020 Trustees Report*

Song, J.G. and Manchester, J., 2007. How have people responded to changes in the retirement earnings test in 2000. *Social Security Bulletin*, *67*, p.1.

Sun, W. and Webb, A., 2009. "How Much Do Households Really Lose by Claiming Social Security at Age 62?" *Available at SSRN 1408729*.

APPENDIX

A. Obtaining Mortality Profiles in Expected Present Value Simulations

In Section IV, we simulate the expected present value of claiming at different ages to illustrate the effect of DRC increases and uncover heterogeneity in responses. In particular, we focus on the how mortality factors into the EPV. For these simulations, mortality profiles are drawn from Bosley et al. (2018). In this paper, the authors relate lifetime earnings to mortality, identifying how mortality rates vary for those with different AIME. In Appendix A, Tables 4 through 8, the authors show the relative mortality ratios for retired worker beneficiaries, by sex and age group. The relative mortality ratio is the death rate of a particular AIME group relative to all individuals of that sex and age group. For example, in Table 4, men in the lowest AIME quintile and of age 62-64 in 1995 have 1.65 times the death rate of an average male in 1995 of age 62-64. Subsequent tables show these same results for different calendar years.

Ideally, we would like to have the mortality ratios by birth cohort, AIME quintile, and age. Unfortunately, these tables are not by birth cohort, but rather, calendar year. These tables track the 1995 through 2015 calendar years, for groups of ages up to age 84. Therefore, we chose the 1933 cohort such that we can track the mortality ratios for this cohort for different age groups across the different tables. We use the mortality ratios in the 1995 table for the 62-64 age group (since the 1933 birth cohort would be 62 in 1995), the 2000 table for ages 65-69 (the 1993 group would be 67 in 2000), and so on. Given that the tables only track ages up to 84, the 80-84 mortality ratios are used for all ages 80+. The final adjusted death probabilities are the product of the probability of death at a particular age and the mortality ratio for that age. For example, the quintile adjusted probability of death for a 63 year old born in 1933 is the death probability in 1996 for a 63 year old (derived from the 2018 Trustees Report) multiplied by the ratios in the 1995 table for the 62-64 age group.

As for characterizing "low mortality" in Figure 3 and Table 4, we again use the 1933 birth cohort mortality profile. We define "low mortality" as the mortality profile of the highest AIME quintile. For simplicity, we average across all the age groups, producing that on average, the highest AIME quintile has 64% of the death rate compared with the average. Therefore, we multiply all the single year of age death probabilities by 0.64.Unfortunately, we cannot back out estimates of the mortality ratios for the 1923 cohort, since the tables of Bosely et al. (2018) are only calculated for the 1930 cohorts and later. For purposes of this simulation, we also give the 1923 birth cohort the same 64% value.

Figure A1. Social Security Claiming Patterns for Birth Years 1924-1933 (Men)

