The Labor Supply Effects of Unemployment Insurance for Older Workers*

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Abstract

Extending Unemployment Insurance (UI) benefits can affect labor supply along two margins: it can lengthen the unemployment duration of an individual who is entering UI – the intensive margin – and it can alter the inflows into UI – the extensive margin. We study both margins of the labor supply effects of UI for older workers using German Social Security data and policy variation over 3 decades. We document extensive margin responses in the form of sharp bunching in UI inflows at various age discontinuities in UI eligibility among workers in their late 50s who use UI as a pathway into early retirement. We present evidence of intensive-margin responses among similarly aged workers using regression discontinuity designs at multiple age-based UI thresholds. To quantify the effects of UI extensions on time out of work along both margins, we use the reduced form bunching and RDD evidence to estimate a dynamic life-cycle model of labor supply in which individuals face retirement and search intensity decisions. We estimate and validate the model using moments that exploit extensive policy variation in UI and retirement institutions. Preliminary calculations suggests that both margins are important: for example a 6 month UI extension for men above age 50 increases non-employment durations by around 0.84 months on the intensive margin, while the extensive margin effect is around two to three times as large.

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1 Introduction

Unemployment Insurance (UI) benefits are an important policy tool for helping workers smooth their consumption after job-loss. A large literature has studied the effects of UI extensions on labor supply using quasi experimental methods (see Schmieder and Von Wachter, 2016, for a review). This literature has typically found that UI extensions have sizable effects on the non-employment duration of individuals who become unemployed - the intensive margin, while not having an effect on the inflow rates into unemployment - the extensive margin. This can be most clearly seen in papers based on regression discontinuity designs around age or experience thresholds, where a standard validity check is to show that the density of inflows into UI does not change at the threshold (Card et al., 2007; Centeno and Novo, 2009; Schmieder et al., 2012; Lalive et al., 2015). However, this literature is largely based on relatively young workers in their 30s, 40s and early 50s, who are highly attached to the labor force. Older workers, in their late 50s and onwards, are much closer to retirement and may use UI as a stepping stone into retirement (Dlugosz et al., 2014; Giesecke and Kind, 2013). This may be reinforced by firms that seek to reduce employment in response to a negative shock by buying out workers with relatively high outside options, thanks to the possibility of going into early retirement via an intermittent UI spell. Understanding the total effect of UI extensions on older workers is particularly important given the common goal of extending the work life of the elderly and reducing the burden on the social security system.

In this paper, we study the labor supply effects of UI extensions for older workers in Germany using the social security data from 1975 to 2013. Numerous reforms to Germany's UI and retirement system over this period altered both the payoffs to entering UI at different age thresholds and the search incentives of the unemployed. Workers in their 50s responded sharply to these policy changes on both the intensive and the extensive margin. We observe and quantify the *intensive margin* effect of UI extensions at 12 different age cutoffs that discontinuously extend UI for workers in their 40s and 50s using regression discontinuity designs. Our evidence suggests that the intensive margin effect is at least as large for workers in their early and late 50s as it is for workers in their 40s.

At the same time, we document large *extensive margin* effects of UI. There is sharp bunching of UI inflows at precisely the age that allows workers to claim their pension right after UI expiration. The age at which workers can enter unemployment and subsequently receive a pension without any uninsured period can be thought of as a kink in a lifetime budget set relating income to exit age. The extensive margin responses to UI policies at this bridge-to-retirement kink are quantitatively meaningful. The bunching mass of inflows is large and yields estimates of the elasticity of exit age with respect to the net return to work that are comparable to other settings in which individuals choose when to retire without having to go through UI (Brown, 2013). Furthermore, UI inflows respond as expected to a series of UI extensions and pension rule changes.

While the reduced form evidence clearly shows that workers responded to potential UI benefit durations on both margins, it is difficult to quantify the total non-employment effect of UI extensions in a manner that allows for comparisons of the relative size of the intensive and extensive margin responses. Doing so requires making implausibly strong assumptions like assuming that the entire bunching mass shifts in response to a UI extension. Instead, we specify and estimate a life cycle model of labor supply that explicitly captures transitions between employment, unemployment with search, and unemployment as retirement. Our model is similar to a basic labor supply model of retirement decisions, but with two key extensions. First we incorporate endogenous job loss into the model in a way that explicitly allows for a role of firms. Second we model unemployment in more detail than past life cycle models. While previous research with life cycle models typically allows for one or two states of unemployment with no dynamics, we model unemployment as a fully dynamic process that allows us to capture the duration of UI benefits and the response to their exhaustion in a natural way.[1] This also has the added benefit that our parameter estimates for the job search part can be compared with previous estimates of job search models (e.g. Paserman, 2008; DellaVigna et al., 2017). We estimate our model using a methods of moments estimator and use the resulting best-fit parameters to decompose the total non-employment effect of counter-factual UI extensions into its extensive and intensive margin components.

Germany provides a particularly interesting context for studying UI extensions for older workers, since there has been a tremendous amount of policy variation over the past decades. In the early 1980s, the maximum potential benefit duration (PBD) of UI was capped at 12 months regardless of age. Throughout the 1980s, maximum PBDs increased dramatically for older workers, reaching up to 32 months of UI benefits for the oldest group. Between 1999 and 2007, Germany reversed track. Maximum PBDs were reduced for older workers and Germany began the process of eliminating early retirement at age 60 following unemployment. This increase (and later decrease) in UI generosity is matched by a sharp increase (decrease) in the unemployment rate among older workers. Previous authors, such as Buchholz et al. (2013), have attributed this to a variety of policy changes aimed at reducing the labor supply of older workers, but these papers have not attempted to isolate the impact of UI.

While Germany provides many compelling advantages for studying the effects of UI for workers it also offers a number of challenges. The main complication is that in addition to UI there are significant number of other policies that changed over the past decades and that

[1]Other structural life cycle papers typically assume workers receive UI forever or model UI as a Markov process with a fixed transition probability to exhaustion (Haan and Prowse, 2010; García-Pérez and Sánchez-Martín, 2015).

may affect inflows into UI and unemployment durations. Some of these changes are about regular and early retirement rules and are relatively easy to understand, but there are also many rules based on collective labor agreements (CLAs) that are on the sectoral level or even specific to individual firms. Such CLAs may themselves take policy induced age discontinuities into account, for example by encouraging workers to exit firms at those age thresholds with severance packages. In this case one can view CLAs as a mechanism of how age discontinuities lead to extensive margin responses. On the other hand, CLAs may also lead to bunching at age thresholds that are not directly related to retirement or UI institutions. This complicates our setting and we consider a variety of approaches to obtain meaningful estimates in light of such confounding factors.

Our setting also raises interesting methodological issues. While several papers have estimated regression discontinuity designs in the presence of manipulation of the forcing variable (see for e.g.: Card and Giuliano (2014); Gerard et al. (2015); Barreca et al. (2016); Hoxby and Bulman (2016)), this manipulation has typically been treated as a nuisance, with researchers attempting to avoid bias using techniques like excluding observations close to the threshold (donut-hole regressions). However, whether and when to enter UI is itself an important outcome and in practice individuals (together with firms) can influence this decision. When UI is used as a pathway to retirement, it essentially constitutes a labor supply decision with respect to a budget set defined by wage rates, the UI system and retirement rules. The UI system creates kinks in this budget set and individuals choosing to enter UI as a step towards retirement should bunch at these kink points. We could thus use bunching techniques to back out labor supply elasticities for these workers, based on the amount of bunching around such kinks (Saez, 2010; Kleven, 2016) but these estimates would be based on a simple static life-cycle model that ignores the possible interactions of responding along the intensive and extensive margin and is difficult to reconcile with inflows into unemployment that are not driven by voluntary quits. Furthermore these kind of static bunching estimators are sensitive to ad hoc restrictions about the counterfactual distribution that are not obvious how to specify in our context (Blomquist and Newey, 2017).[2] While we provide some bunching estimates of these elasticities for comparison with other papers, we focus on the estimating the dynamic model to quantify the reduced form evidence.

While bunching can help recover extensive margin decisions, it complicates identification of intensive margin effects. Ideally, we would use the discrete changes in potential benefit duration at the age thresholds to estimate intensive margin responses. Yet, extensive margin

[2] Note that not all bunching around UI age discontinuities is necessarily related to early retirement. It may also be that firms postpone lay-offs or workers postpone claiming UI benefits until they reach the threshold. This is likely to be most important at ages further away from the retirement age, such as the age threshold at age 54 in the 1990s and the threshold at age 55 in the 2000s.

responses at or around these thresholds lead to direct violation of the RD assumption that there is no manipulation of the running variable and individuals on both sides of the cutoff are therefore comparable. We use three approaches to circumvent this challenge and obtain plausible estimates of intensive margin responses: first, we use donut-hole regressions to exclude the range where most of the bunching occurs. This is most credible when the bunching is not too extreme and there does not appear to be an overall shift in the density outside of a sharp window around the threshold. Second, we include a series of individual level controls to help absorb selection effects. Third, we estimate intensive margin responses at slightly younger age thresholds, where bunching is less of an issue, in particular a threshold at age 54 during the 1990s.[3]

Finally, we view our setting as particularly apt for a structural approach. First, our setting provides clear economic incentives and interactions between incentives that can be captured in a fairly standard model. Second, we have clean policy variation and empirical moments to identify the key parameters of the model, disciplining the estimation approach. Third, there are several reforms that allow us to test the out of sample performance of our model. And fourth, there is a clear value added to having a model for counterfactual and out of sample analyses. A model makes it straightforward to predict how intensive and extensive margin responses determine responses to any given policy reform and what the contribution of each channel is. It also allows us to quantify how much of the changes in retirement age over the past decades in Germany can be explained by the reduction of UI generosity for older workers over this period.

Our paper is related to a large literature on retirement decisions. Several methodologically related papers have analyzed bunching in retirement age to derive labor supply elasticities, for example Brown (2013) looks at bunching at the regular retirement age for teachers and Manoli and Weber (2016) analyze permanent exits from the labor force around tenure thresholds in Austria that lead to discrete increases in severance payments. Unlike these papers we look specifically at entries into UI, rather than exits from the labor force.

A handful of papers examine the effects of UI extensions on older workers (e.g. Kyyrä and Ollikainen (2008); Bennmarker et al. (2013)). Among these, Riphahn and Schrader (2017) and Dlugosz et al. (2014) show that a 2006 reform that shortened UI benefits for older German workers reform increased employment. A small literature explicitly examines interactions of the UI system with retirement decisions. For example, Kyyrä and Pesola (2017) show that postponing eligibility by two years for a retirement-via-UI pathway in Finland increases

[3] We can also estimate intensive margin responses on a sample of individuals who later return to the labor market, which likely obtains a lower bound of the intensive margin response for these workers. Finally, we can follow the approach in Gerard et al. (2015), who explicitly provide a framework to estimate bounds in RD settings in the presence of sorting.

employment by 7 months. Similarly, Kyyrä and Wilke (2007) show that increasing the age threshold of early retirement via UI benefits from 53 to 57 in Finland significantly reduced unemployment durations. Lalive (2008) analyzes the effect of UI extensions for older workers around a discontinuity at age 50 in the Austrian UI system, as well as at a border discontinuity, and finds relatively large disincentive effects, especially for women. Using partially the same variation as Lalive, Inderbitzin et al. (2016) show that much of this was due to early retirement responses. Hairault et al. (2010) provide evidence based on French survey data that job search behavior of the unemployed depends on the distance to retirement age.[4] Several papers analyze the interaction between various retirement rules and labor supply in Germany (see Giesecke and Kind (2013); Boersch-Supan et al. (2004); Boersch-Supan and Hendrik (2011), among others). We focus on quantifying the overall effect of UI extensions on labor supply for older workers, accounting for both extensive and, the well documented (Card et al., 2007; Schmieder et al., 2012), intensive margin behavior.[5]

Finally, our work suggests that German firms play a role in regulating how worker inflows into UI respond to UI extensions. Jaeger et al. (2017) study job destruction following improvement in workers' outside options using variation in UIB in Austria, finding that low surplus jobs are destroyed. A few studies have estimated the sensitivity of layoffs of older workers to monetary incentives with some finding little sensitivity (Behaghel et al., 2008; Johnston, 2017), but not others (Schnalzenberger and Winter-Ebmer, 2009). While we cannot cleanly and separately identify the role of firms from that of workers, we present several pieces of suggestive evidence consistent with a non-negligible role for firms.

This paper proceeds in five steps. We first provide a very general framework for decomposing the effect of UI extensions on time out of work in the presence of intensive and extensive margin labor supply responses, which highlights the importance of imposing additional structure to fully estimate these responses. Section 3, describes the core features of the German UI and retirement institutions and discusses our data. In Section 4, we present reduced form evidence of both intensive and extensive margin responses. We document bunching in UI inflows at the bridge-to-retirement kink and at age cutoffs that discontinuously increase PBDs. We estimate intensive margin effects from RDs at all the older age cutoffs available to us, using various approaches to handle sorting at some of these cutoffs. In Section 5, we develop a full life cycle model of labor supply, job search, and retirement decisions. We estimate this model in Section 6 using a method-of-moments estimator and perform counter-factual simulations. Section 7 concludes.

[4] In contrast, Coile and Levine (2007) find that UI generosity has little impact on retirement in the U.S.

[5] While our focus is to quantify the overall effect of UI extensions rather than discussing optimal policy, our analysis can be viewed as an important input into welfare computations. For papers on the optimal design of UI for older workers see, for example: Hairault et al. (2012); Michelacci and Ruffo (2015); Inderbitzin et al. (2016).

2 Conceptual Framework: The Effect of UI Extensions on Total Time out of Work

To fix ideas and terminology, we present a simple framework that describes how potential UI benefit duration affects time out of work in the presence of extensive margin responses. There is a mass of workers N, who enter the workforce at age 1 and reach a mandatory retirement age at T^R. Let potential UI duration be P. In practice this can be a function of the age of entry into UI, but for simplicity of exposition we take P to be constant. In each period t (meaning at each age), a worker is either working or not working. We do not distinguish between unemployment and non-employment and use the terms interchangeably. The fraction of workers entering unemployment at age t is denoted as $g_t(P)$. If an individual becomes unemployed, the duration of non-employment is defined as the time between entering unemployment and either starting a job again or when the individual retires. We denote the expected non-employment duration of individuals becoming unemployed at age t as $D_t(P)$. The expected total time out of work (T^u) for an individual is given by $T^u(P) = \sum_{t=1}^{T^R} g_t(P)D_t(P)$.

Without specifying micro-foundations of this labor market, we can write the relationship between inflows and durations on benefit durations as reduced form functions and decompose the effects of an increase in potential benefit into intensive and extensive margin components. A change in P can be decomposed into:

$$\frac{dT^u}{dP} = \underbrace{\sum_{t=1}^{T^R} g_t \frac{\partial D_t}{\partial P}}_{\text{Intensive Margin}} + \underbrace{\sum_{t=1}^{T^R} D_t \frac{\partial g_t}{\partial P}}_{\text{Extensive Margin}} \tag{1}$$

The first term represents the standard intensive margin effect of UI extensions on non-employment durations that most of the UI literature has estimated. The second term represents the changes in inflows into unemployment.

The central question of this paper is how to credibly estimate this total effect for older workers. This is challenging to estimate using purely reduced form techniques. Note that $\frac{\partial g_t}{\partial P}$ is never likely to be 0. In practice, if P increases D, employment falls, changing the pool of people at risk of becoming unemployment. Hence, future g_t might decrease, violating $\frac{\partial g_t}{\partial P} = 0$ for some t.

It is instructive to consider two simple cases. Let us assume that such effects on the pool of at-risk people are negligible and that we are focused on younger workers. Schmieder et al. (2012) show that younger workers in Germany do not significantly alter their entry probabilities into UI in response to changes in P. In this case $\frac{\partial g_t}{\partial P} \approx 0$. In such cases $\frac{dT^u}{dP}$ can be recovered from RD estimates of the intensive margin effect at age cutoffs in P as in

Schmieder et al. (2012).[6]

Now consider a case at the other extreme, with extensive but no intensive margin effects. Suppose older workers only use UI as a bridge-to-retirement, and never become unemployed except by their own choice. Once they exit they stay non-employed until the age at which they can claim their pension (T^R). So at each age, the expected non-employment duration is fixed, such that $D_t = T^R - t$. However, suppose that these workers time their exit date (into UI) and that this responds to P. For example, suppose there is a mass $\tilde{N}$ of workers that time their entry to maximize unemployment coverage before retirement by entering UI at age $T^R - P$. In that case there are not intensive margin responses, $\frac{\partial D_t}{\partial P} = 0$, but the density of inflows changes for $t = P$: $\frac{\partial g_t}{\partial P} \neq 0$.[7] Suppose that the mass of entries at all ages is otherwise constant. Then an increase in P to $\hat{P}$ would decrease total non-employment duration by $D_{T^R-P} \times \frac{\tilde{N}}{N}$ at age $T^R - P$ and increase non-employment duration by $D_{T^R-\hat{P}} \times \frac{\tilde{N}}{N}$ at age $T^R - \hat{P}$. The total non-employment effect of the increase in P in this pure, extensive margin setting would thus be given by $\frac{dT^u}{dP} = (D_{T^R-\hat{P}} - D_{T^R-P})\frac{\tilde{N}}{N} = (\hat{P} - P)\frac{\tilde{N}}{N}$.

