The Effect of Unemployment Insurance Extensions on Reemployment Wages [*]

Johannes F. Schmieder[†]
Boston University
NBER, and IZA

Till von Wachter[‡]
UCLA, NBER,
CEPR, and IZA

Stefan Bender[§]
Institute for Employment
Research (IAB)

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Abstract

Does the search subsidy provided by unemployment insurance (UI) help workers find better jobs by or does the resulting increased time out of work lead to skill depreciation and lower reemployment wages? This paper investigates this question by exploiting strict age thresholds in the German UI system that determine workers' maximum potential UI benefit duration. Using a large administrative data set to impolement a regression discontinuity (RD) design we show that longer potential benefit durations lead to sharp increases in nonemployment durations while lowering post-unemployment wages. In order to interpret this finding, we present a new theoretical result that shows how the average effect of UI extensions on reemployment wages can be decomposed into a reservation wage effect and an effect coming from changes in the wage offer distribution throughout the nonemployment spell. This decomposition can be implemented using information on how reemployment wages conditional on non-employment durations are affected by UI extensions. We show empirically that reemployment wages conditional on time out of work are not affected by increases in potential durations. Our theoretical result implies that in this case the negative effect of UI extensions on average wages is entirely due to changes in the wage offer distribution over time. Furthermore we can estimate the change in mean offered wages over time, by regressing reemployment wages on nonemployment durations and instrumenting for time out of work with the increase in potential UI durations at the age discontinuity. This IV estimate implies that each month out of work reduces wage offers (and reemployment wages) by 0.9 percent, pointing to very high costs of long-term unemployment. Furthermore about half of the average wage loss of 25 percent of the unemployed in our sample is explained by time spent out of work.

[†] johannes@bu.edu

‡ tvwachter@econ.ucla.edu

§ stefan.bender@iab.de

1 Introduction

Do unemployment insurance (UI) benefits help workers find better jobs? While critics have long held that unemployment insurance creates a disincentive to find work, proponents of more generous UI systems, apart from arguing for the beneficial insurance effect of UI, often point out that taking more time to search for a suitable job may produce a positive effect on job match quality. Although these effects seem natural when UI benefits are simply viewed as a subsidy to search efforts, another view holds that long periods of unemployment, possibly induced by generous UI benefits, lead to lower reemployment wages and job quality, either due to skill depreciation (Ljungqvist and Sargent, 1998) or stigmatization (Blanchard and Diamond 1994). While estimates of the effect of UI extensions on job quality are important in their own right, they also offer an opportunity to disentangle the search subsidy effect from skill depreciation during non-employment. Estimates of the causal effect of time out of work on reemployment wages (through real or perceived decline of human capital) are important to gauge the long run effects of long-term unemployment.

This paper sets out with a model that highlights how the effect of UI extensions on reemployment wages can be decomposed into an effect that comes through the slope and the shift of the reservation wage path and a second component that stems from wage offers declining throughout the spell of nonemployment. We provide a new theoretical result that shows that if the path of of reemployment wages conditional on the time of exiting nonemployment is not affected by UI extensions, then the average effect of UI extensions on reemployment wages is only due to changes in the wage offer distribution throughout the nonemployment spell. Furthermore, in this case, extensions in UI durations can be used as an instrument to estimate the change of the wage offer distribution with the duration of nonemployment using a 2SLS estimator.

To identify a causal effect of maximum UI durations on reemployment wages we examine the system in Germany where the generosity of UI durations varies with the age at which an

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individual claims UI benefits. These rules lead to sharp increases in potential UI durations at several age thresholds where individuals of ages within a few days of each other face very different potential UI durations. We exploit this variation using a regression discontinuity design (RD) and a large administrative dataset covering unemployed workers in Germany. This design allows us to precisely estimate the effect of longer potential UI durations on various measures of job quality, such as the reemployment wage, whether an individual moved to a new location, switched industry or occupations, or the duration of the post-unemployment job. We then go on to analyze how reemployment wages conditional on nonemployment duration vary with potential UI benefit durations.

