Universitat Autònoma de Barcelona

Essays in Labor Economics

Alexey Filatov

Universitat Autònoma de Barcelona

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Author: Alexey Filatov

Advisor: Nezih Guner

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Contents

Chapter 1

Introduction

Individual labor supply decisions and factors thet shape them are among the most studied topics in economics. This thesis focuses on the two important topics in this broad field. The first is the relative importance of government policies and non-policy factors in shaping the labor supply of elderly people. Second topic is the effect of a major policy intervention, the introduction of the statutory minimum wage in Germany, on the individual wage expectations and durations of unemployment spells.

In the second chapter, I study the increase in the labor force supply of elderly in the US. In particular, since the mid-1980, the labor force participation and hours worked by the US seniors, those above 62, has been steadily increasing. The labor force participation was at the minimum of 16% in 1985 and reached 26% in 2013. This is in contrast to the steady decline of the labor force supply of seniors between 1950 and 1985. The labor supply behavior of seniors has very important implications for the sustainability of the Social Security system, which is a problem faced by many developed countries. As a result, understanding the driving factors behind the changes in the labor supply of seniors is critical.

The potential factors that caused the change in labor supply behavior of seniors can be policy changes, such as the increase in the Normal Retirement Age, or factors not related to policies, such as an increase in longevity. I focus on the differences in the labor supply of the two cohorts of seniors in the US, those born between 1915 and 1934, the Great Depression Kids, and those born between 1945 and 1964, the Baby Boomers. Both labor force participation rate and hours worked of Baby Boomers are higher than those of the Great Depression Kids. In order to identify the key factors that affect the labor supply behavior of seniors, I build a life-cycle model of labor supply, retirement, and asset accumulation. I then bring this model to the data. The calibrated model does a very good job in accounting for the differences between the two cohorts. Finally, by the means of counterfactual exercises, I assess the importance of each driving factor in isolation and do counterfactual policy experiments.

In the third and the fourth chapters we study the effects of the introduction of statutory minimum wage in Germany on January 1, 2015. This reform was one of the major policy interventions in Germany since the beginning of 2000s, and its potential effects were highly debated prior to the implementation. We use the German Socio-Economic Panel dataset, a representative longitudinal survey of the German population, to assess the two questions: how the reform affected individual reservtion wages, and what is the effect of the reform on durations of unemployment spells.

In the third chapter, we empirically study the effect that the introduction of the federal minimum wage had on the distribution of individual reservation wages in Germany. To establish a causal link between reservation and minimum wages, we treat the reform as a quasi-natural experiment. The principal idea is that in the medium and long-run, reservation wages are likely to be not static and may themselves adjust to the economic environment and changes in the minimum wage. We estimate the effect of the reform on the reservation wages using a difference-in-difference framework with continuous treatment.

In the fourth chapter, we are looking at the effect of the minimum wage reform on the duration of unemployment. Since longer unemployment spells are associated with individual economic costs beyond the loss of financial well-being, this topic deserves special attention. In our empirical study, we use a survival analysis in discrete time to demonstrate the causal effect the reform had on the unemployment durations of various population groups.

Chapter 2

Why Are Older Americans Working More Nowadays?

Both labor force participation and hours per worker of seniors, individuals above age 62, have been growing steadily in the US since the mid 1980s. This is in contrast to the steady decline in the labor supply of seniors that began as early as in the 1950s. This paper uses data from the Health and Retirement Study (HRS) to estimate a life-cycle model of labor supply, retirement, and wealth accumulation in order to contrast the labor supply behavior of two cohorts in the US: individuals born after World War I ("the Great Depression Kids"), and those born after the World War II ("the Baby Boomers"). The paper focuses on the differences between these two cohorts in earning and health dynamics as well as policies that they face, a gradual increase in Normal Retirement Age and the elimination of the earnings test in 2000, as potential sources of change. The results show that the effects of policies and policy-unrelated factors are of similar magnitude. The elimination of the earnings test had the biggest impact of all policies. Jointly, the rise in out-of-pocket medical expenditures and the increase in life expectancy are the dominant factors among the non-policy factors.

Around three decades ago labor force participation rate of seniors in the US hit a historical minimum.[1] Since then it has been growing steadily and nearly doubled by 2014. This striking trend was accompanied by the rise of annual hours per worker of seniors, which increased by about 50% during the same time period (Figure 2.1 summarizes these patterns). A clear understanding of factors behind such a significant behavioral change, as well as its economic implications, are important from various perspectives.

First, the trend is the aggregate result of the policies implemented by the US government and policy-unrelated factors. One key policy change was a gradual increase of the Normal Retirement Age (NRA) for those born after 1937 that started in 2002. Another important change occurred in 2000 when The Senior Citizens Freedom to Work Act was enacted, eliminating the earnings test for those who have reached Normal Retirement Age. The non-policy factors include, among others, the increase in life expectancy of younger cohorts, and changes in the earnings dynamics (driven by factors such as the college premium growth and the manufacturing sector decline). The relative importance of these forces and interactions between them are

[1] Particularly, I refer to individuals of 62 years of age and above.

ambiguous (see, e.g., a recent review by Blundell, French and Tetlow (2016)) and disentangling their contributions remains an open question.

Second, the large cohort of Baby Boomers is currently either approaching to or are already in their sixties. Together with the fact that the labor force participation rate of prime-age individuals has been declining since the beginning of the new millenium, the seniors will constitute an increasingly larger part of the US labor force.[2] The effect of this change on the aggregate savings behavior and welfare of the elderly is unclear. There is also a continuing debate about the effect of increase in the Normal Retirement Age on the welfare of the poorest sections of population (e.g. Biggs (2016), Klein (2015), Mermin and Steuerle (2006)). Hence, it is crucial to be able to asses how policy changes affect the labor supply behavior of the elderly.

Finally, some argue that the working seniors can alleviate the increasing cost of the social security system posed by population ageing.[3] Working seniors pay income taxes; furthermore, their social security benefits are reduced due to the earnings test. This means that pressure of aging population on social security system is alleviated by seniors who choose to work instead of retirement. As a result, it is critical to understand factors behind the labor supply decisions of the seniors.

The goal of this paper is to quantify the relative effects of different factors driving the increase in participation rate and labor supply of the US seniors. The model economy has a life-cycle structure and distinguishes between agents of different cohorts and genders. Every period agents choose how much to work, how much to save and whether or not to claim social security benefits if they are eligible. There is no aggregate uncertainty in the model, but there are several sources of idiosyncratic uncertainty. Individual health, wage and medical expenditures are subject to stochastic shocks. Furthermore, every period individuals face a risk of death. In the model economy, the elderly save for two reasons: first, they have a precautionaty savings motive since they face higher riks of experiencing significant medical expenditures as they age. Second, the elderly in the model economy have a warm-glow bequest motive. These two forces prevent the elderly from deaccumulating their assets sharply at the end of their life as it would be predicted by a standard life-cycle model.

I focus on two relatively large cohorts of individuals: those born between 1915 and 1934 (post-WWI period and Great Depression, "Great Depression Kids") and those born between 1945 and 1964 ("Baby Boomers"). Figure 2.2 shows the differences in annual labor supply and participation rates between the two chosen cohorts as observed in the March supplement of the Current Population Survey. Both men and women of the Baby Boomers cohort have higher participation rates and work longer hours than the Great Depression Kids of the same age. What are the potential differences between these two cohorts that can account for the differences in their labor supply behavior? I integrate into the model several potentially important explanations of the labor force participation trends that have been put forward in the literature.

First, the government policies that are aimed at, directly or indirectly, keeping the seniors in the labor force include the increase in the Normal Retirement Age, changes in the early retirement penalty and increase of delayed retirement credit, the elimination of the earnings test

[2]See, for example, https://www.bls.gov/emp/ep_table_303.htm

[3]See, for example, https://www.ssa.gov/policy/docs/ssb/v66n4/v66n4p37.html

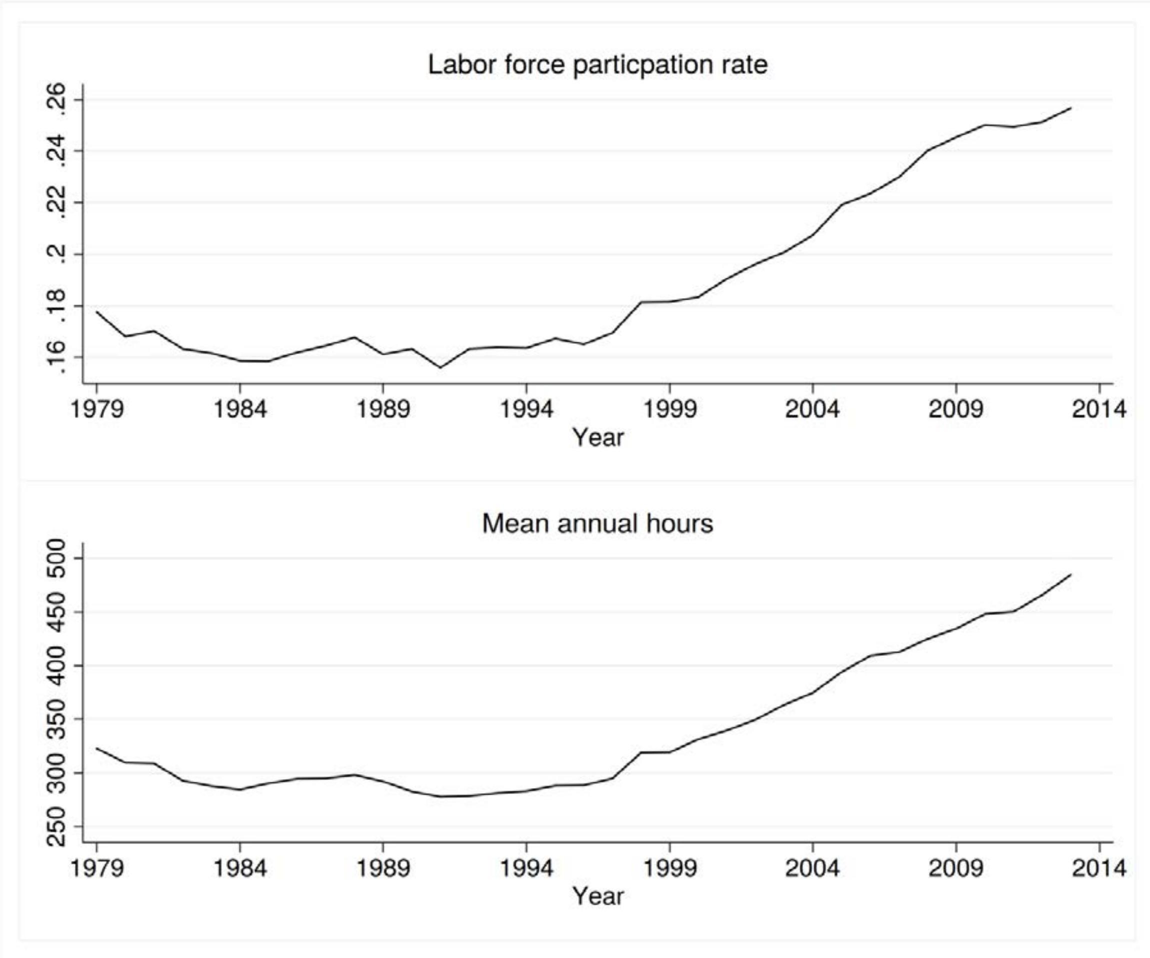

Figure 2.1: Labor force participation rate and mean annual hours of the seniors aged 62 and older. Data sample is taken from the March supplement of the US Current Population Survey (CPS) for the years from 1980 to 2014. The access to the data is provided by the Integrated Public Use Microdata Series (IPUMS), hosted by the Minnesota Population Center.

at the NRA in 2000, and changes in the effective income taxes. Wedges created by these changes provide Baby Boomers with additional incentives for staying in the labor force longer. The quantitative results suggest that the elimination of the earnings test plays the most significant role among the policies. Indeed, the earnings test creates an artificial earnings threshold, and both in the data and in the model the bunching of individuals with earnings just below the threshold is observed. That is, many individuals choose to work the amount of hours which will ensure their annual earnings do not exceed the threshold imposed by the earnings test.

Second, the life expectancy of the current pre-retirees is higher. From individual prospective this means longer post-retirement years to finance, which increases the risk of running down individual savings to maintain the desired quality of life. Furthermore, one might perceive that living longer bears higher risk of severe health conditions at older ages, which is inevitably tied to very high medical spendings. I find this to be the most important factor driving the change in the behavior of the Baby Boomers as compared to Great Depression Kids. Indeed, the quantitative results of the baseline model suggest that increase in the survival rate together with the rise in medical spendings are the most important driving factor.

The benchmark economy is populated by the large number of forward-looking individuals.

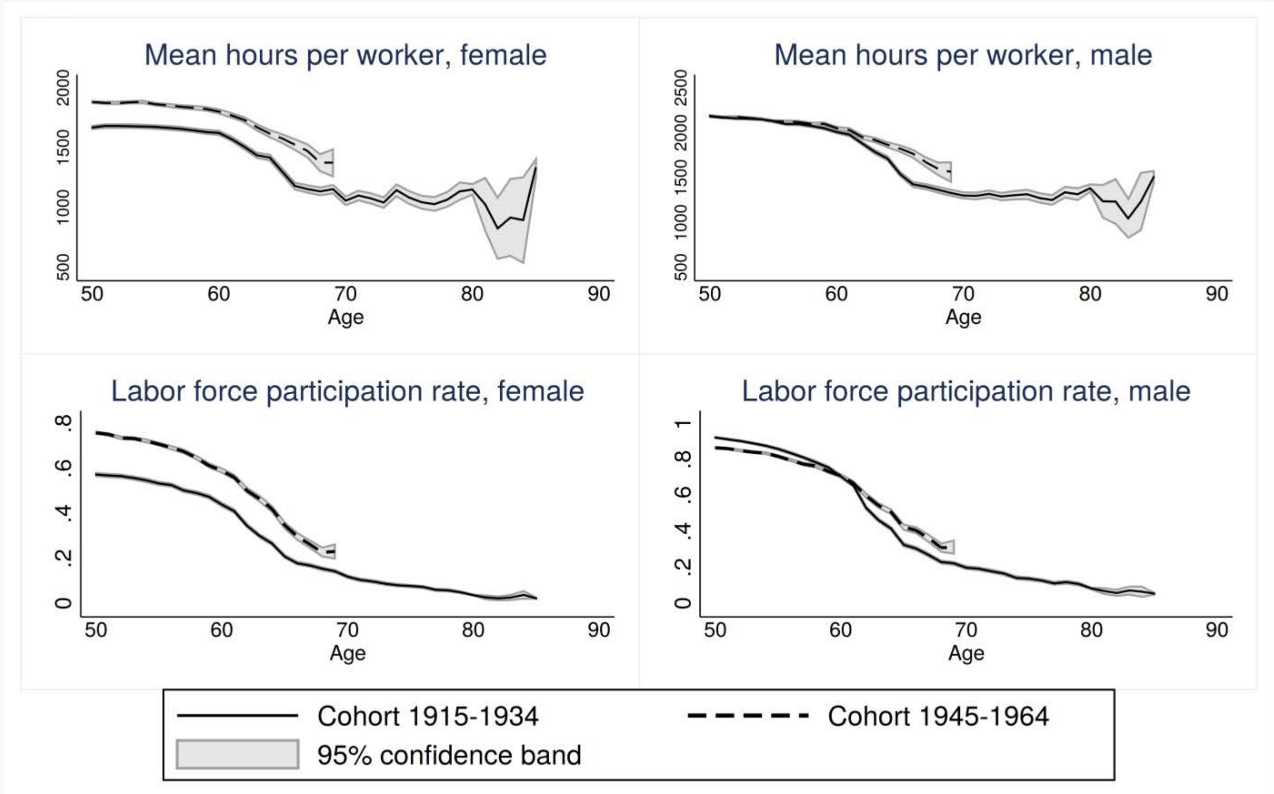

Figure 2.2: Mean hours per worker and labor force participation rates of the two cohorts of the seniors: those born after World War I ("Great Depression Kids"), and those born after World War II ("Baby Boomers"). Data sample from the IPUMS-CPS, March supplement, 1980 to 2014).

They face the exogenous set of policies, stochastic health dynamics, wages and out-of-pocket medical expenditures, as well as per-period survival rates of the older cohort. Given these exogenous factors, they choose how much to work, how much to save and when to apply for social security benefits if eligible.

I use this model to implement two exercises. First, I study the effect of different factors, such as the increase in longevity, the increase in out-of-pocket medical spendings and the inrcrease in the Normal Retirement Age, using counterfactual exercises. To this end, I first estimate the parameters of the model to match the behavior of the older cohort. Then, I subject the individuals in the benchmark economy to exogenous factors (health, wage and medical expenditire shocks) as well as policies that are faced by inidividuals born between 1945 and 1964 ("Baby Boomers"). The model does a very good job in accounting for changes in the labor supply behavior of these two cohorts. Then I open up channels of the differences between these two cohorts one by one and look at the resulting labor supply behavior of the Great Depression Kids. For example, in order to highlight the contribution of the earnings test, I simulate a variation of baseline economy for the Great Depression Kids where the earnings test is eliminated at the NRA. In the similar way, I uncover the interactions between the driving factors, opening up corresponding channels simultaneously.

The results show that the effects of policies and policy-unrelated factors are of similar magnitude. The elimination of the earnings test at the NRA has a strongest effect on both participaton and hours worked of individuals between 62 and 67, explaining in isolation almost half of the overall increase in participation and one-third of the incease in hours worked.

Jointly, the rise in out-of-pocket medical expenditures and the increase in life expectancy are the dominant factors among non policy-related factors, i.e. if the Great Depression Kids had both life expectancy and OOP medical expenditures of the Baby Boomers, the model predicts very large increase in both participation and hours worked. This reflects the fact that longer life expectancy implies higher chances to live up to a very old age, where extreme medical shock is very likely. As a result, seniors continue to work to hoard additional resources to insure against this shock. Finally, the results show that the medical expenditure risk and bequest motive play an equal role in accounting for the slow deaccumulation of assets by the elderly as they age.

2.1 Literature

The paper relates to the two major strands of literature.

First, it draws on empirical literature that investigates the increase in labor force participation of elderly in US since mid-80's. A number of papers document the trend and put forward possible explanations, e.g. Clark and Quinn (2002), Friedberg (2007), Juhn et al. (2006), DiCeccio et al. (2008), Maestas and Zissimopoulos (2010). Blau and Goodstein (2010) claim social security changes can account for up to 18-20% of the effect. Hurd and Rohwedder (2011) use HRS data to study the effect of pensions (in particular, moving from DB to DC pension plans) on labor force participation of elderly, and find the effect to be modest, at around 2.5% of overall increase. Schirle (2008) identifies a coordination in retirement schedule among spouses as another important reason and finds that husband's response to wives' labor force participation can explain up to 25% of the increase in married males' participation.

Second, it is closely related to the papers that use life-cycle models to explore various aspects of behavior of seniors. The excellent review of this literature is Blundell, French and Tetlow (2016). The models of the first wave were concerned primarily with timing of retirement (e.g. Gustman and Steinmeier (1986), Burtless (1986)). In these models, retirement was an abrupt process, and they abstracted from any uncertainty. The models failed to replicate early retirement puzzle. Rust and Phelan (1997) resolved the puzzle by introducing realistic uncertainty and individual inability to save and borrow, but they also abstracted from the individual choice of hours to work.

French (2005) extends these models by allowing agents to save, borrow to a certain limit and choose hours to work. He finds that the tax wedge embedded in the earnings test to be one of the most important determinants of the early retirement. Yavuzoglu (2016) investigates the labor force attachement of the senior single males beyond the Normal Retirement age, and finds that the elimination of the earnings test explains one-forth of the increase in the labor force participation of the individuals between 66 and 70. Similarly, I find the elimination of earnings test at the NRA to be the most important policy change that affected the inctrease in labor force participation of Baby Boomers aged 62-67. However, I still find that joint effect of increase of out-of-pocket medical expenditures and increase in longevity is dominating. Van der Klaauw and Wolpin (2008) estimate a model of retirement and savings on the sample of households from the first three waves of the HRS in order to simulate the response of these households on the changes in various hypothetical policies (social security benefit levels, earnings test, Early and Normal Retirement Ages). They predict significant responses in average annual hours of work and participation rate to the changes in policies.

De Nardi, French and Jones (2010) model savings of retired singles under uncertain medical expenditures and heterogenous life expectancies and find that risk of extreme medical spendings at older age to be one of main reasons why seniors decumulate their assets slowly. Kopecky and Koreshkova (2014) highlight the expexctation of extreme medical spendings for nursing homes as the main determinant of slow asset decumuation of US seniors. Lockwood (2010) investigates the relative effect of precautionary and bequest motives on speed of asset decumulation of seniors, and finds them to be of relatively the same magnitude. This is also the case in the current paper. Blundell, Britton, Dias and French (2016) investigate the effects of health dynamics on the employment of seniors in the US and the UK, and find that permanent health shocks have a strong and lasting negative effect on employment. French and Jones (2011) study the effects of employer-provided health insurance, Medicare, and social security on retirement behavior. Casanova (2010) studies the effects of leisure complimetarity in spousal retirement decision, and finds it to be an important determinant of retirement timing. Bairoliya (2015) finds that moving from Defined Benefit to Defined Contribution pension plans results in a significant increase of labor force participation rates of 57 to 69 years old.