In practice, though, neither of these cases fully captures the complexity of reality. Older workers are likely to still have strong intensive margin responses to changes in P, and some older workers, even at later ages, will find themselves unemployed not by their own choice. Note further, that a simple two type model (one representing each case above) is problematic in its artificiality – in practice workers may choose to transition between states as a function of P. For these reasons, after presenting our reduced form evidence, we opt to specify and estimate a fully specified, dynamic life cycle model that explicitly accounts for these types of transitions.

3 Institutional Background and Data

3.1 Unemployment Insurance

The German unemployment insurance system provides income replacement to eligible workers who lose their job. Prior to 1985, eligible workers were entitled to at most 12 months of benefits. Replacement rates for UI are relatively stable over the period of study (1980–2015)

[6]Note also that if one is interested in more complicated changes in potential benefit durations (that is not just an increase at a single age level k), then it is still relatively straightforward to estimate the intensive margin effect by aggregating estimates of $\frac{\partial D_k}{\partial P_k}$ at different age levels.

[7]This could be the case if we think of workers as maximizing lifetime utility over consumption and leisure subject to their budget constraints. Depending on institutional parameters, such a lifetime budget constraint might exhibit a kink at the 'bridge-to-retirement-via-UI' age of $T^R - P$, as we will show is the case in Germany. Extending P moves this kink, and hence moves UI exit mass. If this were the correct model, we could calibrate its key parameters using bunching techniques in a manner similar to Brown (2013) and then simulate $\frac{dT^u}{dP}$.

(67-68% for an individual with children and 63-60% for an individual without children).[8] Beginning in 1985, numerous reforms changed the maximum UI potential benefit duration (PBD) in a manner that tied the maximum PBD to recipients' exact age at the beginning of their UI spell.[9]

Reforms in 1985 and 1987 increased maximum PBDs for workers above age 42. The most generous PBD -- up to 32 months -- became available to workers aged 54 and above following the 1987 reform. Reforms in 1999 and 2006 gradually decreased the generosity of the system. In 1999, age thresholds were increased, and then, beginning 2006, maximum PBD was reduced from 32 to 18 months for workers above age 55, while everyone else could only receive 12 months. There was a modest reversal of this trend in 2008 when workers above 58 could attain a maximum PBD of 24 months.

Figure 1 plots maximum PBD by age for older workers in each different institutional regime.[10] Appendix Table D.1 provides details about each reform. These policy changes provide highly useful empirical variation, both at the age thresholds, and by changing incentives on when to enter unemployment if using unemployment as a bridge-to-retirement, as we elaborate on in the next section.

3.2 Pension System and Early Retirement Via Unemployment

Germany has a generous pay-as-you-go public pension insurance with high effective replacement rates. Participation is mandatory, with the exception of civil servants and the self-employed, which are not covered by our data. Pension benefits depend on workers' earnings, years of contributions, an adjustment factor, and the type of pension claimed. Benefits are roughly proportional to lifetime income at an average replacement rate of 50% (Deutsche Rentenversicherung (2017)).

The statutory retirement age (SRA) for a regular old age pension remained at 65 throughout our sample period, with the only prerequisite being 5 years of contributions. Earlier retirement was possible under several circumstances. Several alternate pathways made receiving a pension before 65 an option.[11] Our focus is on the pathway into retirement via

[8]Individuals who exhausted UI benefits prior to 2005 and whose net liquid wealth fell below a certain threshold were eligible for unemployment assistance (UA) benefits with an effective average replacement rate of around 30%. In principle, replacement rates were between 50% and 57% but lower in practice due to deductions like spousal income. See Schmieder et al. (2012) for a discussion. From 2005 on, UA was replaced by unemployment insurance benefits 2 (UIB II), a completely means tested program. Both UA and UIB II are unlimited in duration.

[9]See Hunt (1995); Fitzenberger and Wilke (2010) for an analysis and discussion of these reforms.

[10]We omit the short 1985 regime in the interest of brevity and because it appears that some individuals who entered UI in 1985 retroactively benefited from the UI extensions in later years.

[11]The four alternative pathways to retirement were old-age pensions for long-term insured, old-age pensions for women, old-age pensions due to unemployment (and, later, part-time work) and old-age pensions for

unemployment.[12]

The unemployment pathway (UI pathway) provided eligible workers with an option to retire at the age of 60.[13] The eligibility requirements for this pathway were: 1) at least 15 years of contributions, at least 8 of which must have occurred in the past 10 years, and 2) being unemployed for at least 1 year after the age of 58 and a half. The generosity of UI benefits, combined with lenient job search requirements for older workers, made old-age pensions due to unemployment attractive. After the late 1980s, unemployed individuals aged 58 and older could receive generous unemployment benefits without actively looking for a job or other obligations.[14] We note, further, that entering UI voluntarily is highly feasible in Germany and at most lightly penalized.[15]

This system incentivizes workers considering early retirement to time their entry it to UI around the age that allows workers to transition directly from UI to pension, without any uncovered period. Put differently, the possibility of using UI as a bridge-to-retirement introduces a kink in a lifetime budget constraint relating lifetime income to year of exit into UI. Individuals retiring before $60 - P$, with P being the maximum UI PBD, are forced to spend time reliant on a spouse or on unemployment assistance (UA/UIB 2) before their pension, whereas individuals who leave at or after $60 - P$ can take the full UI duration and transfer directly into pensions. This reduces the value of an extra year of work after the kink, decreasing the slope of the budget constraint. In general, the size of the kink is exacerbated by the generosity of the UI system, the size of the drop comparing UI to UA/UIB 2, and how generously time on UI is counted towards pension contributions.[16] We will show that UI entries react to this kink at age $60 - P$.

The unpenalized retirement age (NRA) and the earliest allowable retirement age (ERA) via the UI pathway remained at 60 until a 1992 reform. Cohorts born between January 1937 and December 1941 saw their NRA increase in steps by birth month from 60 to 65. While they could continue to retire at the ERA of 60, they now faced an actuarial adjustment in the form of a 0.3% pension reduction per each month they retired in advance of the NRA.

severely disabled persons (see for example Boersch-Supan and Wilke (2005)). Appendix Table D.2 documents the earliest possible retirement age for each of these pathways over the past 4 decades.

[12]While early retirement due to disability is also quantitatively important, Riphahn (1997) argues that in practice they are not close substitutes and that retirement due to disability is in fact usually associated with a health shock.

[13]For the first 3 cohorts we will focus on, the unpenalized/normal retirement age (NRA) as well as the earliest possible retirement age (ERA) via the UI pathway was age 60. Persons satisfying the requirements could retire at 60 with no penalty other than the loss of additional years of pension contributions. Later, these ages increase.

[14]This so-called "58er-Regelung" was formally introduced end of 1985 and in place until end of 2007.

[15]A worker may be sanctioned if he or she quits a job voluntarily. These sanctions take the form of losing the first few weeks of benefits and vary from a 4-12 week penalty over the study period. These sanctions, which are not always applied, are insufficient to offset the appeal of using UI as a pathway into retirement.

[16]In practice, unemployment counts as an 80% contribution year calculated on pre-unemployment wages.

Cohorts born in or after 1946 saw their ERA for the UI-pathway increase in steps by birth month from 60 to 63, ending with cohorts born in December 1948. This meant that these cohorts could no longer claim their pensions at age 60, even with a penalty. Cohorts born after 1952 (after our sample) saw this pathway into retirement via UI entirely abolished.

Figure 2 plots the evolution of stylized lifetime budget constraints for select cohorts experiencing different UI and pension regimes. Appendix B contains detailed descriptions of how these budget sets are constructed. For simplicity, we assume that the max PBD is fixed over time for each cohort at the level that prevailed when they were close to the kink.[17] In panels (a)-(c), representing the 1924, 1929, and 1935 cohort respectively, the the NRA and ERA for retirement via unemployment was age 60, but maximum PBD varied. In panel (d), representing the 1941 cohort, the ERA remained at 60 but the un-penalized NRA was increased to around 64, with slight variation by month of birth. This amounted to a financial penalty for retiring at age 60 of approximately 18% of gross lifetime pension benefits. In panel (e), representing the 1949 cohort, the ERA was increased to 63 and the NRA was 65.[18] The penalty for retiring at age 63 via unemployment was 7.2%. In panel (f), representing the 1952 cohort, the pathway into retirement via unemployment was abolished, leaving the earliest possible retirement age as 63 for long-term insured workers with over 35 years of qualified contributions.[19]

Throughout the rest of this paper, we focus primarily on the first kink induced by using UI as a bridge-to-retirement.[20] In practice, agents might also use UI as a bridge to the long-term contribution retirement age of 63 or the regular retirement age of 65.[21] Since we cannot credibly calculate whether or not a person is eligible for the long-term contribution rate, examining bunching at these kinks is challenging. Note also, that changes in other pathways may create alternative substitutes for workers aiming to retire early. Appendix Table D.3 summarizes the reforms for all of the different pathways over our study period.

3.3 The Role of Firms and Collective Labor Agreements

Firms clearly play an important role in workers' early exit from the labor force over our time period. After labor shortages in the 1960s and 1970s and extremely low unemployment rates

[17] The dashed line shows the realized lifetime budget set that takes into account all the UI policy-induced changes and age cutoffs.

[18] Retiring at 63 via the long-term insured pathway is slightly more costly.

[19] The un-penalized NRA for the long-term insured is 65.5, making the financial penalty for retiring at 63 9%.

[20] Note that the large, discontinuous increases in PBD at the various age cutoffs in the PBD duration schedule (see Figure1) could also induce selection into UI. This could occur for both people who only plan to temporarily be on unemployment and among people planning to retire. We do not focus on this behavior directly, but discuss it as appropriate.

[21] Individuals cannot receive UI past 65, and cannot receive UI and pensions simultaneously.

(~1%) the German economy was hit with the 1982 recession and shrinking labor demand that led to fast rising unemployment. Facing employment protection laws and powerful unions and work councils, firms sought to downsize employment through voluntary ways by negotating collective labor agreements (CLAs) with their workforce. These CLAs typically offered severance packages to older workers to voluntarily quit the firm, and the severance packages were often tied to specific age threshold. These severance packages essentially constituted a way to buy workers out, and the represent a form of a mutually agreed upon ending to the employment relationship.

Clearly whether or not a worker would be willing to accept a severance package would depend on outside option of the worker when leaving his job. In a labor market with high unemployment rates, that prevailed in the 80s and 90s, exiting a job as a worker in their 50s was essentially equivalent accepting never to find work again and the availability of unemployment benefits was a crucial factor in this process. Firms and labor unions who negotiated were clearly aware of the institutional setting and would take the structure of UI benefits into account when negotiating workforce reductions and exit packages as part of CLAs.

CLAs on early retirement were often implemented at the sectoral level but could be specific to individual firms. The details of these CLAs, including the earliest retirement age and the corresponding severance package, vary. For example, workers in the printing sector can enter early retirement via CLAs as early as age 57, while workers in the construction sector can going into early retirement starting at age 55. The metal industry provided compensation for deductions from retirement benefits of DM 450 per month for a maximum of two years, while the chemical industry provided compensation between DM 450 and 750 per month for a maximum of four years (see Trampusch, 2009).

Such CLAs may themselves take age discontinuities induced by the UI and public pension system into account, for example by encouraging workers to exit firms at those age thresholds with severance packages. In this case one can view CLAs as a mechanism of how age discontinuities lead to extensive margin responses. On the other hand, CLAs may also lead to bunching at age thresholds that are not directly related to retirement or UI institutions. This complicates our setting and we consider a variety of approaches to obtain meaningful estimates in light of such confounding.

The government also supported CLAs on early retirement in other forms, such as subsidizing employers' costs of buying-out older workers through the so-called partial retirement law (Altersteilzeitgesetz).[22] This partial retirement law provided a maximum public subsidy for

[22]This partial retirement law (Altersteilzeitgesetz) was enacted in mid-1990s and was suspended in 2009. Most CLAs on early retirement based on this law were not renewed. This partial retirement was realized by halving older workers' working time (either via part-time work or early retirement). The employer paid 50% of the previous full-time income and the state government provided the remaining 50% to the employers,

up to five years. Combined with the ERA being at age 60, this requirement meant that the CLA early retirement option applied most directly to employees age 55 and older (Trampusch, 2005). Age 55, and to a lesser extent age 56, was a common cutoff used in CLAs.

3.4 Data

We use rich administrative data from the German Social Security system, assembled by the Institute for Employment Research (IAB) into the Integrated Employment Biographies data file (IEB) (see also Card et al., 2013; Jäger, 2016; Schmieder et al., 2012). This data contains information on all employment periods covered by social security and on all periods of UI receipt between the years 1975 and 2013. The employment information covers approximately 80% of the regular workforce, with the self-employed and civil servants being the most common exceptions. The data on UI receipt stem from administrative UI records and contain information on the exact duration of UI-receipt and the amount of daily benefits.

This version of the paper focuses on men, since early retirement rules differ for men and women. Results for women are available upon request and yield a qualitatively similar picture. We select all male UI-entries between 1980 and 2010 who qualify for their age-specific maximum PBD based on their working histories. This gives us a five year window before the first year in the data (1975) and a three year window after the last (2013), allowing us to calculate UI eligibility for all individuals and unemployment durations for up to three years after UI entry. Since some of the requirements for maximum PBD eligibility, such as the duration over which claims could be accumulated, changed over the study period, the restrictions set on this duration differ slightly over time. We summarize these restrictions in Appendix Table D.4. Additionally, we exclude mining and steel construction from our analysis, since both sectors are known to have specific early-retirement rules for at least some of the periods. For other specific subgroups which face some, but less clear or pronounced early retirement rules we do not exclude cases a priori, but address them throughout the analysis. For the selected individuals, we construct detailed biographical information such as experience tenure or past exposure to unemployment.

but only under the condition that the vacancy was replaced by an unemployed person or a freshly trained apprentice. In addition, the government supported this early retirement option by topping up the pension contribution of the workers who entered early retirement.

4 Reduced Form Evidence

4.1 Graphical Evidence by Cohorts

This Section documents the behavior of older individuals entering UI over three decades. We present evidence of sizable extensive margin UI responses at the bridge-to-retirement kink and show that UI inflows react to UI and retirement policy changes. Specifically, we document spike in UI inflows at each bridge-to-retirement age: at 59 when the ERA was 60 and maximum PBD was 1, at 58 when maximum PBD was extended to 2, and at age 57 and 4 months when maximum PBD was extended to 32 months. As the NRA increases this bunching is reduced, and eventually as the ERA increases it dissipates. The next Section quantifies the bunching mass and estimates regression discontinuities at each of the PBD age cutoffs to quantify responses on the intensive margin.

We will also see evidence of clear bunching at various other thresholds, not all of which corresponds to kinks or notches in our stylized budget sets. For example, beginning with the 1929 cohort, we see bunching into UI entries at age 55. While some of this could be round number bunching or bunching at reference points, much of this is driven by specific collective labor agreements at the firm or sectoral level that specified retirement packages and ages. Indeed, this type of bunching is almost entirely absent in the years leading up to and including 1982, consistent with the timing of the first major CLAs specifying retirement ages (see Trampusch et al., 2010). Our sample drops the mining and steel sectors which have clearly defined CLAs, but inevitably picks up other sectors and firms with CLAs.

As mentioned in the section 3, the importance of these CLAs fades throughout the late 90s and early 2000s. Generally, the bunching at the kink into retirement exceeds bunching at these alternative thresholds. Nevertheless, the data points to an active role for firms, together with workers, in governing responses to UI extensions. Regardless of the source, it will be clear that changes in UI durations generate extensive margin responses that should be taken into account when designing policy.

Figure 3 shows the number of individuals entering UI by age for 6 select cohorts in our sample, each chosen to represent a different institutional regime. We opt to display these annual cohort-level graphs to keep retirement rules constant within-figure. In practice the retirement rules vary by month of birth (see Appendix Table D.3), but fixing year of birth is a good approximation and increases sample sizes. When constructing cohort-by-cohort figures, the state of the economy is not fixed at one point in time, so we also plot the prevailing unemployment rate at the time for reference. Furthermore, since UI rules changed over time (and not by cohort) UI entrants at different ages in the same cohort can have different PBDs (see Appendix Table D.1).[23] Graphs of UI entrants by calendar year offer different trade-offs but ultimately yield a similar picture and are available upon request. Figures 4 and 5 complement Figure 3 by plotting mean UI benefit receipt duration and mean non-employment duration (capped at 36 months) by age for each cohort. We now discuss each cohort in turn.