While a substantial body of research has documented the disincentive effect of UI benefits (for example, Solon 1979; Moffitt 1985; Katz and Meyer 1990; Meyer 1990; Hunt 1995), and the consumption smoothing effect of UI (for example, Gruber 1997), the evidence is weaker on how UI affects match quality. The early literature found mixed results based on research designs using observational studies (see Addison and Blackburn 2000 and Meyer 2002 for reviews of this literature). More recent studies by Lalive (2007) and Card, Chetty and Weber (2007)) used regression discontinuity designs to more clearly identify the effects and find negative impact on wages. However, results are relatively imprecisely estimated and not statistically significantly different from zero, while the confidence intervals contain possible negative and postive values that are economically meaningful.

We add to this literature in several ways: First, thanks to a large sample size and treatment variation, we obtain small confidence intervals that allow for meaningful economic interpretation. Second, we are able to investigate a number of alternative measures of job and match quality, such as whether a job requires the employee to move to a new location, job stability, and industry or occupation mobility. Third, we investigate various long-term job outcomes such as wages and employment status five years after the start of unemployment. Finally, we provide a careful analysis to ensure that our effects are not driven by selection effects or unobservables.

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We find strong effects of increased potential UI durations on job finding hazards. One month of additional potential UI benefits increases nonemployment durations by about 0.14 months. Contrary to the view that this subsidy improves match quality, an additional month of potential UI duration decreases wages at the post-unemployment job by about 0.12 percent. This decrease is statistically significant and robust to many alternative specifications. Furthermore, match quality appears to be worse along many other measures of job quality, such as long term employment and wage outcomes or region, industry, and occupation mobility. This finding supports the view that long term unemployment may put the unemployed at a disadvantage by lowering their skills. Turning to our dynamic results, we show that reemployment wages conditional on time out of work do not change at the age discontinuity for most points of support. Together with evidence on how selection of individuals at different non-employment durations change at the age discontinuity, our theoretical results imply that increases in potential UI duration can serve as an instrument for the effect of nonemployment durations on reemployment wages. Our estimates imply that 1 month out of work lowers wages by about 0.7 percent. This implies that the perceived productivity of the average UI recipient in our sample declines by about 10 percent due to skill depreciation during the unemployment spell.

The next section shows in a search model how the effect of UI extensions on reemployment wages is the sum of a reservation wage effect and an effect coming from changes in the wage offer distribution throughout the unemployment spell. In addition we derive conditions under which using UI extensions as an instrument for nonemployment durations provides valid estimates for the change in the wage offer distribution. Section 3 describes the institutional setting, the data, and the empirical methods used in this paper. In section 4 we present the main results of how potential UI durations affect average match quality, and we extensively verify the robustness of our findings. Section 5 presents dynamic results of how reemployment wages and selection conditional on nonemployment duration vary with potential UI increases and implements the IV estimator. Section 6 discusses these results and concludes.

3

2 Theory

2.1 Setup of Model

The model is a discrecte time non-stationary search model, based on van den Berg (1990), with the extension of allowing for endogenous search intensity. The model describes the job search process of a single unemployed individual and thus does not explicitly model heterogeneity. We will discuss identification in the presence of heterogeneity in the methods section.[1]

Unemployed individuals are risk neutral and maximize the present discounted value of income. Workers become unemployed in period $t = 0$ and immediately start looking for jobs. In each period t workers receive UI benefits b_t and choose search intensity λ_t which is normalized, so that the probability of receiving a job offer in that period is equal to λ_t. The cost of job search $\psi(\lambda_t)$ is an increasing, convex and twice differentiable function and the cost is incurred at the beginning of the period.

Jobs offer a wage w and wage offers are drawn from a distribution with CDF F_t, which may vary with the duration of unemployment t. If a job is accepted, the worker starts working at the beginning of the next period. Otherwise the worker remains unemployed for the remainder of the period.

Employment is an absorbing state, i.e. once employed a worker does not get laid off or move to a better jobs. Since workers discount the future at the common subjective discount rate ρ, the value of being employed V^e satisfies:

$$V^e(w) = \frac{1}{\rho}w.$$

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[1]Essentially we can think of this model as allowing for heterogeneity if we simply allow individuals to have different parameters, so that all observed moments are integrated over these parameters.

