The Rise in The Full Retirement Age and Social Security Disability Insurance Enrollment

A thesis submitted by

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In partial fulfillment of the requirements for the degree of

Master of Science

In

Economics

Tufts University

May 2020

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Abstract

In recent years, the percentage of the population in the U.S. receiving Social Security Disability Insurance (SSDI) has increased rapidly. An explanation for this rapid increase is: Social Security's 1983 Amendments reduced the generosity of claiming retired worker benefit before reaching the full retirement age, made SSDI relatively more generous, and encouraged the older individuals to apply for SSDI. I exploit this exogenous variation in retired worker benefits with a cohort discontinuity analysis to estimate the impact of raising the full retirement age on SSDI enrollment using the March supplement of the Current Population Survey from 1992 to 2007. My results suggest a 2-month increase in the full retirement age increased SSDI enrollment by 8.7 percentage points. I also find that the SSDI enrollment increase was driven by single men and non-single women: a 2-month FRA increase led to a 29.4 percentage points increase in SSDI enrollment for single men and a 9.9 percentage points increase in SSDI enrollment for non-single women. My results suggest that raising the full retirement age might cause a significant spillover effect on SSDI enrollment.

Acknowledgement

I truly appreciate the guidance and support from Professor Melissa McInerney who not only helped me tremendously on my thesis but also greatly supported me to pursue my future study in Economics. I also greatly benefited from the discussions with Professor Elizabeth Setren and Professor Gilbert Metcalf. I would like to thank Professor Marcelo Bianconi, Professor Federico Esposito, Professor David Garman, Professor Alan Finkelstein-Shapiro, Professor Gilbert Metcalf, and Professor Jeffrey Zabel for their well-taught lectures that helped me to develop my interest in Economics and build a solid foundation for future studies. Thanks to my mon, dad, and Yuefan for their support during my study at Tufts.

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1. Introduction

Social Security Disability Insurance (SSDI) enrollment increased dramatically in recent decades. For example, in 1970, only 1.7 percent of individuals aged 25-64 were receiving SSDI benefits, but in 2010, the fraction of individuals aged 25-64 receiving SSDI increased to 5 percent (Duggan and Imberman, 2009). Previous work has suggested several explanations for this dramatic increase. However, there is an explanation that has not been widely studied in previous work: the impact of raising the full retirement age (FRA) on the SSDI enrollment.

This approach is motivated by Social Security Administration (SSA)'s 1983 amendments: in 1983, SSA implemented amendments which increased the FRA[1] gradually from 65 years to 67 years, starting for individuals who were born in 1938. This rise in the FRA was trying to improve the financial sustainability of the Social Security retired worker benefits program.

However, the increase in the full retirement age simultaneously increased the penalty of early retirement: an individual born in 1937 or before could get 80 percent of her "full" retirement benefits[2] if she would retire at 62 years old, while an individual born in 1943 could only get 75 percent of her "full" retirement benefits at age 62. While the generosity of claiming the retired worker benefit early fell, the generosity of other social insurance programs, such as Social Security Disability Insurance (SSDI), did not change. Since SSDI's benefit amount is calculated

[1] Full retirement age is the age that an individual can start claiming "full" retirement benefits. If an individual wants to claim benefits before the FRA, her benefits will be discounted based on her age. Full retirement benefit means an individual can get 100% of her primary insurance amount (PIA). I discuss how to calculate full retirement benefit and the PIA below.

[2] In this paper, an individual gets "full" retirement benefits means that she gets 100 percent of her PIA and an individua gets 80 percent of the "full" retirement benefits means she get 80 percent of her PIA.

the same way as how the retired workers benefits are calculated[3] and SSDI's eligibility is not determined by age, as the FRA increased, SSDI became relatively more generous. This might motivate individuals to apply for SSDI instead of applying for the retired worker benefits and increased SSDI enrollment.

Bound (1989) first mentioned there could be possible substitution between social insurance programs. He observed that 65.4 percent of rejected applicants of SSDI enrolled in other public income programs afterward suggesting that these rejected applicants viewed these public income programs as a substitute for SSDI. Also, more recent research such as Duggan et. al. (2007), Li and Maestas (2008), and Coe and Haverstick (2010) also suggested that generosity reduction in one social insurance program might cause spillover effects on other social insurance programs' enrollment. With this evidence, it is possible that the decline in generosity of claiming retired worker benefits early might encourage more individuals to apply for SSDI and increase SSDI enrollment.

Therefore, my research question is whether and to what extent the rise in the full retirement age encourages older workers to apply for SSDI. To estimate the impact of the FRA increase on SSDI enrollment, I follow a cohort[4] discontinuity analysis that is based on the exogenous variation in retired worker benefits brought by the 1983 Amendments. This exogenous variation allows me to identify the causal relationship between the FRA increase and SSDI enrollment.

[3] I introduce more about how benefits are calculated in these programs below.
[4] All the cohort in this paper means the year of birth.

My sample contains individuals who were born between 1930 and 1943 and were 62 to 64 years old from 1992 to 2007. I divide my sample into two groups: individuals born between 1930 and 1937 are in the control group (because they were not affected by the 1983 amendments) and individuals born between 1938 and 1943 are in the treatment group. I focus on individuals aged from 62 to 64 because the early eligibility age (EEA)[5] is 62 and the FRA is 65 for individuals in the control group. Finally, I estimate a logit model with 13 binary variables (each binary variable indicating if an individual is born in a certain year) and other control variables to investigate if individuals in the treatment cohorts (1938-1943) have a higher probability of SSDI enrollment than individuals born in 1937.

My regression results suggest that raising in the FRA had a significant impact on SSDI enrollment. My estimates indicate that a two-months increase in the FRA led to a 8.7 percentage points increase in SSDI enrollment (29.4 percentage points for single men and 9.9 percentage points for non-single women), which implies that the SSDI enrollment increase was driven by single men and non-single women.

My work contributes to the growing literature that studies the rapid growth in SSDI beneficiaries and the decline in older workers' labor force participation in the United States. My work contributes to the literature in two aspects. Firstly, this study is the first one that uses the cohort discontinuity method to study the impact of the FRA increase on SSDI enrollment in the U.S. I depart from the commonly used relative generosity method, which uses the changes in the

[5] The early eligibility age (EEA) is the earliest age that an individual can claim retired worker benefits. However, claiming retired worker benefits at the EEA might cause a reduction in monthly benefits. The detailed rules for punishing early retirement are discussed below.

present value of retired worker benefits brought by the 1983 Amendments as a source of exogenous variations to identify the impact of the FRA increase on SSDI enrollment. Secondly, I include more years of data and more recent data than most previous studies. My sample contains yearly data from 1992 to 2007[6] while most previous research has fewer years of data or less recent data in their sample. Including more years of data and more recent data can help me to better capture the direct behavior responses of individuals facing the FRA increase.

My work has strong policy implications. The Social Security Administration (SSA) proposed the FRA increase to encourage older workers to stay longer in the labor force and contribute more to social insurance programs by paying social security taxes. However, if the FRA increase actually encouraged older workers to switch to other social insurance programs like SSDI, the FRA increase might offset the expected gains in the SSA's financial sustainability

2. Institutional Background

Social Security retired worker benefits and Social Security Disability Insurance (SSDI) are both social insurance programs. These two programs are similar in many aspects but also have crucial differences.

Social Security retired worker benefits are the main source of income for many retired workers. A worker is eligible for Social Security retired worker benefits after being employed and paying

[6]Duggan et. al. (2007) use the SSA administrative data from 1983 to 2005. Li and Maestas (2007) use the biennial HRS from 1992 to 2006. Though Coe and Haverstick (2010) use the HRS from 1992 to 2008, since HRS is biennial, I still have data from more years in my sample.

Social Security taxes for 10 years[7] (40 quarters). The amount of benefits is determined through several steps and the first step is to calculate the Average Indexed Monthly Earning (AIME). The AIME is first calculated by averaging the highest 35 years of earnings adjusted by a national wage index. After computing the AIME, a progressive formula is applied to obtain the Primary Insurance Amount (PIA). The PIA increases at a decreasing rate as the AIME increases (Figure 1). For the part of the AIME that is less than or equal to $816, every dollar in AIME would be counted as $0.90 towards the PIA. For the part of AIME that is higher than $816 but lower than $4,917, every dollar in the AIME would be counted as $0.32 in the PIA. For the part of the AIME that exceeds $4,917, every dollar in the AIME would only be counted as $0.15 in the PIA. Finally, the PIA is adjusted based on if an individual claims the retired worker benefits before the FRA, which is 65 years for workers born in or before 1937. If an individual chooses to retire at 62, the PIA would be discounted by 25 percent. However, as Figure 1 shows, individuals who earn more during their working life tend to have a higher retirement benefit, but the increase in the marginal benefit is decreasing. The retired worker's benefit is around 40 percent[8] of an individual's pre-retirement earning.

However, different from retired worker benefits, which is available to all workers who have worked long enough in covered employment, SSDI has stricter eligibility requirements. First, to be eligible, a worker has to have worked in at least 5 of the last 10 quarters. Also, there is an additional requirement on employment history: depending on age of disability onset, a

[7] If this individual was born in 1929 or after.
[8] The replacement rate is 60 percent for individuals with low income and 40 percent for individuals with high income.

worker has to have 40 quarters employment history over lifetime to be eligible.[9] Moreover, a worker must have disabilities defined as "inability to engage in substantial gainful activity (SGA)[10] and have physical or mental impairments that have been or will last at least 12 months or would result in death." After fulfilling these requirements, an individual could file an application for SSDI to a local Social Security office and the application would be reviewed by an examiner from the state Disability Determination (DDS) office. If an individual's application is rejected, she could have four levels of appeals. Around two thirds of individuals' applications are approved after different levels of appeals, though the initial acceptance rate is around 37 percent (Duggan and Imberman, 2009; Maestas et. al., 2013). Successful SSDI applicants then begin receiving benefits five months after disability onset and are eligible for Medicare after two years. However, a SSDI beneficiary can lose her SSDI if she earns more than the SGA, which is $2,110 for blind individuals and $1,260 for non-blind individuals.