Benchmark: 1924 Cohort Figure 3 Panel (a) shows UI inflows for the 1924 cohort. Note that UI entries pre-age 59 track the official West German male unemployment rate at the time (the dashed line).[24] When this cohort was less than 61 years old, their PBD was 12 months.[25] Cohorts born before 1937, including this cohort, could retire early and without penalties at age 60 following a year of unemployment insurance. Since the maximum PBD was 12 months for this cohort, the 'bridge-to-retirement' pathway, in which individuals will be covered by UI or pensions without gaps, has individuals entering unemployment at age 59. This is indicated by the red and blue shaded areas under the figure (see also Figure 2 panel (a)).

We observe clear bunching in UI entries at age 59, precisely the age at which individuals can transition into retirement immediately following UI expiration. There is no comparable bunching elsewhere. Figure 4 panel (a) shows average UI benefit duration for the individuals in Figure 3 Panel (a). The average UI benefit receipt of 11.8 months around age 59 is close to the maximum PBD of 12 months, supporting the idea that entrants are predominantly using UI as a bridge-to-retirement. UI durations increase at older ages, in step with the UI reforms in those years, and exhibit declining patterns before the retirement age for the long-term insured at 63 and the standard retirement age at 65. Figure 5 panel (a) plots average non-employment duration (capped at 36 months) for the individuals in Figure 3 Panel (a). This peaks at age 59, averaging 35.4 months, again supporting the idea that most entrants at this age retire. Together, this is clear evidence of sizable, extensive margin responses to UI policy. This view is reinforced below, where we examine UI entries for later cohorts facing longer PBDs and hence kinks at different, earlier ages.

1929 Cohort Figure 3 Panel (b) shows UI entries for the 1929 cohort. This cohort faces the same potential retirement ages as the 1924 cohort, but has longer PBDs in their late

[23]From the perspective of a single cohort, UI can change for two reasons. It can change at known age cutoffs (represented by the dashed red lines in Figures 3 – 5), for example the 1941 cohort would have turned 54 in 1995, amid a UI policy regime that had maximum PBD of 26 months for workers entering UI below age 54 and 32 for workers above age 54. These age-cutoffs would be known to the individual years before turning 54. Alternatively, UI can change for workers above a certain age in a cohort due to a policy change in the future. These policy changes would not necessarily be know to the individual in advance.

[24]We seasonally adjust reported UI rates using X-13 ARIMA - SEATS.

[25]On January 1 1985, UI was extended to 18 months. This means that when a person born Dec 31st 1924 turns 60 and a day, they would be eligible for 18 months of PBD. By age 61, everyone in the 1924 cohort is eligible for 18 months of PBD. This 'entire-cohort eligibility' point is indicated by the change in the lower, grey-shaded bars, which also show the later UI reforms.

50s. Specifically, those who enter UI at age 58 have 24 months maximum PBD. This shifts the 'bridge-to-retirement' age to 58, and indeed, we see extensive bunching at around age 58, while we continue to note some excess mass at 59.

This figure also clearly shows bunching in UI entries at other, non-kink points, particularly at age 55 and 57. As discussed, these likely represent firm-specific collective bargaining agreements to release or buy out workers once they turn 55. This also suggests that the bunching at the bridge-to-retirement age is driven by joint decisions between firms and workers.

Panel (b) of Figures 4 and 5 show average UI benefit receipt and average non-employment durations for this cohort. Figures 4 (b) reveals that average UI duration at the kink is 23.0 months, close to the full 24 around the 58 cutoff. There are also clear spikes in UI durations at age 55 and 57, mirroring bunching in UI entries at those ages. Figure 4 Panel (c) shows that average non-employment duration at the bridge-to-retirement age reaches 34.3 months.

1935 Cohort The 1935 cohort continues to face the same potential retirement ages as the prior cohorts, but even more generous UI. Workers entering UI at or after age 54 had a maximum PBD of 32 months. Accordingly, Figure 3 Panel (c) shows that UI entries exhibit strong bunching at precisely age 57 and 4 months, or 32 months before the early retirement age of 60. We continue to see some excess bunching at age 58 and 59, as well as some at 55 and 56. Panel (c) of Figures 4 and 5 confirms once again that people entering at the bridge-to-retirement age take UI for close to the maximum duration (29.7 months) and have a 35.3 month average capped non-employment duration. These figures also show discrete jumps at age 54 and 55 in average UI duration and non-employment duration. The jump at age 54 is consistent with the July 1987 reform that extended maximum PBD from 24 to 32 months (26 months) for workers above 54 (between 49 and 54). The jump at 55 in both figures continues to reflect the fact that layoffs after age 55 differ in composition and likely reflect firm-level CLAs.

1941 Cohort This is the first cohort for which retirement rules change. The 1941 Cohort could still retire at age 60 following a year of unemployment, but a 1992 reform introduced actuarial adjustments for retirement before age 65. These were introduced gradually by month and year of birth for cohorts born between January 1937 and December 1942, resulting in an approximate 18% penalty for anyone in the 1941 cohort retiring at 60. The maximum PBD remained at 32 months for workers above age 54. Figure 3 Panel (d) reveals that we continue to see bunching at age 57 and 4 months, but it is now more muted relative to entries below this age. Moreover, consistent with the larger penalties, we see in Figure 4 panel (d) that average UI benefit duration no longer reaches 32 months at this bridge-to-retirement age, but instead averages just 25.3 months. Similarly, average non-employment durations are also lower, around 30.5 months, suggesting that some workers are returning to work instead of retiring at the penalized ERA. Spikes at age 55 and 56 continue to be visible in entries, UI receipt, and duration. Interestingly, this figure displays what looks very much like a discrete jump in the level of UI entries after age 55.

Additionally, this cohort faced a stable PBD schedule in their 50s, with a known age cutoff at 54 (for which maximum PBD jumped from 26 to 32 months). We see some bunching at this cutoff, which could arise from people expecting long unemployment timing their entry into UI or from those considering early transitions to retirement. The sorting around this age cutoff poses a challenge to standard RD estimates of the effects of PBD extensions on non-employment duration, as we discuss further below.

1949 Cohort The 1949 cohort faced both reduced PBD if retiring at later ages and a stricter retirement law. Individuals born in 1949 could no longer retire early via unemployment at 60, but instead could only draw pensions at age 63 at the earliest. They had to wait until age 65 to draw pensions without actuarial adjustments (7.2% for retiring at 63). Figure 3 Panel (e) shows some bunching at 61, consistent with an early retirement age of 63 and the 2 years maximum allowable PBD, but it is not extensive. Importantly, now that the bridge-to-retirement at 60 has been removed, we now no longer see bunching between ages 57 and 59. We continue to see some age-55 bunching. Panel (e) in Figures 4 and 5 shows that average UI durations at the new bridge-to-retirement reach 13.7 months, well below 24, and non-employment durations average 33.2 months.

1952 Cohort This cohort is no longer allowed to retire early via unemployment, although if they are eligible for the long-term, old age pension, they could retire at age 63. Unfortunately, we run out of data past age 59 (as we need 3 years post-2010 to calculate non-employment durations). Nevertheless, the distribution of UI entries continues to look relatively smooth. The 1952 cohort would have known about the age 58 PBD cutoff extending maximum PBD from 18 months to 24 months starting in 2008 (i.e. when they turn 56). As with the 1941 cohort at age 54, we see evidence of sorting into UI to take advantage of this UI extension. We continue to see some bunching at age 55.

Overall, we observe clear bunching into UI at the bridge-to-retirement age. The bunching mass responds to UI extensions. We have also seen evidence of sorting into UI at earlier age cutoffs where PBDs are extended discontinuously, including age 54 and 58. Bunching at other points in the distribution related to CLAs, suggest that firms play an important role. While we cannot easily identify the extent to which responses come from workers or firms, it is clear that a full accounting of the effects of UI extensions on non-employment need to take into account this extensive margin to avoid downward bias.

4.2 Estimating Intensive Margin Responses using Regression Discontinuity Estimators

We have documented large, extensive margin UI responses. Individuals also respond to UI along the intensive-margin. Here, we obtain reduced-form, intensive-margin estimates of the non-employment effects of UI extensions using a Regression Discontinuity Design (RD) that exploit discontinuous PBD increases at numerous age cutoffs (as in Schmieder et al. (2012)). Figure 1 displays each age cutoff. These estimates require that there is no sorting into UI around age cutoffs. This is satisfied at younger ages, but not at the oldest ages.

We estimate RDs pooling all years under each UI regime, starting with the 1987-1999 period (see Appendix Table D.1).[26] This gives us 12 age cutoffs at which we can estimate the non-employment effect of UI extensions. In practice the age cutoffs provide between an extra 3–6 months of PBD, and we will divide each estimate by the number of months PBD was extended to get the marginal non-employment effects of an extra month of PBD.

At each age cutoff we estimate the following RDD specification:

$$y_i = \delta\, \mathbf{1}(a_i \geq A)\, \Delta PBD + f(a_i) + X_i\beta + \varepsilon_i \tag{2}$$

y_i is non-employment duration (capped at 36) for individual i, a_i is the age at time of UI entry (measured on the daily level) and $\mathbf{1}(a_i \geq A)$ is an indicator function which equals one when individuals age is above the age cutoff A where benefits are extended discontinuously by ΔPBD months. In this specification, δ measures the effect of a one month increase in PBD. The function $f(a_i)$ is set to be a linear function with different slopes on each side of the cutoff in the baseline specification. X_i is a vector of additional controls. We use a local polynomial regression with rectangular kernel and cluster standard errors on the daily level. We set the bandwidth to two years, but restrict it to one year on the right side of the 49 and 54 age cutoffs during the 1987-1999 period due to the presence of other discontinuities at 50 and 55.

This specification is well-identified for the age cutoffs where UI entries and other pre-determined outcomes are smooth around the cutoff. This is case for all of the younger ages (Schmieder et al. (2012)). Sorting at the cutoff is a concern for some of the older cutoffs (this can be seen, for example, at the age 54 cutoff in Figure 3 panel (d)). The degree of sorting varies between cutoffs and is usually most pronounced within the first 1 to 2 months around the cutoff. We apply an imperfect solution to address this concern by excluding 2 months on each side of the cutoff – the donut hole – in all our regressions. Second, we add detailed

[26] We omit the period 1986-1987 due to it being a short transition period. There is no first stage in this period.

individual controls such as education, tenure and other pre-unemployment characteristics to help address some of this selection.[27]

In Figure 6 we present figures that correspond to 2 of our 12 RDs estimates by way of example. These figures plot mean non-employment durations (capped at 36 months) by age around the age 49 cutoff for the 1987-1999 period and around the age 52 cutoff for the 1999-2006 period. Appendix Figure D.1 plots UI inflows around the age cutoffs during those two periods.

The results of these and the other 10 RD estimations are depicted in Figure 7 and reported in Table 1. The results at the cutoffs at or below age 50 are relatively insensitive to the inclusion of controls and average 0.089 meaning an extra month of PBD results in an extra 0.089 months – or about 3 days – of non-employment (as in Schmieder et al. (2012)).[28] We obtain a similar estimate, if slightly higher, at age 52 and 54. Results for the oldest cutoffs are biased upwards by sorting. Given the sorting it is difficult to draw strong conclusions, but the evidence suggests that the intensive-margin, non-employment effect of UI extensions may increase slightly for workers in their mid-50s relative to workers in their 40s. While these patterns alone are of interest, the policy relevance of these results is compounded by the presence of extensive margin effects.

4.3 Quantifying Total Responses: A Back of The Envelope Calculation

The graphical evidence from the previous sections shows that workers respond to the potential UI benefit duration both on the intensive and the extensive margin. While this reduced form evidence is clear, it is hard to quantify the total effect of a UI extension and the relative importance of the intensive and extensive margin responses. For the intensive margin effect alone — which has been estimated in many other contexts and is a well-defined behavioral elasticity — one could use a linear extrapolation using the marginal effects $\frac{\partial D}{\partial P}$, but it is not clear that such a linear prediction will do well close to the retirement age where interaction effects become important (e.g. the value of finding a job will go down the closer a person is to retirement and thus the shorter the expected duration of holding a job). For the extensive margin response, it is clear that changes in PBD would shift the bunching mass (as we see empirically), but it is not clear how many individuals would be affected by this, as not everyone

[27]We are also exploring using Gerard et al. (2015)'s method, which bounds RDDs estimates in situations where the running variable is manipulated.

[28]Note that there are differences in sample restrictions between this paper and that in that here we restrict to men who tend to be less responsive to UI, omit some industries, and have different pre-unemployment sample restrictions.

would find it optimal to move to the new kink point in the lifetime budget set.[29]

Furthermore, there is no simple way to credibly combine bunching and regression discontinuity estimates to arrive at the total non-employment effect of UI changes. If one were to, for example, specify a two-type model where one type always times their entry into retirement via UI and the other type never does but can become unemployed randomly, one could make progress. But the assumption that individuals can be so easily categorized is unrealistic. Moreover, modeling this behavior purely from the individual perspective is at odds with the evident role that firms play in both generating un-desired layoffs and agreed-upon exits. Instead, in Section 5, we develop a life cycle model of labor supply, job search and retirement decisions.

Before proceeding, we perform a preliminary calculation to highlight the relative contribution of each margin under stark assumptions. Taking the 1935 cohort as a baseline, we consider the effect of reducing maximum unemployment duration by 8 months (from 32 months to 24 months) for workers aged 50 to 60. We assume that this would shift the entire bunching mass at the 57.33 kink rightward to age 58, without otherwise affecting UI entries. This is somewhat consistent with the fact that the bunching mass is relatively similar for the 1929 and 1935 cohorts (under the polynomial counter-factual).[30] We consider two scenarios for how the bunching mass might respond to this change: in the first, we assume each individual belonging to the entire bunching mass from 57.33 to 60 delays entry by the full 8 months, in the second we assume only individuals in the reduced bunching mass between 57.33 and 58 delay entry, and the rest do not move at all.[31]

We imagine that all non-bunchers (aged 50 to 60) are fired involuntarily and opt to search for jobs. We ignore the idea that involuntary exits closer to the kink might switch from searching to not-searching. We take the counter-factual polynomial estimate to represent the number of people aged 50 to 60 who respond to this intensive margin estimate. Further, we assume that, conditional on unemployment, all these workers respond to a 1 month UI extension by increasing their non-employment durations by 0.089, the average RD estimate for workers aged 40-50.

Under this set up, the intensive margin effect acting on workers aged 50 to 60 is a reduction

[29]In Appendix C.2, we obtain measures of the labor supply elasticities using bunching estimators, as in Brown (2013), but these estimates are based on a simple static life-cyle model that ignores the possible interactions of responding along the intensive and extensive margin and is difficult to reconcile with inflows into unemployment that are not driven by voluntary quits.

[30]Figure 3 shows the estimated amount of bunching at each retirement-via-UI kink. Table C.1 contains our estimates of the bunching mass, obtained using standard bunching approaches as described in Appendix C.2.

[31]In practice the conservative approach yields a distribution closer to the actual 1929 distribution. It nevertheless over-estimates bunching between 58 and 60. In addition, the 1935 counter-factual under-estimates the mass of people in UI between 52 and 56 for the 1929 cohort. Together, this means that we may be over-estimating the importance of the extensive-margin effect.

of 0.712 months of non-employment. The counter-factual accounts for 65% of workers, so the expected intensive margin effect is -0.463 months. The full bunching mass accounts for 35% of workers, so the expected extensive margin effect is -2.795 months. The more conservative bunching mass estimate (in which only those between 57.33 and 58 respond to the change) accounts for 16% of workers, making the expected extensive margin effect -1.277 months. The total effect of time out of work for this 8 month reduction is thus -3.258, or -1.740 using the conservative version. That is, the total non-employment effect is between 2.44 and 4.58 times as large in magnitude as the pure intensive margin. While a number of things would deflate this estimate – including using a larger estimate of the true intensive-margin effect, using the 1929 counter-factual instead of the 1935 one, and assuming that some bunchers respond by less than the full 8 months – it is apparent that the extensive margin plays a non-negligible role.

5 Dynamic Labor Supply Model

In this section, we turn to using the reduced form moments to estimate a life cycle model of labor supply, job search and retirement decisions using a methods of moments estimator. Our setting strikes us particularly promising for such a structural approach for a number of reasons: first, the setting provides clear economic incentives and interactions between incentives that can be captured in a fairly standard model. Second, we have clean policy variation and empirical moments to identify the key parameters of the model and that discipline the estimation approach. Third, there are several reforms, that allow us to test the out of sample performance of our model. And fourth, there is a clear value added to having a model for doing counterfactual and out of sample analysis, as well as to quantify the economic importance of our parameter estimates. For example a model will make it straightforward to predict how the intensive and extensive margin response will determine the response to a policy reform and what the contribution of each channel is. It will also allow us to quantify how much of the changes in retirement age over the past decades in Germany can be explained by the slow reduction of UI generosity for older workers over this time period.