The Bellman equation for an unemployed worker is given as:

$$V^u(t) = b_t + \max_{\lambda_t}\left[-\psi(\lambda_t) + (1-\lambda_t)\frac{1}{1+\rho}V^u(t+1)\right.$$
$$\left.+\lambda_t\frac{1}{1+\rho}\int_w \max_{accept,reject}\left[V^e(w),V^u(t+1)\right]dF_t(w)\right]$$

Since $V^e(w)$ is increasing in w, the optimal search behavior of the worker is described by a reservation wage ϕ_t, so that all wage offers $w \geq \phi_t$ are accepted. This allows for writing the Bellman equation as:

$$V^u(t) = b_t + \max_{\lambda_t}\left[-\psi(\lambda_t) + \frac{1}{1+\rho}\left(V^u(t+1) + \lambda_t\int_{\phi_t}^{\infty}V^e(w) - V^u(t+1)\,dF_t(w)\right)\right]$$

2.2   Optimal Reservation Wage and Search Intensity Paths

Suppose that the environment becomes stationary for some $t \geq T$. In particular UI benefits and the wage offer distribution become constant after T: $b_t = b$, $F_t(w) = F_T(w)$. This implies that the optimal search strategy is a constant: reservation wage ϕ_T. Using the fact that $V^u(t) = V^u(t+1)$ in the stationary environment, it follows that the stationary reservation wage and the optimal search intensity are given by the follwowing system of equations:

$$\frac{1}{1+\rho}\phi_T = b_T - \psi(\lambda_T) + \frac{\lambda_T}{\rho}\int_{\phi_T}^{\infty}w - \phi_T dF_T(w) \tag{1}$$

$$\rho\psi'(\lambda_T) - \int_{\phi_T}^{\infty}w - \phi_T dF_T(w) = 0 \tag{2}$$

In the nonstationary environment, $t < T$, we use the fact that: $\frac{1}{\rho}\phi_t = V^u(t+1) + b_t$ and $\frac{1}{\rho}\phi_{t-1} = V^u(t) + b_{t-1}$. Therefore knowledge about the reservation wage ϕ_t and the optimal search intensity λ_t in period t will allow us to find the reservation wage in period $t-1$ using this equation:

$$(1+\rho)\phi_{t-1} = (1+\rho)\rho\left(b_{t-1} - \psi(\lambda_t)\right) + \phi_t + \lambda_t\int_{\phi_t}^{\infty}w - \phi_t dF_t(w) \tag{3}$$

Once we have found the reservation wage ϕ_{t-1} in period $t-1$ we can directly solve for the optimal search intensity in the same period:

$$\rho\psi'(\lambda_{t-1}) - \int_{\phi_{t-1}}^{\infty} w - \phi_{t-1}dF_t(w) = 0 \qquad (4)$$

In our empirical application we consider a system where UI benefits are at a constant level b up to the maximum potential duration of receiving UI benefits P. After benefit exhaustion, indivduals receive a second tier of payments indefinitely. We therefore have that $b_t = b$ for all $t \leq P$ and $b_t = \underline{b}$ for all $t > P$. Consider how the reservation wage path and the search intensity path is affected by a change in potential UI durations P. Using the first order conditions we get that:

$$\frac{d\phi_t}{dP} = \frac{dV_{t+1}^u}{dP}\rho \qquad (5)$$

and

$$\frac{d\lambda_t}{dP} = -\frac{dV_{t+1}^u}{dP}\frac{1 - F_t(\phi_t)}{\psi''(\lambda_t)} \qquad (6)$$

If there is at least a small chance that individuals might not find a job until UI exhaustion at $t = P$, then increasing P will increase the value of remaining unemployed for all $t \leq P$, so that $\frac{dV_{t+1}^u}{dP} > 0$. Therefore increasing P will increase the reservation wage ϕ_t and lower search intensity λ_t.

Since the hazard of leaving unemployment is given as $h_t = \lambda_t(1 - F_t(\phi_t))$, we get that

$$\frac{dh_t}{dP} = -\frac{dV_{t+1}^u}{dP}\left[\frac{(1 - F_t(\phi_t))^2}{\psi''(\lambda_t)} + \rho\lambda_t f(\phi_t)\right] \qquad (7)$$

Therefore if the extension in UI benefits affects the value of being unemployed in period t, then it will lower the probability of leaving unemployment in that period.