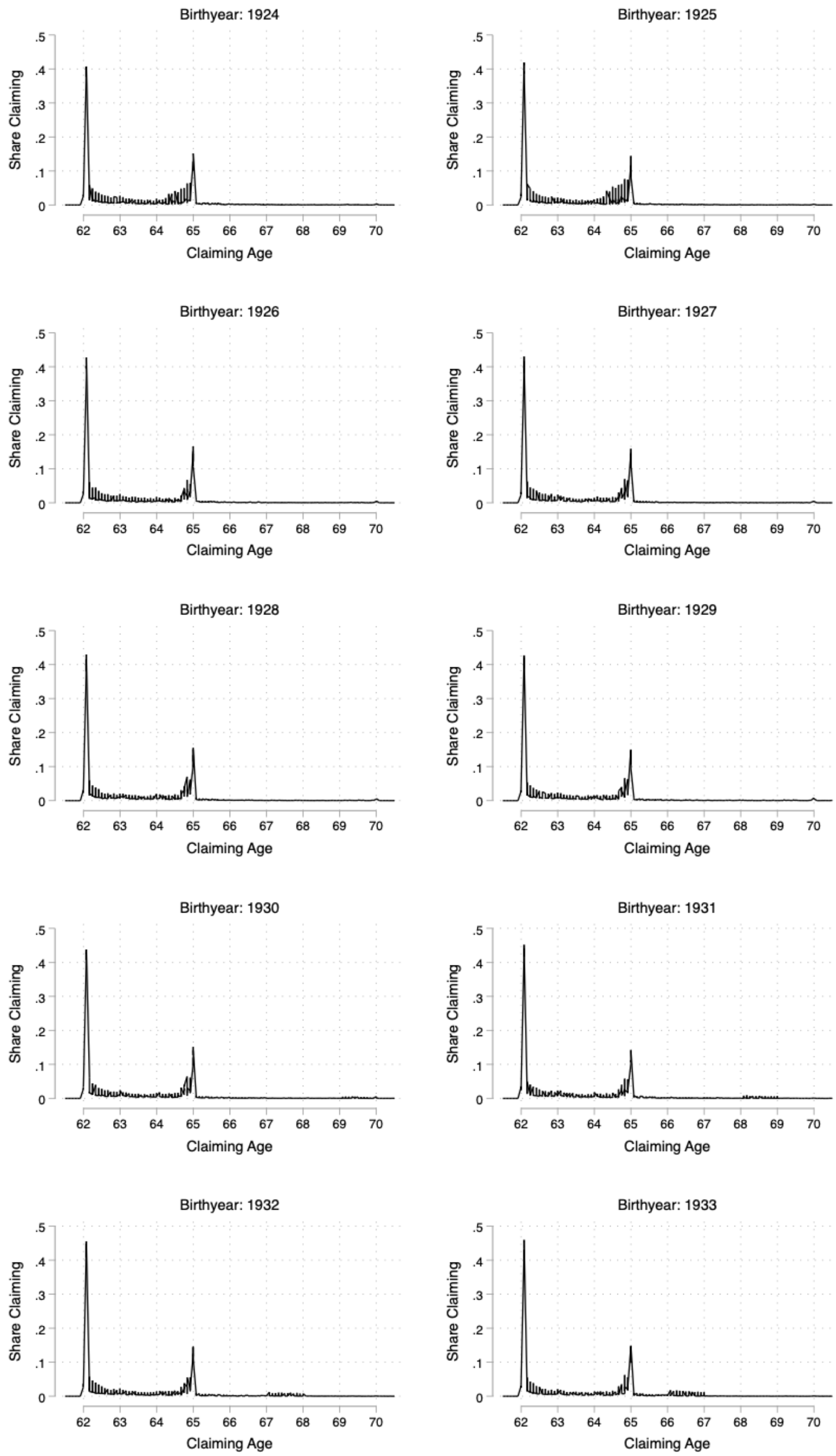

Table A1. Summary Statistics - Women

	Mean	St Dev	Min	Max	N
Claim at 66+	0.051	0.220	0	1	286345
Birth Year	1928	2.913	1923	1933	286345
Birth Window Number	9.572	2.872	5	14	286345
DRC Changes in Window	0.501	0.500	0	1	286345
Birth Month: Jan, Feb, Mar	0.508	0.500	0	1	286345
Black	0.089	0.285	0	1	286345
Other Race	0.023	0.150	0	1	286345
PIA	300.67	151.84	10.61	1086.12	286345

Note: Summary statistics for our main regression sample, derived from the 10% Social Security administrative data. Includes women in birth cohorts 1923-1933 for those born in months Jan-March and Oct-Dec. A Birth Window is a 6-month window around January of a calendar year (Oct-March). The PIA is the Primary Insurance Amount, derived from the age 62 AIME.

Table A2. The Effect of Increase in Delayed Retirement Credit on Claiming Behavior of Women

Outcome: Claim 1+ Year Post FRA

Sample	Full				Below Median PIA	Above Median PIA	Top PIA Decile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Change Window x Post Jan	-.0034**	-.0034**	-.0035**	-.0036**	-.0028	-.0042	.0011
	(.0016)	(.0016)	(.0016)	(.0016)	(.00190)	(.0027)	(.0056)
Change Window	.0050***	.0050***					
	(.0012)	(.0012)					
Post Jan	.0023**						
	(.0012)						
Black				.0278***	.0336***	.0208***	.0249***
				(.0017)	(.0021)	(.0027)	(.0062)
Other Race				.0227***	.0240***	.0216***	.0471***
				(.0032)	(.0038)	(.0053)	(.0126)
Constant	.0482***	.0518***	.0400***	.0053***	.0103***	.0399***	.0553***
	(.0008)	(.0012)	(.0015)	(.0017)	(.0019)	(.0027)	(.0056)
Month FE		x	x	x	x	x	x
Window FE			x	x	x	x	x
PIA Decile FE				x	x	x	x
Mean Dep. Var	.0510	.0510	.0510	.0510	.0336	.0684	.0603
N	286,345	286,345	286,345	286,345	142,967	143,378	28,633