5.1 Set-up

Our model is similar to a basic labor supply model of retirement decisions, but with two key extensions. First, we incorporate endogenous job loss into the model in a way that explicitly allows for a role of firms. This strikes us as important given the evidence that firms are key for the extent of bunching in the previous section and the general importance of CLAs. Second, we model unemployment and job search in more detail than past life cycle models.

While previous research with life cycle models typically just allows for a single (or sometimes two) states of unemployment with no dynamics, we model unemployment as a fully dynamic process that in particular allows us to capture the duration of UI benefits and the response to their exhaustion in a natural way.[32] This also has the added benefit that our parameter estimates for the job search part can be compared with previous estimates of job search models (e.g. Paserman, 2008; DellaVigna et al., 2017).

We assume that there are firms and workers. Firms hire a single worker who then produces output p_t per period. If workers are not working they are non-employed N, which can either mean that they are unemployed U or out of the labor force O. We assume that once a worker drops out of the labor force she will never return to the labor force. If a worker is employed, she and the firm bargain over the wage w_t in each period that the job continues to exist. An important state variable in our model is the total unemployment duration of a worker d^U. In practice we will estimate our model starting at age 50, so that d^U will be the duration in unemployment since then. To keep the state space manageable, we also assume that workers initially are eligible to the maximum benefit duration but do not reaccumulate benefit eligibility if they are reemployed after losing a job. Under this assumption d^U is sufficient to both calculate remaining UI benefit durations for each individual as well as the pension of an individual if the person retires.[33]

We can therefore write the value functions for the firm and worker as functions ofp_t and d^U. d^U is deterministic, while p_t is uncertain.

Wages will be determined by Nash bargaining, so we can write this as $w_t(p_t, d^U)$. The value function of the firm will be denoted as: $J_t(p_t, d^U)$, the value of employment for a worker is $V_t^E(p_t, d^U)$, the value of non-employment is given as: $V_t^N(d^U)$.

5.2 Value functions and job search

Firms employ a single worker who produces p_t in each period and pay this worker a wage w_t. Firms can always decide to close (lay off the worker) if the value falls below 0 or workers can quit, in which case the value of the firm falls to 0.

[32] Other structural life cycle papers (Haan and Prowse, 2010; García-Pérez and Sánchez-Martín, 2015) typically assume workers receive UI forever or model UI as a Markov process with a fixed transition probability to exhaustion.

[33] A full accounting of the benefit eligibility in the presence of multiple unemployment spells would require to separately keep track of d^U as well as the remaining benefit duration in each unemployment spell and employment duration in each employment spell. This quickly becomes computationally very challenging due to the curse of dimensionality. As long as repeated unemployment spells with long in-between employment spells are rare, which they are in practice, our approach is only a very minor simplification that vastly reduces the computational complexity.

The value function of the firm as a function of p_t is given as:

$$J_t(p_t, d^U) = p_t - w_t(p_t, d^U) + \delta_t E\left[\max\left\{J_{t+1}(p_{t+1}, d^U), 0\right\}\right] \tag{3}$$

where δ_t is the firm's discount factor and the expectation is taken over p_{t+1}.

Employed workers earn wage w_t and experience disutility from working ϕ_t. The value of employment is V_t^E will also depend on the productivity in the present period since it affects the wage and whether a job continues and can be written as

$$V_t^E(p_t, d^U) = u(w_t(p_t, d^U)) - \phi_t + \beta\, E\left[\max\left\{V_{t+1}^E(p_{t+1}, d^U), V_{t+1}^N(d^U)\right\}\right] \tag{4}$$

where the expectation is taken over p_{t+1} and V_t^N is the value of non-employment in period t.

Individuals who become non-employed at time t can choose to go into unemployment or to drop out of the labor force

$$V_t^N(d^U) = \max\left\{V_t^U(d^U), V_t^O(d^U)\right\} \tag{5}$$

A worker who drops out of the labor force receives home production y_h and eventually, when reaching the early retirement age, can claim pension benefits according to the institutional rules. The expected value from this income path is given as $V_t^O(d^U)$.

While an individual still has benefits remaining ($d^U < P$), she receives UI benefits b and she also receives some level of home production y^h. An unemployed individual searches for a job and chooses an optimal level of search effort s which is normalized to the probability of finding a job. Generating search effort comes at a cost $\psi(s)$ which is increasing and convex and may depend on age. Finally, whether or not an individual receives a job offer she can decide to retire at the end of the period. If she remains unemployed d^Uincreases by one period. The value function is thus:

$$\begin{aligned} V_t^U(d^U) \;&=\; u(b\,1_{d^U<P} + y^h) + \max_s \left\{-\psi_t(s) + \beta s \max\left[V_{t+1}^E(d^U+1), V_{t+1}^O(d^U+1)\right]\right. \\ &\qquad \left. +\beta(1-s)\max\left[V_{t+1}^U(d^U+1), V_{t+1}^O(d^U+1)\right]\right\} \end{aligned} \tag{6}$$

Individuals choose search effort so that the marginal return to search equals the marginal cost up to the constraint that$s \leq 1$. For an interior solution we have the first order condition for the optimal level of search effort s^*:

$$\psi'(s^*) = \beta \max\left[V_{t+1}^E(d^U+1), V_{t+1}^O(d^U+1)\right] - \beta \max\left[V_{t+1}^U(d^U+1), V_{t+1}^O(d^U+1)\right]$$

5.3 Wage bargaining and job loss

We assume wages are determined by Nash bargaining as long as there is some level of the wage so that both workers and firms are better off than their outside option:

$$w_t = argmax \left(V_t^E(d^U) - V_t^N(d^U)\right)^\kappa \left(J_t(d^U)\right)^{1-\kappa}$$

Employment relationships end when there is no more surplus, that is there is no wage such that:

$$J_t(p_t, d^U) \geq 0 \text{ and } V_t^E(p_t, d^U) - V_t^N(d^U) \geq 0 \tag{7}$$

We also assume that the wage setting process is such that if there is no wage such that condition(7) holds, then the wage will be such that both $J_t(p_t, d^U) < 0$ and $V_t^E(p_t, d^U) - V_t^N(d^U) < 0$, so that both sides prefer to end the relationship.

This makes it clear that jobs end endogenously and that the probability of jobs ending is both determined by the distribution of p_t as well as the evolution of the outside option for the worker that determines the value of non-employment $V_t^N(d^U)$. For example an increase in retirement benefits will push up $V_t^N(d^U)$ for workers close to the retirement age and can increase the rate of jobs ending. Similarly an extension in UI benefits will raise the value of $V_t^N(d^U)$, for all workers but in particular for those in the range where they are able to bridge the full time to early retirement with UI with the extension.

Note that this formulation explicitly does not take a stance on whether a job ending is initiated by the firm (a layoff) or the worker (a quit), but instead is the outcome of essentially a bargaining process. Though of course a job ending after a bad realization of p_t will seem like a lay-off to a worker. This seems realistic given the prevalence of CLAs that guided job reductions and early retirements in Germany over this period.

5.4 Functional form assumptions for estimation

For simplicity, and this is something we will relax in future iterations, we also assume that productivity takes on only two values $p_t \in \{p_H, p_L\}$. Furthermore all jobs start such that$p_t = p_H$ and in each period there is a probability λ that productivity permanently drops to $p_t = p_L$.

Following DellaVigna et al. (2017), we also assume that $\psi(s) = k\frac{s^{1+\gamma}}{1+\gamma}$ so that $\psi'(s) = ks^\gamma$ and the FOC for search effort becomes:

$$s^* = \left(\frac{1}{k}\left(\max\left[V_{t+1}^E(d^U+1), V_{t+1}^O(d^U+1)\right] - \max\left[V_{t+1}^U(d^U+1), V_{t+1}^O(d^U+1)\right]\right)\right)^{\frac{1}{\gamma}}$$

Finally we assume that workers have log utility $u(.) = \ln(.)$ and that the disutility of work can be written as: $\phi(t) = (t/a)^{1/e}$, where a is a parameter capturing ability and e describes how fast the disutility of working is increasing with age. This parameter essentially corresponds to the labor supply elasticity in a static retirement model, such as Brown (2013).

The model has many attractive features. It can clearly capture intensive margin responses through the job search part of the model, where individuals would respond to benefit extensions by adjusting their search effort. The intensive margin response would mainly be captured by the curvature of the job search function γ. The model can also capture extensive margin responses as job endings will respond to the outside option as determined by the institutional setting of the UI and pension system.

6 Model Estimation and Simulations

Note, while we have programmed the basic version of the model and simulated it, we have not yet estimated it. The discussion below is preliminary and based on these simple simulations.

6.1 Moments and Estimation Strategy

We estimate the model from Section 5 structurally, using a minimum distance estimator to match the empirical reduced form moments from Section 4. Denote as ξ the parameters of the structural model. Furthermore, let $m(\xi)$ be the vector of moments predicted by the model as a function of the parameters ξ, and by $\hat{m}$ the vector of observed moments.

The moments $m(\xi)$ we use for matching are the monthly transition probabilities of workers between age 50 and 65 between employment and non-employment.[34] To the extent that most workers who enter non-employment below the early retirement age then enter UI, this essentially corresponds to the UI inflow figures in Section 4. These transition probabilities mainly pin down the extensive margin response. In order to pin down the intensive margin response and to guarantee that the model makes reasonable predictions regarding the intensive responses, we use the RD estimates of $\frac{\partial D}{\partial P}$ as additional moments.

The estimator chooses the parameters $\hat{\xi}$ that minimize the distance $(m(\xi) - \hat{m})' W (m(\xi) - \hat{m})$ where W is a weighting matrix. Since each transition probability is calculated using the full cohort and thus the estimate of the transition probability has a small variance that is approximately constant by age, we simply use a diagonal weighting matrix with weights ω_t on the diagonal. For the intensive margin RD moments, we use a larger weight (to be determined) since this is a causal estimate that we have significant confidence in given the research in this

[34]While we observe UI receipt, we cannot distinguish unemployment from OLF after UI benefits are exhausted. For this reason we simply distinguish between non-employment and employment, which we can easily generate from the model predictions by pooling the unemployed and OLF states.

paper and many other well identified estimates from the literature and we want to make sure our fitted model generates realistic predictions for intensive margin responses. As a baseline we estimate the model using 2 cohorts with different policy regimes: the 1924 and the 1929 birth cohorts.

With a deterministic p_t and a single type, the model above would generate a unique age for when jobs end and workers drop out of the labor force depending on the parameters of the model. In order to generate bunching in UI entries, we need essentially a distribution of individual employment paths that are subject to different shocks or different preference parameters. There are various alternatives for doing this. One obvious one is to allow for randomness in p_t in a more flexible way than the binary distribution assumed above. If p_t is random and follows for example a log-normal distribution, then as the outside option for the worker is increasing with age (due to approaching ERA and the option of UI as a bridge to retirement), it becomes more likely that a given shock in p_t will push the job into the negative surplus territory and thus leads to a job ending. Since the combination of UI / ERA leads essentially to a jump in the outside option, there should be a much larger number of jobs that end right at the bridging age, thus leading to bunching in UI entries.

An alternative way to allow for bunching is to have different types of workers, for example by allowing for heterogeneity in the disutility of working ϕ_t and then integrating over types. Each type will then have a different preferred retirement age and given a smooth distribution of this heterogeneity the jump in the outside option will lead to bunching. In the preliminary version presented below we use this approach, but we are working on a more flexible distribution of p_t, which we think is very much in the spirit of the exercise.

6.2 Preliminary Simulations

To illustrate our approach, we present some simple simulations of our model, based on parameters that can roughly generate the features we observe in our data.

The simulations are based on the 1924 cohort only, where the institutions are very simple with an ERA of 60 and a maximum of 1 year of UI benefits. We use quarterly transition probabilities and simulate the model on the quarterly level for this exercise. Figure 7 (a) shows the empirical UI inflow rates from Figure 3 (a) collapsed to the quarterly level (the dots). The line shows the simulated UI inflows from our model where we set the institutional parameters to the actual values faced by the 1924 cohort and furthermore set the preference parameters to be: $e = 0.8$, $k = 200,000$, $\gamma = 4$, $\lambda = 0.03$, $\kappa = 1$ and $\beta = \delta = 0.995$. We model heterogeneity of types using the parameter a that determines the level of the disutility of work path. We allow for 100 different values (from an evenly spaced grid between 0.01 and 50) that we then integrate over.

While this simulation is only meant to be illustrative, Figure 7 (a) shows that it can replicate at least the basic feature of bunching of UI entries at the UI-bridge-to-retirement age of 59. Figure 7 (b) shows the predicted UI entry rates from the model with the same set of parameters but changing PBD to 2 years and thus lowering the UI-bridge-to-retirement age to 58 years. The empirical moments are kept the same for comparison. The simulation clearly shows that the spike in UI entries moves to age 58 and is slightly smaller than the previous bunching mass at 59. This is because for younger people retirement is generally less attractive so that they are less likely to be pushed into UI / ERA at an earlier age by a jump in the outside option.

The simulation thus mimics the main features from our cohort level analysis of UI entries, namely that a) there is bunching at the UI-bridge-to-retirement age, b) the bunching mass responds clearly to changes in PBD, c) bunching is smaller at younger ages.

While the parameters used in this simulation are largely arbitrary and for example do not try to match more than one cohort or the intensive margin responses, it illustrates that at least in principle the model is able to capture the main mechanisms for the intensive and extensive margin responses to UI benefit duration for people close to retirement.

7 Conclusion

In this paper we document the labor supply effects of UI benefit extensions for workers approaching retirement age. We show that extensive margin responses, that is UI-induced inflows into non-employment, play an important role, and operate in addition to the standard intensive margin UI responses for younger workers that most of the previous literature has focused on. The combination of intensive and extensive margin responses, as well as voluntary and involuntary inflows into UI, complicates the application of standard non-parametric estimators such as RD designs and bunching estimators, but we argue the discontinuities, kinks, and notches induced by the UI and retirement institutions can still be used to learn about labor supply responses.

Our evidence reveals sizable labor supply responses on both the intensive and extensive margin. A naive, back of the envelope calculation suggests that using standard, intensive-margin estimates of the non-employment effects PBD extensions for workers aged 50 to 60 will severely underestimate the non-employment effects of UI. However, such a calculation draws an unrealistic distinction between types who bunch into UI for retirement and types who only get fired involuntarily and search for a job no matter what. These types are unrealistically separate, with reality surely being more fluid: individuals might voluntarily leave employment but not go into retirement and instead look for a job, and individuals involuntarily fired at later ages might choose to retire after attempting job search.

To address this, we specify and estimate a life-cycle labor supply model that explicitly accounts for transitions between employment, unemployment, and retirement and how they are affected by labor demand as well as the structure of UI benefits and parameters of the old age pension system. Simulations of our model show that we can match the key patterns observed in the data. In ongoing work, we estimate this model using a minimum distance estimator. The data in Figures 3–5 provide compelling moments to match. Additionally we explicitly match our model to our reduced form RD estimates. Once estimated, we can use our model to perform out-of-sample simulations and compare this to what actually happened over our 3 decades of policy variation. We will also be able to answer questions like how much of the stark decrease in employment among older workers in Germany throughout the late 80s and 90s was caused by the UI extensions.

Overall, this paper helps provide policy-relevant advice to better aid policymakers about the true, total non-employment effects of UI and retirement policy changes for older workers in Germany.