As for programs' generosity, the SSDI benefits are calculated in almost the same way as the retired workers benefits are calculated and SSDI beneficiaries can automatically get full retirement benefits when they reach their FRA. There are only two main differences. The first is that the calculation of AIME in SSDI is based on average income of 10 highest years, instead of 35 years. The other difference is that the SSDI benefit is not subject to penalty for claiming at an early age. An individual could apply and claim full SSDI benefits whenever her disability is onset. An individual born in 1943 could get 75 percent of her PIA if she chose to claim the

[9] For individuals aged 62 or more, the work history requirement is 10 years which is the same for the work history requirement of retired worker benefits.
[10]The SSA's 2019 report defines SGA as "A person who is earning more than a certain monthly amount (net of impairment-related work expenses) is ordinarily considered to be engaging in SGA." The amount is $2,110 for blind individual and $1,260 for non-blind individuals.

retired worker benefits at 62 while the FRA is 66. However a individual born in 1937 could get 80 percent of her PIA if she retired at 62 while the FRA is 65. These differences in age requirement and similarities in benefit amount are the main reasons for the possible substitution between these two programs.

3. Literature Review

Previous research has suggested that social insurance programs, such as SSDI, can have a significant impact on older workers' labor supply decision (Kruger and Pischke, 1992; Staubli and Zweimuller, 2013; Blau et. al., 2007). For example, Maestas et. al. (2013) use the variations in allowance rates among SSDI application examiners to study the impact of SSDI benefit receipt on workers' labor supply decisions. Using the Disability Operational Data Store (DIODS) from 1995 to 2009, they find that for the 23 percent of beneficiaries on the margins of SSDI entry in 2005 and 2006, the employment rate of these beneficiaries would have been 28 percentage points higher two years later if they had never received SSDI.

However, though these works help us better understand the impact of social insurance programs on workers' labor supply decisions, most of them do not consider the potential spillover effects between social insurance programs, which might underestimate their results.

To fill this gap, several previous studies focusing on the spillover effect and substitutions between social insurance programs.

Bound (1989) first implied there might be substitutions between SSDI and other social assistance programs. In his research focusing on health and labor market outcomes of SSDI's rejected

applicants, he found that 65.4 percent of rejected applicants of SSDI enrolled in other public income programs afterward, which suggested that these rejected applicants viewed these public income programs as a substitute for SSDI. However, Bound did not estimate the level of substitution between SSDI and other social assistance or social insurance programs, which inspired other researchers to address this issue.

Moreover, in 1983, the Social Security Administration (SSA) implemented the 1983 amendments to retired worker benefits, which promoted more work to utilize the exogenous variations brought by the 1983 amendments to estimate the level of substitutions between social insurance programs. However, the first several works did not find significant substitutions between social insurance programs (Mitchell and Phillips, 2000; Bound et. al., 2004).

Mitchell and Phillips (2000), to my knowledge, first estimate the extent of substitution between the social security retired worker benefit and SSDI. Using the first four waves of the HRS (1992 – 1998), they estimate a conditional logit model with three choices of retirement pathways: claiming at the early retirement age, claiming at the full retirement age, or applying for SSDI benefit to study the impact of an increase in the FRA on SSDI enrollment. Their simulated results suggest that a $25,000 reduction[11] in the present value of early retirement benefits, which is around a 15-percent reduction in monthly payment for the median HRS respondent, will increase SSDI enrollment by 0.6 percentage points.

[11] Duggan et. al. (2007) suggest that this reduction in benefits is around 40% higher for a 62-year old when the full retirement age is 67, implying this amount is hard to compare to.

Moreover, Bound et. al. (2004) construct a dynamic programming model and also use the first four waves of the HRS to simulate single men's behavioral response towards raising the normal retirement age, eliminating early retirement, and introducing universal health insurance. They suggest a simulated increase in the FRA to 67 will increase SSDI enrollment by 0.2 percentage points. Their results agree with Mitchell and Phillips (2000) showing the FRA increase has a small effect on one-year SSDI enrollment.

Both Mitchell and Phillips (2000) and Bound et. al. (2004) suggest that the FRA increase would have a limited impact on SSDI enrollment and that the level of substitution between Social Security retired worker benefits and SSDI is relatively small. However, one major limitation of these two studies is that they only used the first four waves of HRS (from 1992 to 1998, while the first affected cohort only reached 62 in 2000) to conduct out-of-sample predictions on SSDI enrollment outcomes. Lacking the ability to observe the direct behavior responses of the older individuals might undermine their results' credibility.

As more individuals in the affected cohorts reached their early retirement age, later studies started to use the survey responses of these individuals to directly study the impact of the increase in the FRA on SSDI enrolment. Different from previous works, these later works tended to find that the increase in the FRA would have a relatively larger impact on SSDI enrollment (Duggan et al., 2007; Li and Maestas, 2008; Coe and Haverstick, 2010).

Duggan et al. (2007) is the first work I know that uses the observed behavior responses of workers to study the impact of the FRA increase on SSDI enrollment. Using aggregate social

security administrative data from 1983 to 2005 and focusing on men and women between the age of 45 and 64,[12] they use a linear regression model with year/age-specific DI enrollment as their dependent variable and the one-year change in the present value of age-specific retired worker benefits as their main explanatory variable. They find a larger effect than Mitchell and Phillips (2000) and Bound et. al. (2004) do. Duggan et. al. (2007) suggest that the 1983 Social Security amendments increased SSDI enrollment by 0.58 percentage points for men and by 0.89 percentage points for women. In 2005, this corresponds to an increase in SSDI enrollment of 8.7 percent. Moreover, they also find strong heterogeneous effects between men and women: they suggest that each $5,000 reduction in the average present value of retired worker benefits causes SSDI enrollment to increase by 0.4 percentage points for men and 0.8 percentage points for women (0.6 percentage points for men and women pooled), which is relatively larger than Mitchell and Phillips (2000)'s estimations.[13]

Though Duggan et. al. (2007) were the first to directly quantify the effect of the FRA increase on SSDI enrollment, there were still limitations in their method. For example, Li and Maestas (2008) point out that Duggan et al. (2007) did not restrict the population to individuals who are covered by SSDI, which might cause bias in their estimations.

To solve these limitations, Li and Maestas (2008) make several improvements based on Duggan et. al. (2007)'s approach. Li and Maestas (2008) merge the first eight waves of the HRS (1992-

[12] Duggan et. al. (2007) did not specify the birth cohorts they included. However, from the age groups and years they covered, it implied they include birth cohorts from 1918 to 1960. Also, they mentioned they focused on estimating the effects for birth cohorts (1938-1943) who are in the first phase of the FRA increases.
[13] Mitchell and Phillips (2000) suggest that a $25,000 reduction in retired worker benefits would cause 0.6% increase in SSDI enrollment. Their simulated benefits reduction is 5 times the amount in Duggan et. al. (2007), but the obtained the same 0.6 percentage points increase in SSDI enrollment, implying that Duggan et. al. (2007)'s results are much larger than Mitchell and Phillips (2000)'s.

2006) with the Social Security administrative earning history data to accurately identify individuals who are aged from 55 to 64, are covered by SSDI, and were born between 1930 and 1941. Moreover, instead of using SSDI enrollment as dependent variable, they aggregate their data by age and use the age-specific SSDI application rate as their outcome variable.[14] Finally, they identify three different ways of measuring the relative generosity between retired worker benefits and DI and use these three measurements[15] as their main explanatory variables respectively to estimate a probit model. With these differences, they obtain different results from Duggan et. al. (2007). Their estimations suggest that a 4-month increase in the FRA caused the 2-year DI application rate to increase by 0.30 percentage points for the whole sample, by 1.47 percentage points for individuals reaching early eligibility age (EEA), and by 0.89 percentage points for individuals with work limitations. Furthermore, to compare their results with Bound et. al. (2004), they restrict their sample to the single working men and estimate the impact of a 2-year FRA increase on the 1-year DI application rate[16]. Their results suggest a 2-year increase in the FRA would result in a 0.12-0.46 percentage points[17] increase in 1-year DI application rate, which align with Bound et al. (2004)'s estimations.

Adopting Li and Maestas (2008)'s method, Coe and Haverstick (2010) use the HRS from 1992 through 2008 with SSA detailed earning data, restrict their sample to 56 – 63 years old individuals who born between 1935 and 1943, and exclude married women[18] from their sample.

[14] Li and Maestas (2008) define SSDI application rate as the percentage of individuals who apply to DI in two years after being eligible for SSDI. They use the two-year DI application rate because of the HRS is biennial.
[15] The three ways include: 1. Measuring by the increase in FRA. 2. Measuring by the ratio between retired worker benefits and DI benefits for each individual. 3. Measuring by the ratio between the present value of retired worker benefits to the present value of DI for each individual.
[16] However, they do not articulate how to transfer 2-year DI application to 1-year DI application rate.
[17] Because they use three different ways of measuring relative generosity.
[18] They claim that the increase in labor force participation among married women made more married women eligible for SSDI, which might confound their results.