2.3 The The Reemployment Wage Path

Let w^* be the reemployment wage of an individual. The expected reemployment wage conditional on exiting at time t is then $E[w^*|t] \equiv E[w|t, w \geq \phi_t] = \frac{\int_{\phi_t}^{\infty} w dF_t(w)}{1 - F_t(\phi_t)}$.

To see what determines the evolution of the reemployment wage throughout the unemployment spell, one can decompose the change in the reemployment wage $E[w^*|t+1] - E[w^*|t] \equiv \Delta_t E[w^*|t]$. To simplify notation we will assume that F_t is fully described by one parameter, the mean μ_t.[2] Using a first order taylor approximation:

$$\Delta_t E[w^*|t] \approx \frac{dE[w^*|t]}{d\phi_t} \Delta_t \phi_t + \frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t \tag{8}$$

where $\Delta_t \phi_t = \phi_{t+1} - \phi_t$ and $\Delta_t \mu_t = \mu_{t+1} - \mu_t$. This breaks up the slope of the reemployment wage path into two components. The first part, $\frac{dE[w^*|t]}{d\phi_t} \Delta_t \phi_t$, is the change in the reemployment wage that is due to a change in the reservation wage over time. The second part, $\frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t$, is the part that is due to changes in the wage offer distribution over time.

We are interested in applying this decomposition, however this is complicated by the fact that neither μ_t nor ϕ_t are directly observable. To get around this it is helpful to consider the effect of extending UI benefits on the reemployment wage at time t:

$$\frac{dE[w^*|t]}{dP} = \frac{dE[w^*|t]}{d\phi_t} \frac{d\phi_t}{dP} = \frac{dE[w^*|t]}{d\phi_t} \frac{dV_t^u}{dP} \rho, \tag{9}$$

where the second equality uses equation (5) above. If $\frac{dh_t}{dP} < 0$, then equation (7) implies that $\frac{dV_{t+1}^u}{dP} > 0$. Therefore if the hazard rate shifts at t in response to an increase in P, we can solve this equation for $\frac{dE[w^*|t]}{d\phi_t}$. Plugging this into equation (8) yields:

$$\Delta_t E[w^*|t] = \frac{dE[w^*|t]}{dP} \left(\frac{dV_{t+1}^u}{dP} \rho \right)^{-1} \Delta_t \phi_t + \frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t \tag{10}$$

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[2]This is easily generalizable to more flexible distribution functions characterized by a vector of parameters $\boldsymbol{\mu_t}$.

The advantage of this formulation is that $\Delta_t E[w^*|t]$ and $\frac{dE[w^*|t]}{dP}$ are in principle observable. Furthermore equation (7), suggests that we can learn something about $\frac{dV_{t+1}^u}{dP}$ from observing $\frac{dh_t}{dP}$. In particular equations (7) and (10) directly imply the following proposition.

Proposition 1. *If* $\frac{dE[w^*|t]}{dP} = 0$ *and* $\frac{dh_t}{dP} < 0$, *then*

$$\Delta_t E[w^*|t] = \frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t,$$

i.e. the decline in reemployment wages with time is entirely due to the decline in the wage offer distribution over time.

In the empirical part we will estimate both $\frac{dE[w^*|t]}{dP}$ and $\frac{dh_t}{dP}$ and argue that the conditions for Proposition 1 seem to hold empirically.

2.4 The Effect of Increasing Potential UI Durations on the Reemployment Wage

The expected reemployment wage can be calculated by integrating the reemployment wage conditional on exiting unemployment at t over the distribution of nonemployment durations. Thus if $g(t)$ is the probability mass function of the distribution, we have that $E[w^*] = \sum_0^\infty E[w^*|t]\, g(t)$. An extension in potential UI durations P affects the expected reemployment wage through two components:

$$\frac{dE[w^*]}{dP} = \sum_{t=0}^\infty \left[\frac{dE[w^*|t]}{dP} g(t) \right] + \sum_0^\infty \left[E[w^*|t] \frac{dg(t)}{dP} \right] \tag{11}$$

The first term can also be written as $\sum_{t=0}^\infty \left[\frac{dE[w^*|t]}{dP} g(t) \right] = \sum_{t=0}^\infty \left[\frac{dE[w^*|t]}{d\phi_t} \frac{d\phi_t}{dP} g(t) \right]$, where we use the fact that $E[w^*|t]$ only depends on ϕ_t and not on the search intensity. This term represents the change in the average wage due to the shift in the reservation wage at every point. The second term, $\sum_{t=0}^\infty \left[E[w^*|t] \frac{dg(t)}{dP} \right]$, represents the change in the average wage due

to the shift in the distribution of spells along the expected reemployment wage path.