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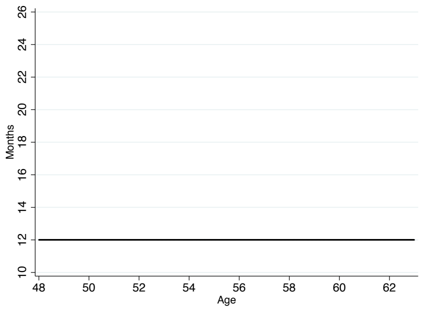
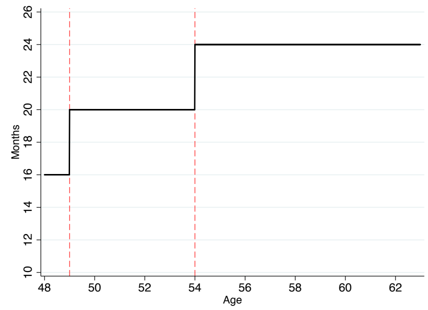
Figures

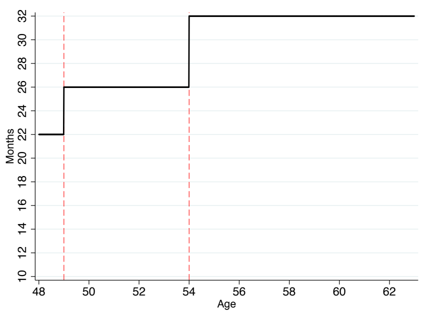
Figure 1: Maximum Potential UI Benefit Durations (PBDs) by Age for Different Time Periods in Germany



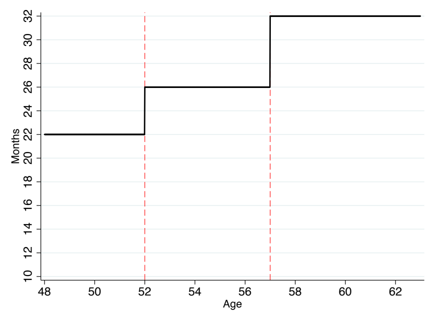
(a) PBD, 1980-1984



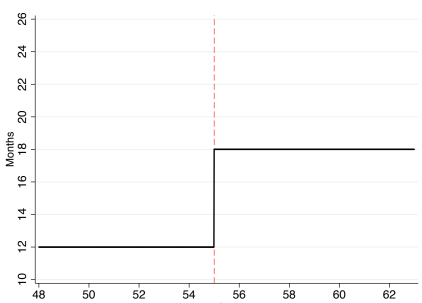
(b) PBD, 1986-June 1987



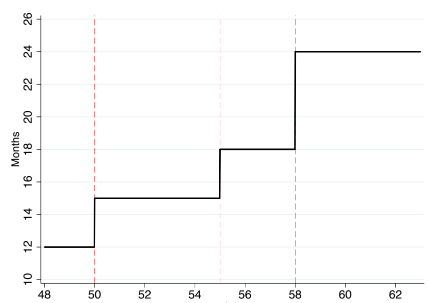
(c) PBD, Jul 1987- Mar 1999



(d) PBD, Apr 1999 - Jan 2006



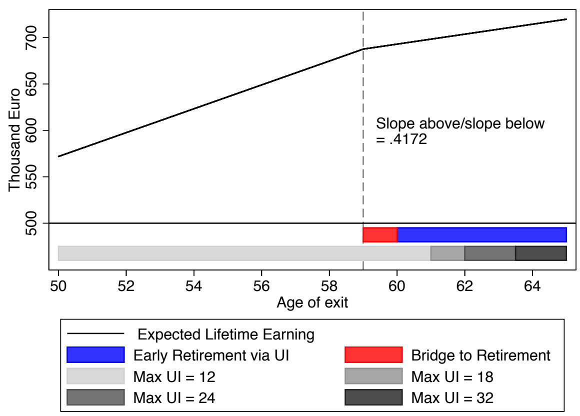
(e) PBD, Feb 2006 - Dec 2007



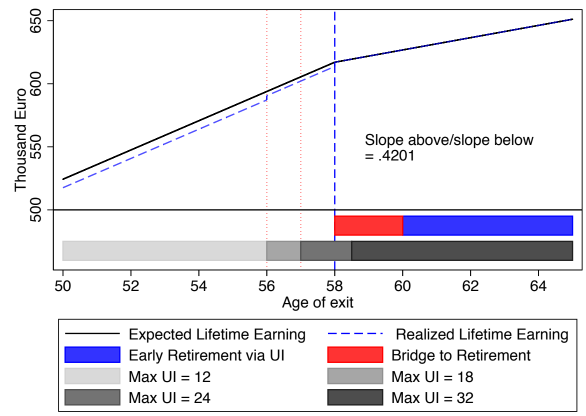
(f) PBD, Jan 2008 - 2010

Notes: This table plots maximum potential benefit durations for unemployment insurance in Germany between 1980 and 2010. We drop the brief 1985 regime in the interest of brevity. Appendix Table D.1 contains more detailed information on each institutional regime, including eligibility requirements and benefit levels.

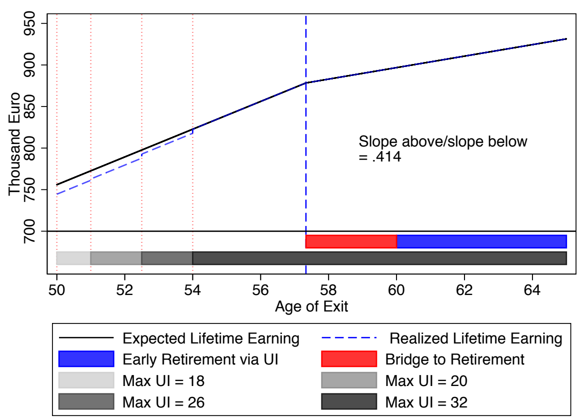
Figure 2: Stylized Budget Sets by Exit Age for Different Cohorts



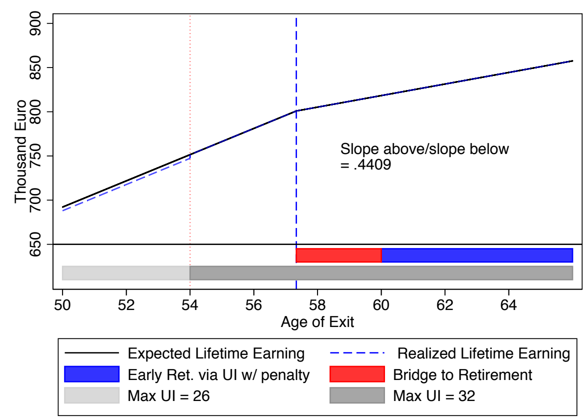
(a) Lifetime Earnings w/ 1 yr PBD, 1924 Cohort



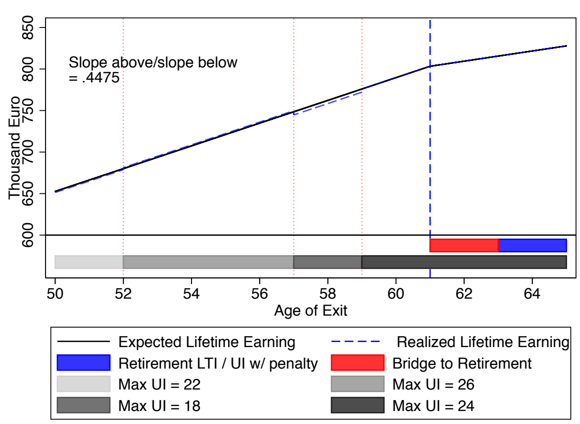
(b) Lifetime Income w/ 2 yr PBD, 1929 Cohort



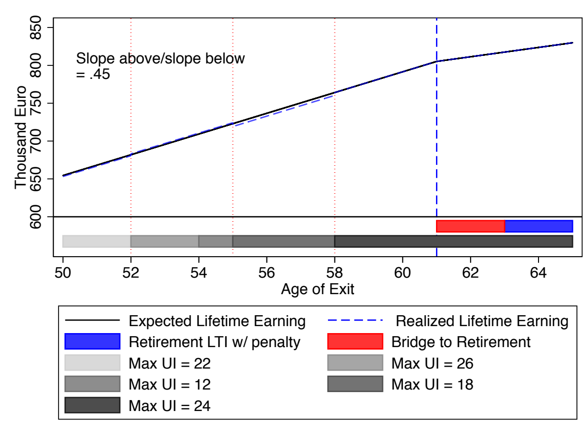
(c) Lifetime Income w/ 32 m PBD, 1935 Cohort



(d) Lifetime Income w/ 32 m PBD, 1941 Cohort



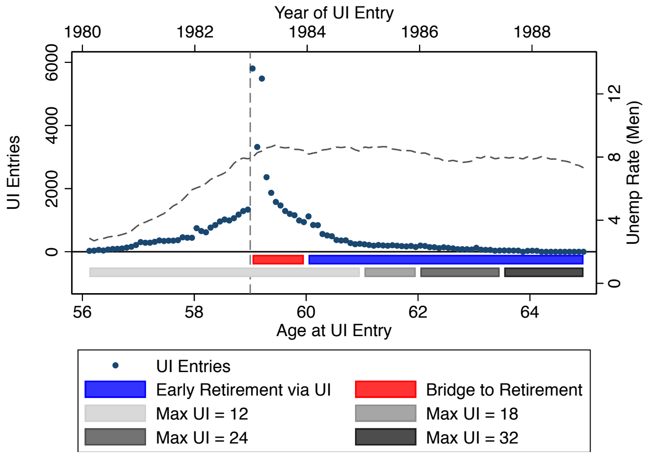
(e) Lifetime Income w/ 2 yr PBD, 1949 Cohort



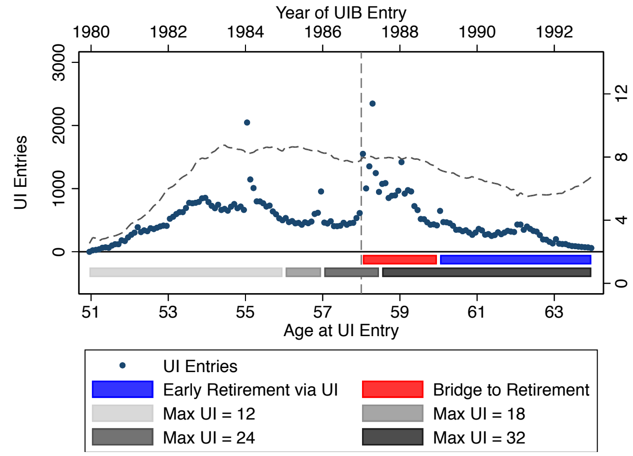
(f) Lifetime Income w/ 2 yr PBD, 1952 Cohort

Notes: These figures contain lifetime budget sets as a function of exit age (into UI and eventually retirement). The ERA is 60 for the first 4 cohorts and 63 for the last two. The NRA (un-penalized retirment age) for retirement via UI is 60 for the first three cohorts, approximately 64 for the second cohort, and 65 for the last cohort. We assume PBD are fixed at 1 year, 2 years, 32 months, 32 months, 2 years, and 2 years respectively.

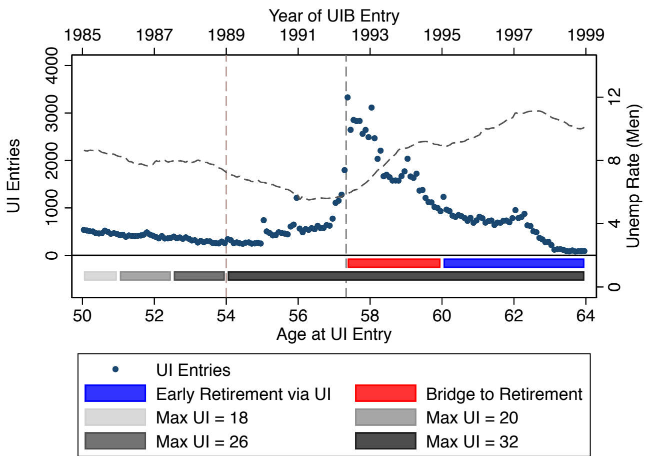
Figure 3: UI Inflows by Age for Different Cohorts in Germany, Men



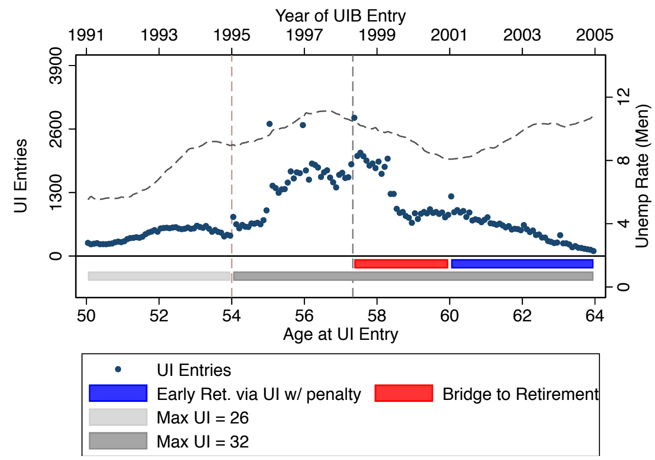
(a) UI Inflows, 1924 Cohort



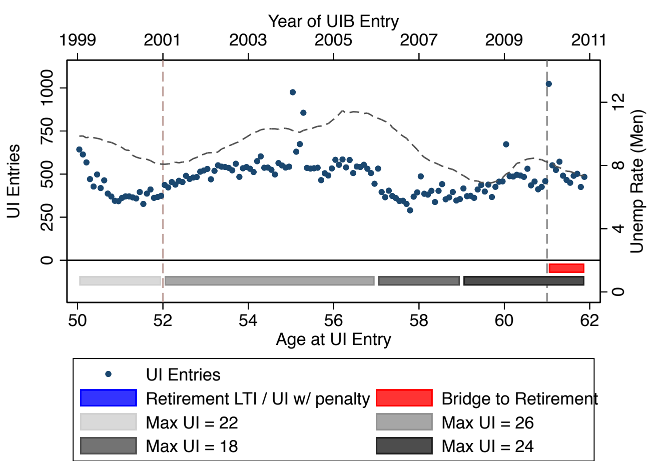
(b) UI Inflows, 1929 Cohort



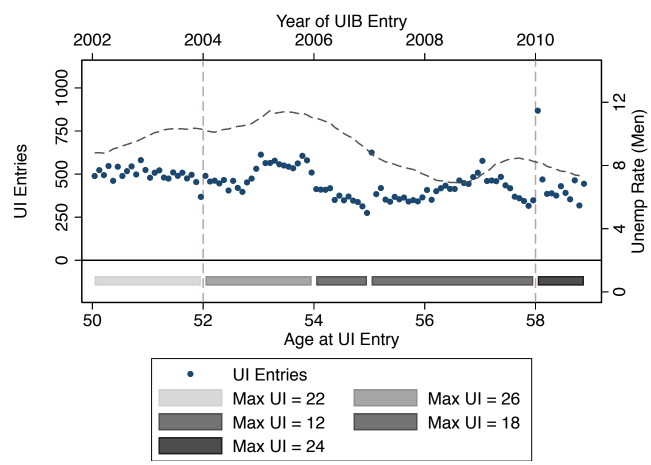
(c) UI Inflows, 1935 Cohort



(d) UI Inflows, 1941 Cohort



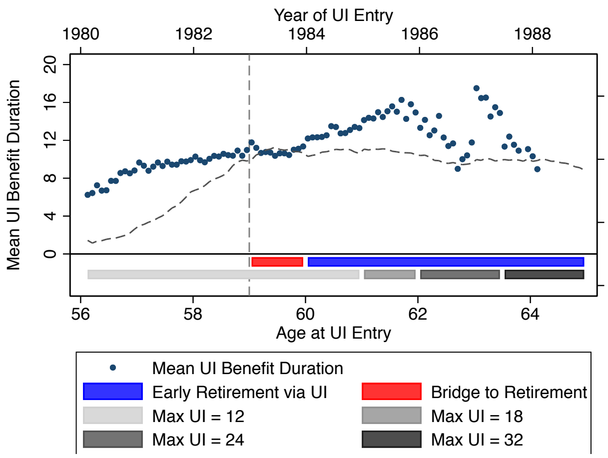
(e) UI Inflows, 1949 Cohort



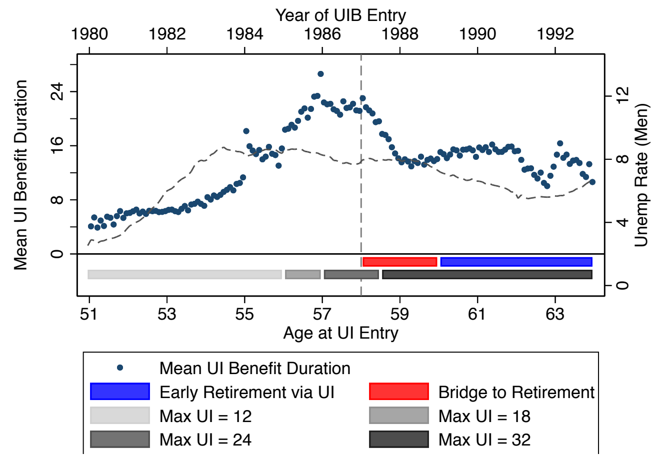
(f) UI Inflows, 1952 Cohort

Notes: These figures plot UI inflows by age for different cohorts of West German Men with full UI eligibility. The red bar under the figure indicates the period over which an individual would receive UI before drawing pension (the blue bar). The different shades of grey represent different maximum PBD eligibility for UI, which can change because of an existing age-cutoff (the red dashed line) or because of an overall UI policy change enacted in that year.