They use the same outcome and main explanatory variable as Li and Maestas (2008) to estimate the impact of an increase in the FRA on the 2-year SSDI application rate[19]. Moreover, they also investigate if the spillover effect would affect the solvency of SSDI. Their estimations suggest that a 1 percentage point decrease in the retirement-to-disability benefit ratio leads to a 0.25 percentage point increase in the 2-year SSDI application rate, which represents an 8 percent increase in the 2-year SSDI application rate. However, they also find, controlling for health and other SSDI coverage determinants, the increase in SSDI applicants only leads to decreasing acceptance rate, which keeps the total recipients at a constant level and would not have a huge impact on SSDI's solvency.

Though the later studies (Duggan et. al., 2007; Li and Maestas, 2008; Coe and Haverstick, 2010), using a similar method, provide strong evidence suggesting an increase in FRA has a strong impact on DI enrollment or DI application rate, there is one concern in all these works: all these works require individuals' detailed earning history to calculate the present value of the retired worker benefit or present value of DI benefit and calculating present value needs strong assumptions like the probability of surviving at a certain age, which might result in serious measurement error bias.

One possible solution for this measurement error bias is to use the cohort discontinuity method proposed by Mastrobuoni (2009). Using the monthly CPS from January 1989 to January 2007 and restricting the sample to individuals born between 1928 and 1941 and aged between 61 and 65, Mastrobuoni (2009) first use this cohort discontinuity method to study the impact of the FRA

[19] Similar definition as Li and Maestas (2008).

increase on the older workers' retirement decisions. This cohort discontinuity method utilizes the fact that the level of the FRA increase is only based on birth cohorts. Therefore, by including one dummy variable for each cohort and dropping the dummy variable of cohort 1937, this method is comparing the probability of retirement of individuals from each cohort with the probability of retirement of individuals from cohort 1937. Controlling for several individual characteristics, like education level and ethnic group, we should be able to isolate the impact of rising FRA on workers' behavior. Using this method, Mastrobuoni (2009) suggests that the older workers react dramatically to the increase in the FRA: a two-month increase in the FRA caused workers to delay retirement by one month. This method has two main advantages. Firstly, it can avoid the measurement error bias by not using the present value of retired worker benefit as the main explanatory variable, which can be estimated in error because life expectancy is unknown. Secondly, it does not require information on detailed earning history to calculate present values, which means it is easier for others to replicate.

Adopting this cohort discontinuity method, Atalay and Barrett (2015) study the 1993 Australian age pension reform, which increased the eligibility age from 60 to 65 years for woman, using administrative data from 1994 to 2010. Their results suggest the increase in the eligibility age for women increased the disability support program's enrollment by 13 to 23 percentage points.

Compared to the U.S.-based literature I list above, my study is different in two ways. Firstly, I am using the cohort discontinuity method focusing on the discontinuity in SSDI enrollment between individuals born in 1937 and individuals born between 1938 and 1943. To my knowledge, there has been no research using this method to study the impact of the FRA increase

on SSDI enrollment in U.S. Compared to Duggan et. al. (2007), Li and Maestas (2008), and Coe and Haverstick (2010), this method requires less information on individuals' earning history and does not suffer from measurement error bias brought by calculating the present value of retired worker benefits or disability insurance. Secondly, I include more recent data[20] and data from more years, which could help me to better capture the observed behavior responses of the affected individuals.

4. Data

I use the Annual Social and Economic Survey (ASEC) of the Current Population Survey (CPS) from 1992 to 2007 to conduct a cohort discontinuity study in SSDI enrollment. I only include older individuals who are 62-64 years old and were born between 1930 to 1943. I exclude cohorts after 1943 because the FRA stays at 66 for 12 cohorts from cohort 1944 through cohort 1955. In my sample, individuals born from 1930 to 1937 are in the control group and individuals born from 1938 to 1943 are in the treatment group because the FRA increase started at cohort 1938.

I only include individuals aged between 62 and 64 because only individuals in this age group have choices between filing for retired worker benefits and applying for SSDI. For individuals who are younger than 62, they can only retire through SSDI because they are younger than the early entitlement age and are not able to get early retired worker benefits.

[20] I use the ASEC- CPS from 1992 to 2007. Duggan et. al. (2007) use administrative data until 2005, and Li and Maestas (2008) use the HRS until 2006. Though Coe and Haverstick (2010) use HRS from 1992 to 2008, I still have more years in my sample because the ASEC-CPS is yearly while HRS is biennial.

For individuals in the control group who are older than 64, they would choose retired worker benefits because SSDI is not eligible for individuals at full retirement age; however, if they are in the treatment group, the FRA ranged from 65 years and 2 months to 66 years for each year implying they can still choose between SSDI and retired worker benefits when they are older than 64 years old. Nevertheless, to make it easier to compare between cohorts, I only include individuals aged between 62 and 64 from the treatment cohorts.

The outcome variable I am going to examine is a binary variable which equals one if an individual enrolled in SSDI and equals zero otherwise. I identify SSDI enrollment by enrolling in Medicare, receiving income from Social Security, and being younger than 65 years old[21].

I put several restrictions on my sample. I exclude individuals who are not disabled[22] or did not quit their job because of disability[23] since only the disabled can apply for DI. If I do not put this restriction, the impact of the FRA increase on SSDI enrollment might be underestimated: most of the individuals in my sample would not be eligible for SSDI and could not decide between DI and retired worker benefits. Also, I exclude all the individuals who have missing data on the control variables. Finally, I merge my sample with the yearly state-level unemployment rate data from the U.S. Bureau of Labor Statistics to obtain a measure of the state-level labor market conditions. After these sample restrictions, I obtain a sample with 9,553 individuals (4,763 individuals belonging to the control group and 4,590 individuals belonging to the treated group)

[21] DI recipients would get Medicare coverage 24 months after receiving DI benefit. Otherwise, Medicare is only for individuals above FRA or have certain types of disabilities. Therefore, most of the individuals, who are covered by Medicare, aged below FRA, and are receiving income from Social security, are DI recipients.
[22] The questionnaire asks: "Do you have a health problem or disability which prevents you from working or which limits the kind or amount of work you can do?"
[23] The questionnaire asks: "Did you ever (retire or leave/ retired or left) a job for health reasons?

I am using the ASEC - CPS because this dataset contains detailed information on individuals' Medicare participation and Social Security Income, which is necessary for me to identify SSDI[24] enrollment. Also, the ASEC-CPS has information on each worker's age, which can help me to identify their birth cohorts. Although it can be used as a longitudinal dataset, I only use it as a repeated cross-sectional dataset because linking through different waves will lose a large portion of my data points.

I use the ASEC- CPS instead of the HRS, which has been widely used in previous literature, for three main reasons**.** Firstly, the ASEC-CPS has much more observations on individuals of my interest, compared to HRS[25]. Secondly, the ASEC-CPS is a yearly survey, while HRS is biennial. This difference is important because individuals from the treatment cohorts reached their full retirement age between 2003 and 2008. However, the HRS only has data every two years, which means I cannot observe the short-term behavior response of some of these individuals if I use the HRS. Thirdly, the previous works use the HRS because they need to calculate the present value of retired worker benefits for each individual and the HRS can be merged with Social Security administrative earning history to accurately identify eligibility and benefit amount. However, in my study, I use the cohort discontinuity method which does not require these present-value calculations. Therefore, HRS's potential to be merged with detailed earning data is not necessary for my approach.

[24] In this work, all the SSDI enrollment is identified as covered by Medicare, receiving Social Security income, and younger than FRA.
[25] After putting all the restrictions, I list above, I obtain a sample with around 200 observations using the HRS. However, using the ASEC-CPS, I get 9,553.

Even though the ASEC-CPS has many advantages, it still has limitations. The first limitation is that it does not contain information on an individual's past employment and earning history and if an individual's employment is covered by SSDI. However, as the Social Security Administration's 2019 report shows, around 89% of workers have enough employment history to fulfill the employment requirement of DI (Social Security Administration, 2019).

Another concern of using the ASEC-CPS is that it does not contain information on an individual's birth cohort, which means I have to use her age and year of the survey to calculate her birth cohort. However, this calculation might result in misclassification, because a worker who was born in April 1937 would have a different age from a worker who was born in January 1937 in the ASEC-CPS (which is fielded in March of each year). If I used this method to generate the birth cohort variable, these two workers would be assigned to different birth cohort groups. However, this misclassification problem could be easily adjusted by adding one year to the age of an individual. Krueger and Pischke (1992) solved this problem by adding one year to all the individuals in their sample and they suggested this method could largely reduce the probability of misclassifying from 0.8 to 0.2. I apply this method on my sample. With 0.2 probability of being wrongly classified, this misclassification problem would not have a significant impact on my results.

Finally, the ASEC-CPS does not have as accurate of an identifier of DI enrollment as the HRS does. However, as I mentioned above, I can solve this difficulty by using being covered in Medicare, receiving income from Social Security, and being younger than FRA to identify DI enrollment.

The summary statistics are shown in Table 1. As Table 1 shows, both the control cohorts and the treated cohorts are similar in sample size and in most control variables. After adjusting for inflation[26], there is a small difference in real household income: the difference is $2,300, which means the individuals in the treatment group have income that is 10.3 percent more than the individuals in the control group.