*** p<0.01, ** p<0.05, * p<0.1

Note: Data derived from the 10% Social Security Administration data, with restrictions outlined in Table 2. Columns (1) - (4) include the full sample of women in the 1919-1933 birth years and subsequent columns restrict to samples as indicated. Full Sample refers to the inclusion of all PIA deciles. All regressions utilize a 6-month window around January, thus including only those born in October- December and January-March. Those born in months without a full DRC window are also removed; for example, we remove those born in 1919 January - March, since there are no corresponding Oct-Dec observations. Mean Dep. Var is the average fraction claiming at 66+ for those in windows where DRC changes, born in Oct-Dec for the relevant samples. Robust standard errors in parenthesis.

Table A3. Effect of Change of DRC on Claiming Behavior of Women

Outcome: Claim 1+ Year Post FRA										
Change in DRC	3 → 4.5					4.5 → 5.5				
Birth years with DRC Change	1924-25, 1926-27, 1928-29					1930-31, 1932-33				
Control birth years	1923-24, 1925-26, 1927-28, 1929-30					1929-30. 1931-32, 1933				
Sample	Full		Below Median PIA	Above Median PIA	Top PIA Decile	Full		Below Median PIA	Above Median PIA	Top PIA Decile
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Change Window x Post Jan	-.0048**	-.0049***	-.0022	-.0076**	.0020	.0005	.0002	-.0042	.0048	-.0043
	(.0019)	(.0019)	(.0022)	(.0031)	(.0066)	(.0028)	(.0028)	(.0032)	(.0045)	(.0093)
Change Window	-.0048**	-.0049***	-.0022	-.0076**	.0020	.0005	.0002	-.0042	.0048	-.0043
	(.0019)	(.0019)	(.0022)	(.0031)	(.0066)	(.0028)	(.0028)	(.0032)	(.0045)	(.0093)
Post Jan	.0030*	.0031*	.0010	.0051*	-.0036	-.0009	.0000	-.0012	.0012	.0056
	(.0018)	(.0018)	(.0021)	(.0030)	(.0064)	(.0028)	(.0028)	(.0033)	(.0045)	(.0092)
Black		.0286***	.0307***	.0258***	.0328***		.0261***	.0416***	.0106***	.0073
		(.0020)	(.0024)	(.0032)	(.0079)		(.0028)	(.0038)	(.0041)	(.0084)
Other Race		.0218***	.0242***	.0193***	.0417***		.0252***	.0237***	.0277***	.0627***
		(.0038)	(.0047)	(.0063)	(.0153)		(.0048)	(.0055)	(.0085)	(.0199)
Constant	.0373***	.0047***	.0109***	.0348***	.0589***	.0584***	.0172***	.0178***	.0633***	.0526***
	(.0016)	(.0018)	(.0021)	(.0029)	(.0061)	(.0022)	(.0025)	(.0028)	(.0041)	(.0071)
Month FE	x	x	x	x	x	x	x	x	x	x
Window FE	x	x	x	x	x	x	x	x	x	x
PIA Decile FE	x	x	x	x	x	x	x	x	x	x
Mean Dep. Var	.0458	.0458	.0299	.0617	.0558	.0610	.0610	.0406	.0814	.0671
N	197,553	197,553	98,478	99,075	19,826	117,629	117,629	58,884	58,745	11,651

*** p<0.01, ** p<0.05, * p<0.1

Note: Data derived from the 10% Social Security Administration data, with restrictions outlined in Table 2. Columns (1) - (5) includes all birth cohorts for which DRC increased from 3 to 4.5% (1923-1930) and columns (6)- (10) includes all birth cohorts for which DRC increased from 4.5 to 5.5% (1929-1933). Full Sample refers to the inclusion of all PIA deciles. All regressions utilize a 6-month window around January, thus including only those born in October-December and January-March. Those born in months without a full DRC window are also removed; for example, we remove those born in 1919 January - March, since there are no corresponding Oct-Dec observations. Mean Dep. Var is the average fraction claiming at 66+ for those in windows where DRC changes, born in Oct-Dec for the relevant sample. Robust standard errors in parenthesis.