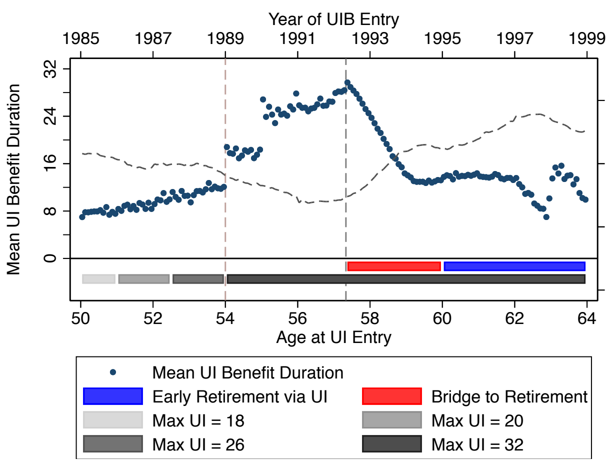
Figure 4: Mean UI Benefit Receipt Duration by Age for Different Cohorts in Germany, Men



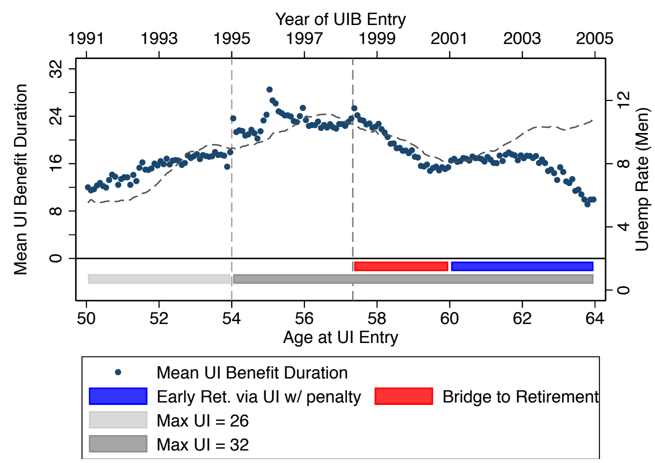
(a) Mean UI Benefit Receipt, 1924 Cohort



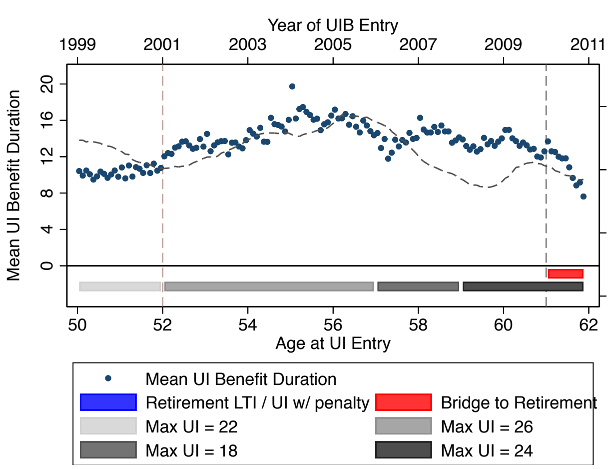
(b) Mean UI Benefit Receipt, 1929 Cohort



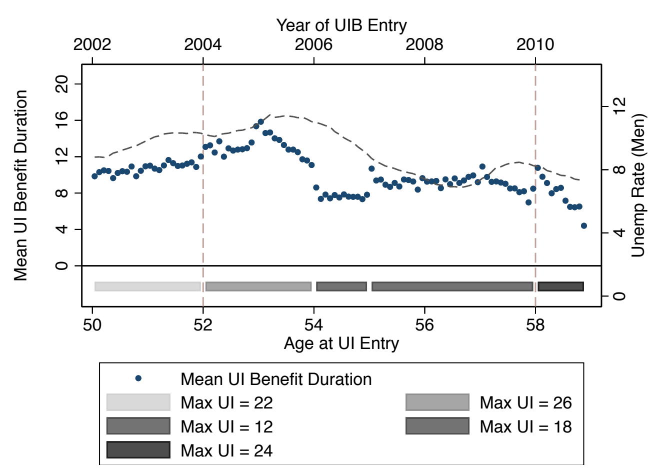
(c) Mean UI Benefit Receipt, 1935 Cohort



(d) Mean UI Benefit Receipt, 1941 Cohort



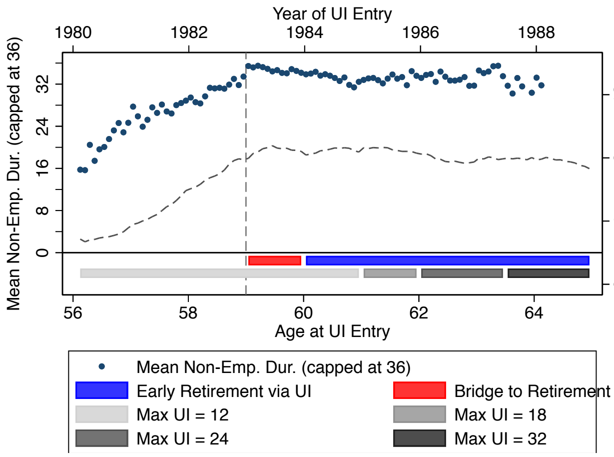
(e) Mean UI Benefit Receipt, 1949 Cohort



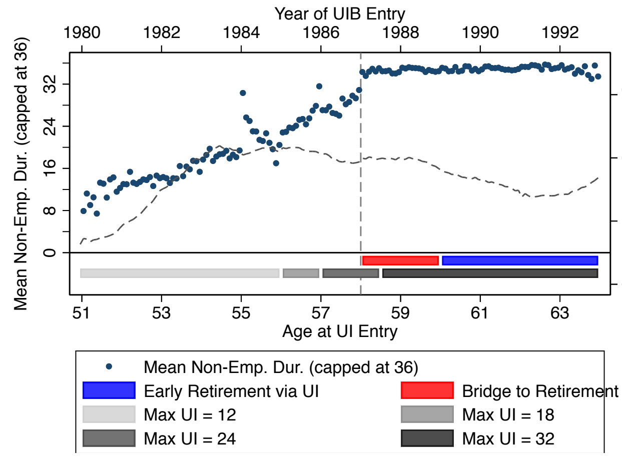
(f) Mean UI Benefit Receipt, 1952 Cohort

Notes: Red dashed lines represent ages at which UI benefit duration increases discontinuously; the black dashed line shows the earliest bridge-to-retirement kink.

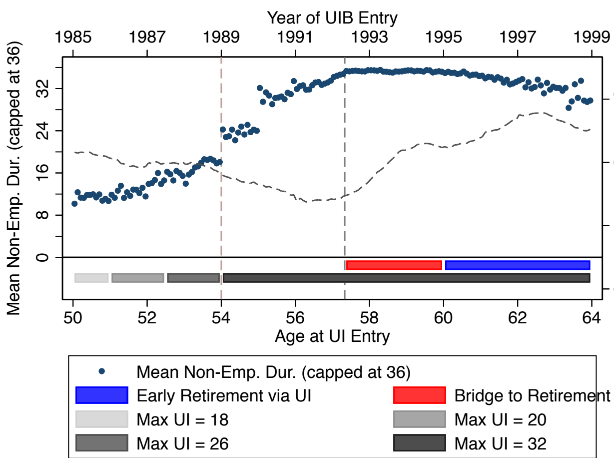
Figure 5: Mean Capped Non-Emp. Duration by Age for Different Cohorts in Germany, Men



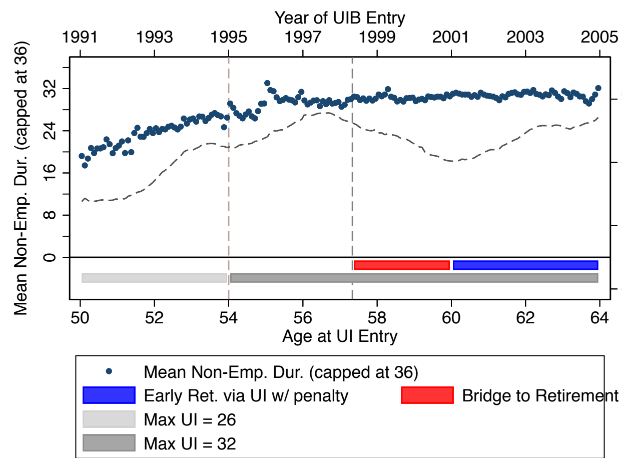
(a) Mean Non-Emp. Duration, 1924 Cohort



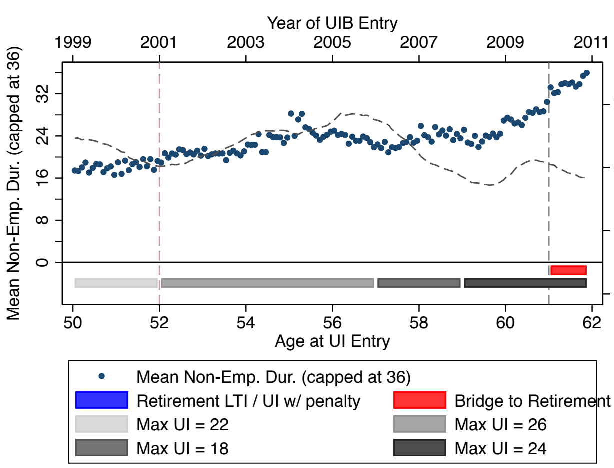
(b) Mean Non-Emp. Duration, 1929 Cohort



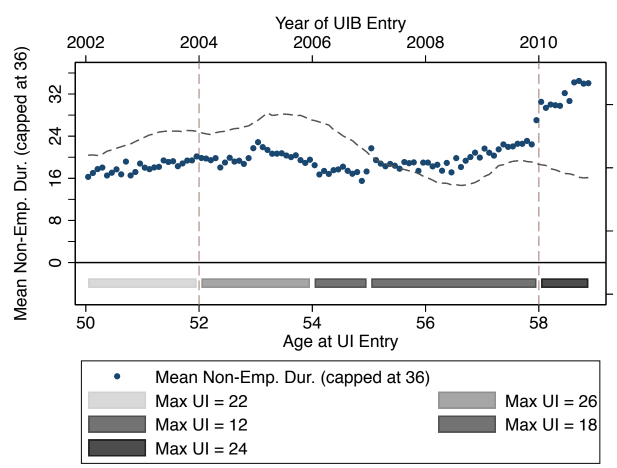
(c) Mean Non-Emp. Duration, 1935 Cohort



(d) Mean Non-Emp. Duration, 1941 Cohort



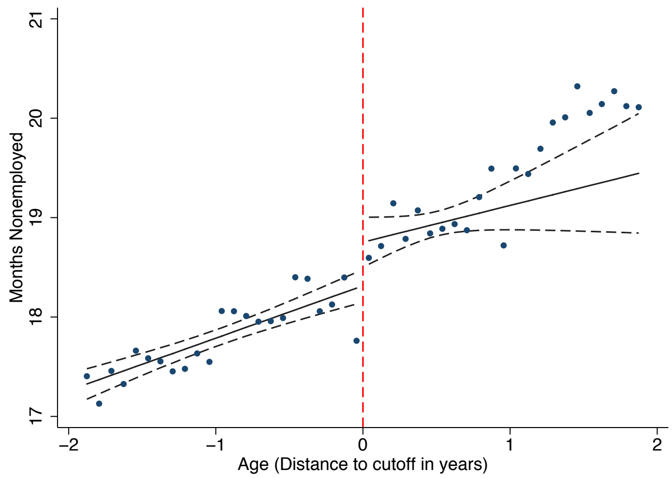
(e) Mean Non-Emp. Duration, 1949 Cohort



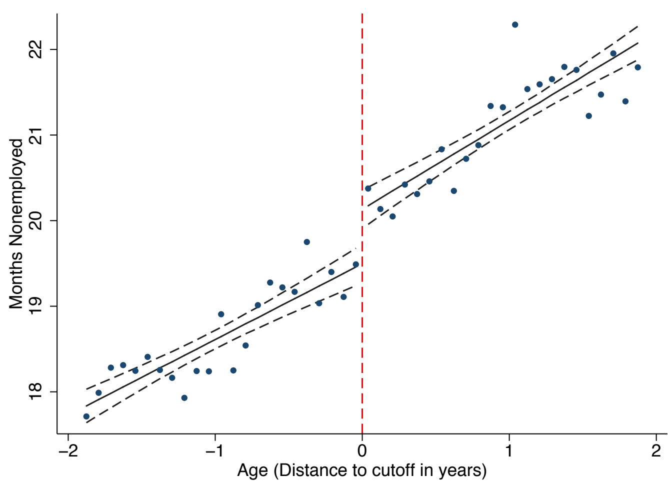
(f) Mean Non-Emp. Duration, 1952 Cohort

Notes: These figures plot mean non-employment duration (capped at 36 months) by age for different cohorts of West German Men with full UI eligibility, excluding mining and steal construction. The bin width is 1/12 of a year.

Figure 6: Sample RD Estimates of the effect of PBD extensions on Non-Emp. Duration



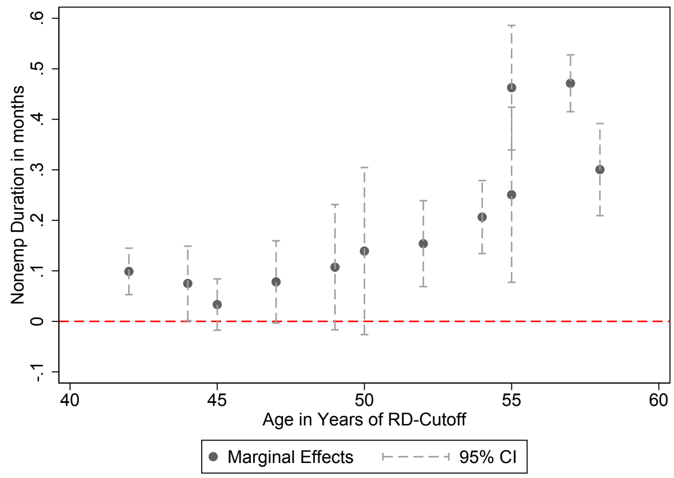
(a) Mean Non-Emp. Duration, Jul 1987 - Feb 1999, cut-off: age 49, Δ PBD = 4



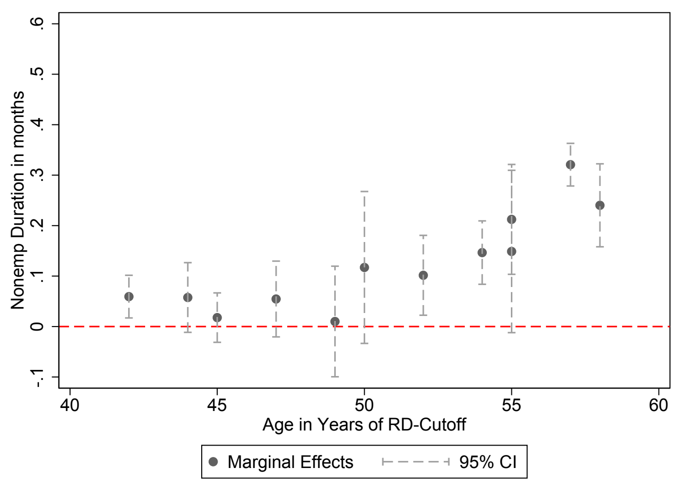
(b) Mean Non-Emp. Duration, Mar 1999- Jan 2006, cut-off: age 52, Δ PBD = 4

Notes: These example RD figures show how mean non-employment duration changes at two sample age-cutoffs that extend potential unemployment benefit duration. Subfigure (a) shows the effect of a 4 month maximum PBD extension (from 22 to 26 months) at the age 49 cutoff in Jul 1987- Feb 1999, while subfigure (b) shows the effect of a 4 month maximum PBD extension (from 22 to 26 months) at the age 52 cutoff in Mar 1999-Jan 2006. We estimate separate, local linear regressions on each side of the cutoff using a rectangular kernel (the solid line). Dashed lines show the 95% CI. The bandwidth is 2 years on each side of the cutoff except in subfigure (a), where it is 1 year to the right of the cutoff due to an additional discontinuity at age 50. We also omit 2 months on each side of the cutoff — the donut hole — to partially address sorting. Estimates at all the age cutoffs in our data, including the above cutoffs, can be found in Table 1 and visualized in Figure 7.