Finally, I also plot the percentage of the population enrolling in DI enrollment for individuals born between 1930 and 1950. In Figure 2 a), the percentage of SSDI enrollment is higher for the treated cohorts (1938-1943) comparing to cohort 1937. In Figure 2 b), for women, the percentage of SSDI enrollment is also higher for the treated cohorts. However, for men, there is no clear evidence showing higher percentage of SSDI enrollment for the treated cohorts. Moreover, directly comparing SSDI enrollment between cohorts might suffer from the fact that cohorts have differences in observable characteristics. Therefore, to control for these observable characteristics, I need to estimate regressions to prevent them from confounding my results.

5. Empirical Strategy

I use a cohort discontinuity method that follows from Mastrobuoni's (2008) research into the impact of an increase in FRA on the retirement decision of older workers. This cohort discontinuity method has two advantages. Firstly, I avoid the measurement error bias caused by calculating the present value of retired worker benefits for each individual. Secondly, instead of focusing on the coefficient of changes in the present value of retired worker benefit, it is more

[26] Based on dollar value in 1988.

informative to observe the discontinuities and to contrast differences in SSDI enrollment behavior between individuals in the control group and individuals in the treatment group.

Figure 2 a) illustrates a large increase in the percentage of population enrolling in SSDI between individuals born in 1937 and individuals born in the treated cohorts (1938-1943). However, my results might be confounded by other observable characteristics between individuals including the differences in educational attainment. Therefore, to parametrically control these confounding effects, I also include labor market conditions (measured by the state-level unemployment rate), socioeconomic characteristics (age, educational attainment, household size, household income marital status, ethnic group, and gender), and geographic characteristics (State).

There are three main assumptions behind my study. Firstly, I assume that self-reported working limitation is an accurate measure of fulfilling SSDI's disability requirement. Secondly, I assume that all the disabled individuals would have enough employment history to be eligible for SSDI. Thirdly, I assume that after controlling individuals' characteristics, labor market characteristics and geographical characteristics, the observed increase in SSDI enrollment was caused by the FRA increase.

The regression function is showed below:

$$P(SSDI)_{i,c} = \beta_0 + \sum_{c \neq 1937} \beta_c \, 1(C_i^* = \mathrm{c}) + age_i + State_i + \gamma' \boldsymbol{X}_i + \varepsilon_{i,c} \quad (1)$$

In this empirical model, $P(SSDI)_i$ is a binary variable indicating if an individual enrolled in SSDI. Enrollment in SSDI is identified by enrolling in Medicare, receiving income from Social security, and being younger than 65.[27]

6. Results

6.1 The effect of the FRA increase on SSDI enrollment for the pooled sample

I estimate a set of logit models of SSDI enrollment for individuals who were aged from 62 – 64 and were born between 1930 and 1943 using Equation (1) and data from the ASEC-CPS (1992-2007). The pooled results are shown in Table 2; In column (1), I only include the cohort dummies and in column (2) I additionally control for state-level unemployment rate, household size, ethnic group, educational attainment, marital status, gender, age fixed effects, and state fixed effects. The first thing to note in Table 2 is that after including all the control variables, the logit coefficient estimates of cohort 1938, 1939, 1940, 1941 are large, positive, and statistically significant. Also, for the pre-reform cohorts (1930-1937), all the logit coefficient estimates are smaller and are not statistically significant. These results suggest that individuals born between 1938 and 1941 had a higher probability of enrolling in SSDI comparing to individuals born in 1937 implying that the FRA increase had a large impact on individuals' SSDI enrollment behavior such that individuals who were born several years after 1937 had significantly higher probabilities of SSDI enrollment than individuals born in 1937. Moreover, I calculate the average marginal effects[28] and my calculations suggest that a 2-month FRA increase led to an 8.7

[27] Individuals who enrolled in SSDI can apply for Medicare after two years of receiving SSDI benefits. This is the main way that a worker can get Medicare before 65.
[28] I calculate the average marginal effects using the *margin, dydx ( )* commend in STATA.

percentage points increase in SSDI enrollment. Based on the average SSDI enrollment rate[29] for the pre-reform cohorts, this represents a 22.3 percent increase in SSDI enrollment for individuals in pre-reform cohorts. This is consistent with Li and Maestas (2008), who find that for individuals with working limitations, the FRA increase led to a 20.7 percent increase in SSDI application rate.

6.2 The effect of the FRA increase on SSDI enrollment for men

I next estimate my logit models with men and women separately. I present the logit coefficient estimates for men in Table 3 and for women in Table 4.

In Table 3, column (1) presents the results for all men, column (2) present the results for single men, and column (3) present the results for non-single men. As column (1) presents, when I focus on all men, the impact of the FRA increase on SSDI enrollment is not very clear. For the post-reform cohorts (1938 -1943), only cohort 1938 has statistically significant coefficient estimate, which implies that the FRA increase only affected SSDI enrollment for individuals born in 1938 and larger FRA increases did not affect SSDI enrollment for individuals born in later cohorts (1939 -1943). Also, for the individuals in the pre-reform cohorts, cohort 1936 also has an increased likelihood of receiving SSDI relative to the 1937 cohort. These results imply that for men, there might be some other factors confounding my results.

Therefore, to reduce the effect of the confounding factors, I further divide my sample into single men and non-single men and estimate my logit model again. When I focus on single men, the impact of the FRA increase on SSDI enrollment is clear again. As column (2) presents, cohort

[29] The SSDI enrollment rate here means the cohort-level percentage of population enrolling in SSDI. Around 38.8 percent of individuals born between 1930 and 1937 enrolled in SSDI during 1992 to 2001 (when they were aged from 62 to 64).

1938, 1939, and 1940 have positive, large, and marginally significant logit coefficient estimates while none of the pre-reform cohorts has statistically significant or positive coefficients. Though the coefficient estimates are only marginally significant and have large standard errors because of my sample size for single men is relatively small,[30] using a small sample to find marginally significant and large coefficient estimates suggest that the FRA increase did have a strong impact on single men's SSDI enrollment behavior. Finally, I calculate the average marginal effects of the FRA increase on SSDI enrollment for single men. My results suggest that, for single men, a 2-month FRA increase led to a 29.4 percentage points SSDI enrollment increase. This large percentage points increase further support my finding that the FRA increase had a strong impact on single men's SSDI enrollment behavior.

However, for non-single men, the results are still confusing. As column (3) presents, the post-reform cohorts (except cohort 1938) still do not have statistically significant coefficient estimates and cohort 1936 still have statistically significant coefficients. Moreover, comparing with column (2), all the post-reform cohorts in column (3) have smaller coefficient estimates indicating that the FRA increase might have larger effects on single men than on non-single men.

6.3 The effect of the FRA increase on SSDI enrollment for women

In Table 4, I report the logit coefficient estimates for women. Column (1) presents the result for all women, column (2) presents the results for single women, and column (3) presents results for non-single women. As column (1) shows, for women, the impact of the increase in the FRA is clear: cohort 1938, 1939, 1940, and 1941 have large, positive, and statistically significant

[30] Single men constitute only a small portion (2.9%) of my whole sample. This might be the reason why the coefficient estimates are only marginally significant and have large standard errors.

coefficient estimates and all the pre-reform cohorts have coefficient estimates that are not statistically significant, implying that the FRA increase had a significant impact on SSDI enrollment. Also, in column (3), for non-single women, the impact of the increase in the FRA is also clear: for the post-reform cohort, cohort 1938, 1939, 1940, and 1941 have large, positive, and statistically significant coefficient estimate while none of the pre-reform cohorts has statistically significant coefficient estimates. Finally, I calculate the average marginal effect of the FRA increase. My results suggest that a 2-month FRA increase led to a 9.9 percentage points SSDI enrollment increase.

For single women, the impact of the FRA increase on SSDI enrollment is not as clear as the results for non-single women, since as column (2) presents, none of the post-reform cohorts has statistically significant coefficient estimates. However, since I only have 321 observations of single women and the logit coefficient estimates have large standard errors, I cannot rule out the possibility that the FRA increase did have impact on single women's SSDI enrollment. Therefore, I calculate the upper bound of the 95 percent confidence interval of logit coefficient estimates for single women. For all the post-reform cohorts, the coefficient estimates have large and positive 95 percent upped bounds. The 95 percent upper bound of coefficient estimate of cohort 1938 is 2.1 which implies that I cannot rule out effects as large as 40.5 percentage points increase in SSDI enrollment for single women. Therefore, with this large and positive 95 percent upper bound, I cannot rule out the possibility that the FRA increase might also have significant impact on single women's SSDI enrollment.

6.4 Further discussion on reasons for heterogenous effects

I now discuss why the FRA increase had larger effects on single men's and non-single women's SSDI enrollment. There are two possible explanations for the larger behavior responses of single men and non-single women. The first reason is that single men and non-single women have lower AIMEs (and PIAs) than non-single men, which I will show is expected to make them more responsive to the rise in FRA. Though, I do not have information on each individual's AIME, the retired workers' benefits and the SSDI benefits are calculated based on AIME. Therefore, I use social security income as a measure of AIME. I calculate the unweighted sample mean of social security income[31] of single men ($6,122), non-single men ($8,850) and non-single women ($ 6,523). Single men and non-single women have much lower unweighted average Social Security income than non-single men implying that they have lower AIMEs.

This might make single men and non-single women more responsive because compared with individuals with higher AIMEs, individuals with lower AIMEs would suffer from a higher percent reduction in retired worker benefits, which implies that the FRA increase had larger impacts on individuals with lower AIMEs. To illustrate the heterogenous effects of the FRA increase on individuals with relatively higher AIMEs and on individuals with relatively lower AIMEs, I plot what percent of AIMEs individuals could get as retired worker benefits (if they retired at 62) for individuals with different AIMEs as a functions of birth cohort in Figure 3. As Figure 3 presents, compared with individuals with AIMEs equal $10,000, individuals with AIMEs equal $500 experienced larger percentage reductions in retired worker benefits.