However, these papers abstract from differences in behavior of different cohorts, which is the key contribution of the current paper. I provide direct comparison between the two different cohort of seniors. Furhermore, to my knowldge, this paper is the first attempt to disentangle and quantify the effects of the potentially relevant factors on both extensive and intensive margins of labor force supply decisions of seniors.

2.2 Model

I consider a partial equilibrium life-cycle model of labor supply decisions in old ages. Individual of age t, with t = 59, 60…,100, chooses consumption c_t, savings a_{t+1}, hours of work l_t, and an age $\alpha \in \{62, \ldots, 100\}$ at which to apply for social security benefits. Let the amount of social security benefits of an agent of age t who has claimed them at age α to be denoted by b_t^{α}. There is a chance of death every period. Individuals can choose to work zero hours and claiming social security benefits doesn't prevent them from working.

The agents are of different gender $\mathbf{g} \in \{m, f\}$, are either single or married $j \in \{0, 1\}$, belong to one of the two cohorts $\mathbf{c} \in \{1, 2\}$, and can be healthy or unhealthy, indicated as $h_t \in \{0, 1\}$. As I detail below, belonging to a particular cohort determines the set of government policies related to retirement and social security structure (such as the Normal Retirement Age), effective income tax, and risk-free saving rate. I focus on the two cohorts of individuals: those born between 1915 and 1934 (post-WWI period and Great Depression, which I label as "Great Depression Kids") and those born between 1945 and 1964 (which I label as "Baby Boomers").[4] Combination of cohort, gender, and health status determines average life-cycle profiles of wages

[4] The choice of these cohorts is determined by data limitations. First, the cohorts have to contain sufficiently large number of individual observations. Second, the cohorts have to be separated by the time period long enough for the differences in their labor supply behavior to be significant. Third, the individuals of the older cohort have to reach their sixties in mid 1980s - mid 1990s, whereas the individuas of the younger cohort have to enter these age in mid 2000s and later.

$\bar{w}_t^{\mathbf{g},\mathbf{c}}(h_t)$, out-of-pocket (OOP) medical expenditures $\bar{m}_t^{\mathbf{g},\mathbf{c}}(h_t)$, and survival probabilities $s_t^{\mathbf{g},\mathbf{c}}(h_t)$ for an individual. Furthermore, health status $h_t^{\mathbf{g},\mathbf{c}}$ is stochastic and follows a Markov process that depends on cohort and gender. Both wages and out-of-pocket medical expenditures are a combination of a life-cycle profile and a persistent idiosyncratic shocks, which also follow a Markov process that do not depend on gender, cohort and marital status.

The marriage decision is not modelled explicitly. Instead, marital status is defined at the beginning of an individual life cycle. I assume that an unmarried individual remains single. Marital status of an agent does not change except due to the death of the partner. Survival probability of a spouse is given by $s_t^{1-\mathbf{g},\mathbf{c}}(h_t)$ of an opposite gender. Marital status determines an annual amount y_t^p of spousal income, and taxable share of social security benefits. I provide further details in the following sections.

Each individual is therefore subject to four sources of idiosyncratic uncertainty at any given age: survival probability, health shock, wage shock and medical spendings shock. In addition, married individuals are uncretain about spousal income, which nullifies with the death of the spouse. There is no aggregate uncertainty.

2.2.1 Preferences

Agents maximize expected lifetime utility. The expectation is taken with respect to agent's survival probability, wage, health status and medical expenditure shocks. Future is discounted by a common discount factor β. Individuals derive utility from consumption, leisure and leaving bequests. Starting from the Early Retirement Age of 62, individuals can irreversibly apply for social security benefits. If one applies for the benefits at age α, then the actual annual amount of social security transfers at age $t \geq \alpha$ is denoted by b_t^{α}.

The survival probability $s_t^{\mathbf{g},\mathbf{c}}(h_t)$ depends on individual's current health status h_t, cohort and gender. Surviving agent enjoys consumption and leisure. Upon death, individual derives utility from leaving a bequest $B(a_t)$, which depends on asset level at the time of death. Labor force participation is costly, and the cost of participation increases with age. Each period agents have an opportunity to work and potential individual wage at any given age is determined exogenously. Retirement is modeled as a form of non-participation, i.e. choosing zero hours, with individuals being able to re-enter labor force at no cost. In what follows, I suppress indices of cohort and gender unless necessary.

Per-period utility function takes the form

$$u(c,l,h,t) = \frac{[c^{\gamma_c}(\bar{L} - \delta h - l - I_{l>0}(\kappa + \xi t^{\nu})^{\gamma_l}]^{1-\sigma}}{1-\sigma},$$

where $\bar{L}$ is annual hours endowment of an agent, $h \in \{0,1\} = \{good, bad\}$ is binary health status, and δ is a leisure penalty associated with poor health. Furthermore, $I_{l>0}$ is an indicator that takes value of 1 if an individual participates in the labor market, and 0 oterwise, κ is the fixed cost of participation, and function ξt^{ν} reflects the increase of participation cost with age. Finally, γ_c and γ_l are consumption and leisure weights, respectively, and σ is the parameter of relative risk aversion.

Bequest function $B(a)$ is defined as

$$B(a) = \eta \frac{(a+d)^{(1-\sigma)\gamma_c}}{1-\sigma},$$

where η is the magnitude of the bequest motivation, and d characterizes the extent to which bequests are luxury goods. If d is relatively large, an individual only gets significant utility gain if he or she leaves a sizeable bequest.[5]

2.2.2 Budget Constraint

The timing of the model is as follows. At the start of each period, health, wage, and out-of-pocket medical expenditure shocks are realized and individuals observe the realizations. If an individual is at least 62 and has not applied to social security benefits yet, then he or she chooses whether to apply now or wait. Next, individuals decide how many hours to work, receive corresponding salary and transfers, and pay taxes. Finally, consumption and saving decisions are made. At the very end of the period, survival shocks are realized.

Individual hourly wage is exogenous and modelled as a combination of a life-cycle profile and a persistent Markov process:

$$\log w_t^{\mathbf{g},\mathbf{c}}(h_t) = \log \bar{w}_t^{\mathbf{g},\mathbf{c}}(h_t) + \hat{\varepsilon}_t^w, \tag{2.1}$$

where $\bar{w}_t^{\mathbf{g},\mathbf{c}}(h_t)$ is the average life-cycle profile of individual wages, and $\hat{\varepsilon}_t^w$ is corresponding stochastic process represented by a Markov chain. Similarly, annual out-of-pocket medical expenditures are given by

$$\log m_t^{\mathbf{g},\mathbf{c}}(h_t) = \log \bar{m}_t^{\mathbf{g},\mathbf{c}}(h_t) + \hat{\varepsilon}_t^m, \tag{2.2}$$

with the notation being analogous to that of wage process. I provide a detailed explanation of how these objects are constructed in the sections 2.3.4 and 2.3.3.

Consider a person of age t with wage rate w_t (defined as in 2.1) and savings a_t who chooses to work $l_t \geq 0$ hours. His or her total taxable income net of transfers, y_t, is given by the sum of labor income, spousal income and return on savings

$$y_t = w_t l_t + \mathbb{I}(married) \times y_t^p + r^{\mathbf{c}} a_t, \tag{2.3}$$

where $\mathbb{I}(married)$ is marital status dummy, y_t^p is the spousal income and $r^{\mathbf{c}}$ is cohort-specific riskless rate of return. Let period-t medical expenditure of this individual mt be determined as in equation 2.2. Then, the budget constraint for this person is given by

$$c_t + a_{t+1} + m_t = a_t + T(y_t + \mu(y_t, b_t^{\alpha})) + (1 - \mu(y_t, b_t^{\alpha})) + b_t^{ma}, \tag{2.4}$$

with

$$a_{t+1} \geq 0,$$

[5] According to Lockwood (2016), bequest motives for the seniors are roughly as important as precautionary motives in explaining slow wealth decumulation at old age.

where c_t is period-t consumption and a_{t+1} is savings choice that must satisfy borrowing constraint. Gross individual taxable income is sum of y_t and a taxable portion of social security benefits $\mu(y_t; b_t^\alpha)$. Tax function $T(\cdot)$ maps individual taxable income into a disposable income. The share of the benefits that is subject to the income tax, $\mu(\cdot,\cdot)$, depends on y_t and level of benefits b_t^α. Social security transfer b_t^α is modelled following the actual rules of the Social Security Administration, hence it's a complex object. The next section of the paper is devoted to the description of the retirement benefits calculation in the US and it's implementation in the model. It includes the definition of taxable share of retirement benefits and function $\mu(\cdot,\cdot)$ associated with it.[6]

Finally, I follow Hubbard, Skinner, and Zeldes (1995) and include a government transfer b_t^{ma} as a mean to provide poor or impoverished individuals with a minimal level of consumption, a consumption floor $\underline{c}$, so that

$$b_t^{ma} = \max\{0, \underline{c} + m_t - (a_t + T(y_t + \mu(y_t, b_t^\alpha)) + (1 - \mu(y_t, b_t^\alpha)))\}. \tag{2.5}$$

The transfer b_t^{ma} captures the Medicaid-type insurance: one becomes eligible to it if and only if all other resources are exhausted due to a high medical expenditure shock m_t.

2.2.3 Social Security

This section provides details on the construction of social security benefits b_t^a in the equation 2.4, as well as the description of earnings test rules, and details on taxation of social security benefits, i.e. function $\mu(\cdot,\cdot)$. An individual becomes eligible for social security benefits upon reaching Early Retirement Age of 62. Benefit application is irreversible: as soon as one starts to receive benefits, one can not stop it. The benefits are then paid until death. There are four major factors that affect actual amount of benefits that individual receives.

First, the base for benefit calculation is the Average Indexed Monthly Earnings ($AIME$) of individual, which is the average of earnings over the 35 highest earning years in the labor market. Second, based on the AIME, the amount of benefits that one would get at Normal Retirement Age, Primary Insurance Amount (PIA), is calculated. In what follows, I denote te PIA as b_t, treating it as a potential amount of benefits that individual is entitled to. Third, the PIA is reduced or increased according to early retirement penalty or delayed retirement credit, respectively. Fourth, if an individual elects to receive benefits and works at the same time, the benefits might be subject to the earnings test or income tax.

Average Indexed Monthly Earnings

The amount of social security benefits that one is entitled to is based on the history of individual earnings, adjusted for inflation. I use the following formula to calculate the AIME within the

[6]The model currently does not distinguish between individual savings and employer-provided pensions, despite the illiquidity of the latter until certain age plays a potentially important role.

model:

$$AIME_{t+1} = \begin{cases} AIME_t + \frac{\min\{w_t l_t, 90000\}}{35}, \ \mathcal{T} < 35 \\ AIME_t + \max\{0, \frac{\min\{w_t l_t, 90000\} - AIME_t}{35}\}, \ \mathcal{T} \geq 35 \end{cases} \tag{2.6}$$

where $\mathcal{T}$ is the actual number of the total individual working years prior to age t.

Essentially, it implies that the AIME is strictly increasing as long as an individual has less than 35 years of job experience, and nondecreasing afterwards. Notice also that time-t contribution to the AIME is actually limited from above. The limit, known as the contribution and benefit base, is set annually by the federal Old-Age, Survivors, and Disability Insurance (OASDI) program.[7] All monetary terms in the model are expressed in 2005 US dollars, so any dollar amounts I'm going to present in the paper are already converted to $2005, unless it is explicitly stated otherwise. In 2005, contribution and benefit base was $90000, therefore I use this value as contribution cap, i.e. if individual earnings $w_t l_t$ exceed $90000, I use this limit in AIME calculations instead of actual earnings.

Primary Insurance Amount

Next step is to calculate the Primary Insurance Amount (PIA) at age t as a function of the AIME and the two AIME bendpoints:

$$PIA_t \equiv b_t = \begin{cases} 90\% \text{ of the first } \$7524 \text{ of annual } AIME_t, \text{ plus} \\ 35\% \text{ of annual } AIME_t \text{ over } \$7524 \text{ and through } \$45348, \text{ plus} \\ 10\% \text{ of annual } AIME_t \text{ over } \$45348 \text{ and through the cap of } \$90000. \end{cases} \tag{2.7}$$

The graphical representation of this formula is shown on Figure 2.3.

By definition, the PIA is a value of pre-earnings test social security benefits for those who claim them exactly at the Normal Retirement Age (65 for Great Depression Kids and an average of 66.3 for Baby Boomers). In the case of early or delayed application the amount is adjusted. For each year before the NRA that an individual first collects benefits, they are reduced by 6,7% per year during three nearest years (e.g. 62 to 65, or 63 to 66), and by 5% for any additional year before that but not earlier than 62 (the latter thus only appy to those whose NRA is higher than 65, e.g 62 to 63 if the NRA is 66). In partiular, penalty or credit adjustment depending on claim age α for the two cohorts is calculated by

$$\mathcal{P}_\alpha = \begin{cases} -(NRA - \alpha) \times 0,0667 \text{ if } \alpha \geq 62 \cap \alpha < NRA \cap (NRA - \alpha) \leq 3, \\ -0.2 - (NRA - 3 - \alpha) \times 0.05, \ \text{if } \alpha \geq 62 \cap (NRA - \alpha) > 3, \\ (\alpha - NRA) \times 0.0425, \ \text{if } \alpha \geq 65 \cap \alpha < 70 \cap \mathbf{c} = 0, \\ (\alpha - NRA) \times 0.08, \ \text{if } \alpha \geq NRA \cap \alpha < 70 \cap \mathbf{c} = 1. \end{cases} \tag{2.8}$$

[7] More detailed information on construction of these values can be found on the US Social Security website https://www.ssa.gov/oact/cola/Benefits.html

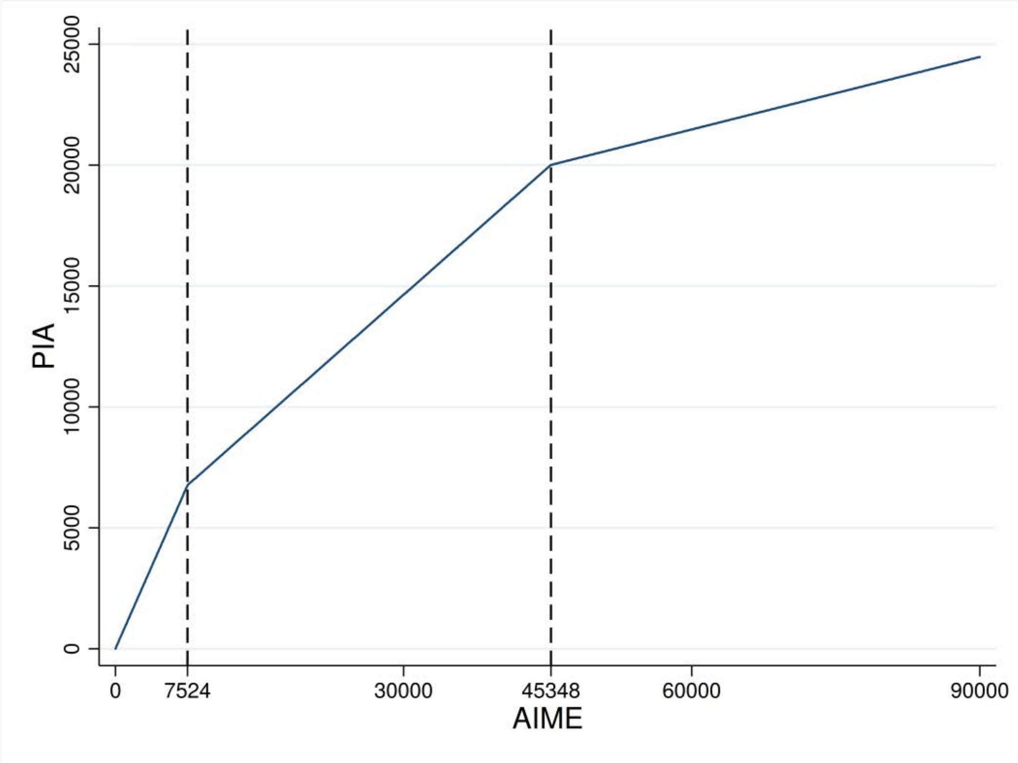

Figure 2.3: The dependence of the annual Primary Insurance Amount on the individual annualized Average Indexed Monthly Earnings and the two AIME bendpoints, expressed in 2005 US dollars.

The baseline pre-eanings test amount of benefits for an individual who claims them at age α is thus

$$b_{Bt}^{\alpha} = b_t(1 + \mathcal{P}_{\alpha}).$$

Figure 2.4 exemplify these rules for Normal Retirement Age of 65 and 66.

To interpret 2.8, first consider a Great Depression Kid who applied for social security at the early retirement age of 62, that is, three years short of the NRA of 65, having a particular value of the AIME on the personal record. The PIA of this individual is then calculated by 2.7. This amount has to be reduced by 6,7% per each year of early retirement, that is, by $6,7\% \times 3 = 20\%$ in total. The reduced amount is then represents an annual social security benefit transfer to the individual. Second, consider a Baby Boomers whose NRA is 66, retiring at 62, that is, four years short of the NRA. The PIA of this individual is going to be reduced by $6,7\% \times 3 + 5\% \times 1 = 25\%$. This difference creates an incentive for Baby Boomers as compared to Great depression Kids to delay benefit draw past 62.

Similarly, for each year of delaying benefit claim past the NRA and until 70, the PIA would increase by additional $3 - 5\%$ (averaging at 4,25%) for the Great Depression Kids, which was actuarially unfair, and by 8% for Baby Boomers, which is roughy actuarially fair (e.g. French 2005). Actuarial unfairness of delayed retirement for Great Depression Kids worked against delayed retirement. Since it was actuarially unfair, an individual was losing in total expected social security wealth if the retirement was delayed, whereas it's not the case for Baby Boomers anymore.

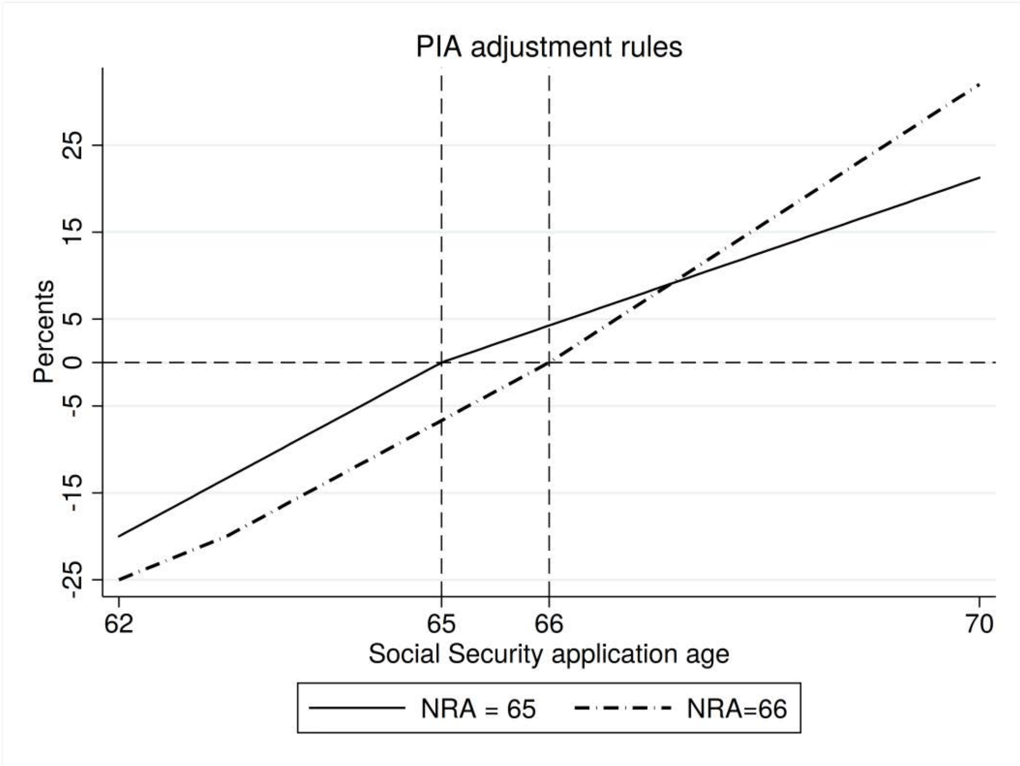

Figure 2.4: Early Retirement Penalty/Delayed Retirement Credit Calculation Procedure.