In the empirical part we argue that it seems a reasonable approximation that the reemployment wage is a linear function of non-employment duration. Suppose that $E[w^*|t]$ is in fact linear in t and can therefore be written as $E[w^*|t] = \pi_0 + \pi_1 t$. We can plug this into equation (11). After some rearranging this yields:

$$
\begin{aligned}
\frac{dE[w^*]}{dP} &= \sum_{t=0}^{\infty} \left[\frac{dE[w^*|t]}{d\phi_t} \frac{d\phi_t}{dP} g(t) \right] + \pi_1 \frac{dD}{dP} \\
&= \sum_{t=0}^{\infty} \left[\frac{dE[w^*|t]}{dP} g(t) \right] + \left[\frac{dE[w^*|t]}{dP} \left(\frac{dV_{t+1}^u}{dP} \rho \right)^{-1} \Delta_t \phi_t + \frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t \right] \frac{dD}{dP} (12)
\end{aligned}
$$

Where $\frac{dD}{dP}$ is the marginal effect of an increase in P on the expected non-employment duration D.[3] For the second equality we use the fact that $\pi_1 = \Delta_t E[w^*|t]$ and equation (10). Together with Proposition 1 this yields:

Proposition 2. *If $E[w^*|t]$ is linear in t, $E[w^*|t] = \pi_0 + \pi_1 t$, and the conditions for Proposition 1 hold, then the first 2 components in equation (12) are zero and $\frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t = \frac{\frac{dE[w^*]}{dP}}{\frac{dD}{dP}}$*

This proposition is useful, since it suggests an alternative way of estimating the change in the wage offer distribution $\frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t$ if it is not possible to directly estimate $E[w^*|t]$, for example because of unobserved heterogeneity. We come back to this in Section 3 when we discuss estimation.

2.5 The Change in the Wage Offer Distribution over Time

An important goal of this paper is to provide a direct estimate of the change in the wage offer distribution with the duration of non-employment. In the empirical section we argue that our results, combined with Proposition 1 and 2, allow us to provide an estimate of $\frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t$, the change in the expected reemployment wage over time that is due to a change in the wage

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[3]Note that $D = \sum_{t=0}^{\infty} [t\, g(t)]$ and $\frac{dD}{dP} = \sum_{t=0}^{\infty} \left[t \frac{dg(t)}{dP} \right]$. Furthermore π_0 cancels out because the changes in the probability mass function have to sum up to 0, so that $\sum_{t=0}^{\infty} \frac{dg(t)}{dP} = 0$.

offer distribution. Since we only observe reemployment wages above the reservation wage, this is the same as the change in the expected wage offer conditional on the wage offer being above the reservation wage.

The following result links this to the change in the expected wage offer:

Proposition 3. *If the support of the wage offer distribution is convex and $\frac{dE[w^*|t]}{dP} = 0$, then $1 - F_t(\phi_t) = 1$, i.e. the reservation wage is not binding. Furthermore*

$$\frac{dE[w^*|t]}{d\mu_t}\Delta_t\mu_t = \Delta_t E[w|t], \tag{13}$$

i.e. the decline in the reemployment wage over time due to is the same as the change over time in the mean of the wage offer distribution.

If the conditions for Proposition 3 hold, then workers are never rejecting job offers because they are below their reservation wage. In this case we observe every offered wage, which yields equation (13). We will discuss the plausibility of the convexity assumption as well as implications of this below.

2.6   Empirical Content of Model

The goal of the paper is to estimate the effect of potential UI durations on reemployment wages and decompose the effect into the different components of equation (12). Furthermore we are particularly interested in obtaining an estimate of the decline in the wage offer distribution throughout the non-employment spell. Estimating the effect of UI extensions on average reemployment wages is relatively straightforward, as long as there is exogenous variation in potential UI durations. Yet, as the discussion in this section showed, it is difficult to decompose the effect $\frac{dE[w^*]}{dP}$ into its components since neither wage offers nor the reservation wage are directly observed.