[31] Since social security income is calculated based on income, it can be used as a compromised measure of life-time income. To calculate the unweight average social security income, I exclude individuals who receive zero income from SSA.

Therefore, it is possible that single men and non-single women were more responsive to the FRA increase because they had lower AIMEs than non-single men and faced larger percentage reductions in retired worker benefits if they claimed early, which might encourage them to response more to the FRA increase.

Secondly, the difference in educational attainment could be another reason for difference in SSDI enrollment between single men and non-single women and non-single men. As I mentioned above, SSDI is only eligible for people who has been determined as disabled and are not able to engage in SGA. As Michaud et. al. (2018) suggest, low educational attainment may obstruct an individual's ability to find new work, which makes poorly educated individuals more likely to get a disability determination and enroll in SSDI. Therefore, it is possible that single men and non-single women are relatively less well-educated than non-single men, which made it easier for them to get a disability determination and SSDI enrollment. To support this claim, I calculate the unweighted sample mean of educational attainment for single men, non-single men, and non- single women separately. In my sample, around 49.7 percent of single men and 42.7 percent of non-single women do not have high school diplomas, while only 39.4 percent of non-single men do not high school diplomas. Therefore, this difference in the educational attainment might make single men and non-single easier to get SSDI and make their SSDI enrollment higher facing the FRA increase.

In addition to these two reasons I mentioned above, there is another reason that could explain the difference in SSDI enrollment between non-single men and non-single women is the spousal age difference within couples. As Rolf and Ferrie (2008) suggest, men were on average 2 to 3 years

older than their spouses between 1970 and 2000. Also, Coile (2004) showed that couples tend to retire together. These two studies suggest that when non-single men reached their full retirement age, their spouse was still in their early 60s and far from their full retirement age. Therefore, for non-single women, if they want to retire with their spouses, they have to retire before their full retirement age, which means they would be affected more by the reduction in the generosity of claiming retired worker benefits early brought by the FRA increase.

Finally, I test for if the difference in SSDI enrollment between men and women is statistically significant. As Table 5 suggests, the estimated coefficients of the interaction terms between female and cohort dummies from 1938 to 1943 are all positive but not statistically significant. This suggests that though there are differences in SSDI enrollment between men and women in post-reform cohorts, the differences are not statistically significant.

7. Robustness Check & Falsification Check

As for robustness checks, I first use a different method to define SSDI enrollment. In the ASEC-CPS, WHYSS1 documents the reason why an individual is receiving income from SSA. There are 8 reasons listed including retired, disabled, widowed and others. Since the only SSA disability insurance program asked about in this question is SSDI[32], I can use that variable to identify who are enrolling in SSDI. However, there is one limitation of using this variable to identify SSDI enrollment: this variable starts in 2001 when only two control group cohorts are 65 years old or younger. Therefore, I only use this variable to conduct my robustness check. I present the logit regression coefficient estimates in Table 6. As Table 6 shows, my method is

[32] Supplemental Security Income (SSI) can also provide income for individuals with disabilities. However, SSI is not included in this variable.

robust under another definition of SSDI enrollment: the logit coefficient estimates of cohort 1938 is still positive and statistically significant.

One concern of my empirical approach is that the cohort discontinuity depends on which birth cohort I drop from my logit regression. Therefore, I omit different cohorts to test if the discontinuity in SSDI enrollment happens exactly between cohort 1937 and cohort 1938. I report my results in Table 7. I omit from cohort 1931 to cohort 1941 respectively and in no case does the cohort immediately following the placebo omitted cohort have a positive and significant logit coefficient estimate. Moreover, cohort 1938, 1940, and 1941 still have positive and statistically significant coefficient estimates in most regressions regardless which cohort I drop.

The other falsification check is that I apply the cohort discontinuity method on the enrollment of other disability income programs to figure out if this cohort discontinuity is resulted by an increase in the numbers of workers who are disabled. These disability income programs include: worker's compensation; company or union disability; federal government civil service disability; U.S. military retirement disability; state or local government employee disability; U.S. Railroad Retirement disability; accident or disability insurance; black lung miner's disability; or state temporary sickness payments. However, programs funded by Social Security or the Veterans' Administration are excluded. I report the logit coefficient estimates in Table 8. I use this falsification to test if the FRA increase is only encouraging substitution between retired worker benefits and SSDI, or a more general shift to other disability programs. Almost all of these other disability insurance programs have stricter eligibility requirements: for example, the U.S. Railroad Retirement disability requires applicants to have at least 10-year service at the rail

system and the U.S. military retirement disability is only available for individuals who are referred by doctors to the disability determination board.[33] For these programs, comparing to cohort 1937, the post-reform cohorts (1938- 1943) do not have positive and statistically significant coefficient estimates. This might because these programs have stricter eligibility requirements. This falsification test suggests that compared to cohort 1937, there is no significant increase in the total amount of disabled population in the post-reform cohorts and the SSDI enrollment increase was not caused by the increase in disabled population.

8. Conclusion

In this study, I estimate the impact of raising the full retirement age on SSDI enrollment among older individuals aged from 62 to 64. My results show that a 2-months increase in the FRA increased SSDI enrollment by 8.7 percentage points. I find this SSDI enrollment increase is driven by single men (29.4 percentage points) and non-single women (9.9 percentage points).

It is of interest to compare my estimations with previous studies. I try to compare my results with Duggan et. al. (2007) and Coe and Haverstick (2010). However, Duggan et. al. (2007) did not limit his sample to individuals with working limitations and Coe and Haverstick (2010) did not include non-single women in their sample. Therefore, I only compare my results with Li and Maestas (2008). Based on the average SSDI enrollment rate for the pre-reform cohorts, the 8.7 percentage point increase in SSDI enrolment indicates a 22.3-percent increase in SSDI enrollment for individuals in pre-reform cohorts. This is consistent with Li and Maestas (2008),

[33] An Individual cannot apply for U.S. military retirement disability. This program is only available if an individual is examined by a doctor and the doctor would refer this individual to be reviewed by a board to determine eligibility.

who find that for individuals with working limitations, the FRA increase led to a 20.7-percent increase in SSDI application rate.

My study contributes in two aspects. Firstly, this study is the first one that uses the cohort discontinuity method to study the impact of FRA increase on SSDI enrollment in the U.S. Secondly, I include more years of data and more recent data than most of the previous studies. My sample contains yearly data from 1992 to 2007, while most previous research has fewer years of data or less recent data in their sample.

However, though my results suggest that the FRA increase led to increase in SSDI enrollment, there is no evidence showing that the SSDI enrollment increase would significantly harm the SSA's financial sustainability. For example, Coe and Haverstick (2010) show that SSDI recipients are only a small portion of the whole population and the increase in SSDI enrollment would not have a huge impact on SSA's financial sustainability because the SSA could change the SSDI acceptance rate to reduce its future spending.

One major limitation of this work is that the ASEC-CPS does not have detailed information on each workers' earning and employment history, or disability, which makes it hard to accurately identify SSDI eligibility. Therefore, future research might adopt better data to accurately identify who are eligible for SSDI to generate more reliable results.

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Appendix

Table

Table 1: Comparing SSDI enrollment and other observable characteristics between individuals born between 1930 and 1937 (control group) and individuals born between 1938 and 1943 (treatment group) using the ASEC-CPS from 1992 to 2007

	FRA = 65 (Control)	FRA > 65 (Treatment)	
	1930 – 1937 Cohorts	1938 – 1943 Cohorts	Total
Demographic Characteristics	**Mean**	**Mean**	**Mean**
Percentage of Population Enrolling in SSDI	39.0%	44.0%	43.0%
Real Household Income	$22,400	$24,600	$23,500
Age	63.05	63.07	63.06
Female	44.0%	49.0%	47.0%
Married	94.0%	92.0%	93.0%
Ethnic Group			
Black	14.0%	19.0%	17.0%
White	82.0%	74.0%	78.0%
Indian	2.0%	2.0%	2.0%
Asian or Pacific	2.0%	1.0%	1.0%
Education Attainment			
Less than High School	41.0%	31.0%	36.0%
High School	33.0%	36.0%	35.0%
Some College	17.0%	20.0%	18.0%
College and Above	9.0%	13.0%	11.0%
Total Observations	4,763	4,590	9,553

Note: The sample is from the ASEC-CPS between 1992 and 2007, covering individuals aged from 62 to 64 and born between 1930 and 1943. The means are unweighted. All the numbers in this table shows the mean value of a certain variable. For example, in column 1/row 1, 39.0% means 39.0 percent of population who were born between 1930 and 1937, aged from 62 to 64, enrolled in SSDI. * FRA stands for the full retirement age

Table 2: Estimated average SSDI enrollment using logit regression with the ASEC-CPS from 1992 to 2007, 1937 is the omitted cohort (unweighted and SSDI enrollment as dependent variable)