Figure 7: RD Results: 1 month PBD Extension on Non-Employment, Men



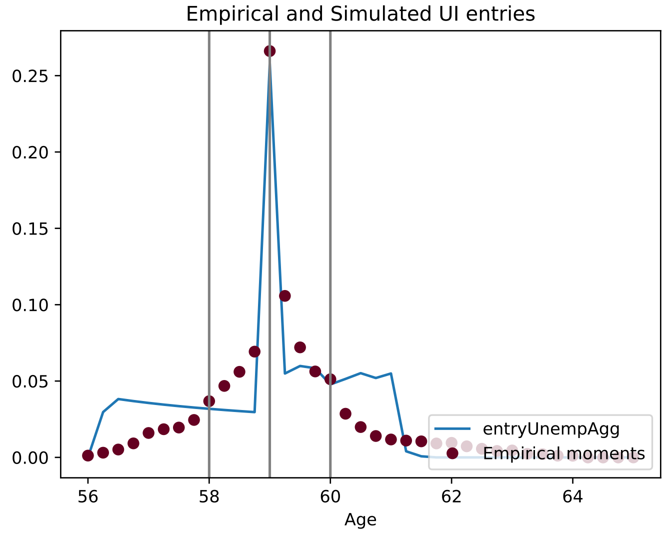
(a) Donut-Hole RD Results: Non-Emp. Duration, Men w/out controls



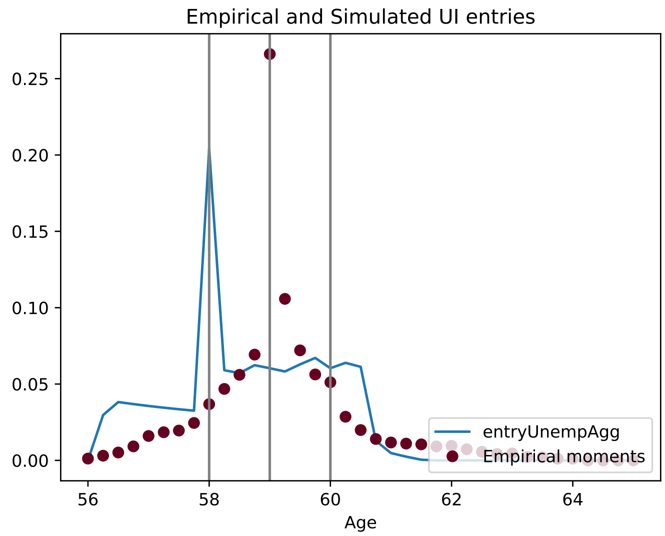
(b) Donut-Hole RD Results: Non-Emp. Duration, Men w/ controls

Notes: These figures contain Regression Discontinuity estimates of the effect of UI potential benefit durations at each age cut-off beginning July 1987. See Table 1 for the estimates. We pool all years under the same UI regime. We employ a local polynomial regression with a rectangular kernel and cluster standard errors at the daily level. 95% CI are plotted. All results are divided by the number of months PBD was extended. The bandwidth is 2 years except for the '87-'99 age 49 and 54 cutoffs where it is 1 year on the right due to other discontinuities. We exclude 2 months on each side of the cutoff – the donut hole – to partially address sorting. We also include detailed individual controls. Controls include: pre-unemployment wage, gender, nationality (non-german), experience, wage/occupation/firm-tenure, education, industry (3-digit), firm-size, month and year. Sample Restrictions: West German Men With full eligibility, excluding mining and steal construction.

Figure 8: Simulated UI Entry Rates from our Life-Cycle Model



(a) ERA = 60, SRA = 65, P = 12 months



(b) ERA = 60, SRA = 65, P =24 months

Notes: For this figure, we simulate our model at the quarterly level. The blue line in subfigures (a) and (b) plots the empirical UI entry rates as a fraction of UI entries for the 1924 cohort. The red dots in subfigure (a) correspond to the simulated UI entry rates from our model for the following institutional parameters: ERA at age 60, SRA at 65, PBD = 12 months. We use a linearly spaced grid of a-types (where a is the inverse disutility of work) between 50 and 100 in intervals of 0.01. Subfigure (b) simulates the effect of increasing PBD to 24 months. Other parameters are chosen as follows: First period is 56, last period is 68, $p_H = w_{ret} = 3000$, $\rho = 0.009$, home production = 100, the UI replacement rate is 0.6, $e = 0.8$,$k = 200{,}000$, $\gamma = 4$, $\lambda = 0.03$,$\kappa = 1$, and $\beta = \delta = 0.995$ annually.

Tables

Table 1: The Effect of PBD on Non-employment Durations

		All Exits	
		(1) No Controls	(2) Controls
Jul 1987 - Feb 1999			
Age 42, P: (12-18), ΔP: 6	$\frac{dy}{dP}$	0.099	0.059
		[0.023]***	[0.022]***
	Observations	205478	205478
	Mean of Dep. Var.	15.2	15.2
Age 44, P: (18-22), ΔP: 4	$\frac{dy}{dP}$	0.075	0.058
		[0.038]**	[0.035]
	Observations	200089	200089
	Mean of Dep. Var.	16.1	16.1
Age 49, P: (22-26), ΔP: 4	$\frac{dy}{dP}$	0.107	0.010
		[0.063]*	[0.056]
	Observations	118965	118965
	Mean of Dep. Var.	17.6	17.6
Age 54, P: (26-32), ΔP: 6	$\frac{dy}{dP}$	0.206	0.147
		[0.037]***	[0.032]***
	Observations	150812	150812
	Mean of Dep. Var.	23.9	23.9
Mar 1999- Jan 2006			
Age 45, P: (12-18), ΔP: 6	$\frac{dy}{dP}$	0.033	0.018
		[0.026]	[0.025]
	Observations	181770	181770
	Mean of Dep. Var.	15.2	15.2
Age 47, P: (18-22), ΔP: 4	$\frac{dy}{dP}$	0.078	0.055
		[0.041]*	[0.038]
	Observations	170340	170340
	Mean of Dep. Var.	16.3	16.3
Age 52, P: (22-26), ΔP: 4	$\frac{dy}{dP}$	0.154	0.102
		[0.043]***	[0.040]**
	Observations	149850	149850
	Mean of Dep. Var.	19.9	19.9
Age 57, P: (26-32), ΔP: 6	$\frac{dy}{dP}$	0.471	0.321
		[0.029]***	[0.022]***
	Observations	208831	208831
	Mean of Dep. Var.	28.5	28.5
Feb 2006- Dec 2007			
Age 55, P: (12-18), ΔP: 6	$\frac{dy}{dP}$	0.463	0.212
		[0.063]***	[0.056]***
	Observations	35124	35124
	Mean of Dep. Var.	17.8	17.8
Jan 2008- Dec 2010			
Age 50, P: (12-15), ΔP: 3	$\frac{dy}{dP}$	0.139	0.117
		[0.084]	[0.077]
	Observations	85107	85107
	Mean of Dep. Var.	16.2	16.2
Age 55, P: (15-18), ΔP: 3	$\frac{dy}{dP}$	0.251	0.149
		[0.088]***	[0.082]
	Observations	67199	67199
	Mean of Dep. Var.	19.2	19.2
Age 58, P: (18-24), ΔP: 6	$\frac{dy}{dP}$	0.300	0.240
		[0.047]***	[0.042]***
	Observations	62228	62228
	Mean of Dep. Var.	22.7	22.7

Notes: This table contains Regression Discontinuity estimates of the effect of UI potential benefit durations at each age cut-off beginning July 1987.Standard errors are in brackets and clustered on day level (* P<.1, ** P<.05, *** P<.01)). We pool all years under the same UI regime. We employ a local polynomial regression with a rectangular kernel and cluster standard errors at the daily level. 95% CI are plotted. All results are divided by the number of months PBD was extended. The bandwidth is 2 years except for the '87-'99 age 49 and 54 cutoffs where it is 1 year on the right due to other discontinuities. We exclude 2 months on each side of the cutoff – the donut hole – to partially address sorting. We also include detailed individual controls. Controls include: pre-unemployment wage, gender, nationality (non-german), experience, wage/occupation/firm-tenure, education, industry (3-digit), firm-size, month and year. Sample Restrictions: West German Men With full eligibility, excluding mining and steal construction.

Online Appendix

A Data Appendix

The Integrated Employment Biographies (IEB) in Germany contains information on all social security reliable employment periods and periods of UI receipt between the years 1975 and 2013. The employment information comes from the employers who are required to report information on all their employees annually, with the exact duration of the employment periods and corresponding individual information. A new employment period starts with a new year, the beginning of a new job or changes at the current job that require notification, such as a switch of the health insurance or from minor employment to social security reliable employment. For each of those periods, individual characteristics such as the birth date, gender and nationality and employment information such as daily gross wage, occupation, educational status and several employer characteristics are reported. The data on UI receipt stems from administrative UI records which are used in the local UI agencies to determine eligibility and to govern the payment process to the UI recipients. It entails information on the exact duration of UI-receipt, the daily benefits.

Due to the daily character of our data, we can exactly determine, whether an individual is regularly employed, on UI benefits, or – when currently not in the data – non-employed. The structure of the data allows furthermore constructing detailed biographical information such as experience or tenure or past exposure to unemployment. We select all UI-entries between 1980 and 2010, which qualify based on their working history for their age-specific maximum PBD. The period-specific calculations are shown in table D.4.

B Budget Set Construction

We assume individuals earn a constant (after tax) wage w and at retirement receive total pension payments $y^R(S)$ and UI payments $y^{UI}(S)$. This yields a budget constraint of the form

$$C = w(E - s) + y^{UI}(E) + y^R(E)$$

Here we detail how we compute the budget set. We denote p as the gross pension replacement rate per year of pension contribution[1]. In other words, Each year of work with wage of w will increase pension benefits $y^R(E)$ by pw. We also denote UI provides income support of $0.68w$. Each year spent on UI increases pension benefits $y^R(E)$ by $0.8 \times pw$. We assume individuals take their full UI duration upon exit and then rely on UA retire, this too can be modified. For illustration purpose, here we assume UA provides zero income. In the simulation, we assume UA yields $0.30w$ and workers spend $T^R - E - P$ on UA.

The budget constraint is thus given by:

$$C = w(E - s) + \underbrace{bD + 0.8 \times pwD \times [T - max\{T^R, E - s + T^u]}_{y^{UI}(E)}$$

$$+ \underbrace{pw(E - s) \times [T - max\{T^R, E - s + T^u\}]}_{y^R(E)}$$

Where D is UI duration, T^u is unemployment duration, P is maximum potential UI duration, b is UI benefit level, m is the UA benefit level. The retirement type r, by definition, $T^u = D \geq P$.

Therefore,

$$C = Y = \begin{cases} w(E - s) + bP + pw \times (E - s + 0.8P) \times [T - T^R] & \text{if } E < T^R - P \\ w(E - s) + bP + pw \times (E - s + 0.8P) \times [T - (E - s + T^u)] & \text{if } E \geq T^R - P \end{cases}$$

The stylized budget sets in Figure 2 make an assumption that worker always retire at the earliest possible retirement age. Lets take as example the 1924 cohort (where $P = 1$ and $T^R = 60$). Therefore, the budget set is

$$C = Y = \begin{cases} w(E - s) + bP + pw \times (E - s + 0.8P) \times [T - 60] & \text{if } E < 60 - P \\ w(E - s) + b(60 - E) + pw \times (E - s + 0.8 * (60 - E)) \times [T - 60] & \text{if } E \geq 60 - P \end{cases}$$

$$\frac{dY}{dE} = \begin{cases} w + pw[T - T^R] & \text{if } E < T^R - P \\ w - b + pw(1 - 0.8)[T - T^R] & \text{if } E \geq T^R - P \end{cases}$$

[1] On average, the net pension replacement rate for an average earner with 45 years of insurance is 70%

Parameters in the budget sets

The baseline busget set by cohort is constructed for the sample of married couple without dependent children. Given that in our sample, around 80% are married and around 15% have dependent children, it is representative to construct the life time budget constrain for married couple without children. We use the following parameters: $s = 20$, $T = 80$, $a = 0.8$ and $B = 0.68w$. The tax rate of married individual with average wage income and whose spouse makes average wage income for cohort 1924, 1929, 1935, 1941, 1949 and 1951 are 0.22 0.24, 0.22, 0.22, 0.22 and 0.18, respectively [2]. The gross average wage are 19456, 17779, 24886, 24886, 22477 and 22423; and the pension replacement rate p also varies by cohorts. Moreover, we use a linear approximation to the curved budget set to measure the changes in slope at the kink point.

The pension replacement rate p

The public pension is calculated on a complex formula of individual career earnings, average pay, revaluation, and insurance periods. The main determinant of pension payments is the sum of individual accumulated earnings points (Entgeltpunkte). One pension earnings point (EP) represents annual pension contributions made by a contributor earns average income. The gross lifetime pension income of a worker who claims old age pension without financial adjustment[3] and insured for $E - s$ years is the following:

$$Y^R_{gross} = \sum_{t=T^R}^{T} AR_t \times \sum_{\tau=s}^{E} \frac{w_\tau}{\bar{w_\tau}}$$

where AR_t is aggregate pension base of year t, w_τ is gross individual income in year τ, $\bar{w_\tau}$ is the average income of all insured people in the pension system. AR_t also represents the pension value of one EP[4]. If we assume constant wage and take the mean of AR_t and $\bar{w_\tau}$,

$$Y^R_{gross} = (T - T^R)\frac{AR}{\bar{w}}(E - s)w_{gross} = (T - T^R)(E - s)pw_{gross}$$

where $p = \frac{AR}{\bar{w}}$ is the gross pension replacement rate per year of pension contribution. A person with 45 years of contribution year has a gross pension replacement rate around 50%.

Prior to 1982, gross pension is the same as net pension benefit. After 1982, pension is subject to health care contribution (KVdR). This percentage of contribution ranges between 6.8% and 8.5%.[5]

[2]The tax rates are obtained from https://www.bmf-steuerrechner.de/ekst

[3]See section for detailed pension calculation when pension types and financial adjustments are considered.

[4]Both AR_t and $\bar{w_\tau}$ are public available information. Table lists AR_t and $\bar{w_\tau}$ of our sample period.

[5]This contribution includes health care insurance contribution and long term care contribution. From April 1, 2004 pensioners have to pay the full contribution (1.7%) for long-term care insurance instead of only half of it.

$$
\begin{aligned}
Y_{net}^R = Y^{gross}(1 - KVdR) = (T - T^R)(E - s)p(1 - KVdR)w_{gross} \\
\simeq (T - T^R)(E - s)p(1 - 8\%)w_{gross} \\
p_{net} = \frac{AR}{\bar{w}}(1 - KVdR)
\end{aligned}
$$

Each additional year of S increases life time income by w_{net} and $p(1 - KVdR)w_{gross}$. The $p(1 - KVdR)$ of married individual with average wage income is 0.01128, 0.01077, 0.01173, 0.00969, 0.00953, 0.00945 of the six cohorts, respectively.

C Labor Supply Elasticity Estimates from a Static Life Cycle Model

This appendix uses the static life cycle model in Brown (2013) and bunching techniques (Saez, 2010; Kleven, 2016) to estimate a labor supply elasticity from the bunching mass at the bridge-to-retirement UI kink. We provide these estimates for comparison with other papers. Section C.1 explains the model and Section C.2 explains the estimation.

C.1 Static Life Cycle Model

Here we describe a static life cycle model as in Brown (2013) and how it can be used to estimate how UI entries vary with maximum PBD duration P ($\frac{dg_t}{dP}$). We assume that all workers maximize lifetime utility subject to their lifetime budget set. In particular, let T be the last period of life, C be total consumption and E be the year of exit from the labor force (which equals total years of work S plus years of schooling s). We assume no discounting and that T is known with certainty. We assume the lifetime utility function take the following function form:

$$U(C,E) = C - \frac{a}{1+\frac{1}{e}}\left(\frac{E}{a}\right)^{1+\frac{1}{e}}$$

where e is the labor supply elasticity and a is ability. The heterogeneity is captured by a a density distribution $\mu(a)$. This quasi-linear, iso-elastic utility function rules out income effect. This model predicts perfect consumption smoothing over the lifecycle: $c_t = \frac{C}{T}$.

We assume individuals earn a constant (after tax) wage w and at retirement receive total pension payments $y^R(S)$ and UI payments $y^{UI}(S)$ which both will depend on years worked (in potentially discontinuous and non-differentiable ways). Note that this yields a budget constraint: $C = w(E-s) + y^{UI}(E) + y^R(E)$.

Note that we can write the elasticity of exit age with respect to the change in effective net wage of working an additional period is $e = \frac{dE}{E} \times \frac{w^{net}}{dw^{net}}$ where $w^{net} \equiv w + \frac{\partial y^{UI}}{\partial E} + \frac{\partial y^R}{\partial E}$. This elasticity will be obtained by a bunching estimator.

The FOC of this problem is given by

$$E = a[w^{net}]^e$$

If the distribution of ability $\mu(a)$ is smooth, this implies a smooth distribution of exit age with density $h_0(E)$. We know that a constant potential benefit duration of P induces a convex kink at age $T^R - P$ where the slope of the individual's budget set exhibits a discrete decline from w^{net}_H to w^{net}_L. Bunching at this kink can be used to recover the elasticity e (see e.g. Saez, 2010; Kleven, 2016). In particular, under the given utility function, we can use the fact that the marginal buncher with ability $a^* + \Delta a^*$ is indifferent between locating at her optimal point under w^{net}_{above} and locating at the kink to get an exact formula for the elasticity of

$$e = -\frac{ln\left(1+\frac{\Delta E^*}{E^*}\right)}{ln\left(\frac{w^{net}_{Above}}{w^{net}_{Below}}\right)}$$

The total amount of observed bunching B is given by $B = \int_{E^*}^{E^*+\Delta E^*} h_0(E)dE$.

Thus, observed bunching and an estimate of $h_0(E)$ can be used to recover e and $\mu(a)$.