Earnings Test

Earnings test applies to individuals who receive social security benefits and income from labor at the same time. For those who didn't yet attain NRA there is an annual exempt amount (e.g. $12000 in 2005), and for each $2 of annual earnings on top of this amount $1 of annual benefits are withheld. That is, for sufficiently high-earning individuals all of their annual benefits could be withheld by the earnings test. However, part of the benefits withheld by the earnings test is not lost for good, but instead individual annual benefits are permanently adjusted upwards to account for the time during which a part of benefits was withheld. The amound witheld by the earnings test at age-t of an individual of cohort $\mathbf{c}$ in the model is calculated by

$$b_t^{ET}(\mathbf{c}, w_t l_t) = \begin{cases} \min\{b_{Bt}^{\alpha}, \frac{w_t l_t - 12000}{2}\}, & \text{if } (t < NRA \cap \mathbf{c} = 1) \cup (t < 70 \cap \mathbf{c} = 0), \\ 0, & \text{otherwise}, \end{cases} \tag{2.9}$$

where $12000 correspond to exempt amount in the year 2005. If part of the benefits $b_t^{ET} > 0$ was withheld, the amount of baseline pre-earnings test benefits is permanently adjusted by

$$b_{Bt+1}^{\alpha} = b_{Bt}^{\alpha} + b_t^{ET}(\mathbf{c}, w_t l_t)(1 + \mathbb{P}_t), \tag{2.10}$$

where

$$\mathbb{P}_t = \begin{cases} 0.05, & \text{if } t \geq 62,\ (NRA - t) > 3, \\ 0.067, & \text{if } t \geq 62,\ (NRA - t) \leq 3, \\ 0.0425, & \text{if } t \geq 65,\ t < 70,\ \mathbf{c} = 0, \\ 0.08, & \text{if } t \geq NRA, t < 70,\ \mathbf{c} = 1. \end{cases} \tag{2.11}$$

The dependence on cohort in 2.9 requires clarification. Prior to 2000, individuals were earnings tested until age 70. In 2000, however, earnings test system underwent a significant reform, which eliminated earnings test at NRA, meaning that Great Depression KIds were subject to earnings test until age 70, whereas Baby Boomers have to take this wedge into account only between the early and normal retirement age. Therefore, this policy difference poses another distinction between the two cohorts. French (2005) and Yavuzoglu (2016) show that this reform is a significant determinant of labor supply of seniors aged 62 and older.

To provide a bit of intuition to 2.10-2.11, consider an individual who have claimed benefits at age 62, thus being entitled to $0.8PIA$ in social security transfers, but continues to work on his very well paid job. Imagine that his or her annual earnings at the age of 62 exeed the annual exempt amount by so much that all of the annual benefits he of she is entitled to are fully withheld, that is, $b^{ET} = b^{\alpha}_{Bt}$. Basically, this individual doesn't receive any social security this year (despite having applied for it). Next, let this individual quit the job at age 63 and retire in conventional sense. This situation will exactly coincide with the scenario where an individual have applied for the social security benefits at age 63, being entitled to the higher amount of $0.867PIA$ for the rest of life.

Finally, the post-earning test benefits b^{α}_t received by an individual of age t, claim age α and cohort $\mathbf{c}$ is just pre-earnings test benefits net the earnings tested amount:

$$b^{\alpha}_t = \begin{cases} b^{\alpha}_{Bt} - b^{ET}_t(\mathbf{c}, w_t l_t), \text{if } t \geq \alpha \geq 62, \\ 0, \ otherwise. \end{cases} \tag{2.12}$$

This is the formula I use to construct b^{α}_t in the model.

Social Security Benefits Taxation

This section provides the details about taxation of social security benefits. According to Internal Revenue Service rules, up to 85% of benefits can be subject to income taxation under certain conditions. In particular, this is the case if individual combined income (sometimes also called *provisional income*), y^{cmb}_t, defined as the sum of a gross individual income (or gross income of married couple), net of social security transfers, plus half of a social security benefit (net of the earnings test), $y^{cmb}_t = y_t + 0.5b^{\alpha}_t$, exceeds either of two statutory thresholds $\hat{y}_l$ and $\hat{y}_u$ that depend on marital status.[8] In particular,

$$(\hat{y}_l, \hat{y}_u) = \begin{cases} (\$25000, \$34000) \text{ for singles,} \\ (\$32000, \$44000) \text{ for married couples.} \end{cases}$$

If combined income does not exceed the first threshold, one pays no taxes on the social security benefits. In the case when combined income exceeds the first threshold, up to 50% of the social security benefits are subject to income tax. If combined income exceeds second threshold, up to 85% of benefits may be taxable. More precisely, the taxable portion of benefits,

[8]The detailed description and historical background of the corresponding legislation can be found, for example, at https://fas.org/sgp/crs/misc/RL32552.pdf

$\mu(y_t, b_t^\alpha)$, is given by

$$\mu(y_t, b_t^\alpha) = \begin{cases} 0, \text{ if } y_t^{cmb} < \hat{y}_l, \\ \max\{0.5b_t^\alpha, 0.5(y_t^{cmb} - \hat{y}_l)\}, \text{ if } y_t^{cmb} \in [\hat{y}_l, \hat{y}_u], \\ \max\{0.85b_t^\alpha, 0.5(y_t^{cmb} - \hat{y}_l) + 0.35(y_t^{cmb} - \hat{y}_u)\}, \text{ if } y_t^{cmb} > \hat{y}_u. \end{cases} \tag{2.13}$$

2.2.4 Dynamic Programming Problem

This section provides a recursive formulation of an individual problem. Agent's state is fully determined by cohort $\mathbf{c}$, gender $\mathbf{g}$, marital status, age t, current level of assets a_t, health status h_t, wage and medical expenditure shocks $\hat{\varepsilon}^w$ and $\hat{\varepsilon}^m$, and social security receipt status. In particular, an agent can be in different states with respect to social security status. First, nobody can apply for social security until the age of 62. During this period, one keeps track of the one's AIME and the corresponding potential benefits b_t. Second, starting from 62, a person faces a choice between drawing the benefits now or face the choice again next year, with higher benefit level. Third situation describes individual who claimed benefits at some age α, and currently receives a fixed annual amount b_t^α, properly adjusted for the earnings test.

Let $V_t^{\mathcal{A},0}(\mathbf{g}, \mathbf{c}; a, h, \hat{\varepsilon}^w, \hat{\varepsilon}^m, b^\alpha)$ denote the value function of an age-t single individual who has already applied for the social security (a superscript $\mathcal{A}$ stands for "an applicant"), with age-t social security benefits b_t^α. It is given by

$$\begin{aligned} &V_t^{\mathcal{A},0}(\mathbf{g}, \mathbf{c}; a, h, \hat{\varepsilon}^w, \hat{\varepsilon}^m, b^\alpha) = \max_{c,l,a'} u(c, l, h, t) + \\ &\beta\{s_t^{\mathbf{g},\mathbf{c}}(h)\mathbb{E}_t V_{t+1}^{\mathcal{A},0}(\mathbf{g}, \mathbf{c}; a', h', (\hat{\varepsilon}^w)', (\hat{\varepsilon}^m)', (b^\alpha)') + (1 - s_t^{\mathbf{g},\mathbf{c}}(h)B(a'))\}. \end{aligned} \tag{2.14}$$

Similarly, let $V_t^{\mathcal{A},1}(\mathbf{g}, \mathbf{c}; a, h, \hat{\varepsilon}^w, \hat{\varepsilon}^m, b^\alpha)$ be a value function for married individual. It is given by

$$\begin{aligned} &V_t^{\mathcal{A},1}(\mathbf{g}, \mathbf{c}; a, h, \hat{\varepsilon}^w, \hat{\varepsilon}^m, b^\alpha) = \max_{c,l,a'} u(c, l, h, t) + \beta\{s_t^{\mathbf{1-g},\mathbf{c}} \times \\ &[s_t^{\mathbf{g},\mathbf{c}}(h)\mathbb{E}_t V_{t+1}^{\mathcal{A},1}(\mathbf{g}, \mathbf{c}; a', h', (\hat{\varepsilon}^w)', (\hat{\varepsilon}^m)', (b^\alpha)') + (1 - s_t^{\mathbf{g},\mathbf{c}}(h)B(a'))] + \\ &(1 - s_t^{\mathbf{1-g},\mathbf{c}})V_{t+1}^{\mathcal{A},0}(\mathbf{g}, \mathbf{c}; a', h', (\hat{\varepsilon}^w)', (\hat{\varepsilon}^m)', (b^\alpha))'\}. \end{aligned} \tag{2.15}$$

Both 2.14 and 2.15 are subject to

$$\begin{aligned} &c + a' + m = a + T(y + \mu(y, b^\alpha)) + (1 - \mu(y, b^\alpha)) + b^{ma}, \\ &y = r^{\mathbf{c}}a + wl + \mathbb{I}(married)y^p, \\ &w = \bar{w}^{\mathbf{g},\mathbf{c}}(h)exp(\hat{\varepsilon}^w), \\ &m = \bar{m}^{\mathbf{g},\mathbf{c}}(h)exp(\hat{\varepsilon}^w), \\ &b^{ma} = \max\{0, \underline{c} + m - (a + T(y + \mu(y, b^\alpha)) + (1 - \mu(y, b^\alpha)))\}, \\ &a' \geq 0 \end{aligned}$$

and subject to equations 2.9 - 2.13. The functions $V_t^{\mathcal{A},j}$ describe the phase of an individual life cycle following the social security application and until death. During this phase, the social security base for the individual is only updated due to the earnings test. The expectation is taken with respect to the distributions of wage, medical expenditures, and health shocks. Using the available information, an agent chooses consumption, leisure and annual hours optimally. Notice that the choice of hours in turn determines individual earnings at t, which define the effect of the earnings test and a portion of the benefits subject to income taxation.

Consider now a value function of an individual aged $t \geq 62$ who is not currently applying for social security, and has a baseline potential benefit b ("a non-applicant"). First, for single individual, this problem reads as

$$V_t^{\mathcal{NA},0}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b) = \max_{c,l,a'} u(c,l,h,t) + \beta\{s_t^{\mathbf{g},\mathbf{c}}(h)\mathbb{E}_t V_{t+1}^{\mathcal{E},0}(\mathbf{g},\mathbf{c};a',h',(\hat{\varepsilon}^w)',(\hat{\varepsilon}^m)',b') + (1-s_t^{\mathbf{g},\mathbf{c}}(h)B(a'))\}, \tag{2.16}$$

and for a married individual it becomes:

$$V_t^{\mathcal{NA},1}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b) = \max_{c,l,a'} u(c,l,h,t) + \beta\{s_t^{\mathbf{1-g},\mathbf{c}}\times [s_t^{\mathbf{g},\mathbf{c}}(h)\mathbb{E}_t V_{t+1}^{\mathcal{E},1}(\mathbf{g},\mathbf{c};a',h',(\hat{\varepsilon}^w)',(\hat{\varepsilon}^m)',b') + (1-s_t^{\mathbf{g},\mathbf{c}}(h)B(a'))] + (1-s_t^{\mathbf{1-g},\mathbf{c}})V_{t+1}^{\mathcal{E},0}(\mathbf{g},\mathbf{c};a',h',(\hat{\varepsilon}^w)',(\hat{\varepsilon}^m)',b')\}. \tag{2.17}$$

$$\begin{aligned}
&c + a' + m = a + T(y) + b^{ma},\\
&y = r^{\mathbf{c}}a + wl + \mathbb{I}(married)y^p,\\
&w = \bar{w}^{\mathbf{g},\mathbf{c}}(h)exp(\hat{\varepsilon}^w),\\
&m = \bar{m}^{\mathbf{g},\mathbf{c}}(h)exp(\hat{\varepsilon}^w),\\
&b^{ma} = \max\{0, \underline{c} + m - (a + T(r^{\mathbf{c}}a + wl))\},\\
&a' \geq 0
\end{aligned}$$

subject to 2.6 - 2.8. Notice that continuation value of the equation 2.16 includes value functions $V_t^{\mathcal{E},j}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b)$ for marital statuses $j \in \{0,1\}$, which denote the value of an individual who is eligible for social security benefits and decides whether to draw them now or wait and face the decision next year. This function simply chooses the maximum current expected value of the two functions defined before, namely, between the value of not applying as opposed to applying to the social security exactly at current age t:

$$V_t^{\mathcal{E},j}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b) = \max\{V_t^{\mathcal{NA},j}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b); V_t^{\mathcal{A},j}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b^t)\}, \tag{2.18}$$

which provides an adjustment factor for delayed or early retirement, when $t \neq NRA$. Finally, the value function that closes the system is the problem of an agent who is not eligible for social security benefits yet, that is, for the agent with $t < 62$. Notice that for a single individual of the age up to 60 this function takes the form

$$V_t^{\mathcal{NE},0}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b) = \max_{c,l,a'} u(c,l,h,t) + \beta\{s_t^{\mathbf{g},\mathbf{c}}(h)\mathbb{E}_t V_{t+1}^{\mathcal{NE},0}(\mathbf{g},\mathbf{c};a',h',(\hat{\varepsilon}^w)',(\hat{\varepsilon}^m)',b') + (1-s_t^{\mathbf{g},\mathbf{c}}(h)B(a'))\}, \tag{2.19}$$

whereas for 61-year-old person, than is, for an individual who becomes eligible at $t+1=62$, it reads

$$V_t^{\mathcal{NE},0}(\mathbf{g},\mathbf{c};a,h,\hat{\varepsilon}^w,\hat{\varepsilon}^m,b) = \max_{c,l,a'} u(c,l,h,t)+ \tag{2.20}$$
$$\beta\{s_t^{\mathbf{g},\mathbf{c}}(h)\mathbb{E}_t V_{t+1}^{\mathcal{E},0}(\mathbf{g},\mathbf{c};a',h',(\hat{\varepsilon}^w)',(\hat{\varepsilon}^m)',b') + (1-s_t^{\mathbf{g},\mathbf{c}}(h)B(a'))\}$$

subject to 2.1-2.2, 2.6-2.8 and subject to constraints associated with problem 2.17. The corresponding value functions of married individuals are similar to 2.15 and 2.17, that is, they take into account the possibility of spouse's death.

2.3 The Data and Exogenous Profiles

The Health and Retirement Study is an individual level panel data collected every two years from 1992 to 2013. It follows individuals from age 50 until their death. The data consists of several waves of respondents starting from age 50, as well as their spouses regardless of age. The dataset contains very detailed individual information on demographic characteristics, employment history, income, medical expenditures, health status and education attainment among others. I use version O of the RAND files of the HRS, which covers 13 waves of respondents from 1992 to 2013.[9]

I restrict the sample to individuals born between 1915 and 1934 (post-WWI, Great Depression) and between 1945 and 1964 (post-WWII, Baby Boomers), aged 50 to 90. The choice of these two cohorts is not arbitrary. First, cohorts have to be large enough to be able to capture intended heterogeneity with enough data. Second, they have to be separated in time enough for the between-cohort differences to be significant. In the sample, I observe the Great Depression Kids between ages 58 and 90, while the youngest Baby Boomers are 50 and the oldest are 68 (however, there are only eight 68-years-old individuals left in the sample).

I further restrict the sample to those individuals for whom I observe either self-reported age of social security application or the age at which social security benefit are first received. I use total non-housing financial wealth of a houshold from the HRS as a measure of assets.

The RAND version of the HRS imputes a consistent measure of out-of-pocket medical expenditures across all waves of respondents.[10] Out-of-pocket medical spendings combine expenditures on hospital and nursing home stays, doctor visits, home health care, prescription drugs, outpatient and special care that were not covered by medical insurance, and insurance premia paid. In order to make medical expenditures and wages comparable across years, I deflate nominal dollars reported in the data by corresponding Consumer Price Indices, and express all the nominal values in terms of 2005 US dollars.[11] The sample summaries for some important variables are reported in the Table 2.1. Due to the data limitations the two cohorts

[9] http://hrsonline.isr.umich.edu/modules/meta/rand/index.html

[10] The details of imputation procedure can be found at http://hrsonline.isr.umich.edu/modules/meta/rand/randhrso/randhrsO.pdf

[11] The source of Consumer Price Indices is Bureau of Labor Statistics website: http://data.bls.gov/cgibin/surveymost?cu

are only comparable directly for individuals aged between 57 and 67. Hence, the statistics in the table are presented for both the full sample and the age-restricted subsample.

As seen from the table, Baby Boomers are more educated and somewhat healthier. More of them are active in the labor force, and on average they work longer hours. The increase in the labor force participation rate of Baby Boomers are especially stark for women and unhealthy individuals.

Out-of-pocket medical spendings of the two cohorts is the most interesting feature presented in the table. First, sample means of medical expenditures of Baby Boomers and Great Depression Kids aged between 57 and 67 are almost identical. Second, the distributions of the medical expenditures over the percentile scale are very different. That is, medical expenditures of Baby Boomers in lower, middle and higher quartiles are significantly higher than the expenditures of Great Depression Kids in the same quartiles. In the top expenditure percentiles, from p95 and higher, the situation is remarkably different. In particular, there is a small number of Great Depression Kids whose medical bills are extremely high. These few individuals spend significantly more on healthcare than Baby Boomers in their top expenditure percentiles. In fact, they spend so much that the means of medical spendings of the two cohort are similar despite the significantly higher expenditures of the absolute majority of Baby Boomers.

One of the reasons behind this phenomenon might be that there are more Great Depression Kids in this age group with medical conditions that require expensive treatment. Indeed, an individual of this cohort, as compared to Baby Boomer, has higher chances to fall victim to age-related health problems relatively early. Figure 2.5 plots the shares of healthy individuals aged between 59 and 70 in each cohort from the HRS sample. Both profiles demonstrate that the share of unhealthy individuals increase with age. Importantly, Great Depression Kids start to experience rapid age-related health decline as early as at the age of 62, whereas Baby Boomers likely to have a few more years of healthy life. In other words, Baby Boomers have less health problems than Great Depression Kids in the age interval between 57 and 67. Furthermore, if amount of out-of-pocket medical expenditures reflects the seriousness of health condition, then Baby Boomers are unlikely to develop such a condition before reaching the age of 67.

The following sections describe the estimation of exogenous life-cycle profiles that are used as the model inputs.

2.3.1 Health Evolution

Health status h_t of an age-t individual is one of the state variables of the model, as well as one of the sources of idiosyncratic uncertainty. As a measure of health, I use self-reported health status from the HRS, which is coded into 5 categories: excellent, very good, good, fair and poor. I create a binary variable that takes value of 0 ("healthy") if self-reported health is excellent, very good or good, and 1 ("unhealthy") othervise. The evolution of individual health status is fully described by the set of 2×2 transition matrices. Matrices are uniquely defined for each possible age from 59 to 99. The dependence of transition probabilities on age captures age-related health decline. Next-period health status of an individual depends on current health, age, cohort and gender, but doesn't affected by his or her medical expenditures.

Table 2.1: Sample summary statistics, data from the RAND HRS, 1992-2013.

Variable	Great Depression Kids	Baby Boomers
	Full sample	
Male	0.429	0.488
College degree	0.175	0.336
Minimum age in sample	57	50
Maximum age in sample	97	67
Mean sample age	75.18	56.83
Mean Normal Retirement Age	65	66.1
N of observations	73062	33260
N of individuals	11599	9908
	Individuals 57 to 67	
Male	0.481	0.484
College degree	0.195	0.343
Share of healthy individuals	0.801	0.830
Percentiles of annual OOP medical expenditures:		
p25	266.99	389.51
p50	934.71	1208.94
p75	2483.01	2884.23
p99	25133.02	21697.99
In labor force	0.508	0.728
In labor force, unhealthy	0.334	0.572
In labor force, male	0.597	0.762
In labor force, female	0.416	0.648
Mean (SD) annual hours worked	895.34 (1094.30)	1438.55 (1121.23)
Annual hours per worker, male	1965.85	2138.461
Annual hours per worker, female	1626.13	1838.01
N of observations	14740	14674
N of individuals	4320	5742

The HRS is biannual survey, so for any given individual I observe health status every two years. Therefore, I first estimate biannual rates of transition between healthy and unhealthy states using a logistic regression of the form

$$Pr(h_{t+2} = 1 \mid t, h_t, \mathbf{g}, \mathbf{c}) = \Lambda(\beta_h h_t + \beta_c D_c + \beta_g D_g + \beta_1 t + \beta_2 t^2),$$

where $\Lambda(\cdot)$ is the logistic function defined at some x by

$$\Lambda(x) = \frac{e^x}{1 + e^x}.$$

I annualize health transition rates by finding two stochastic matrices such that their product is equal to biannual transition matrix. Figure 2.6 and Table 2.2 provide the results of a logistic estimation. Despite the fact that the estimated health transition profiles of the two cohorts are very close, they do not coincide. The coefficient at cohort dummy is positive and highly significant, meaning that Baby Boomers are healthier.

Figure 2.5: Change in the share of healthy individuals for the two cohorts. Data sample from the RAND version of the HRS, 1992-2013.

2.3.2 Survival Rates

In the data, I observe whether a person has died since the previous interview. Individual probability of survival between age t and $t+1$ depends on t, health status at t, cohort and gender. I estimate biannual survival rates using a logistic model:

$$(s_t^{\mathbf{g},\mathbf{c}}(h_t))^2 = Pr(alive_{t+2} = 1 \mid t, h_t, \mathbf{g}, \mathbf{c}) = 1 \mid t, h_t, \mathbf{g}, \mathbf{c}) = \Lambda(\beta_h h_t + \beta_c D_c + \beta_g D_g + \beta_1 t + \beta_2 t^2).$$

To construct annual survival rates s_t from biannual s_t^2 , I take a square root of biannual rates. Figure 2.7 and Table 2.3 summarize the results.

In order to construct spousal survival rates, I run the same regressions without health dummies in order to obtain survival rates since I do not model the health status of a spouse. Furthermore, I assume that spouse is always of the same age as an agent. The survival probability of a spouse is then equivalent to the one of an individual of the same age and of the opposite gender to an agent, averaged over health statuses.

2.3.3 OOP Medical Expenditures

Out-of-pocket (OOP) medical expenditures are defined as the expenses for medical care that aren't reimbursed by insurance. They include deductibles, coinsurance, copayments, insurance premia paid, expenditures on hospital and nursing home stays, doctor visits, home health care, dental care, prescription drugs, outpatient surgery and special facility services. I use an imputed

Table 2.2: Results of the logistic estimation of the health transition rates.