	(1)	(2)
Sample includes ages:	62-64	62-64
Dependent variable:	SSDI Enrollment	SSDI Enrollment
1930	-0.05*	-0.13
	(0.03)	(0.14)
1931	-0.04	-0.09
	(0.03)	(0.13)
1932	-0.04	-0.08
	(0.03)	(0.12)
1933	-0.05*	-0.15
	(0.03)	(0.12)
1934	-0.01	-0.01
	(0.03)	(0.12)
1935	-0.04	-0.15
	(0.03)	(0.12)
1936	0.04	0.16
	(0.03)	(0.12)
1937	--	--
1938	0.07**	0.36***
	(0.03)	(0.12)
1939	0.03	0.22*
	(0.03)	(0.11)
1940	0.04	0.25**
	(0.03)	(0.11)
1941	0.05*	0.28**
	(0.03)	(0.11)
1942	0.01	0.14
	(0.03)	(0.11)
1943	0.01	0.11
	(0.03)	(0.11)
Unemployment Rate		-0.03
		(0.03)
Household Size = 1		-0.00
		(0.05)
Household Size > 3		0.11*
		(0.06)
Black		-0.02
		(0.06)
Indian		0.06
		(0.17)
Asian or Pacific		-0.34*
		(0.20)
Highschool		-0.02
		(0.05)
Some College		-0.22***
		(0.06)
College and Above		-0.45***
		(0.08)
Single		-0.00
		(0.09)
Female		-0.10**
		(0.04)
Constant	0.41***	-0.23
	(0.02)	(0.23)
Observations	9,553	9,553
State FE	No	Yes
Age FE	No	Yes

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains both men and women who are aged from 62 to 64, born between 1930 and 1943, and in the ASEC-CPS from March 1992 to March 2007. The estimates are not weighted and I only report logit coefficient estimates in this table. "Unemployment rate" indicates the monthly state-level unemployment rate. 1930-1943 are cohort dummies which indicate the year of birth *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Table 3 Estimated average SSDI enrollment for men using logit regression with the ASEC-CPS from 1992 to 2007, 1937 is the omitted cohort (unweighted and SSDI of men, SSDI enrollment of single men, and SSDI of non-single men as dependent variables)

	(1)	(2)	(3)
Sample included ages:	62-64	62-64	62-64
Dependent Variables:	SSDI Enrollment of men	SSDI enrollment of single men	SSDI enrollment of non-single men
1930	-0.03	0.72	-0.08
	(0.18)	(0.97)	(0.19)
1931	-0.02	0.73	-0.05
	(0.17)	(1.03)	(0.17)
1932	0.02	0.29	-0.02
	(0.17)	(0.96)	(0.17)
1933	-0.15	1.13	-0.21
	(0.17)	(0.86)	(0.17)
1934	-0.11	0.51	-0.17
	(0.16)	(0.76)	(0.17)
1935	-0.17	-0.93	-0.20
	(0.16)	(0.94)	(0.17)
1936	0.28*	-0.31	0.30*
	(0.16)	(0.73)	(0.17)
1937	--	--	--
1938	0.34**	1.54*	0.29*
	(0.16)	(0.86)	(0.17)
1939	0.20	1.36*	0.17
	(0.16)	(0.74)	(0.17)
1940	0.20	1.44*	0.14
	(0.16)	(0.80)	(0.16)
1941	0.19	0.64	0.18
	(0.16)	(0.77)	(0.16)
1942	0.15	0.03	0.12
	(0.15)	(0.73)	(0.15)
1943	-0.02	-0.21	-0.03
	(0.15)	(0.72)	(0.16)
Unemployment Rate	-0.02	-0.22	-0.01
	(0.04)	(0.24)	(0.04)
Household Size = 1	-0.06	-1.05*	-0.04
	(0.08)	(0.58)	(0.08)
Household Size > 3	0.02	-0.71	0.02
	(0.07)	(0.74)	(0.07)
Black	0.06	-0.13	0.05
	(0.09)	(0.45)	(0.10)
Indian	-0.16	-2.88**	-0.01
	(0.24)	(1.29)	(0.26)
Asian or Pacific	-0.39	-1.32	-0.41
	(0.28)	(1.44)	(0.29)
High school	-0.15**	0.15	-0.17**
	(0.07)	(0.40)	(0.07)
Some College	-0.42***	0.91*	-0.48***
	(0.09)	(0.47)	(0.09)
College	-0.56***	0.32	-0.62***
	(0.10)	(0.49)	(0.10)
Single	0.19		
	(0.13)		
Constant	-0.19	0.39	-0.15
	(0.31)	(1.65)	(0.32)
Observations	5,096	281	4,795
State FE	Yes	Yes	Yes
Age FE	Yes	Yes	Yes

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains men who are aged from 62 to 64, born between 1930 and 1943, and in the ASEC-CPS from March 1992 to March 2007. The estimates are not weighted and I only report logit coefficient estimates in this table. “Unemployment rate” indicates the monthly state-level unemployment rate. 1930-1943 are cohort dummies which indicate the year of birth *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Table 4 Estimated average SSDI enrollment for women using logit regression with ASEC-CPS from 1992 to 2007, 1937 is the omitted cohort (unweighted and SSDI of women, SSDI enrollment of single women, and SSDI of non-single women as dependent variables)

	(1)	(2)	(3)
Sample included ages:	62-64	62-64	62-64
Dependent Variables:	SSDI enrollment of women	SSDI enrollment of single women	SSDI enrollment of non-single women
1930	-0.23	-0.28	-0.13
	(0.21)	(0.93)	(0.22)
1931	-0.23	0.12	-0.19
	(0.20)	(1.02)	(0.21)
1932	-0.19	-0.55	-0.15
	(0.19)	(0.91)	(0.19)
1933	-0.14	0.78	-0.15
	(0.19)	(0.97)	(0.19)
1934	0.15	-1.78*	0.23
	(0.18)	(0.93)	(0.19)
1935	-0.09	-1.86*	-0.04
	(0.18)	(0.99)	(0.19)
1936	0.07	-0.96	0.06
	(0.18)	(0.91)	(0.19)
1937	--	--	--
1938	0.42**	-0.08	0.46***
	(0.17)	(0.97)	(0.17)
1939	0.30*	-0.56	0.37**
	(0.17)	(0.81)	(0.17)
1940	0.34**	-0.29	0.40**
	(0.17)	(0.81)	(0.18)
1941	0.42**	-0.09	0.44**
	(0.17)	(0.82)	(0.17)
1942	0.16	-0.82	0.23
	(0.16)	(0.79)	(0.17)
1943	0.26	-0.47	0.28*
	(0.16)	(0.73)	(0.16)
Unemployment Rate	0.04	-0.62	0.04
	(0.07)	(0.39)	(0.08)
Household Size = 1	0.23***	-0.75	0.24***
	(0.09)	(0.48)	(0.09)
Household Size > 3	-0.06	-0.11	-0.11**
	(0.05)	(0.20)	(0.05)
Black	0.28***	-0.40	-0.09
	(0.04)	(0.33)	(0.09)
Indian	-0.09	0.38	0.22
	(0.08)	(1.52)	(0.24)
Asian or Pacific	0.20	--	-0.14
	(0.23)		(0.29)
High school	-0.24	0.91***	0.10
	(0.28)	(0.34)	(0.08)
Some College	0.15*	0.46	-0.07
	(0.08)	(0.42)	(0.09)
College	-0.04	0.56	-0.37***
	(0.09)	(0.43)	(0.13)
Single	-0.29**	--	--
	(0.12)		
Constant	-17.93***	0.96	0.25
	(2.56)	(1.27)	(0.34)
Observations	4,457	321	4,091
State FE	Yes	Yes	Yes
Age FE	Yes	Yes	Yes

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains women who are aged from 62 to 64, born between 1930 and 1943, and in the ASEC-CPS from March 1992 to March 2007. The estimates are not weighted and I only report logit coefficient estimates in this table. “Unemployment rate” indicates the monthly state-level unemployment rate. 1930-1943 are cohort dummies which indicate the year of birth *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Table 5 Testing the difference in SSDI enrollment between men and women using logit regression with ASEC-CPS from 1992 to 2007, 1937 is the omitted cohort (unweighted and SSDI enrollment as the dependent variable)

	(1)
Sample included ages:	62-64
Dependent Variable:	SSDI enrollment
Female * 1930	-0.30
	(0.24)
Female * 1930	-0.30
	(0.24)
Female * 1930	-0.28
	(0.24)
Female * 1933	-0.06
	(0.24)
Female * 1934	0.24
	(0.24)
Female * 1935	0.04
	(0.24)
Female * 1936	-0.22
	(0.24)
Female * 1938	0.05
	(0.23)
Female * 1939	0.07
	(0.22)
Female * 1940	0.09
	(0.22)
Female * 1941	0.18
	(0.22)
Female * 1942	0.01
	(0.22)
Female * 1943	0.27
	(0.22)
Constant	-0.25
	(0.28)
Observations	9,553
State FE	Yes
Age FE	Yes
Cohort Dummies	Yes
Control variables	Yes

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains both men and women who are aged from 62 to 64, born between 1930 and 1943, and in the ASEC-CPS from March 1992 to March 2007. The estimates are not weighted and I only report logit coefficient estimates in this table. 1930* female to 1943*female are interaction terms between cohort dummies and female dummy. *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Table 6 Robustness Check: Alternative definition of SSDI enrollment (using whyss1 to define SSDI enrollment)

	(1)
Sample includes genders:	Men & Women
Dependent Variables:	SSDI enrollment (by an alternative identifier (WHYSS1*))
1936	0.200
	(0.132)
1937	--
1938	0.316**
	(0.144)
1939	0.266*
	(0.139)
1940	0.135
	(0.141)
1941	0.156
	(0.123)
1942	0.248**
	(0.120)
1943	0.614***
	(0.120)
Constant	-12.755
	(653.976)
Observations	4,906
State FE	Yes
Age FE	Yes
Control Variables	Yes