C.2 Bunching Estimators

Table C.1 shows the estimated amount of bunching at each retirement-via-UI kink.[6] In order to estimate bunching mass at each kink we need to fit a counter-factual to the data. We use two approaches. First, we use the polynomial approach suggested in (Chetty et al. (2011)). In this approach, we exclude a region around the kink and fit a seventh degree polynomial to the data, including dummies for the excluded bins. The normalized bunching mass is then given by the difference between the actual density and the counter-factual distribution, divided by the average of the counter-factual over this bunching region.[7] Column (1) of Table C.1 shows this normalized bunching mass; the footnote lists the chosen excluded regions.[8] As one would expect, the normalized bunching mass is largest for the 1935 cohort and smallest for the 1949 cohort.

Since this counter-factual approach relies heavily on the shape of the counter-factual (Blomquist and Newey (2017)), we also use an alternative approach. For this second approach, we take advantage of the fact that UI entries closely track the male unemployment rate (UR). We scale the UR by the ratio of the mean number of UI entries between age 49 and the kink to the mean UR in this area, and use this re-scaled UR as a counter-factual. This approach yields a similar qualitative picture, but suggests less bunching for the 1929 cohort and significantly more in 1935 and 1941 cohorts (see Table C.1 column (2)).

To get a sense of how these magnitudes compare to other contexts, we make the strong assumption that all workers around this age behave as if they were following a simple lifetime labor supply model, as in Brown (2013). That is, we assume workers choose their entry date strategically to maximize utility over lifetime consumption and leisure subject to a lifetime budget set. As long as workers abilities are drawn from a continuous distribution, the distri-

[6]There are 5 estimates because our data does not extend far enough yet to observe the kink for the 1952 cohort.

[7]We choose to count the bunching mass as the mass between the cutoff and the right excluded point. Including the imprecision on the left of the cutoff is also an option and would lead to larger estimates.

[8]The exclusion region is chosen visually. For the first three cohorts we simply exclude a region slightly to the left of the cutoff up to age 60. The most difficult and potentially arbitrary choice is for the 1941 cohort, which displays a discrete jump in UI inflows at age 55. We opt to exclude all of the region post-age 55, but only count entries post-age 57.33 in the bunching mass.

bution of UI entries will be smooth. The introduction of a kink in the budget set results in bunching at the kink point, and the amount of bunching allows estimation of a labor supply elasticity, using by now standard techniques (see e.g.: Saez (2010); Kleven (2016)).

Under this model, the normalized bunching mass allows us to recover the elasticity of exit age with respect to the net return to an extra year of work. These estimates are contained in Table C.1. The elasticity estimates for the first four cohorts range from 0.026–0.069 (0.026–0.101 under the UR-counterfactual). While sensitive to the exact specification of the counter-factual, these fall squarely in the range of Brown (2013)'s estimates for Californian teachers timing their retirement as a function of pension benefits.[9] It is striking that we are seeing retirement-via-UI responses that are comparable to standard retirement decisions that do not pass through UI. This should be understood in context, and is likely a function of the lack of serious penalties for entering voluntarily and the relatively low requirements whilst on UI as an older worker.[10]

[9]Brown (2013)'s preferred estimate is 0.04. She examines the retirement behavior of Californian teachers whose normal retirement age is 60 and whose average pension replacement rate is 59%.

[10]It also lines up with findings in other contexts, such as (Kyyrä and Pesola, 2017).

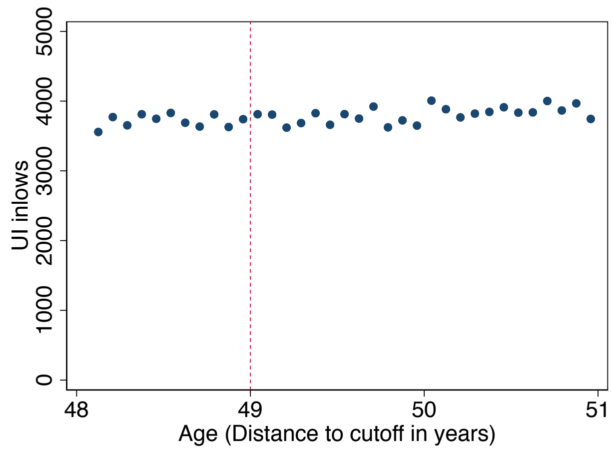
Table C.1: Estimates of the Bunching Mass at each Bridge-to-Retirement Kink

		(1) Estimated Counter-Factual	(2) UR as Counter-Factual
		Bunching at the Kink Induced by Early Retirement via UI	
1924 Cohort			
Age 59, P: (12), R: (60)	Kink Size ($\frac{w^{above}}{w^{below}}$)	0.417	0.417
	Normalized Bunching ($\frac{B}{h_o(0)}$)	15.826	16.343
	Elasticity (e)	0.026	0.026
1929 Cohort			
Age 58, P: (24), R: (60)	Kink Size ($\frac{w^{above}}{w^{below}}$)	0.420	0.420
	Normalized Bunching ($\frac{B}{h_o(0)}$)	37.914	22.294
	Elasticity (e)	0.061	0.036
1935 Cohort			
Age 57.33, P: (32), R: (60)^A	Kink Size ($\frac{w^{above}}{w^{below}}$)	0.414	0.414
	Normalized Bunching ($\frac{B}{h_o(0)}$)	43.033	64.288
	Elasticity (e)	0.069	0.101
1941 Cohort			
Age 57.33, P: (32), R: (60)^A	Kink Size ($\frac{w^{above}}{w^{below}}$)	0.441	0.441
	Normalized Bunching ($\frac{B}{h_o(0)}$)	14.540	23.204
	Elasticity (e)	0.026	0.041
1949 Cohort			
Age 61, P: (24), R: (63)^A	Kink Size ($\frac{w^{above}}{w^{below}}$)	0.447	0.447
	Normalized Bunching ($\frac{B}{h_o(0)}$)	0.686	0.872
	Elasticity (e)	0.001	0.001

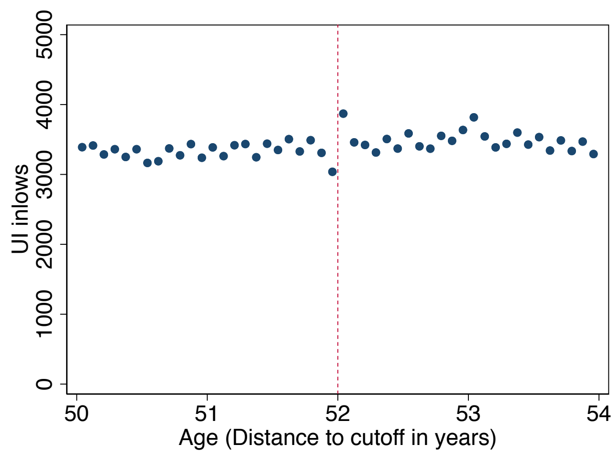
Notes: This table contains estimates of the bunching mass at the bridge to retirement kink for each of the cohorts depicted in Figures 3. The bunching mass is estimated in two ways. First, we fit a 7th degree polynomial to the UI entry data excluding a region around the kink point. For the 1924 cohort we exclude 0.3 years to the left and 1 year to the right of the age 59 cutoff; for the 1929 cohort we exclude 0.3 years to the left and 2 years to the right of the age 58 cutoff; for the 1935 cohort we exclude 0.3 years to the left and 2.66 years to the right of the age 57.33 cutoff; for the 1941 we exclude 2.66 years to the left and 2.66 years to the right of the age 57.33 cutoff; for the 1949 cohort we exclude 0.3 years to the left and 0.66 years to the right of the age 61 cutoff. The normalized bunching mass is given by the difference between observed and counter-factual N between the cutoff and the right exclusion region, divided by the average counter-factual in this region. We also use the unemployment rate as a rough counter-factual by scaling it by the ratio of the mean number of UI entries between age 49 (or lowest available) and the kink to the mean unemployment rate. The normalized bunching mass for this is defined analogously. The bin-width is 1/12 of a year. We also show the point estimate for the elasticity of exit age with respect to the net return to work. This is estimated as described in Appendix C.1.

D Appendix Figures and Tables

Figure D.1: UI Infows by Age, Men



(a) UI Inflows, Jul 1987 - Feb 1999, cut-off: age 49



(b) UI Inflows, Mar 1999- Jan 2006, cut-off: age 52

Notes: These figures plot UI inflows by age for West German Men with full UI eligibility. Panel A shows UI inflows during the 1987-1999 period. The Max PBD jumps from 22 months to 26 months at age 49 during this period. Panel B shows UI inflows during the 1999-2006 period. The Max PBD jumps from 22 months to 26 months at age 52 during this period. The bin width is one month. We set the bandwidth to two years.

Table D.1: Potential Unemployment Insurance Benefit (UIB) Durations as a Function of Age and Months Worked in Previous 7 Years.

Months Worked in prev. X years	January 1983- December 1984	January 1985- December 1985	January 1986- June 1987	July 1987- March 1997	April 1997* - January 2006	February 2006 - December 2007	January 2008 - today (2015)
12	4	4	4	6	6	6	6
16	4	4	4	8	8	8	8
18	6	6	6	8	8	8	8
20	6	6	6	10	10	10	10
24	8	8	8	12	12	12	12
28	8	8	8	14 (≥42)	14 (≥45)	12	12
30	10	10	10	14 (≥42)	14 (≥45)	15 (≥55)	15 (≥50)
32	10	10	10	16 (≥42)	16 (≥45)	15 (≥55)	15 (≥50)
36	12	12	12	18 (≥42)	18 (≥45)	18 (≥55)	18 (≥55)
40	12	12	12	20 (≥44)	20 (≥47)	18 (≥55)	18 (≥55)
42	12	14 (≥49)	14 (≥44)	20 (≥44)	20 (≥47)	18 (≥55)	18 (≥55)
44	12	14 (≥49)	14 (≥44)	22 (≥44)	22 (≥47)	18 (≥55)	18 (≥55)
48	12	16 (≥49)	16 (≥44)	24 (≥49)	24 (≥52)	18 (≥55)	24 (≥58)
52	12	16 (≥49)	16 (≥44)	26 (≥49)	26 (≥52)	18 (≥55)	24 (≥58)
54	12	18 (≥49)	18 (≥49)	26 (≥49)	26 (≥52)	18 (≥55)	24 (≥58)
56	12	18 (≥49)	18 (≥49)	28 (≥54)	28 (≥57)	18 (≥55)	24 (≥58)
60	12	18 (≥49)	20 (≥49)	30 (≥54)	30 (≥57)	18 (≥55)	24 (≥58)
64	12	18 (≥49)	20 (≥49)	32 (≥54)	32 (≥57)	18 (≥55)	24 (≥58)
66	12	18 (≥49)	22 (≥54)	32 (≥54)	32 (≥57)	18 (≥55)	24 (≥58)
72	12	18 (≥49)	24 (≥54)	32 (≥54)	32 (≥57)	18 (≥55)	24 (≥58)
Rahmenfrist - Min emp dur. for new UI eligibility	12	12	12	12	12	12	12
X - Base Period for P≥12	7	7	7	7	7	5	5
X - Base Period for P<12	4	4	4	3	3	2	2
Replacement Rates on Gross Wages in Percent:							
UI (children)	68	68	68	67‡	67	67	67
UI (no children)	63†	63	63	60‡	60	60	60
UA (children)	58	58	58	57‡	57	UIB II	UIB II
UA (no children)	53†	53	53	50‡	50	UIB II	UIB II

Source: Hunt (1995), Bundesgesetzblatt (1983-2015) and Dlugosz etal (2013).

*The reform in 1997 was phased in gradually: For workers who had worked for more than one year during the three years before April 1997, the old rules applied until March 1999 (See Arntz, Simon Lo, and Wilke 2007).

† UI and UA replacement rates were lowered starting in January 1984. Until December 1983, ALG was 68 percent and ALH 58 percent of the previous gross wage, irrespective of whether the recipient had children.

‡ UI and UA were lowered starting in January of 1994.

Table D.2: The schedule of earliest retirement age (ERA) by different retirement pathways

Pathways				
		1957 - 2011		2012 - 2020
Regular old age pension (qualifying period of 5 years)		65		from 65 to 67
	1957 - 2005	2006 - 2011	2012 - 2016	2017 till now
Pension due to unemployment (at least 52 weeks unemployed after $58\frac{1}{2}$) (qualifying period of 15 years)	60	60 to 63	63	same as regular old age pension
		1957 - 2016		2017 till now
Pension for women (qualifying period of 15 years)		60		same as regular old age pension
	1957 - 1972	1972 - 2010	2011	2012 - 2025
Pension for long-term insured (qualifying period of 35 years)	65	63	63 to 62	62

Table D.3: Retirement age by retirement pathways from 1957 till now

Pathways	Time of implementation	Affected cohorts	SRA		Reform
Standard old-age pension	1957 - 2011	<1947 Jan	65	-	
(Years of contribution: 5[11])	2012 - 2030	1947 Jan- 1964 Jan	65 to 67	-	2007 Reform
	> 2031	≥1964 Jan	67	-	
			NRA (no penalty)	ERA (earliest possible)	
Old-age pension for long-term insured	1972 - 1999	1909 Jan - 1936 Dec	63	-	1972 Reform ‡
(Years of contribution :35)	2000 - 2003	1937 Jan - 1938 Dec	63 to 65	63	1992 Reform §
	2004 - 2010	1939 Jan - 1947 Dec	65	63	
	2011 - 2030	1949 Jan - 1964 Jan	65 to 67	63	2007 Reform *
Old-age pension due to unemployment	1972 - 1996	< 1937 Jan	60	-	1972 Reform
or part-time work	1997 - 2006	1937 Jan - 1941 Dec	60 to 65	60	1992/99 Reform
(at least 52 weeks unemployed after $58\frac{1}{2}$, or		1942 Jan - 1945 Dec	65	60	
2 years part-time)	2006 - 2011	1946 Jan - 1948 Dec	65	60 to 63	1992 Reform
(Years of contribution: 15(8 in last 10 yrs))	2012 - 2016	1949 Jan - 1951 Dec	65	63	
	> 2017.1	> 1952 Jan	Phased out	-	2007 Reform
Old-age pension for women	1957 - 2000	<1940 Jan	60	-	
(Years of contribution: 15 (10 after age 40))	2000 - 2009	1940 Jan - 1944 Dec	60 to 65	60	1992 Reform
	2010 - 2016	1945 Jan - 1951 Dec	65	60	
	> 2017.1	> 1952 Jan	Phased out	-	2007 Reform
Old-age pension for disabled workers	1972 - 1977	1911 - 1917	62	-	1972 Reform
(Years of contribution: 35)	1978 - 1980	1918 Jan - 1919 Dec	62 to 60	-	1978 Reform
(Loss of at least 50 percent of earnings	1981- 2000	1920 Jan - 1940 Dec	60	-	
capability)	2001 - 2006	1941 Jan - 1943 Dec	60 to 63	60	1992 Reform
	2007 - 2011	1944 Jan - 1951 Dec	63	60	
	2012 - 2025	1952 Jan - 1963 Dec	63 to 65	60 to 62	2007 Reform
	> 2026	> 1964 Jan	65	62	
Old-age pension for especially long-term insured	<2016	< 1953	63	-	
(qualifying period of 45 years)	2016 - 2028	1953 Jan - 1963 Dec	63 to 65	-	2007 Reform
	> 2029.1	> 1964.1.1	65	-	
Disability pension : independent of age					
	<1985	5 years of contribution			
	> 1985	5 yrs with minimum 3 in last 5 yrs			1984 Reform

† The German public pension system distinguishes "old-age pensions" from "disability pensions": old-age pensions for workers aged 60 and older; and disability benefits for workers below age 60, which at the statutory retirement age are converted to old-age pensions at age 65.
‡ The 1972 reform: "flexible retirement" after age 63 with full benefits became possible for long-term insured; retirement at age 60 with full benefits became possible for women, the unemployed, and older disabled workers.
§ The 1992 reform introduced actuarial adjustment. Since then, we distinguish ERA and NRA. It also increased NRA to 65 for all pathways except for disabled workers. It increased ERA for the unemployed to 63 (See SGBVI appendix 19).
* The 2007 reform increases SRA stepwise between 2012 and 2029 from 65 to 67 for both men and women (see SGB VI 235). For cohorts older born in 1952 and after, retirement pathway for women and the unemployed are phased out.
Sources: Sozialgesetzbuch (SGB) Sechstes Buch (VI), Börsch-Supan and Jürges (2012), Börsch-Supan and Wilke (2006), Giesecke and Kind (2013).

Table D.4: Period Specific Restrictions on Working Histories

Periods	New Eligibility: contributions* >= 12 months during ...	Full Eligibility: contributions* >= ... during ...
1980-1984	previous 4 years	60 months, previous 7
1986-1987	previous 4 years	72 months, previous 7
1987-1999	previous 3 years	72 months, previous 7
1999-2006	previous 3 years	64 months, previous 7
2006-2007	previous 2 years	48 months, previous 5 years
2008-2010	previous 2 years	48 months, previous 5 years

*As contribution duration we count all regular social security reliable employment relationships. For simultaneous employment relationships, we take the one with the highest earnings.