	Estimated coefficients
age	0.00600***
	(3.44)
age^2	-0.000258***
	(-11.46)
Male dummy, D_m	0.000698
	(0.03)
Cohort dummy, D_c	0.107**
	(3.28)
Current health status dummy, D_h	2.654***
	(127.94)
Observations	72222

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

measure of annual individual OOP medical expenditures provided by the RAND version of the HRS. This measure is consistent across all interview years and all waves of respondents.

I model OOP medical expenditures as an exogenous stochastic process, conditional on cohort, gender, and health status. Namely, the model is

$$\log m_t^{\mathbf{g},\mathbf{c}} = \log \bar{m}_t^{\mathbf{g},\mathbf{c}} + \varepsilon_t^m.$$

The process is the combination of a profile $\log \bar{m}_t^{\mathbf{g},\mathbf{c}}(h_t)$, and a stochastic process ε_t^m. I obtain $\log \bar{m}_t$ from the data running the following regression:

$$\log m_t^{\mathbf{g},\mathbf{c}} = \alpha_m + \beta_m t + \beta_{\mathbf{c}}^m D_{\mathbf{c}}^m + \beta_{\mathbf{g}}^m D_{\mathbf{g}}^m + \beta_h^m D_h^m + \varepsilon_t^m.$$

The residuals are then given by

$$\hat{\varepsilon}_t^m = \log m_t^{\mathbf{g},\mathbf{c}} - \log \bar{m}_t^{\mathbf{g},\mathbf{c}},$$

where the life-cycle profile of log OOP medical expenditures $\log \bar{m}_t^{\mathbf{g},\mathbf{c}}$ is defined as

$$\log \bar{m}_t^{\mathbf{g},\mathbf{c}} = \hat{\alpha}_m + \hat{\beta}_m t + \hat{\beta}_{\mathbf{c}}^m D_{\mathbf{c}}^m + \hat{\beta}_{\mathbf{g}}^m D_{\mathbf{g}}^m + \hat{\beta}_h^m D_h^m.$$

Figure 2.8 demonstrates the profiles for the two cohorts.

I assume that $\varepsilon_t^m \sim \mathcal{N}(\mu_t, \sigma_t)$, and use sample average $\hat{\mu}_t$ and sample standard deviation $\hat{\sigma}_t$ of residuals $\hat{\varepsilon}_t^m$ as the estimators of μ_t and σ_t. Furthermore, I approximate ε_t by a discrete Markov process. Namely, for a given age t I build a stochastic transition matrix between the five

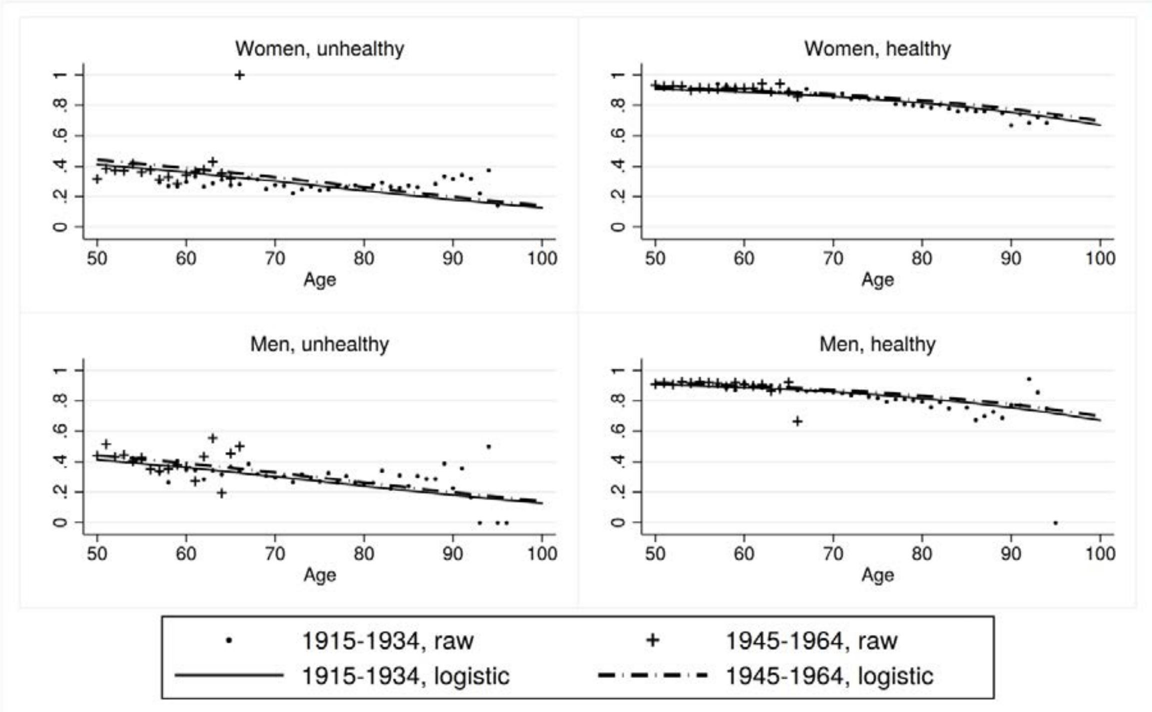

Figure 2.6: Probability of being healthy next year, conditional on gender and current health status, by age. Each panel shows the "raw" number of transitions to "healthy" status, and logistic fit. Data sample from the RAND version of the HRS, 1992-2013.

bins. One of these bins represent extreme medical expenditures (I count medical expenditure as extreme if the distance between the residual and the mean profile exceeds two sample standard deviations), while the rest of the residuals are sorted into four quartiles. I estimate the transition rates across bins using the following multinomial logistic model:

$$Pr(bin_{t+1} = k|t, bin_t = j) = \Lambda(\sum_i \beta_{bin_i} D_{bin_t=bin_i} + \beta_t t),$$

where $i, j, k \in \{1 \dots 5\}$ are identifiers of the five bins, $D_{bin_t=bin_i}$ is an indicator variable that equals to 1 if $\hat{\varepsilon}^m_t$ belongs to bin i at period t, and equals to zero otherwise.

As an example, consider an individual with period-t OOP medical expenditures being such that $\hat{\varepsilon}^m_t$ is contained in some bin j, one of the five. Based on this information and the estimates of the regression coefficients β_{bin_i}, one can calculate the predicted probability of finding $\hat{\varepsilon}^m_{t+1}$ in any particular bin at $t+1$. Five possible initial bins and five possible resulting bins provide 25 different combinations. The probabilities of each combination result in a 5×5 stochastic transition matrix. One example of such a matrix is presented in Table 2.4. The matrices are defined at the ages from 50 to 99. Notably, the resulting autoregressive process defined by these transition matrices is highly persistent.

The transition matrices allow to simulate the evolution of OOP medical expenditures up to the bin it belongs to. On top of this, I add within-bin heterogeneity. Recall that each of the five bins is defined by a continuous interval and have two well-defined borders. A residual

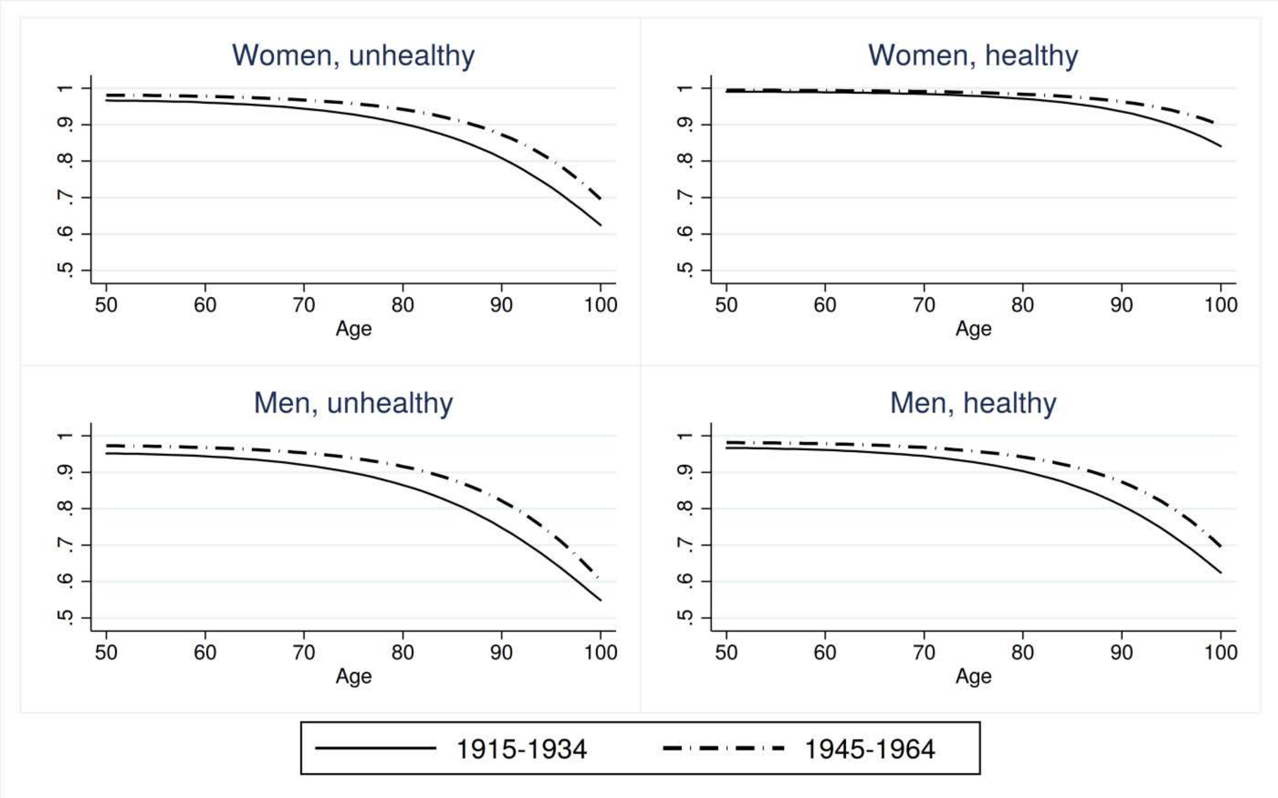

Figure 2.7: Survival probabilities, conditional on gender and health status, logistic fit (data from the HRS sample, 1992 to 2013).

belongs to the certain bin if its numerical value lies inside the interval limited by that bin's borders. Every bin thus covers a range of values. During simulations I draw a random value from the normal distribution truncated by that bins' borders. I treat this random value as an artificial "residual" to construct OOP medical expenditures of an individual at some age t. In particular, I sum up this random draw with the value of predicted profile at t to obtain a counterfactual OOP medical expenditure.

2.3.4 Wages

Wages are constructed by dividing annual labor earnings by annual hours worked. Annual earnings are directly available in the HRS, whereas annual hours worked by individual are computed as a product of weekly hours worked and number of weeks worked during a reference year. Wage profiles and wage evolution is calculated very similarly to the calculation of medical expenditures. I have much fewer observations here, since I only observe wages for those who work, as compared to OOP medical expenditures, which I observe for almost everyone in the sample. The assumed wage process is:

$$\log w_t^{\mathbf{g},\mathbf{c}} = \alpha_w + \beta_w t + \beta_{2,w} t^2 + \beta_{\mathbf{c}}^w D_{\mathbf{c}}^w + \beta_{\mathbf{g}}^w D_{\mathbf{g}}^w + \beta_h^w D_h^w + \varepsilon_t^w.$$

Figure 2.9 demonstrates the wage profiles for the two cohorts.

Table 2.3: Results of logistic estimation of survival rates.

	Estimated coefficients
age	0.116***
	(54.76)
age^2	-0.00122***
	(-46.67)
Male dummy, D_m	-0.397***
	(-15.24)
Cohort dummy, D_c	1.420**
	(20.79)
Current health status dummy, D_h	1.341***
	(51.02)
Observations	105256

t statistics in parentheses

* $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$

Table 2.4: Example of OOP medical expenditures transition matrices.

	age from 59 to 60					age from 80 to 81				
Bins	1	2	3	4	5	1	2	3	4	5
1	.461	.264	.163	.105	.007	.434	.266	.170	.114	.016
2	.242	.343	.253	.153	.009	.222	.338	.260	.162	.019
3	.138	.257	.334	.259	.012	.124	.247	.336	.267	.026
4	.090	.158	.267	.462	.023	.080	.148	.262	.464	.046
5	.103	.132	.235	.475	.056	.087	.119	.223	.461	.046

In order to discretize the autoregressive process of wage transitions I sort residuals into 5 quintile bins. Then I calculate the conditional transition probabilities between bins in the same way as for the OOP medical expenditures. Again, the resulting autoregressive process is remarkably persistent.

Controlling for Wage Selection Bias

The well-known problem with the type of wage profiles I'm using in this paper is that they are subject to selection bias. In the data we only observe the wages of those who work, and we don't observe potential wages for non-workers. At older ages, this bias becomes extremely significant, given the small fraction of working individuals. Thus, it is crucial to correct for this bias to get any credible results from the model. I employ the procedure following French (2005).

The main idea is that we can observe wages of those who doesn't work in the model simulated data. The key assumption is that the difference in wages of those who work and those who don't

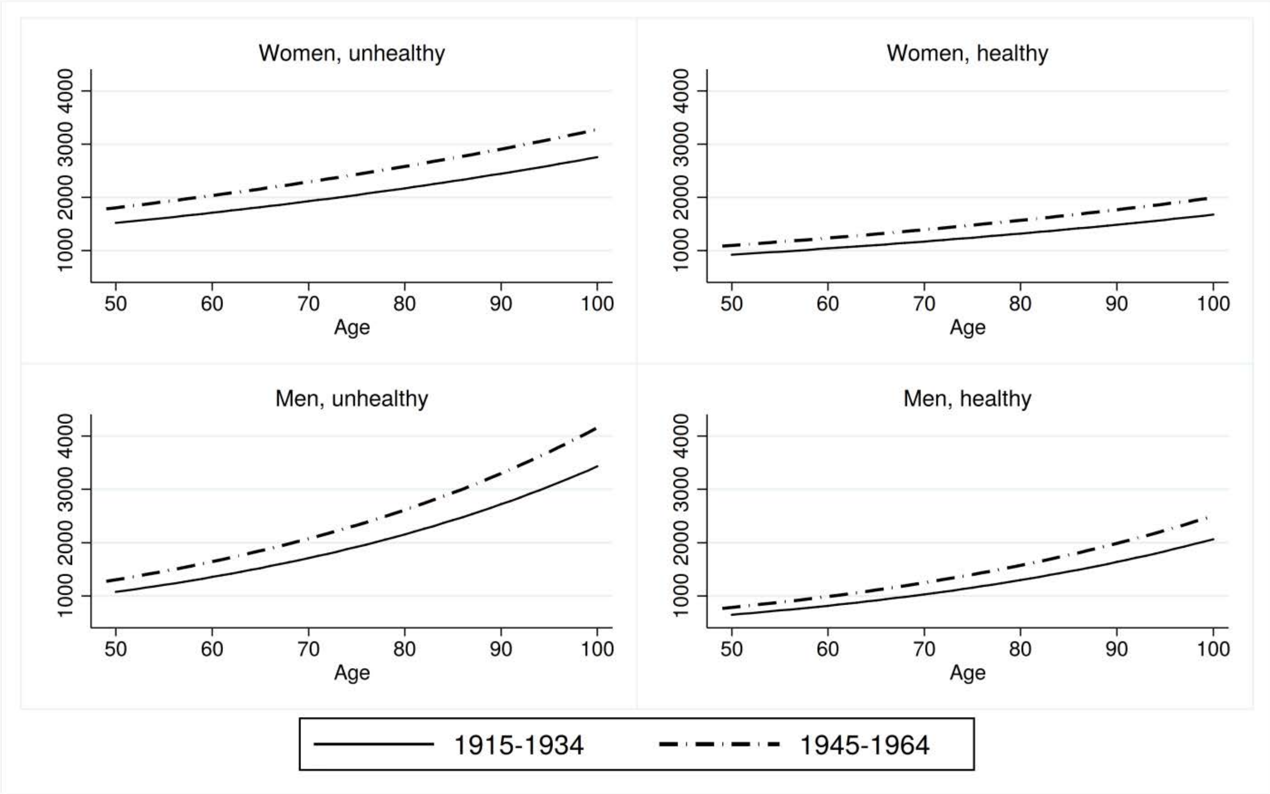

Figure 2.8: Life-cycle profiles of mean OOP medical expenditures, in $2005 (data from the HRS sample, 1992 to 2013).

is similar in the HRS and in the data simulated by calibrated model. I do initial calibration of the model using the biased wage profiles from the HRS. I then use the model to simulate a large number of individual life-cycle decisions. In the simulated data I can observe wage offers and corresponding labor supply decisions for every individual. I use this information to build two life-cycle wage profiles. The first one is constructed using only wages of those who work, so it is analogous to the profile from the HRS. The second profile is constructed using wages of both workers and non-workers. This one represents a potential "true" unbiased average wage profile, which we would like to obtain. The difference between the two profiles provides a sense of an upward bias of the wage profile from the HRS. For example, assume that at a certain age the average wage of simulated workers exceeds the average wage of all simulated individuals by 20%. Then it is likely that the wage in the HRS profile at this age was overestimated by 20%. Therefore, this difference is used to update the HRS profile, decreasing the average wages. This updated profile is then used on the next iteration to produce a new calibration of the model. The updated model is then again used to simulate the two profiles, calculate bias, update the candidate true wage profile again. The procedure is repeated up to the point where the update doesn't change the profile.

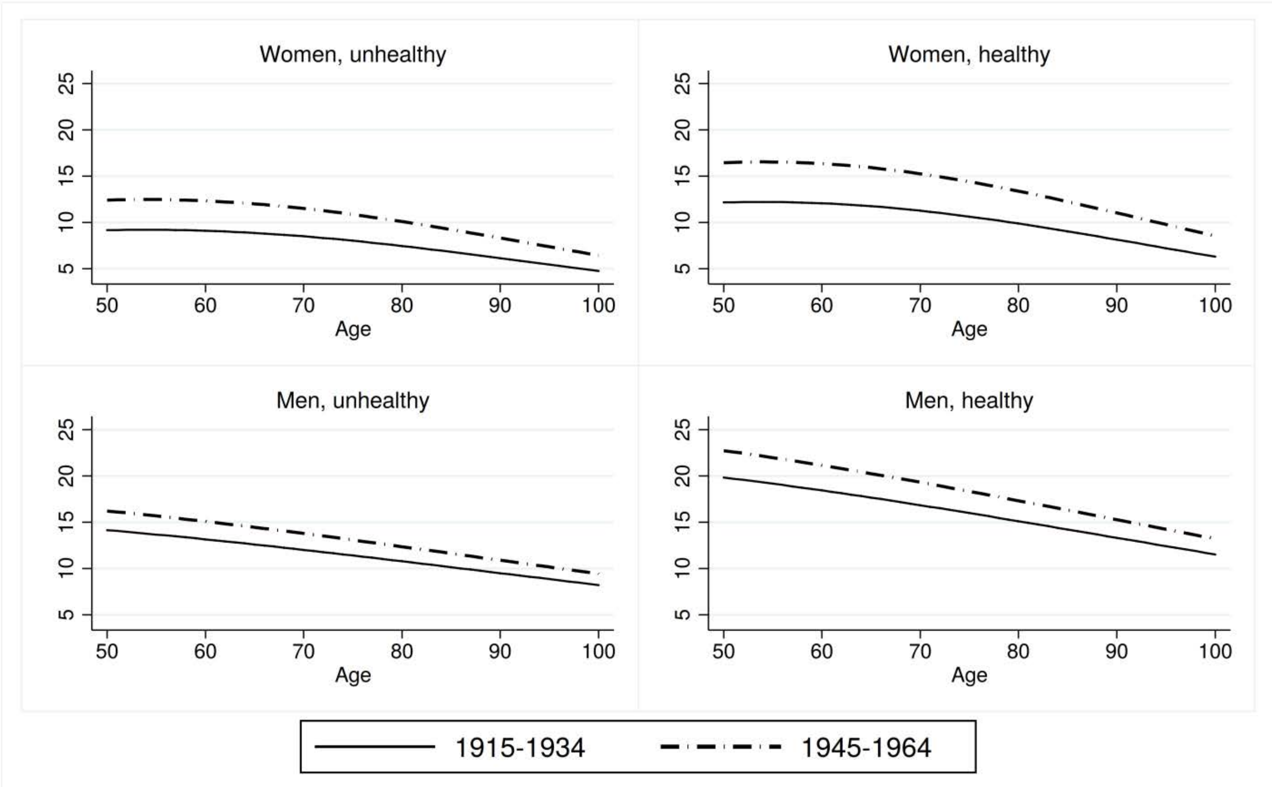

Figure 2.9: Life-cycle profiles of mean hourly wage, in \$2005 (data from the HRS sample, 1992 to 2013).

2.3.5 Spousal Income

I construct a life cycle profiles y_t^p of a married individual regressing log of annual income of the spouse on cohort dummy and age polynomial of the fifth order, conditional on gender:

$$\log y_t^p = \alpha_{yp} + \beta_{\mathbf{c}} D_{\mathbf{c}} + \sum_{i=1}^{5} \beta_i t^i + \varepsilon_t^y.$$

The resulting profiles are displayed on the Figure 2.10.