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains both men and women who are aged from 62 to 64, born between 1936 and 1943, and in the ASEC-CPS from March 2001 to March 2007. The estimates are not weighted and I only report logit coefficient estimates in this table. SSDI enrollment is defined by WHYSS1 = 2 (disabled). 1930-1943 are cohort dummies which indicate the year of birth *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Table 7 Falsification Check: omitting different births cohort from cohort 1931 to cohort 1942 (unweighted and SSDI enrollment as the dependent variable)

Part 1

	(1)	(2)	(3)	(4)
Dropped cohorts:	1931	1932	1933	1934
Dependent Variable:	SSDI enrollment	SSDI enrollment	SSDI enrollment	SSDI enrollment
1930	-0.058	-0.059	-0.057	-0.058
	(0.102)	(0.102)	(0.102)	(0.102)
1931	--	-0.055	-0.053	-0.054
		(0.098)	(0.098)	(0.098)
1932	-0.034	--	-0.033	-0.035
	(0.106)		(0.106)	(0.106)
1933	-0.086	-0.088	--	-0.087
	(0.110)	(0.110)		(0.110)
1934	-0.059	-0.060	-0.057	--
	(0.109)	(0.109)	(0.109)	
1935	-0.089	-0.091	-0.087	-0.089
	(0.109)	(0.109)	(0.109)	(0.109)
1936	0.162	0.160	0.164	0.162
	(0.105)	(0.105)	(0.105)	(0.105)
1937	0.084	0.083	0.086	0.085
	(0.101)	(0.101)	(0.101)	(0.101)
1938	0.263***	0.261***	0.264***	0.263***
	(0.098)	(0.098)	(0.098)	(0.098)
1939	0.081	0.080	0.082	0.081
	(0.094)	(0.094)	(0.094)	(0.094)
1940	0.197**	0.196**	0.198**	0.197**
	(0.087)	(0.087)	(0.087)	(0.087)
1941	0.170*	0.168*	0.171*	0.170*
	(0.095)	(0.095)	(0.095)	(0.095)
1942	0.111	0.110	0.113	0.111
	(0.087)	(0.087)	(0.087)	(0.087)
1943	0.025	0.024	0.027	0.026
	(0.091)	(0.091)	(0.091)	(0.091)
Constant	-11.035***	-11.048***	-11.062***	-11.028***
	(1.476)	(1.475)	(1.475)	(1.476)
Control Variables	Yes	Yes	Yes	Yes
Observations	18,444	18,444	18,444	18,444

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Part 2

	(1)	(2)	(3)	(4)
Dropped cohorts:	1935	1936	1937	1938
Dependent Variable:	SSDI enrollment	SSDI enrollment	SSDI enrollment	SSDI enrollment
1930	-0.057	-0.067	-0.064	-0.072
	(0.102)	(0.102)	(0.102)	(0.102)
1931	-0.053	-0.066	-0.062	-0.072
	(0.098)	(0.098)	(0.098)	(0.098)
1932	-0.033	-0.047	-0.043	-0.054
	(0.106)	(0.106)	(0.106)	(0.106)
1933	-0.084	-0.100	-0.095	-0.106
	(0.109)	(0.109)	(0.109)	(0.109)
1934	-0.056	-0.073	-0.068	-0.080
	(0.109)	(0.109)	(0.109)	(0.109)
1935	--	-0.105	-0.099	-0.112
		(0.109)	(0.109)	(0.109)
1936	0.165	--	0.151	0.138
	(0.105)		(0.105)	(0.105)
1937	0.088	0.069	--	0.061
	(0.101)	(0.101)		(0.101)
1938	0.265***	0.249**	0.253***	--
	(0.098)	(0.098)	(0.098)	
1939	0.083	0.069	0.073	0.063
	(0.094)	(0.094)	(0.094)	(0.094)
1940	0.198**	0.185**	0.189**	0.178**
	(0.087)	(0.087)	(0.087)	(0.087)
1941	0.172*	0.158*	0.162*	0.151
	(0.095)	(0.095)	(0.095)	(0.095)
1942	0.114	0.098	0.103	0.091
	(0.087)	(0.087)	(0.087)	(0.087)
1943	0.028	0.011	0.016	0.004
	(0.091)	(0.091)	(0.091)	(0.091)
Constant	-11.044***	-11.034***	-11.040***	-11.086***
	(1.476)	(1.476)	(1.476)	(1.475)
Control Variables	Yes	Yes	Yes	Yes
Observations	18,444	18,444	18,444	18,444

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Part 3

	(1)	(2)	(3)	(4)
Dropped cohorts:	1939	1940	1941	1942
Dependent Variable:	SSDI enrollment	SSDI enrollment	SSDI enrollment	SSDI enrollment
1930	-0.065	-0.072	-0.068	-0.066
	(0.102)	(0.102)	(0.102)	(0.102)
1931	-0.062	-0.070	-0.066	-0.064
	(0.098)	(0.098)	(0.098)	(0.098)
1932	-0.043	-0.051	-0.048	-0.046
	(0.106)	(0.106)	(0.106)	(0.106)
1933	-0.095	-0.103	-0.100	-0.098
	(0.110)	(0.110)	(0.110)	(0.109)
1934	-0.067	-0.076	-0.073	-0.071
	(0.109)	(0.109)	(0.109)	(0.109)
1935	-0.098	-0.108	-0.104	-0.102
	(0.109)	(0.109)	(0.109)	(0.109)
1936	0.153	0.143	0.147	0.148
	(0.105)	(0.105)	(0.105)	(0.105)
1937	0.075	0.065	0.069	0.071
	(0.101)	(0.101)	(0.101)	(0.101)
1938	0.254***	0.245**	0.248**	0.250**
	(0.098)	(0.098)	(0.098)	(0.098)
1939	--	0.065	0.069	0.070
		(0.094)	(0.094)	(0.094)
1940	0.189**	--	0.184**	0.186**
	(0.087)		(0.087)	(0.087)
1941	0.162*	0.153	--	0.159*
	(0.095)	(0.095)		(0.095)
1942	0.103	0.094	0.098	--
	(0.087)	(0.087)	(0.087)	
1943	0.017	0.008	0.011	0.013
	(0.091)	(0.091)	(0.091)	(0.091)
Constant	-11.045***	-11.066***	-11.066***	-11.014***
	(1.476)	(1.475)	(1.476)	(1.476)
State FE	Yes	Yes	Yes	Yes
Age FE	Yes	Yes	Yes	Yes
Control Variables	Yes	Yes	Yes	Yes
Observations	18,444	18,444	18,444	18,444

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains both men and women who are aged from 62 to 64, born between 1930 and 1943, and in the ASEC-CPS from March 1992 to March 2007. The estimates are not weighted and I only report logit coefficient estimates in this table. 1930-1943 are cohort dummies which indicate the year of birth *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Table 8 Falsification: using other disability income programs' enrollment as dependent variable

	(1)
Sample includes genders:	Men & Women
Dependent Variable:	Other Disability Income Programs Enrollment
1934	-0.422***
	(0.139)
1935	-0.247*
	(0.129)
1936	-0.129
	(0.125)
1938	-0.213*
	(0.117)
1939	0.150
	(0.097)
1940	-0.402***
	(0.117)
1941	-0.366***
	(0.120)
1942	-0.044
	(0.097)
1943	-0.110
	(0.102)
Unemployment Rate	-0.048***
	(0.011)
Age	-0.085***
	(0.022)
Black	0.777
	(1.030)
Indian	0.748
	(1.038)
Asian	0.747
	(1.036)
Pacific	0.400
	(1.119)
High school	0.247***
	(0.047)
Some College	0.409***
	(0.053)
College	0.585***
	(0.058)
Female	-0.475***
	(0.036)
Married	0.289***
	(0.067)
Constant	2.030
	(1.724)
Observations	38,181

Standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

Note: the sample contains both men and women who are aged from 62 to 64, born between 1934 and 1943, and in the ASEC-CPS from March 1992 to March 2007. Other disability programs include worker's compensation; company or union disability; federal government civil service disability; U.S. military retirement disability; state or local government employee disability; U.S. Railroad Retirement disability; accident or disability insurance; black lung miner's disability; or state temporary sickness payments. However, programs funded by Social Security or the Veterans' Administration are excluded. The estimates are not weighted and I only report logit coefficient estimates in this table. "Unemployment rate" indicates the monthly state-level unemployment rate. 1934-1943 are cohort dummies which indicate the year of birth *** indicates significant at 1% level, ** indicates significant at 5% level and * indicates significant at 10% level.

Figures

Figure 1 The Correlation between The AIME and The PIA Under SSA's 2008 Rules

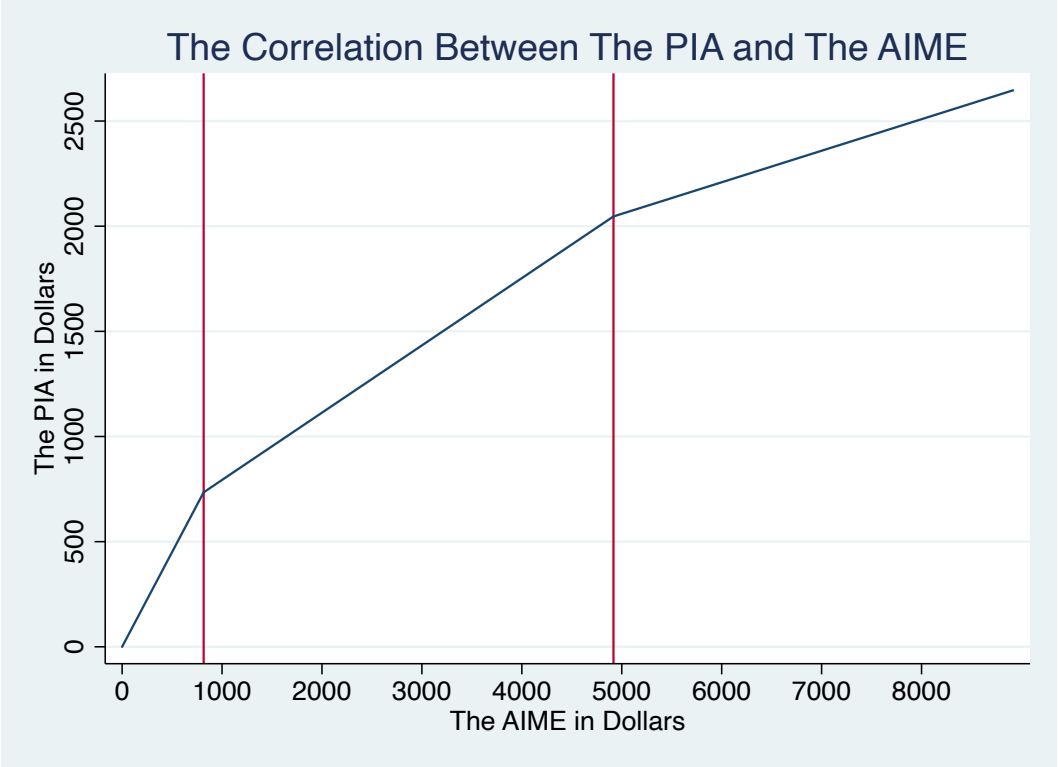

Fig. 1: The ratio between the average indexed monthly earning (AIME) and the primary insurance amount (PIA). The cut offs are $816 and $4,917. (Based on 2008 SSA's rule)

Figure 2

Figure 2 a) The Birth Cohort SSDI Enrollment Rate for The Pooled Sample (Based on Individuals Who Are Aged 62 – 64 and Were Born between 1930 to 1950 Using The ASEC-CPS From 1992 to 2013)

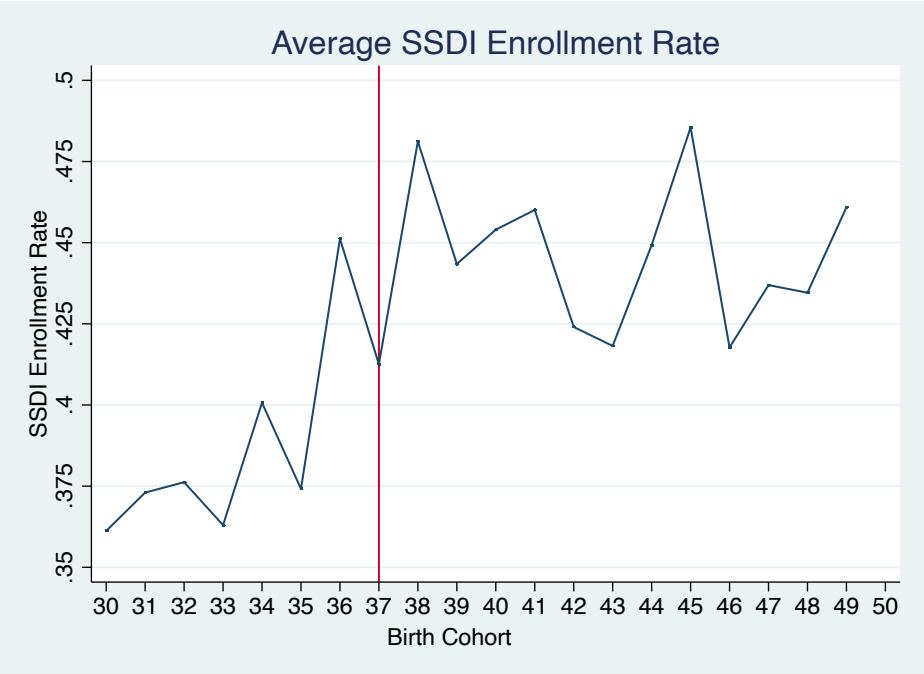

Figure 2 b) The Birth Cohort SSDI Enrollment Rate for Men or Women (Based on Individuals Who Are Aged between 62 and 64 and Were Born between 1930 to 1950 Using The ASEC-CPS from 1992 to 2013)

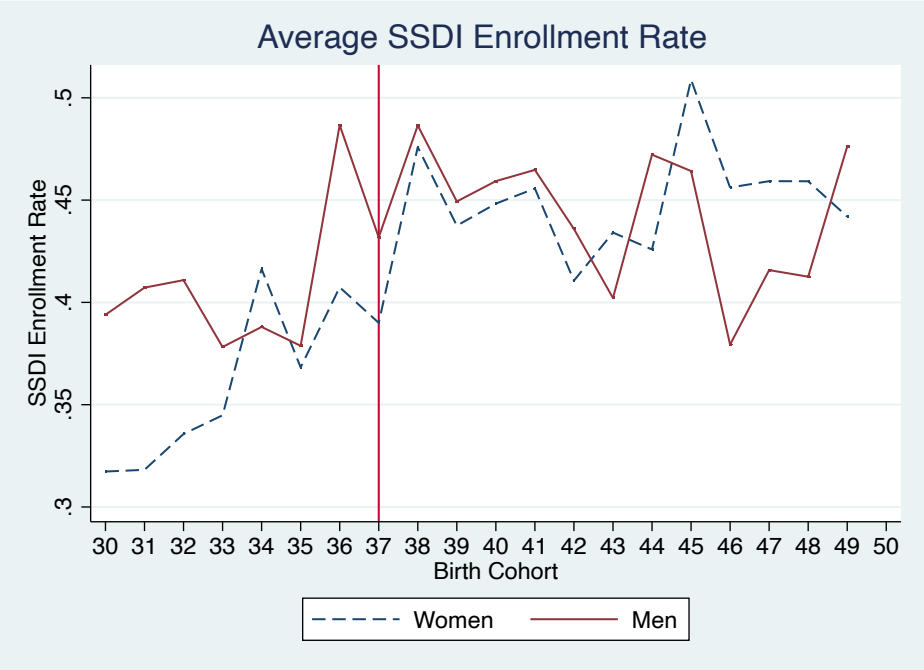

Figure 3 The Ratio between The Retired Worker Benefits and The AIME If An Individual Retired at Age 62 (Including Birth Cohort 1930 to Birth cohort 1950 when the FRA Increased from 65 to 66 Gradually)

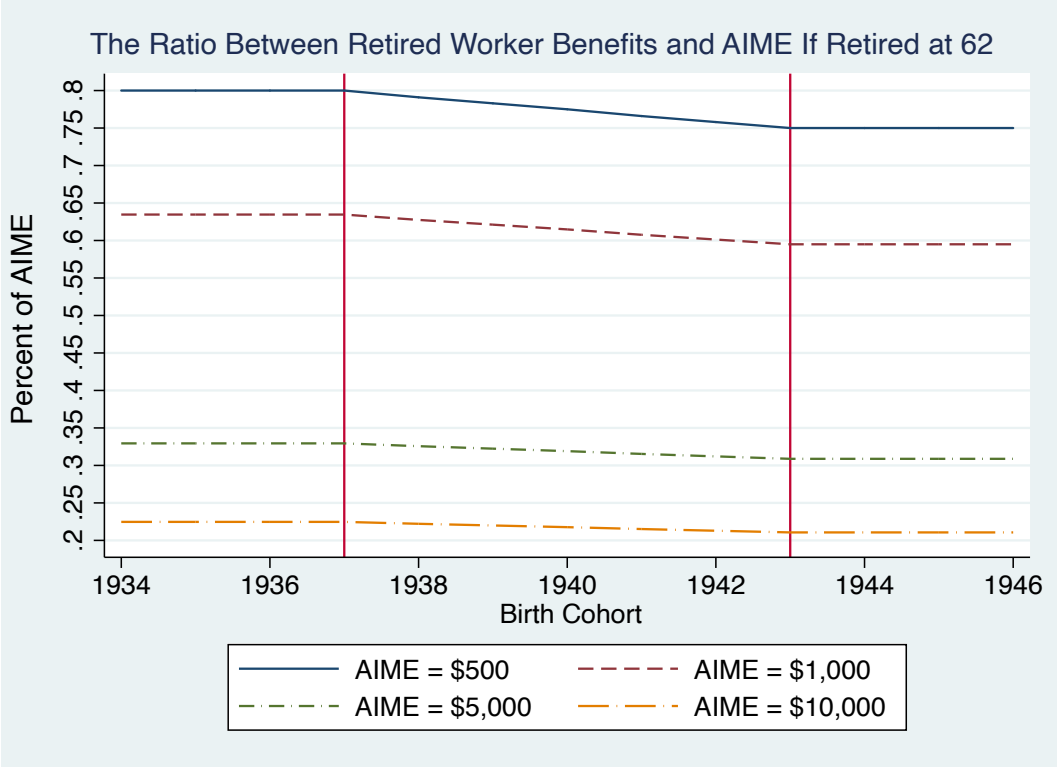

Fig.3: The ratio between the retired worker benefits and the AIME if a worker retired at age 62. (based on individuals born between 1930 and 1950 and SSA's 2008 rules). To clarify, the 25 percent-reduction in benefits for claiming the retired worker benefits early is calculated based on the PIA instead of the AIME. Since the PIA is calculated based on an increasing function of AIME, a 25-percent reduction in PIAs would result in different percentage reduction in AIMES for individuals with different AIMES.