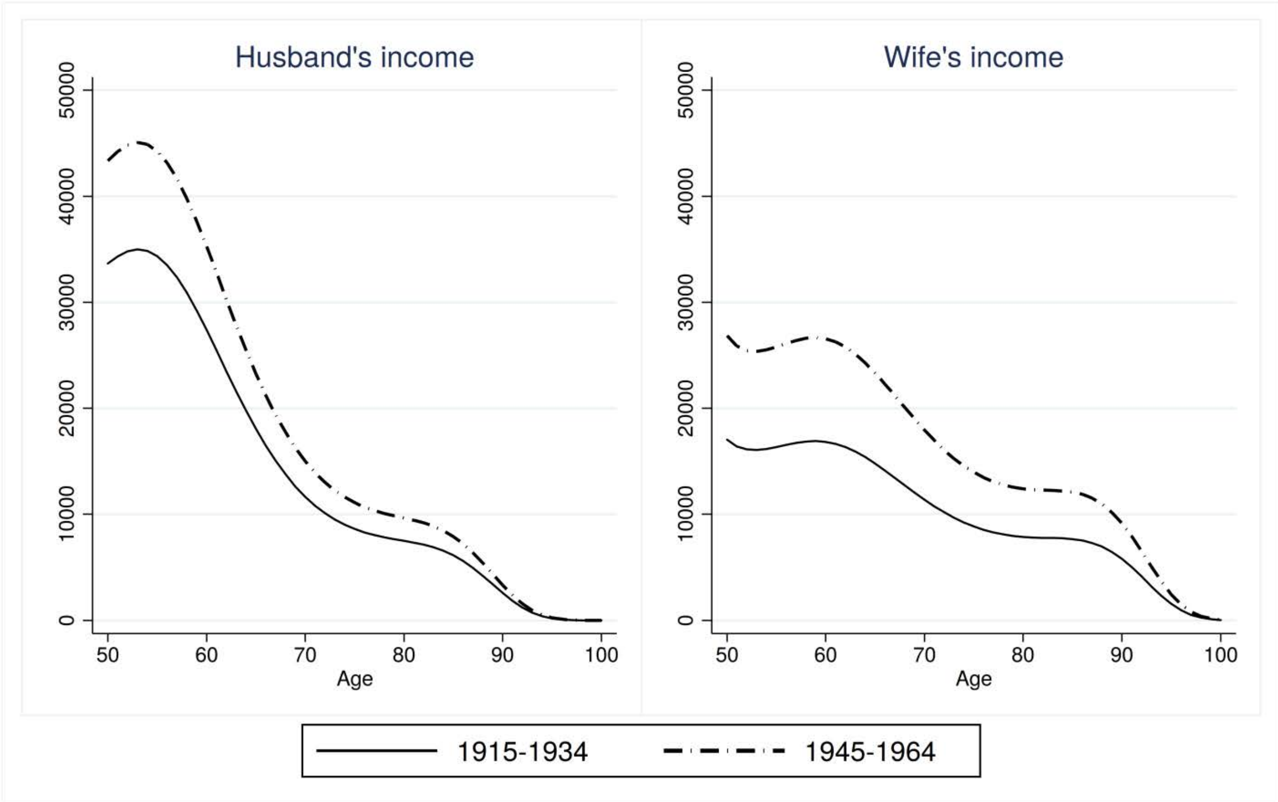

Figure 2.10: Life-cycle profiles of average spousal income of married individuals, in \$2005 (data from the HRS sample, 1992 to 2013).

2.4 Calibration and Estimation

First, I fix labor endowment at age 50 to $\bar{L} = 5000$ and time discounting factor to $\beta = 0.95$. The ultimate goal is to estimate the set of preference parameters $\Upsilon = \{\sigma, \gamma_c, \gamma_l, \delta, \kappa, \xi, \nu, \eta, d, \underline{c}\}$. Main assumption is that Υ is similar between the two cohorts, and all the observed differences in behavior come from the exogenous objects, such as survival rates and policies. In order to estimate Υ, I employ the following two-step approach. On the first step, I estimate health transition matrices, survival rates, wage and medical expenditure processes, and joint distributions of AIME, assets, wages, medical spendings and total yeats worked at age 59 directly from the data. Given these objects and some guess on Υ, I solve an individual dynamic programming problem backwards to obtain a corresponding optimal savings decisions, labor supply decisions, and timing of social security application. Given these set of decisions, I can simulate a large number of individuals, and calculate the difference between observed and simulated labor supply, labor participation, assets and social security timing profiles under a particular distance measure. I then continue to update Υ until the simulated and observed profiles are sufficiently close. I consider the resulting Υ as a parameters of the data-generating process for both cohorts.

After this is done, I can simulate a large number of individuals of the two cohorts with respect to the common Υ and the set of the exogenous processes specific to the each cohort. In an ideal case, this will allow me to see which part of the change in labor force participation decisions can be explained by considered exogenous factors.

Table 2.5: Example of OOP medical expenditures transition matrices.

Cohort	Gender	Positive assets	AIME	Wage	OOP	Years	Assets
BB	Female	No	9.82	2.46	7.09	3.11	
BB	Male	No	10.02	2.75	6.86	3.37	
BB	Female	Yes	10.09	2.75	7.04	3.25	10.45
BB	Male	Yes	10.55	3.03	6.69	3.45	10.45
GDK	Female	No	9.11	2.41	6.96	29.77	
GDK	Male	No	9.32	2.35	6.70	36.52	
GDK	Female	Yes	9.44	2.51	6.85	30.57	11.27
GDK	Male	Yes	9.92	2.92	6.87	40.84	11.44

2.4.1 Initial Distributions

Individuals enter the model at age 59. In order to start simulation, I need to know age-59 joint distribution of AIME, assets, wage, health and medical spendings, and total years worked prior to this age. I estimate this distribution directly from the data. Launching the model simulations, for each individual I draw initial values from this distribution. Out of these, in the sample I directly observe everything apart from AIME for everybody. However, I can recover the initial AIME for Baby Boomers using the Social Security payments for retired individuals together with predicted SS wealth at age 62, 65 or 70 for pre-retirees that is provided in the HRS and the earning histories down to age 59. To estimate the distributions, I assumme that AIME, assets, wage and OOP medical spendings and total years worked prior to 59 are jointly lognormal at age 59, conditional on cohort, gender and having zero assets. I then calculate sample mean vector and variance-covariance matrices. The estimates are presented in the Table 2.5.

Another important initial value to determine is the marital status of individual at the age of 59. For each cohort, I calculate a proportion of married and single individuals at the initial age, conditional on gender. I treat this proportion as a probability of being married, and at the start of simulation I use this probability to make a random draw that determines individual marital status.

2.4.2 Taxation Function

A total amount of taxes paid by individual consists of the two parts: income tax and payroll tax. A payroll tax is a fraction of labor income paid by every worker as a contribution to Social Security system. A payroll tax rate is flat and constitues $\tau_{ss} = 7.65\%$ on an individual labor income. The amount of labor income subject to payroll tax is capped by a contribution and benefit base of $90000.

In order to calculate income taxes, I use the parametic estimates of effective tax functions from Guner et al. (2014). Namely, I use estimates for year 1989 for the Great Depression Kids cohort (right after Raegan tax cuts) and for year 2000 for Baby Boomers cohort, for whom effective income taxes were changed in a way that low-income household paid less taxes that before, whereas richer houshold paid more. The average tax rate is $\tau(y)$, where y is income in \$2005, τ is effective income tax rate, and m, s and p are cohort-dependent coefficients. Thus, the disposable household income is calculated by

$$T(y_t) = (1 - \tau(y_t))y_t - \tau_{ss}w_t l_t, \tag{2.21}$$

where the precise functional form of effective taxation function is given by

$$\tau(y) = m[1 - (sy^p + 1)^{-\frac{1}{p}}]. \tag{2.22}$$

The values of the parameters of taxation function are presented in the Table 2.6.

Table 2.6: Parameters of the taxation function.

Parameter	m	s	p
Great Depression Kids	.258	.0355	.768
Baby Boomers	.264	.0136	.964

2.4.3 Moment Conditions

I estimate the set of structural parameters $\Upsilon = \{\sigma, \gamma_c, \gamma_l, \delta, \kappa, \xi, \nu, \eta, d, \underline{c}\}$ to fit the following set of moments:

- Mean labor force participation, for every age 59 to 75, to identify, to identify κ, ξ and ν.
- Mean labor force participation conditional on having zero assets, for every age 59 to 75, to identify $\underline{c}$.
- Mean labor supply (in hours), for every age 59 to 75, to identify γ_l, γ_c and σ.
- Mean labor supply (in hours), by health status, for every age 59 to 75, to identify δ.
- Mean assets, for every age 59 to 75, to identify η and d.

As a first step weighting matrix I use the inverse of variance of each moment at each given age. The set of calibrated model parameters is presented in Table 2.7.

A few remarks are in order. The cost of participation in the labor force is 302 hours per year at the age 59 and increases with age and health deterioration. This fixed cost of work generates a minimal amount of annual hours that one is willing to work, which depends on gender, social security benefits, health, assets and exogenous profiles. The rise of the fixed cost with age increases the number of individuals on who are almost indifferent between participation and non-participation. Being in bad health additionally increases fixed cost of participation, which further increases the number of the individuals at the participation margin. That is, at a younger age many individuals will choose work over leisure. With the increase of fixed cost, and taking into account relatively high leisure weight γ_l, more individuals at an older age are

Table 2.7: Set of model parameters, male Great Depression Kids. Standard errors are in parenthesis.

Parameter	Value	Description
		Fixed parameters
β	0.95	Time discounting, fixed
$\bar{L}$	5000	Annual labor endowment (hours), fixed
		Estimated parameters
κ	302 (11)	Participation cost (hours)
d	50704 (1229)	Strength of bequest motive
$\underline{c}$	2512 (194)	Consumption floor
δ	211 (24)	Additional participation cost, bad health
η	5.23 (0.46)	Bequest motive magnitude
σ	4.10 (0.25)	Relative risk aversion
γ_c	0.451 (0.029)	Consumption weight
γ_l	0.698 (0.043)	Leisure weight
ξ	1.96 (0.11)	Participation cost due age, magnitude
ν	1.74 (0.13)	Participation cost due age, power

close to indifference between work and leisure. This is reflected in the increase of response of a labor supply to a permanent increase in wage with age.

The estimate of the parameter of relative risk aversion σ lies within the range of estimates reported by French (2005), it is very close to the value of 4.099, reported by Yavuzoglu (2016), but higher than 1.1 reported by Rust and Phelan (1997). The strength of bequest motive of \$50704 can be interpreted as a value of bequest that gives an individual notable utility gains. That is, an agent will not get significant additional utility from leaving a bequest of, say, \$1. On the other hand, an agent gets a tangible utility boost from leaving a bequest over d. This value lies in the range of estimates in the literature, which span from \$11 990 (Yavuzoglu 2016) to \$440 000 (French and Jones 2011).

2.4.4 Model Fit

The calibrated model fits the selected moments reasonably well. Four panels of the Figure 2.11 demonstrate the model fit. The baseline model successfully replicates annual working hours, participation rate and saving decision of male Great Depression Kids. On top of that, the model is able, at least partially, replicate the surge of Social Security applications at the age of 62.

Next, I use the parameters calibrated to fit the moments of the older cohort and use them to simulate the corresponding moments of the younger cohort, using their exogenous profiles and policy regimes. Then I compare these simulated profiles to the actual behavior of Baby Boomers. Figure 2.12 demonstrates the outcomes of this exercise, and shows that the model does excellent job in accounting for the changes in the labr supply behavior of the two cohorts. Table 2.8 summarizes the results for the age interval from 62 to 67.

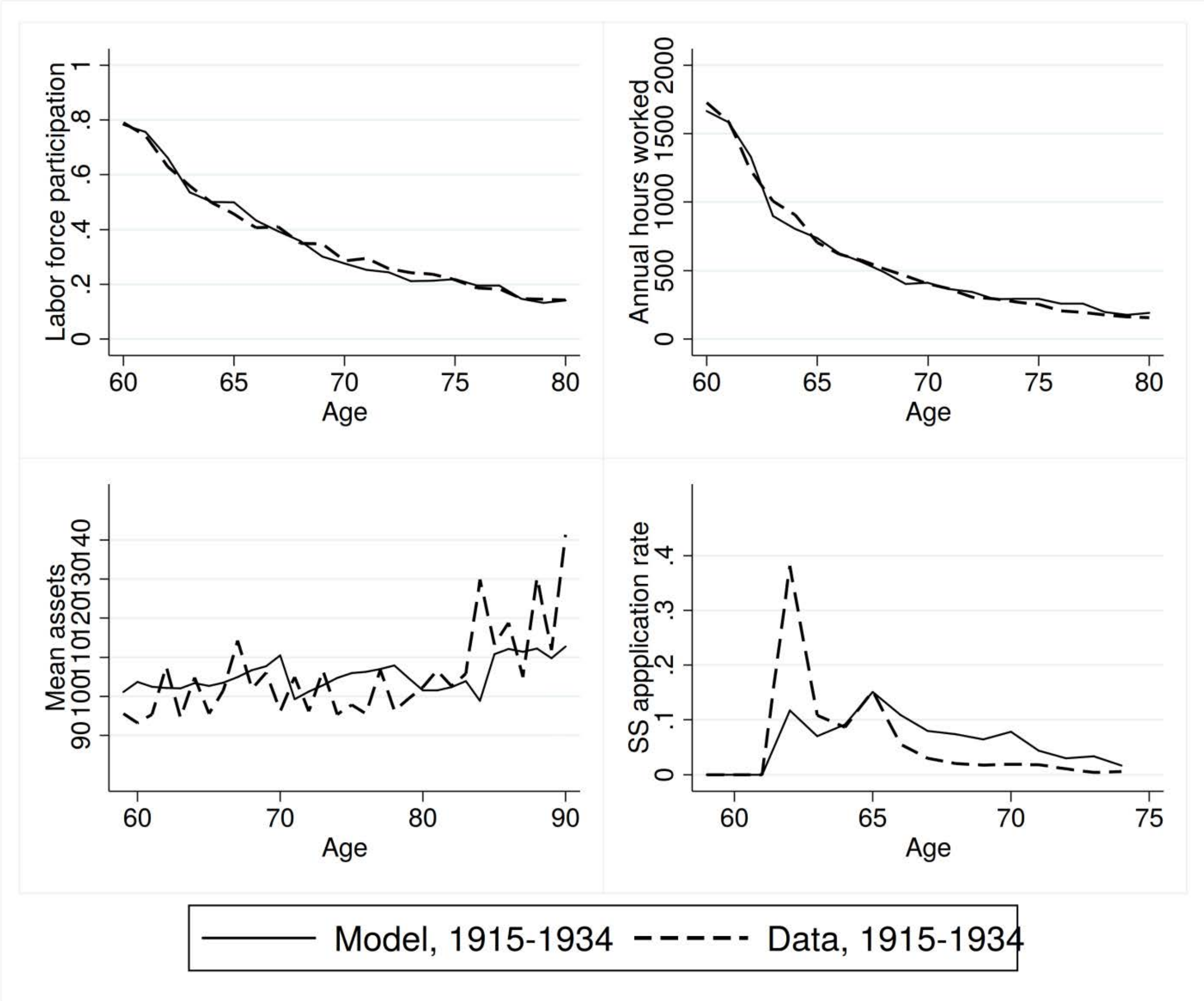

Figure 2.11: Key moments fit: model vs data. The data moments are taken for the male Great Depression Kids.

2.4.5 Implication for Social Security System

With baseline calibrated model at hand, the first exercise I conduct is aimed to calculate difference between total social security benefits received over lifetime and total social security payroll taxes paid after the age of 59.In particular, I calculate these quantities first for the baseline model of the Great Depression Kids cohort. Then I simulate Baby Boomers, using the model with the same structural parameters, but with all the exogenous profiles and policies of the Baby Boomer cohort. Table 2.9 summarize the findings. This exercise shows that the gap between total amount of benefits received and social security taxes paid is smaller as for Baby Boomers than it is for the Great Depression Kids. Notice that due to the increase in labor force supply Baby Boomers pay in total more taxes than their counterparts from the older cohort. Furthermore, they also receive less benefits over their lifetime. This result suggests that an increase in the labor force supply of Baby Boomers indeed alleviates the burden of aging poplation on social security system.

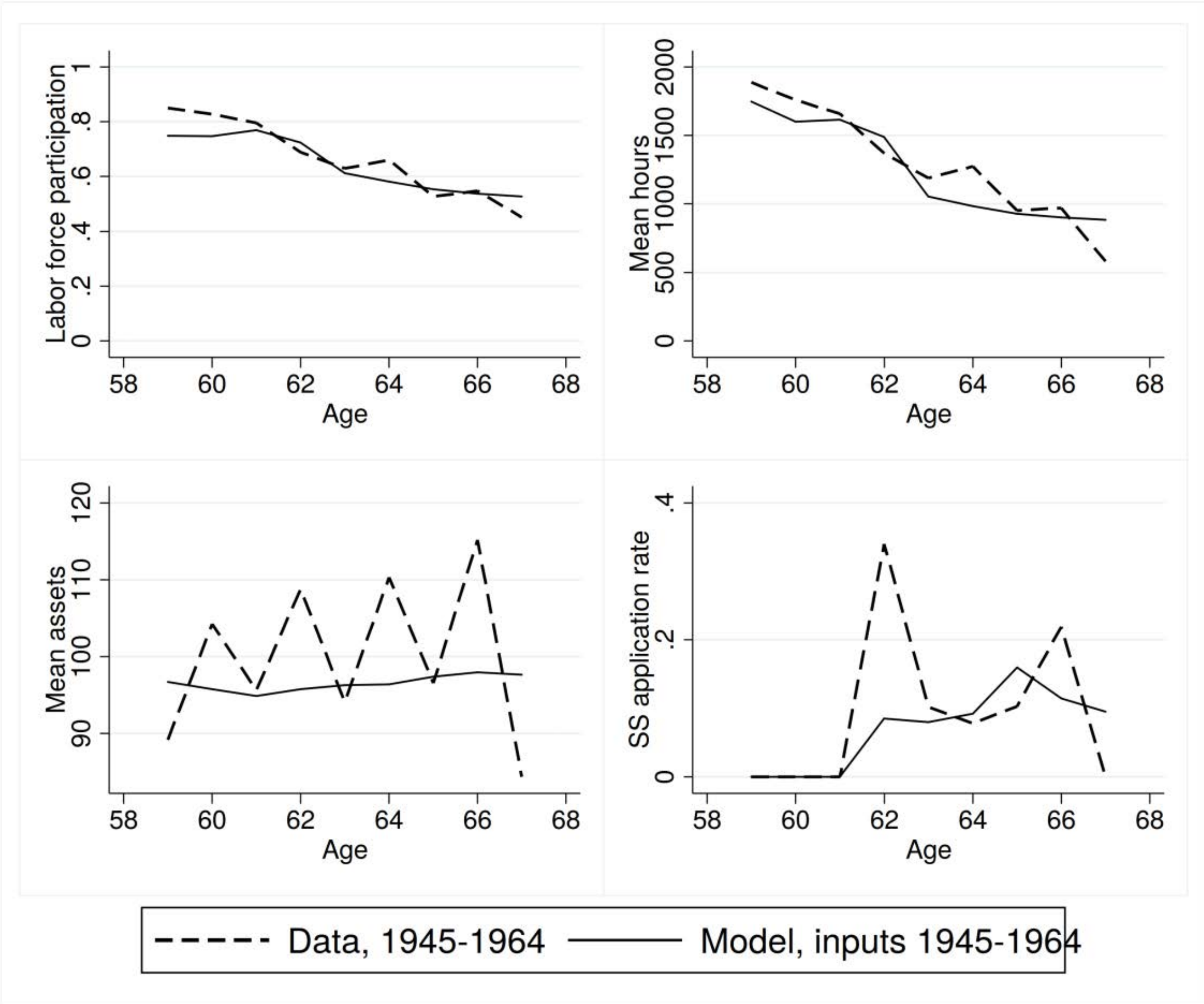

Figure 2.12: Model assessment: comparison of the data from the HRS sample and the model counterfactual data, male Baby Boomers. The parameters of the model are calibrated to fit the data moments of Great Depression Kids, whereas the policy structure and profiles correspond to those of Baby Boomers.

2.4.6 Implied Labor Supply Elasticities

Another exercise I conduct is intended to measure the wage elasticities of labor supply and effects of various policies on labor supply. The primary objects of interest are the percentage increase in average lifetime hours worked and total years worked after 59 in respons to a stimulus. First, I measure responses to a 5% increase in average wage. Second, I measure responses to increase in Normal Retirement Age by one year. Third, exclusively for the Great Depression Kids, I measure responses to the elimination of the earning test at the age of 65.

Table 2.10 summarizes the findings. First column represents the results of the exercise for the baseline model. Second column reports the corresponding exercises applied to results of a baseline calibrated model with all exogenous inputs of Baby Boomers. The results suggest that the labor supply of Baby Boomers is less elastic with respect to the same stimuli.

Table 2.8: Average moments for males, data vs baseline vs model assessment, age 62 to 67.

	Data, GD Kids	Baseline, GD Kids	Assessment, Baby Boomers	Data, Baby Boomers
Mean participation rate	.497	.501	.589	.584
Mean annual hours	855	842	1050	1056
Mean assets, $1000	103.07	102.53	96.89	101.53

Table 2.9: Excercise: total social security taxes paid and total benefits received. Baseline calibrated model vs model with all exogenous inputs and policies as for Baby Boomers.

	Baseline	Model, inputs of Baby Boomers
Δ (SS received - SS taxes paid), in $1000	130.46	119.38
SS received, $1000	160.31	156.39
SS taxes paid, $1000	30.83	37.21
Median SS received, $1000	135.68	130.18
Median SS taxes paid, $1000	29.65	33.58
Kurtosis SS received	3.76	3.03
Kurtosis SS taxes paid	3.11	2.67

Table 2.10: Table documents the simulated percentage increase of mean individual working hours past the age of 59 and percentage increase om mean individual year worked after 59 in response to various stimuli.

	Baseline	Model, inputs of Baby Boomers
% in lifetime hours vs 5% increase in wage	8.90	5.05
% in years worked after 59 vs 5% increase in wage	5.54	4.80
% in lifetime hours vs increase in NRA	5.21	5.63
% in years worked after 59 vs increase in NRA	5.11	1.49
% in lifetime hours vs ET elimination	2.98	–
% in years worked after 59 vs ET elimination	1.96	–

2.5 What Matters For the Increase in Labor Supply?

The primary goal of this paper is to quantify the relative importance of the forces that are, potentially, responsible for the increase in labor force activity of US elderly. The factors under investigation can be thought of either as policy-related or "natural" ones. The policy-related factors include the rise in Normal Retirement Age, elimination of the earnings test, changes in the taxation schedule, changes in the rate of social security early retirement penalty and delayed retirement credit. The policy-unrelated factors include the increase in life expectancy, increase in OOP medical expenditures, changes in wages, changes in health. The main objects of interest are differences in labor force participation and average hours worked between the two cohorts. The goal is to measure which shares of these differences can be attributed to each of the driving forces.

As an example, consider measuring the effect of increase of the Normal Retirement Age on labor force participation rate. In order to isolate the effect, I run the baseline model calibrated to the Great Depression Kids, but use the NRA of Baby Boomers as an input. The simulation results thus reflect the changes in the decisions of agents related to the increase in NRA, while all other factors remain unchanged. I use this approach to assess the effect of every factor of interest. Tables 2.11 and 2.12 represent the results of the exercise. It is clear that individuals react to both policies and non-policy factors.

Table 2.11: Decomposition exercise: policy-related factors, males, age 62-67.

	Baseline	NRA	Taxes	Earnings test	SS rules	Baby Boomers
Mean participation rate	0.521	0.529	0.534	0.552	0.525	0.584
Mean annual hours	842	859	862	906	851	1056
% explained, LFP		12.7	20.63	49.2	6.35	
% explained, hours		7.94	9.35	29.91	4.22	

Table 2.12: Decomposition exercise: other forces, males, age 62-67.

	Baseline	Longevity	OOP med	Wages	Health	Baby Boomers
Mean participation rate	0.521	0.536	0.522	0.556	0.521	0.584
Mean annual hours	842	848	869	836	795	1056
% explained, LFP		23.81	1.59	55.56	0	
% explained, hours		2.80	15.67	-7.22	-21.96	

First, the elimination of the Earnings Test at the NRA has a strong effect on both participaton and hours worked of individuals between 62 and 67, explaining in isolation almost half of the overall increase in participation and one-third of the incease in hours worked. Increase in the NRA and changes in effective tax rates potentially could significant portions of the increase in participation and hours. The change in Social Securty rules, which increased the delayed retirement benefit from 4.25% to 8% per year, has moderate effect, with some 6% more individuals delaying retirement past the NRA.

The factors not directly related to policies have more ambiguous effects. Increase in longevity and increase in OOP medical expenditures, taken in isolation, can only explain a modest part of the incease in labor supply of seniors. Changes in wage structure of Baby Boomers lead to significant increase in participation rate, but at the same time the decline in hours is observed. Hours also decline, if Great Depression Kids are given the health dynamics of the Baby Boomers.

A very important result, hovever, is that some factors interact to produce synergetic effects of high magnitude. One example is a very strong effect that increase in longevity and OOP medical expenditures has on labor force decisions of individuals, as can be seen from Table 2.13. That is, if Great Depression Kids had both life expectancy and OOP medical expenditures of the Baby Boomers, the model predicts very large increase in both participation and hours worked. A potential explanation might lie in the fact that longer life expectancy means higher chances to live up to a very old age, where extreme medical shock is very probable. Thus, seniors continue to work to hoard additional resources to insure against this shock.

Table 2.13: Mutually reinforcing factors, males, age 62 to 67.

	Baseline	Longevity & OOP med	Baby Boomers
Mean participation rate	0.521	0.601	0.584
Mean annual hours	842	970	1050
% explained, LFP	–	126.98	–
% explained, hours	–	59.81	–

2.6 Conclusions

A life-cycle model is built to assess the changes in labor market activity of two different cohorts of US seniors. A baseline calibration of the model to the post-WWI cohort fits the data reasonably well. A validation exercise suggests that the driving factors under investigation can successfully explain most of the differences in the labor force supply of seniors of the two cohorts. The decomposition exercise suggests that both policies and nonpolicy factors play a significant role in shaping changes in labor force behavior of seniors. Furthermore, their effects are of similar magnitudes.

From the policy side, the elimination of the Earnings Test at the NRA plays the most important role. This is due to the fact that individuals who potentially are subject to the Earnings Test may choose to work fewer hours or even stay out of the labor force, so their earnings stay below the Earnings Test threshold. Furthermore, this effect is directly linked to the changes in the rules of how the delayed retirement credit and the early retirement penalties are calculated. In particular, the delayed retirement credit was actuarially unfair for the Great Depression Kids, whereas it is actuarially fair for the Baby Boomers.

From non-policy side, there are interactions between the factors that lead to significant reinforcement of the effects. The most prominent example of such interaction is large effect of increase in longevity tied with increase in the OOP medical expenditures. Indeed, the risks of extremely high medical expenditures increase with age, and are almost certain at the very old age. Individuals nowadays expect to live longer, which means higher probability of living up to a very old age, thus risking to be a subject of an extreme medical condition or a nursing home. Individuals insure against this risk in two ways. First, by choosing to stay in labor market longer, hoarding more resources and simultaneously decreasing the expected retirement time, during which they have to rely on accumulated wealth and social security payments. Second, by avoiding running down of the accumulated assets in the first years after the retirement.

Quantitative exercise suggests that increase in the labor supply of the baby Boomers indeed alleviates a burden of population aging on the social security system. According to the results, the Baby Boomers pay more taxes and receive less social security after the age of 59 than their counterparts from the older cohort. That is, the difference between total social security payments received and total payroll taxes paid by individuals over 59 is about 8% lower for the Baby Boomers.

Chapter 3

Great Expectations: Reservation Wages and the Minimum Wage Reform

In January 2015, a statutory minimum wage of €8.50 per hour was introduced in Germany. This was clearly the most important labor market reform in Germany since the early 2000s. The new minimum wage was set at a high binding level and covered about 10 to 14 percent of all jobs in Germany. The Kaitz index, which defines the ratio of the minimum to median wage was 48%.[1] The reform was debated actively among economists and policy makers. Bruttel, Bauman, Dutsch (2018), Caliendo (2017a,b) provide a detailed discussion of the reform.

In this chapter, we use the German Socio-Economic Panel (SOEP) to study the how the introduction of a statutory minimum wage affected reservations wages in Germany. We find that the reform was associated with an increase in reservation wages of approximately 4 percent at the low end of the distribution. Furthermore, the shifts in reservation wages and observed wages due to the minimum wage reform are comparable in their magnitude. Finally, we show that that German citizens adjust their reservation wages more than immigrants. We also provide suggestive evidence that points to compensation mechanism in which immigrants trade wage growth against job security.

Neoclassical, monopsonistic and search theoretical models all predict negative labor demand reactions to the introduction of a binding minimum wage. When the wages of low-skilled workers increase, employers demand fewer units of labor from that group, and they hire fewer workers if the minimum wage is set higher than the competitive wage. The decrease in the arrival rate of jobs therefore contributes to an increase in unemployment if labor supply remains constant. Models of imperfect markets including search frictions (see Burdett and Mortensen (1998); Flinn (2006); Ahn et al. (2011)) and monopsonistic competition (e.g. Manning (2003)) on the other hand allow for positive labor demand effects. For example, in a traditional model of monopsony with an upward sloping labor supply function an increase in the wage caused by minimum wage will lead to both increase in employment and firm's output (Card and Krueger (1995)).

[1]This is similar to 49% in the UK and 46% in the Netherlands. Among the OECD countries, France has the highest Kaitz index (62%) and Spain has the lowest. See OECD (2016).

In contrast to the large literature on possible positive and negative labor demand effects of minimum wages, the literature on the contribution of labor supply effects to overall unemployment dynamics has been much more limited and has focused almost exclusively on explaining short-term increases in participation. Both competitive and imperfect labor market models predict increases in labor market entry for individuals whose reservation wages are below the new miniumum wage but were above the market wage before the reform. Reservation wages are defined by the threshold at which a potential worker is willing to accept a job offer. Because higher reservation wages lead to longer periods of job search, an increase in reservation wages can lead to more medium-run increases in unemployment and participation levels. However, in the medium and long-run, reservation wages are likely to be not static and may themselves adjust to the economic environment and changes in the minimum wage. If reservation wages react to the minimum wages, this will create a supply effect that will interact with lower demand and can lead to even higher (if reservation wages go up) or lower (if reservation wages go down) unemployment.

Empirically, the link between minimum wages and reservation wages has largely been neglected due to lack of information about individual acceptance thresholds. An important exception is Falk et al. (2006), who find persistent, positive and significant effects within the framework of a laboratory experiment. They demonstrate that minimum wages can set a new standard for what is considered fair pay and create entitlement effects even after the removal of a wage floor. A further possible channel relates to how minimum wages affect relative wage differences between workers, thus increasing wage expectations of workers previously earning slightly above the minimum wage. The latter can help to explain the observed phenomenon of spillover effects throughout the wage distribution following an increase in minimum wages. Spillover effects can emerge either because workers seek to maintain their relative position in the wage distribution or simply because they anticipate increases in prices due to a large-scale minimum wage increase.

Very few papers have investigated the minimum wage effects on reservation wages in a quasi-experimental setting. While the literature on the empirical effects of minimum wages on reservation wages remains scarce, previous literature regarding the wage curve has documented a negative relationship between wage levels and regional unemployment (Blanchflower and Oswald (1995); Blanchflower and Oswald (2005); Bell et al. (2002)) using individual-level microdata. Brown and Taylor (2015) expand this research focused solely on workers earning observed wages to incorporate expected and reservation wages among the unemployed and inactive. Using the British Household Panel Survey (BHPS), they investigate the reservation wage curve for the UK and find a negative relationship between unemployment spell duration and reservation wages. Blien et al. (2012) model the reservation wage curve for Germany. Recent research in this area has tried to explain the inverse relationship by causally identifying a negative impact of unemployment on reservation wages.

Detailed survey information from the Socio-Economic Panel (SOEP), in combination with a natural experiment from the introduction of a legal minimum wage in Germany, provide a unique opportunity to analyze this unexplored relationship. The reform was a major intervention into the German labor market due to the two factors. First, the scope of the refore was almost universal, with only a few legal exemptions. Second, the reform affected a signicant share of labor force. In 2015, when the reform was introduced, around 10 to 16 percent of eligible employees had a wage rate below €8.50 per hour (see e.g. Amlinger et al. (2016);

Brenke (2014)). We show that the introduction of such a high-impact minimum wage induces a substantial increase in reservation wages among non-workers at the low end of the distribution. Specifically, the minimum wage causes an increase of 4.2 percent at the 10th percentile of the reservation wage distribution. Likewise, it induces a 3.9 percent growth at the 25th percentile. Higher percentiles do not exhibit any change.

3.1 Theoretical Framework

The relationship between reservation and minimum wages can be derived from a simple individual search problem. Consider the decision problem of an individual who lives forever and discounts the future at rate ρ. The time is continuos and at each instant jobs arrive according to Poisson process with parameter α. A job is simply a wage w from a known distribution $F(w)$. If the individual accepts the job, she works for this job forever. If she declines the job, she enjoys unemployment benefits b, and is in the same position.

The value of employment is given by

$$\rho W(w) = w.$$

The value of unemployment is given by

$$\rho U = b + \alpha \int_0^\infty \max[0, W(w) - U] dF(w),$$

where rU is the per period (flow) value of being unemployed. It is b plus any expected change in the worker's situation.

Then, the reservation wage rate can be defined as

$$W(w^{res}) = U.$$

Since $W(w) = w/r$, we have,

$$W(w) - U = \frac{w - w^{res}}{\rho}.$$

As a result, given

$$w^{res} = \rho U = c + \alpha \int_0^\infty \max[0, W(w) - U] dF(w),$$

we get

$$w^{res} = c + \frac{\lambda_u}{\rho} \int_{w^*}^\infty (w - w^{res}) dF(w)$$

or, using integration by parts,

$$w^{res} = b + \frac{\alpha}{\rho} \int_{w^*}^\infty [1 - F(w)] dw. \tag{3.1}$$

Hence, in this simple framework, the reservation wage can be expressed as a function of the expected future value of employment. This value depends on the arrival rate of job offers, α, and the observed wage distribution, $F(w)$. More complicated models may include extensions such as job destruction and job-to-job transitions, but these would not change the relevant predictions for minimum wages that motivate this paper.

Given equation (3.1), a minimum wage may affect this optimal reservation strategy in two ways: through a negative impact on the job offer rate or through the rightward (positive) shift of the wage offer distribution. According to this model, a rightward shift of the wage distribution should increase the reservation wage whereas a decrease in the job offer rate should decrease the reservation wage. These relationships are demonstrated by the sign the first order conditions of the reservation wage equation with respect to α and $[1 - F(w)]$. Whether the minimum wage, however, actually affected these two variables and to what extent they served as drivers of increases in reservation wages remains an empirical exercise to explore, which we address in the next section.

3.2 Data

We use the waves 2013 to 2015 of the SOEP, a representative longitudinal survey launched in 1984 that currently features about 30,000 individuals in 15,000 households annually (doi: 10.5684/soep.v32). The data provide detailed information on all household members, consisting of Germans living in both the former West and East Germany, as well as foreigners and immigrants. The range of topics include demographics, labor market and occupational dynamics, employment, housing, earnings, consumption, savings, wealth, health and life satisfaction indicators. The main survey instruments are a household-level and an individual-level questionnaire. The study asks non-workers[2] the questions that allow us to construct a measure of individual hourly reservation wage. First, they are asked, *"What would your net income have to be for you to accept a position? (euros per month)"* Subsequently, they are asked, *"How many hours per week would you have to work for this income?"* Using this information, we calculate net hourly reservation wages.

Mean, inflation-adjusted hourly reservation wages did not grow between 2012 and 2014, but increased markedly by almost 5 % from 2014-2015.[3] At the same time, the hours respondents reported necessary to earn their threshold wage did not change. Thus, the growth in reservation wages stems from the price (wages per hour) rather than quantity (hours) margin. Figure 3.1 depicts the average shift in reservation wages across the distribution for Germany. As expected, the largest increases in wage thresholds occur at quantiles below the minimum wage, with a small spike immediately below the minimum wage. Interestingly, however, the entire distribution of reservation wages shifted to the right, indicating that even the group of non-workers not directly affected by the minimum wage adjusted their wage threshold for transitioning into employment. Those who expressed a willingness to take up employment at a threshold above the minimum wage prior to the introduction of the €8.50 wage floor required a higher wage level after the reform. This observation suggests that spill-over effects do no only exist for wages of workers, but for the reservation wages of non-workers as well.

[2]Including respondents in voluntary military service, voluntary social year, or federal volunteer service, and excluding any kinds of employment, training programs, apprenticeships and partial retirement

[3]Wages have been CPI-adjusted and are displayed in 2015 constant Euros. Inflation during this period remained very low in Germany.

Table 3.1: Sample description (2013-2015)

	2013	2014	2015	Total
Female share	0.54	0.54	0.54	0.54
Age, average	42.03	42.62	43.13	42.54
German share	0.86	0.87	0.89	0.87
Primary education share	0.32	0.30	0.28	0.30
Secondary education share	0.45	0.46	0.47	0.46
Tertiary education share	0.24	0.24	0.25	0.24
Married share	0.62	0.62	0.62	0.62
Living in HH with children below 16, share	0.50	0.48	0.47	0.49
Observations	23,022	19,915	17,922	60,859

Source: SOEP v32, own calculations.

For the three survey years, we have 6,708 observations for which we observe information on reservation wages and socio-demographic characteristics (gender, age, citizenship, highest educational level, marital status and presence of children in the household). We trim the top and the bottom percentile of annual reservation wage distributions. Table 3.1 describes the sample by survey year.

To establish a causal link between reservation and minimum wages, we use the introduction of a high-impact statutory minimum wage in Germany as a quasi-natural experiment. The statutory minimum wage was introduced on January 1, 2015 for the vast majority of German employees and set immediately at a high level – €8.50 per hour (gross). Thus, the new minimum wage became binding for 10 to 16 percent of eligible employees (Amlinger et al., 2016). Despite the nationally uniform introduction of the minimum wage, its impact differs across regions. Figure 3.2 shows the shares of eligible employees with actual gross hourly wages below €8.50 in 2013 in 96 planning regions in Germany. Due to data limitations, we exclude three regions with less than 30 observations.

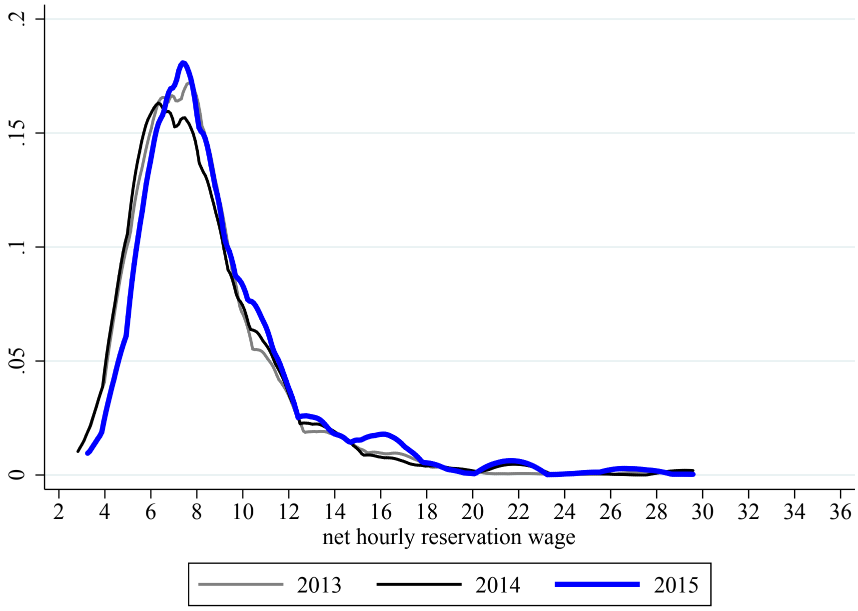

Figure 3.1: Distribution of hourly reservation wages, 2013-2015
Source: SOEP v32, own calculations. Net hourly wages are CPI adjusted.

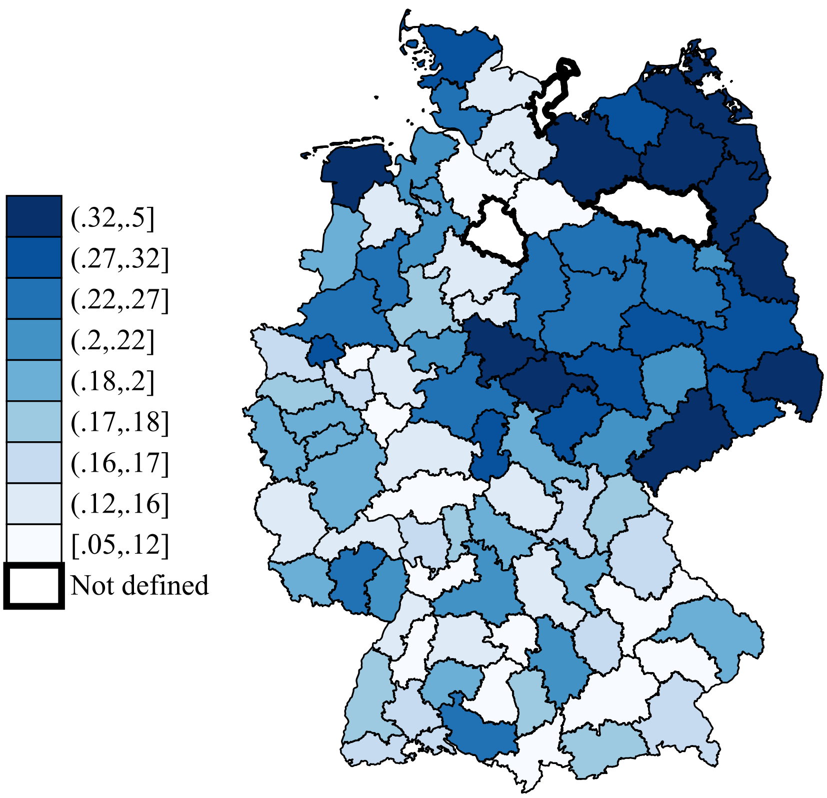

Figure 3.2: Share of employees with actual hourly wages below the minimum wage in 2013
Source: SOEP v32, own calculations.

3.3 Estimation strategy

We estimate the effect of the reform on the reservation wages using a difference-in-differences framework with continuous treatment. Our empirical strategy follows Caliendo (2017a, b). The identification comes from the regional differences in the reform treatment intensity. That is, there is the variation in the average wage across regions, as well as the variation in shares of active employees who are paid below the minimum wage in the year prior to the reform, i.e. how much *bite* the reform has in a region. The dependent variable is the log of the net hourly reservation wage of individual i at time $t \in (2014, 2015)$ residing in region r. The particular specification is as follows:

$$\log(RW_{irt}) = \alpha + \beta \times D_t^{2015} + \gamma Bite_r^{2013} + \delta \left(D_t^{2015} \times Bite_r^{2013} \right) + \mu \mathbf{X}_{irt} + \epsilon_{irt}. \tag{3.2}$$

Here, D_t^{2015} is a dummy variable distinguishing the pre- and post-reform year, with $D_t^{2015} = 1$ if $t = 2015$ and zero otherwise. $Bite_r^{2013}$ denotes treatment intensity (also called the "bite") as captured by the region-specific shares of eligible employees with actual hourly wages below €8.50 (divided by the treatment intensity averaged over all regions in 2013, so that that the average of $Bite_r^{2013}$ over r equals to 1). In order to rule out potential anticipatory effects, we use the lagged treatment intensity for 2013. The coefficient δ on the interaction term $D_t^{2015} \times Bite_r^{2013}$ captures the treatment effect of the reform. Vector $\mathbf{X}_{irt}$ additionally controls for gender, age, marital status, German citizenship, the highest level of education (3 categories), presence of children aged below 16 in the household, as well as the lagged regional unemployment rate and GDP per capita. The latter control for the regional potential to adapt to the reform (Dube et al., 2010; Dolton et al., 2015).

Because we expect the minimum wages to have a differential effect along the distribution of reservation wages, we estimate an unconditional quantile regression based on the re-centered influence function (Firpo et al., 2009).

3.4 Results

Table 3.3 presents estimates of the coefficients δ, β and γ from specification 3.2 for quintiles of the distribution of log hourly reservation wages. At the 10th percentile, the interaction term documents a growth of 4.2 percent due to the reform. Given the reservation wage of €5.24 at the 10th percentile in 2014, this wage growth amounts to €0.22/hour. In the 25th percentile, the growth is 3.9 percent, or €6.24 × 0.039 = €0.24/hour. In higher quintiles, the effect is insignificant. This result confirms that the minimum wage introduction induced an increase in reservation wages exclusively at the bottom of the distribution, where potential low-wage workers are disproportionately located.

The upper panel of Table 3.2 presents the estimates of a placebo regression based on the specification 3.2 using one-year lagged variables and confirms no effect on the distribution of reservation wages prior to the reform, thus, supporting the validity of our identification strategy.

The impact on reservation wages roughly corresponds to the adjustment in the observed hourly wages, which increased by 5.5% in the bottom decile as a result of the reform. Results

Table 3.2: Growth in reservation wages in 2013-2014 (upper panel) and growth in real wages 2014-2015 (lower panel)

	Percentiles of log net hourly reservation wages (2013-2014)				
	P10	P25	P50	P75	P90
$D^{2014} \times Bite^{2012}$	0.014	-0.024	-0.040	0.024	0.053
	(0.039)	(0.047)	(0.049)	(0.042)	(0.038)
D^{2014}	-0.010	0.068	0.093	-0.036	-0.105
	(0.068)	(0.086)	(0.096)	(0.092)	(0.086)
$Bite^{2012}$	-0.120***	-0.150***	-0.122***	-0.127***	-0.092***
	(0.030)	(0.038)	(0.040)	(0.034)	(0.028)
Observations	2205	2205	2205	2205	2205
	Percentiles of log gross hourly wages (2014-2015)				
	P10	P25	P50	P75	P90
$D^{2015} \times Bite^{2013}$	0.055***	0.009	-0.018	-0.004	0.009
	(0.018)	(0.015)	(0.011)	(0.012)	(0.014)
D^{2015}	-0.025	0.016	0.063***	0.038	-0.004
	(0.032)	(0.028)	(0.022)	(0.024)	(0.030)
$Bite^{2013}$	-0.091***	-0.105***	-0.074***	-0.066***	-0.053***
	(0.015)	(0.012)	(0.009)	(0.010)	(0.012)
Observations	20240	20240	20240	20240	20240

Source: SOEP v32, own calculations. Robust standard errors in parentheses.
Significance levels: * p<0.1, ** p<0.05, *** p<0.01

Table 3.3: Causal Effect on Reservation Wages by Percentiles (2014-2015)

	Percentiles of log net hourly reservation wages				
	P10	P25	P50	P75	P90
$D^{2015} \times Bite^{2013}$	0.042**	0.039**	0.023	0.014	-0.064
	(0.020)	(0.020)	(0.019)	(0.025)	(0.050)
D^{2015}	-0.042	-0.031	0.008	-0.009	0.219*
	(0.039)	(0.039)	(0.040)	(0.057)	(0.112)
$Bite^{2013}$	-0.058***	-0.070***	-0.046***	-0.041**	0.005
	(0.018)	(0.016)	(0.015)	(0.020)	(0.034)
Observations	2416	2416	2416	2416	2416

Source: SOEP v32, own calculations. Robust standard errors in parentheses.
Significance levels: * p<0.1, ** p<0.05, *** p<0.01

for the effect on observed wages as estimated from specification 3.2 with log gross hourly wages of eligible employees as the dependent variable are displayed in the lower panel of the Table 3.2. Results confirm that the minimum wage reform similarly affected the distributions of observed and reservation wages.

In order to investigate the channels through which the minimum wage might affect reservation wages, we refer back to the standard search model discussed in the section 3.1. The increase of 5.5% in the bottom quintile of the observed wage distribution should have a positive effect on reservation wages and in fact likely caused the growth of reservation wages at the lower end of the distribution. While an increase in the job arrival rate, α would likewise cause an upward adjustment in reservation wages, it is implausible that the introduction of a minimum wage would induce employers to announce *more* vacancies. Although the lack of individual-level data on job arrival rates prevents us from causally identifying the impact of the reform on α, our results imply that, in the lowest quantile of reservation wages, any decrease in α must have been offset by the rightward shift in $F(w)$.

3.5 Heterogeneous treatment effects

In this section, we investigate heterogeneous effects among two demographic groups with differential exposure to the minimum wage reform: Germans and immigrants[4]. In particular, we re-estimate the regression from Equation (3.2) by interacting the main DiD-terms with a dummy variable equal to unity for Germans and zero for immigrants. Table 3.4 presents this triple interaction estimated using a RIF-regression for unconditional quantiles of the distributions of reservation wages (upper panel) and observed wages (lower panel). The table documents that, below and at the median, Germans exhibit an adjustment in reservation wages of up to 16.5 percentage points higher than immigrants. Above the median, no significant difference exists. Concerning observed wages, the lower panel of Table 3.4 also reveals no differential impact.

Table 3.4: Growth in Reservation and Observed Wages for German Citizens (2014-2015)

Percentiles of:	P10	P25	P50	P75	P90
Log hrly reservation wage					
$D^{2015} \times Bite^{2013} \times German$	0.114*	0.165***	0.132**	-0.005	-0.050
s.e.	0.061	0.059	0.058	0.083	0.155
Log hrly observed wage					
$D^{2015} \times Bite^{2013} \times German$	0.069	-0.028	-0.034	-0.034	-0.081
s.e.	0.090	0.070	0.047	0.048	0.059

Source: SOEP v32, own calculations. Robust standard errors in parentheses.
Significance levels: * p<0.1, ** p<0.05, *** p<0.01

The differential adjustment in reservation wages may stem from different expectations of the two groups concerning the job arrival rate. While we do not have individual-level information on job arrival rates, we approximate α_i for individual i by the respective group-specific transition rate from unemployment into employment: $\alpha_g = UE_{g,t}/U_{g,t-1}$. $UE_{g,t}$ denotes a transition from unemployment to employment in time t and $U_{g,t-1}$ represents the share of job seekers in the given subgroup in $t-1$. Table 3.5 provides descriptive statistics for Germans and immigrants at or below the median reservation wage, including the changes in $\widehat{\alpha}$.[5] Because the approximation of $\widehat{\alpha}_i$ necessarily captures an effect from previously adjusted reservation wages, it must be considered an approximation of job acceptance rates rather than job arrival rates if reservation wages are not static. As such, $\Delta\widehat{\alpha}$ provides suggestive evidence for strategic non-adjustment among immigrants, who may anticipate a lower job arrival rate and adjust their reservation wage downward in order to boost their job finding probability. Whereas natives in 2014 had a higher $\widehat{\alpha}$ than immigrants, it was significantly lower in 2015, indicating a higher change in job

[4]Immigrants are defined as not having German citizenship.
[5]Reservation wage quantiles stem from the 2013 distribution.

acceptance rates among immigrants compared to natives. Thus, if job security motivated the observed negative reservation wage adjustment among immigrants, this search strategy appears to have been successful.

Table 3.5: Characteristics of Germans and immigrants, at or below the median 2013-2015

	Immigrants	Germans	Diff. p-value
Female share	0.52	0.48	0.350
Average Age	39.88	39.30	0.625
Primary education share	0.64	0.50	0.001
Secondary education share	0.25	0.44	0.000
Tertiary education share	0.12	0.06	0.008
Living in HH with children below 16, share	0.41	0.29	0.000
Avg. job acceptance rate 2014, $\widehat{\alpha}_{t-1}$	0.17	0.24	
Avg. job acceptance rate 2015, $\widehat{\alpha}_t$	0.33	0.26	
% Change 2014-2015 job acceptance, $\Delta\widehat{\alpha}$	98.37	11.88	
Observations 2013-2015	8913	72414	

Source: SOEP v32, own calculations.

3.6 Conclusion

In this chapter, we exploit a natural experiment consisting of the introduction of a statutory minimum wage in Germany in 2015 in order to provide novel evidence of a causal impact of minimum wages on reservation wage growth. The contribution of this paper is twofold. Firstly, we show that reservation wages adjust upward in reaction to minimum wage increases and document the absence of supply-side positive employment effects in the short run. Secondly, we provide evidence useful to search theoretical models that reservation wages are not static in the short run. We also show that reservation wages adjust differentially across groups. In particular, natives adjust their reservation wages more than immigrants. One potential explanation is adaptation to increasing productivity requirements on labor and readiness to trade wage growth for job security. These findings have implications for job search and unemployment duration (e.g. Brown and Taylor, 2015).

Chapter 4

German Minimum Wage Reform and Unemployment Duration

Introduction of the federal minimum wage in Germany on January 1, 2015, had a positive effect on the hourly wages of workers at the bottom of the distribution, and virtually no effect on unemployment levels (Caliendo et al., 2017a,b). In the previous chapter, we have studied how this reform affected the reservation wages of workers. In this chapter, we use again the German Socio-Economic Panel data to estimate the effect of the reform on unemployment duration. The importance of this effect comes from the fact that longer unemployment spells are associated with additional costs beyond the loss of financial well-being. These include deterioration of a human capital, a decrease in the probability of re-employment, and social exclusion (Addison et al., 2004). Therefore, the effect of minimum wages on unemployment duration deserves special attention in the light of the ongoing debate on the reform effect. Using a survival analysis in discrete time, we demonstrate that the introduction of the minimum wage is associated with a pronounced increase in duration for younger men. This effect aligns well with the result we obtained in the previous chapter, where we found an increase in the reservation wages for workers who are at the end of the wage distribution. We obtain similar results using two different concepts of unemployment; one that considers employment in mini-job during an unemployment spell as part of that same spell and one that does not. We also provide robustness checks using two alternnative specifications of the baseline hazard, semi-nonparametrical and polynomial.

Existing theories do not give a decisive answer on whether duration should increase or decrease with the introduction of the minimum wage, and how the direction of change depends on the economic environment. Generally, the duration of unemployment depends on the probability that an individual receives a job offer and the probability that the individual will accept it. In the simplest case of a perfectly competitive labor market and exogenous productivity, the introduction of a minimum wage is expected to result in a decrease of the labor demand because firms face higher rental rates of labor, which results in the lower rate of job offers. Simultaneously, the labor supply is expected to increase, since there will be individuals attracted to enter the labor force under the new conditions. These two factors in combination can lead to increase in the unemployment duration since more individuals compete for the vacancies on a tighter market. On the other hand, a higher minimum wage can lead to a decrease in search time for the individuals with reservation wages below minimum wage, since they will accept any job offer. Which effect dominates is not immediately clear.

Once we drop the assumption of a perfectly competitive labor market, things get even less clear-cut. This includes, for example, models of the monopsonistic labor market and efficiency wages. In the case of the "efficiency wage" theory, higher wages can positively affect the expected productivity of a worker, moivating them to work harder. This effect can in principle outweight the additional costs and lead firms to post more vacancies (Schmitt, 2013). In a traditional model of monopsony (see e.g. Card and Krueger, 1995) with an upward sloping labor supply function under certain conditions an increase in the wage caused by minimum wage hike will lead to both increase in employment and firm's output. Therefore, theoretical predictions depend on the model of choice, and the further empirical investigation is necessary.

Furthermore, the analysis need to be adjusted for a specific feature of the German labor market. In particular, unemployed worker in Germany can have a limited amount of work and additional income during their job search without losing unemployment benefits (see e.g. Caliendo et al., 2016). They are allowed to do so by being marginally employed on a "mini-job", where an employee earns up to €450 per month (in 2013) working a few hours per week. Because of this, the data contains a number of overlapping spells, where individuals both report themselves as unemployed and having a mini-job. In order to control for this feature, we conduct our analysis using two concepts of unemployment. In the first case, we treat overlapping spells of this type as employment. In the second case, we treat the overlapping spells of this type as unemployment. Our findings are similar across these two concepts.

First, we find that the introduction of the reform is associated with an increase in expected spell duration. This effect is more pronounced for younger males. Second, we find that young women tend to experience a decrease in their unemployment durations.

There are two reasons our results should be taken with a pinch of salt. First, our sample size shrinks significantly when we condition our estimations both on gender and age groups. Nevertheless, the new wave of SOEP data is coming along in the October 2018 that will allow us to identify the effects with better precision. At this stage, we find significanteffect for the whole sample, but weaker effects for the subsamples. Second, due to the fact that we only observe little more than a year of the post-reform period, a significant part of the spells are right-censored. This is another reason we are expecting the next wave of the SOEP in order to implement our estimates on the data with larger amount of full durations.

4.1 Literature

This study is related to a large literature which has looked at the labor market effect of the minimum wage. The effect of the minimum wage on unemployment levels is perhaps one of the most studied in economics. A review of the literature is presented by Neumark and Wascher (2007). A number of recent studies, such as Dube et al. (2010), Giuliano (2013), have found virtually no negative effects of the minimum wage on unemployment levels. Card and Crueger (1995) offer several possible explanations for the absence of negative employment effects, including monopsonistic labor markets. Schmitt (2013) reviews a literature on the absence of employment effects of minimum wage.

Another strand of the literature is concerned with the factors affecting the duration of unemployment. For example, Blanchard and Diamond (1994) suggest that longer unemployment duration are associated with lower hiring rate. In the same line, using the European Community Household Panel, Addison et al. (2004) find that the arrival rate of the job offers decline, reservation wages and reemployment wages fall with the increase in unemployment spell duration. Addison and Portugal (1992), Fallick (1991), Bover et al. (1996), McCall (1997) find that receipt of unemployment benefits decrease probability of leaving unemployment. Our results are consistent with this paper. Gorter and Gorter (1993) find that levels of unemployment benefits are not very important in ending the search spell. Caliendo, Tatsiramos and Uhlendorff (2009) confirm disincentivising effect of unemployment insurance, but also find a spike in the re-employment hazard which occurs around benefit exhaustion. They also find a positive effect of unemployment benefits extension of susbequent employment outcomes.

Neumark and Wascher (1995) find that higher minimum wage stimulates labor market entry among teenagers. Wessels (2005), on the other hand, finds that higher minimum wage both reduce participation and discourage new participants.

Despite a large body of a literature examinimg effect of minimum wage on unemployment levels, there are only few papers looking at the effect of minimum wage on unemployment spell durations. One example is Pedace and Rohn (2011), who study the effect of minimum wages on duration of unemployment spells using the Displaced Worker Survey of the US. They find that higher minimum wages are associated with shorter spells for most men, while older women experience longer spells. In contrast to this study, we find that in Germany after the reform introduction younger men experience longer duration, whereas younger women experience shorter ones. There is virtually no effect of the reform on the unemployment duration of the older females.

4.2 Data

We use the waves 2012 to 2016 of the Socio-Economic Panel, a representative longitudinal survey launched in 1984 that currently features about 30,000 individuals in 15,000 households annually (doi: 10.5684/soep.v33). The data provide detailed information on all household members, consisting of Germans living in both the former West and East Germany, as well as foreigners and immigrants. The range of topics includes demographics, labor market, and occupational dynamics, employment, housing, earnings, consumption, savings, wealth, health and life satisfaction indicators. Furthermore, the SOEP data is supplemented with personal biography files, which include information on activity status on monthly basis. Each interview, the individuals are asked retrospectively about their activities during the previous year month by month. For the after-reform period, we observe individual histories until the interviews in 2016, which means some unemployment spells are right-censored. We construct complete or right-censored incomplete unemployment spell durations, supplementing spell data with the information from the panel, and drop left-censored spells from the analysis.

The universal legal minimum wage in Germany was introducen on January 1, 2015. Virtually all employees in Germany have been eligible for the minimum of 8.5€per hour since the introduction, with just a few exceptions. First, the long-term unemployed, those who have been

	All age			Age< 25			Age> 50		
	All	Men	Women	All	Men	Women	All	Men	Women
N individuals	2677	1230	1447	449	251	198	521	251	270
Average spell duration (months), MJ as employed	7.977	6.895	8.765	7.574	7.086	7.856	9.751	9.356	10.613
Average spell duration (months), MJ as unemployed	8.351	6.964	9.253	7.817	7.332	8.339	10.283	9.697	11.309
Age	38.27	37.79	38.71	21.68	21.74	21.59	57.2	57.65	56.74
Secondary education, share	0.44	0.39	0.40	0.39	0.35	0.39	0.50	0.48	0.43
Tertiary education, share	0.15	0.12	0.17	0.02	0.006	0.02	0.17	0.15	0.19
East Germany resident	0.32	0.32	0.32	0.28	0.25	0.33	0.37	0.4	0.35
Married	0.4	0.43	0.37	0.05	0.03	0.08	0.61	0.71	0.52
German	0.85	0.86	0.90	0.86	0. 86	0. 87	0.88	0.87	0.92

Table 4.1: Sample characteristics. SOEP v. 33, own calculations.

registered as such for more than 12 month, are not eligible for the minimum wage during their first 6 months on a new job. In practice, however, this exception is rarey used (see vom Berge et al., 2016). Second, minors below age of 18, are not eligible. Finally, the minimum wage doesn't apply to trainees, to voluntary interns during the first three months of their internship, and to entry-level apprentices. The size of the exempted group is small compared to the size of eligible population. In our estimation, we control for the exemptions. Our sample includes individuals who are active on the labor market, above 18 years old, non-istitutionalized, who experienced at least one unemployment spell between January 2012 and an interview date in 2016.

We split our baseline sample by gender and three age groups: individuals of all age, those from 18 to 25 years old, and those above 50. We also construct duration spells using two concepts of unemployment. In the first scenario, we treat overlapping spells of of unemployment and mini-job as employment. In the second case, we treat the overlapping spells of this type as unemployment. The sample characteristics are presented in the Table 4.1.The number of individuals in the subsamples restricted by both age group and gender is small. Women in general expreience longer unemployment durations than men, older individuals stay unemployed longer than young, more singles are unemployed as compared to married, and majority of unemployed have secondary education. Finally, if we treat mini-jobs as unemployment, the average durations are significantly longer.

4.3 Estimation

The typical duration of unemployment spell is few month (see Table 4.1). Therefore, we specify the duration model in discrete time. In particular, let $t \in \{1, 2..\}$ denote the time (in months) elapsed since the beginning of the unemployment spell. The dependent variable is an indicator $\mathbb{I}_{it} \in \{0, 1\}$, that takes a value of 1 if an individual i finds a job in the month t, and 0 otherwise. Let furthermore $h(t, \mathbf{X}_{it})$ be a month-t interval hazard function of an individual i with a vector of covariates $\mathbf{X}_{itr}$ residing in region r. We use random effect model to control for the panel structure of the data.

In our estimations, we use the complementary log-log specification with an interval hazard given by

$$h(t, \mathbf{X}_{itr}) = 1 - \exp(-\exp(\mu(t, \mathbf{X}_{itr}))). \tag{4.1}$$

Our empirical specication follows Addison and Portugal (1992), Pedace and Rohn (2011), Caliendo et al. (2017a, b). These studies, among others, suggested that education, gender, and location of residence were important determinant to use in the empirical analysis of duration. We are adopting the specification of Addison and Portugal (1992) to the case of the German labor market. The independent variables in our specification include:

- a period dummy $\mathbb{I}_t^w$, distinguishing the pre- and post-reform period, with $\mathbb{I}_t^w = 1$ for any calendar month in 2015 or 2016, and 0 otherwise;
- reform treatment intensity on a regional level, $Bite_r^{2013}$. This is so-called "bite" of the reform, defined as a share of employees with hourly wages less than a newly establised minimum in a pre-reform period. In order to control for anticipation effects we use treatment intensity for 2013. As the base regional unit, we use NUTS2 classification of Eurostat;
- regional unemployment rate U_r and a set of year effects D_y to control for changes in aggregate labor market conditions. Since, however, minimum wage effects could be partially captured by the unemployment rate, we also run estimation excluding the unemployment rate from the specification.

Finally, we include a set of individual level controls: gender, age at the beginning of the spell, education level (primary, secondary or tertiary), marital status, presence of kids under the age of 16 in the household, German citizenship, total household income, unemployment benefits receipt indicator, and industry dummies.

We specify the pattern of duration dependence of the baseline hazard $\{\gamma_t\}_{t=1}^T$ in two ways. First, we consider a flexible semi-nonparametric specification with

$$\gamma_t = \sum_{\tau=1}^{T} \gamma_\tau \mathbb{I}\{t = \tau\}.$$

Second, we specify a fifth-order polynomial of the log of t as

$$\gamma_t = \sum_{j=0}^{5} \gamma_j (\log t)^j.$$

Our identification strategy relies on the differential treatment effect of the reform across regions. That is, the effect of minimum wage introduction on a local labor market is likely to be stronger in regions where workers tend to earn relatively less. Analogously, the local labor market would be barely affected by the reform if the majority of workers earn wages higher than the introduced minimum. The measure of the regional treatment intensity we use ("bite") is a share of local eligible employees with hourly wages below the introduced minimum of €8.50. Furthermore, in order to control for the possibility that the impact of minimum wage on unemployment duration across regions varies with the bite of the reform, we interact bite with $\mathbb{I}_t^w$. The estimated coefficients at $\mathbb{I}_t^w$ and the interaction term between regional bite and minimum wage indicator together are intended to capture the effect of the reform on unemployment spell duration. In order to control for the panel structure of the data, we use individual random effects model. The baseline specification then reads

$$\mu(t, \mathbf{X}_{itr}) = \alpha \mathbb{I}_t^w + \phi(\mathbb{I}_t^w \times Bite_r^{2013}) + \gamma_t + D_y\beta_y + U_r + \mathbf{X}_{it}\beta + \tilde{V}_i, \tag{4.2}$$

where $\tilde{V}_i$ indicates individual random effect.

The two parameters α and ϕ capture the effect of minimum wage introduction on the hazard of leaving unemployment. Theis signs of indicate the direction of the duration change. That is, positive coefficients point towards decreased hazard rate and, consequently, an increase in the duration.

4.4 Results

In this section, we present the results of the maximum likelihood estimation of the equation 4.2. The results of the estimation for the two concepts of unemployment are presented in the tables 4.2 and 4.3.

First, let's focus on the results presented in the table 4.2. These results are for the case where mini-jobs are treated as employment. In the baseline specification for the whole sample, there is evidence of a significant negative effect of the minimum wage on the unemployment duration. Note that the signs of the estimated coefficients at the $\mathbb{I}_t^w$ and the interaction term are different. However, only the coefficient at the interaction term is statistically significant. This means, in the regions with the higher share of employees receiving wages less then introduced minimum prior to the reform, the impact of the reform is stronger. In other words, the regions where the bite of the reform is higher, the introduction of the minimum wage is associated with longer unemployment spells. The isolated effect of the reform, as captured by the positive coefficient at the indicator, signals the overall decrease in unemployment duration. However, this effect is not statistically significant. Recall that mini-jobs are treated as employment in this sample, and it is precisely these jobs that pay the lowest wage. In particular, Caliendo et al. (2017a, b) state that about 51% of the mini-job employees were being paid less than the minimum wage prior to the reform. One explanation driving our result, therefore, is lower take-up rate of transition from unemployment to a mini-job. This effect is expected to be the most pronounced in the regions with the highest bite of the reform, since the majority of low-wage earners being employed at the mini-jobs. Caliendo et al. (2017a, b) find that the reform led to the mini-jobs being transformed into regular jobs. This fact is consistent with the potential interpretation where fewer mini-job vacancies being opened. There is no significant effect of the reform introduced in the subsamples of individuals under 25 and above 50 years old. One potential reason for this is relatively small sample size. The signs and magnitudes of coefficients, though, are consistent with the assumption that younger individuals experience an increase in duration.

Next, let's move to the results presented in the table 4.3, where the overlapping spells of registered unemployment and mini-job are treated as unemployment. In this specification, we treat overlapping spells of registered unemployment and mini-job as unemployment. The effect of minimum wage on unemployment duration is also negative. The coefficient at the interaction

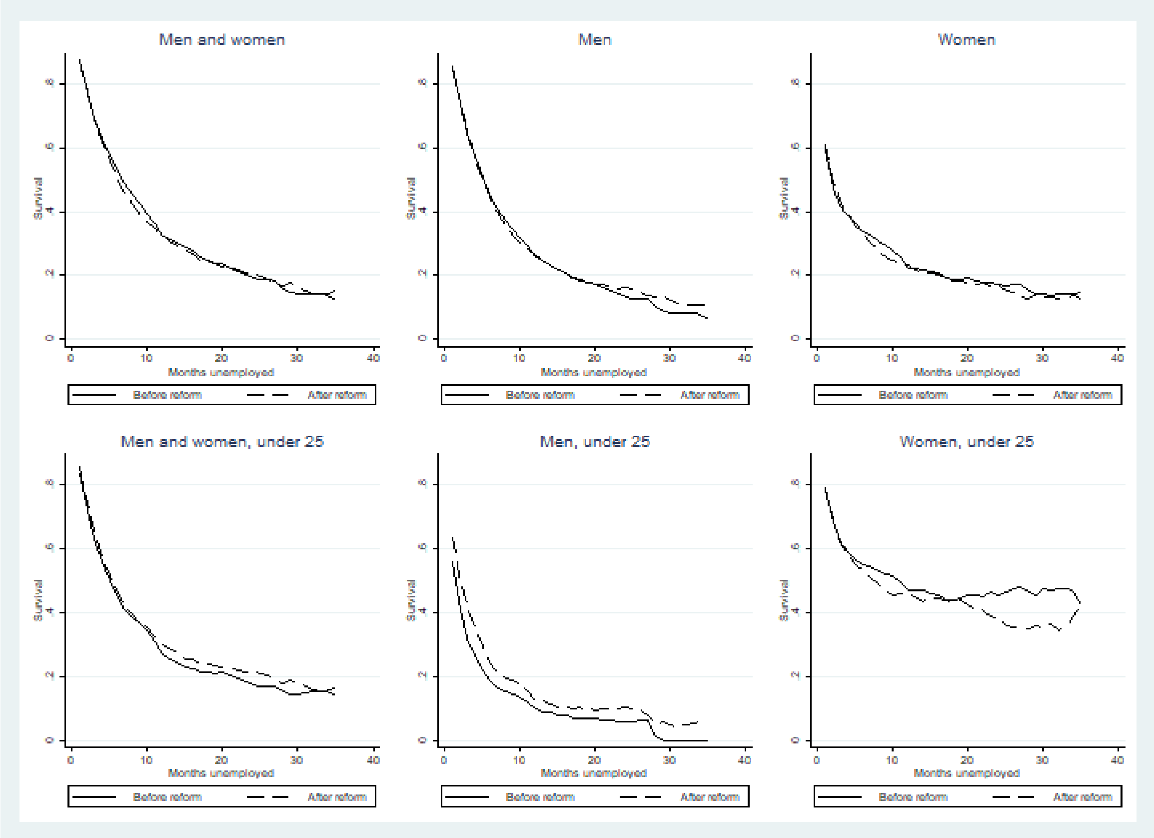

Figure 4.1: Estimated survival functions, mini-jobs treated as employment.

term is positive and not significant anymore, whereas the coefficient at the reform dummy is negative and significant. This means the introduction of the reform is associated with the increase in duration of unemployment spells.

Consistently across all samples, we find strong confirmation that women experience longer durations. Furthermore, the presence of unemployment benefits is associated with longer durations, a result that is consistent with the existing literature. Higher educational attainment and higher household income is associated with shorter unemployment spells. Residence in East Germany seems to be associated with shorter unemployment spells in all of the specifications. Marital status, in turn, seems not to affect durations.

In the table 4.4 we report the effects of the minimum wage introduction on the separate samples of men and women. We do not report the estimated coefficients apart from the α and ϕ, snce their signs are consistent with the estimations in the Tables 4.2 and 4.3. Despite the small samples, we find very strong association of the reform with an increase in unemployment durations across young men. On the opposite, young women seem to experience the decrease in durations. Furthermore, we find virtually no effect in the samples of the older individuals. The result contradicts to Pedace and Rohn (2011), who find theat increase in minimum wage is associated with shorter durations of young men and longer durations for aged women.

The empirical survival functions are presented on the figures 4.1 and 4.2. The figures visually demonstrate the strong increase in the survival probabilities for young men and some decrease for young women.

	All age		Age< 25		Age> 50	
	Non-parametric	Polynomial	Non-parametric	Polynomial	Non-parametric	Polynomial
Minimum wage $\mathbb{I}_t^w$	0.250	0.283	0.131	0.00841	0.101	-0.00800
	(0.182)	(0.185)	(0.411)	(0.405)	(0.413)	(0.426)
$\mathbb{I}_t^w \times Bite_t^{2013}$	-1.218*	-1.276*	-1.882	-1.670	-0.424	-0.224
	(0.694)	(0.739)	(1.611)	(1.617)	(1.476)	(1.526)
Female	-0.286***	-0.414***	-0.0869	-0.139	-0.385***	-0.471***
	(0.0580)	(0.0638)	(0.122)	(0.122)	(0.142)	(0.146)
Age at the beginning of the spell	-0.0202***	-0.0334***	-0.0118	-0.0887***	-0.0177	-0.0505***
	(0.00257)	(0.00275)	(0.0298)	(0.0174)	(0.0177)	(0.00841)
Secondary education	0.302***	0.292***	0.459***	0.454***	0.276*	0.270*
	(0.0604)	(0.0670)	(0.130)	(0.135)	(0.143)	(0.151)
Tertiary education	0.339***	0.363***	0.759*	0.829**	0.178	0.210
	(0.0818)	(0.0898)	(0.402)	(0.367)	(0.191)	(0.197)
East Germany resident	0.536***	1.223***	0.439	0.696***	0.518*	0.684**
	(0.113)	(0.108)	(0.267)	(0.268)	(0.272)	(0.280)
Married	0.0845	0.162**	0.168	0.221	-0.167	-0.166
	(0.0621)	(0.0677)	(0.286)	(0.290)	(0.139)	(0.146)
Kids under 16 in the HH	-0.246***	-0.402***	-0.293**	-0.332**	-0.367**	-0.478***
	(0.0559)	(0.0601)	(0.131)	(0.132)	(0.173)	(0.175)
German	0.189**	-0.0209	-0.0667	-0.161	0.445	0.329
	(0.0777)	(0.0808)	(0.165)	(0.163)	(0.277)	(0.282)
HH income, EUR	0.000167***	0.000135***	0.000200***	0.000176***	0.000126***	0.000123***
	(0.0000238)	(0.0000257)	(0.0000441)	(0.0000445)	(0.0000419)	(0.0000443)
Unemployment benefits	-0.604***	-0.661***	-1.100***	-1.111***	-0.201	-0.235
	(0.213)	(0.227)	(0.391)	(0.408)	(0.492)	(0.512)
Regional unemployment level	-0.107***	-0.220***	-0.122**	-0.163***	-0.0690	-0.0882*
	(0.0213)	(0.0210)	(0.0492)	(0.0492)	(0.0475)	(0.0500)
$Bite_t^{2013}$	0.440	-2.266***	2.219*	1.494	0.392	-0.258
	(0.523)	(0.530)	(1.240)	(1.262)	(1.236)	(1.293)
AIC	14660.1	14757.4	2224.6	2233.7	3053.7	3052.0

Standard errors in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Table 4.2: Full sample, mini-job treated as employment, unrestricted durations. Complementary log-log specification. SOEP v. 33, own calculations.The robust standard errors are reported in parentheses.

	All age		Age< 25		Age> 50	
	Non-parametric	Polynomial	Non-parametric	Polynomial	Non-parametric	Polynomial
Minimum wage $\mathbb{I}_t^w$	-0.408*	-0.460**	0.116	0.00630	-0.182	-0.319
	(0.179)	(0.183)	(0.409)	(0.406)	(0.395)	(0.412)
$\mathbb{I}_t^w \times Bite_t^{2013}$	0.353	0.243	-1.656	-1.513	0.575	0.816
	(0.680)	(0.731)	(1.590)	(1.603)	(1.432)	(1.496)
Female	-0.325***	-0.466***	-0.183	-0.244*	-0.454***	-0.564***
	(0.0584)	(0.0650)	(0.127)	(0.129)	(0.143)	(0.147)
Age at the beginning of the spell	-0.0210***	-0.0349***	-0.0208	-0.101***	-0.0131	-0.0527***
	(0.00259)	(0.00283)	(0.0313)	(0.0188)	(0.0176)	(0.00844)
Secondary education	0.340***	0.328***	0.519***	0.521***	0.332**	0.317**
	(0.0607)	(0.0680)	(0.135)	(0.145)	(0.142)	(0.150)
Tertiary education	0.391***	0.414***	0.249	0.375	0.182	0.214
	(0.0818)	(0.0907)	(0.481)	(0.446)	(0.189)	(0.194)
East Germany resident	0.443***	1.165***	0.405	0.680**	0.321	0.511*
	(0.114)	(0.107)	(0.275)	(0.277)	(0.270)	(0.274)
Married	0.0912	0.175**	0.210	0.271	-0.183	-0.187
	(0.0636)	(0.0701)	(0.307)	(0.314)	(0.137)	(0.145)
Kids under 16 in the HH	-0.271***	-0.435***	-0.364***	-0.409***	-0.354**	-0.485***
	(0.0562)	(0.0610)	(0.134)	(0.137)	(0.169)	(0.171)
German	0.238***	0.0136	0.0784	-0.0193	0.577**	0.443*
	(0.0798)	(0.0827)	(0.176)	(0.176)	(0.265)	(0.268)
HH income	0.000186***	0.000154***	0.000206***	0.000185***	0.000156***	0.000156***
	(0.0000241)	(0.0000259)	(0.0000450)	(0.0000455)	(0.0000416)	(0.0000443)
Unemployment benefits	-0.624***	-0.671***	-1.034**	-1.056**	-0.265	-0.301
	(0.211)	(0.226)	(0.411)	(0.423)	(0.528)	(0.548)
Regional unemployment level	-0.101***	-0.221***	-0.142***	-0.188***	-0.0610	-0.0870*
	(0.0214)	(0.0211)	(0.0513)	(0.0519)	(0.0473)	(0.0493)
$Bite_t^{2013}$	0.513	-2.347***	2.891**	2.186*	0.710	0.0479
	(0.537)	(0.541)	(1.271)	(1.308)	(1.210)	(1.287)
AIC	15150.1	15256.7	2306.1	2313.2	3160.3	3156.9

Standard errors in parentheses
* p<0.10, ** p<0.05, *** p<0.01

Table 4.3: Full sample, mini-job treated as unemployment, unrestricted durations. Complementary log-log specification. SOEP v. 33, own calculations. The robust standard errors are reported in parentheses.

	All age		Age< 25		Age> 50	
	Non-parametric	Polynomial	Non-parametric	Polynomial	Non-parametric	Polynomial
Men			Minijob treated as employment.			
Minimum wage $\mathbb{I}_t^w$	0.0676	-0.468*	-1.225*	-1.286*	0.286	0.173
	(0.266)	(0.279)	(0.497)	(0.533)	(0.614)	(0.642)
$\mathbb{I}_t^w \times Bite_t^{2013}$	-0.708	0.257	-0.289	-0.401	-0.454	-0.183
	(0.994)	(1.094)	(1.812)	(1.981)	(2.044)	(2.157)
AIC	6892.7	6920.9	1270.2	1292.1	1527.9	1522.0
			Minijob treated as unemployment.			
Minimum wage $\mathbb{I}_t^w$	-0.0396	-0.617**	-1.323*	-1.381*	-0.0644	-0.154
	(0.261)	(0.273)	(0.505)	(0.541)	(0.562)	(0.602)
$\mathbb{I}_t^w \times Bite_t^{2013}$	-0.302	0.788	-0.582	-0.732	0.734	0.990
	(0.972)	(1.073)	(1.856)	(2.033)	(1.880)	(2.033)
AIC	7030.8	7065.2	1298.8	1321.4	1539.0	1534.1
Women			Minijob treated as employment.			
Minimum wage $\mathbb{I}_t^w$	0.373	0.382	1.272*	1.302*	-0.197	-0.213
	(0.246)	(0.245)	(0.738)	(0.731)	(0.578)	(0.579)
$\mathbb{I}_t^w \times Bite_t^{2013}$	-1.421	-1.452	-4.090	-4.195	0.0926	0.136
	(0.959)	(0.957)	(3.375)	(3.069)	(2.205)	(2.197)
AIC	7779.1	7782.2	974.0	965.9	1542.7	1546.6
			Minijob treated as unemployment.			
Minimum wage $\mathbb{I}_t^w$	0.189	0.189	0.712	0.697	-0.378	-0.447
	(0.245)	(0.247)	(0.689)	(0.668)	(0.570)	(0.590)
$\mathbb{I}_t^w \times Bite_t^{2013}$	-0.917	-0.935	-3.237	-3.028	0.738	0.926
	(0.940)	(0.948)	(2.857)	(2.824)	(2.247)	(2.321)
AIC	8131.4	8136.7	1020.2	1016.7	1641.2	1637.4

Standard errors in parentheses

* p< 0.10, ** p< 0.05, *** p< 0.01

Table 4.4: Results of estimation of minimum wage effect on male and female subsamples. The specification is for full durations, using both concepts of unemployment. Complementary log-log specification. SOEP v. 33, own calculations. The robust standard errors are reported in parentheses.

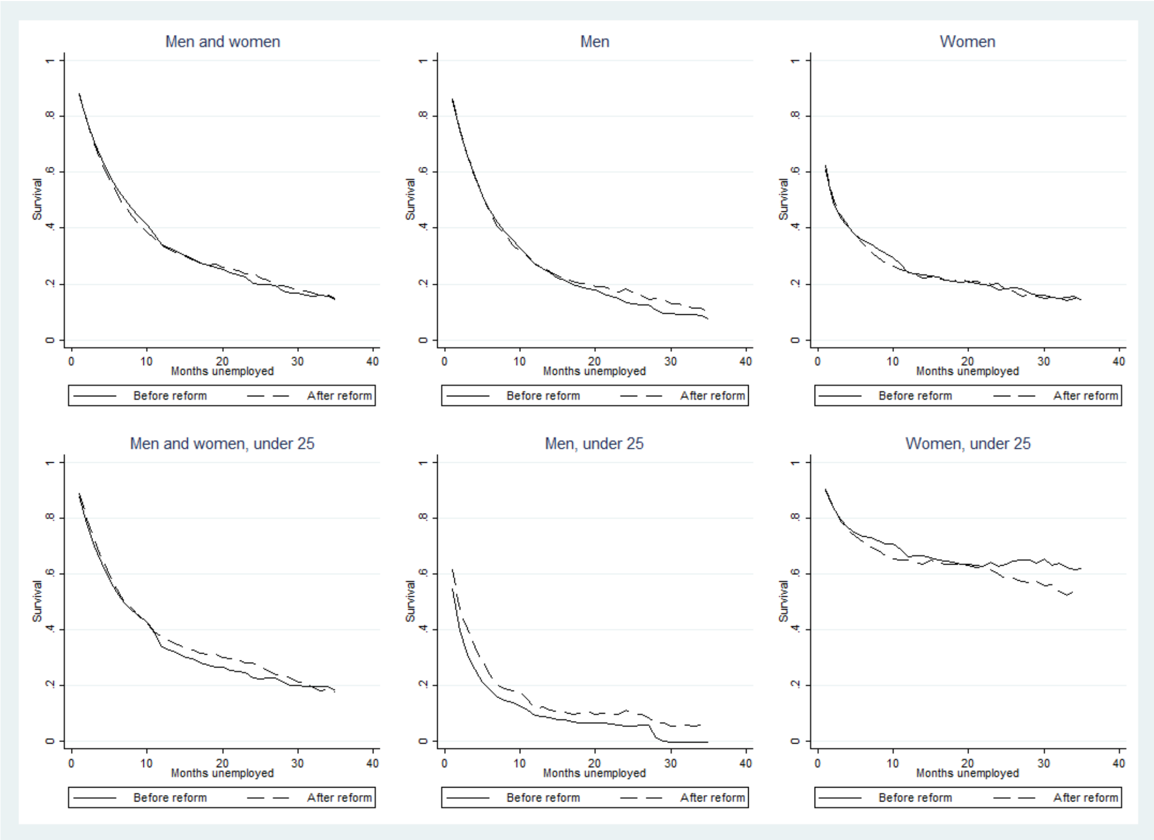

Figure 4.2: Estimated survival functions, mini-jobs treated as unemployment.

4.5 Conclusion

In this chapter we demonstrated that the introduction of the federal minimum wage in Germany is associated with an increase of the average unemployment duration spells below one year. Our estimations suggest that the effect is especially pronounced for younger men. It is robust to the specification of baseline hazard and the definition of unemployment. This result is consistent with the increase of the individual reservation wages. However, the precise mechanism requires structural investigation. One potential explanation of the increase in duration is transformation of the mini-jobs into regular employment.

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