Propositions 1 and 2, however, offer a way to identify the different components under

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some specific conditions, in particular in the case that $\frac{dE[w^*|t]}{dP} = 0$ and $\frac{dh_t}{dP} < 0$. In order to see whether these conditions hold and to apply the decomposition if they do hold, it is crucial to obtain consistent estimates of $\frac{dE[w^*|t]}{dP}$, $\frac{dh_t}{dP}$, $\frac{dD}{dP}$, and $\Delta_t E[w^*|t]$. While these are relationships between observables, estimation of these moments is complicated by the presence of observed and unobserved heterogeneity. We will address these problems carefully when discussing our methods in Section 3.3.

3 Institutions, Data and Empirical Methods

3.1 Institutional Background

After working for at least 12 months in the previous three years, unemployed workers in Germany are eligible for UI benefits that provide a fixed replacement rate of 63 percent for an individual without children. Workers losing a job through no fault of their own are eligible to receive unemployment insurance benefits if they have worked for at least 12 months in the previous three years.[4] Individuals who have quit their jobs voluntarily are subject to a 12 weeks waiting period. To focus on individuals who lost their job involuntarily and minimize selection concerns due to quitting we restrict our sample to individuals who claimed UI benefits within 12 weeks after their job ended. This paper focuses on the time period between 1987 and 1999, during which the maximum duration of benefits was tied to recipients' exact age when they began receiving UI benefits and to their labor force history. Between July 1987 and March 1999, the maximum potential UI duration for workers who were younger than 42 years old was 12 months.[5] For workers age 42 to 43 maximum potential UI duration increased to 18 months; for workers age 44 to 48, the maximum duration further

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[4]Sanctions for not taking suitable jobs exist but appear to be rarely enforced (Wilke 2005). For individuals with children the replacement rate is 68 percent. There is a cap on earnings insured, but according to Hunt (1995) it affects a small number of recipients. Since they are derived based on net earnings, in Germany UI benefits are not taxed themselves, but can push total income into a higher income tax bracket.

[5]For an investigation of the stepwise introduction of these age cutoffs between 1983 and 1987 see Hunt (1995).

rose to 22 months.[6] As we explain further below, to obtain precise measures of potential UI durations, we restrict ourselves to a sample of workers who were, based on their employment history, eligible to the maximum potential UI durations in their age group. At the end of the 1990s a reform was enacted to reduce potential disincentive effects of unemployment insurance. Starting in April 1999, potential UI durations were lowered and the age thresholds were increased by 3 years. Thus, to be eligible for 18 or 22 months of benefits a worker had to be at least 45 or 47 on the claiming date. We will use these alternative thresholds to validate our main research design.[7]

Individuals who exhaust regular UI benefits and whose net liquid wealth falls below a threshold are eligible for unemployment assistance (UA), which does not have a limited duration. The nominal replacement rate is 53%, but UA payments are reduced substantially by spousal earnings and other sources of income. For example, for a worker earning as much as 10% less than his or her spouse, the UA benefits are zero. Given about 80% of individuals in our cohort and age range are married, based on average earnings levels, UA benefits average about 35% for males and 10% for females.[8] Among all new periods of UI receipt in our sample, about 10 to 15% eventually receive UA benefits. We study the potential effect of UA on our findings in our empirical analysis.

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[6]There are additional thresholds at older ages. For example, at age 49 potential UI durations increase to 26 months and at age 54 to 32 months. Since relatively few individuals reach these higher thresholds and the increases are smaller relative to the level of potential UI benefits at the left side of these thresholds the match quality estimates are quite noisy and not very informative; therefore, we do not present results on them here. Furthermore at the age 54 threshold there is a more substantial effect on permanently leaving the labor force which makes the match quality estimates harder to interpret due to selection concerns.

[7]The reform was enacted in 1997 but phased in gradually, so that for people in the highest experience group, which constitutes our analysis sample, it took effect in April 1999 (See Arntz, Lo, and Wilke 2007). To avoid confusion we refer to this period as the post 1999-regime in the text. In 2003 and 2004, the entire German social security system underwent a comprehensive series of reforms (the so-called Hartz reforms). We use individuals who started receiving UI benefits between April 1999 and December 2004 as a second sample, thus excluding workers who became unemployed after the Hartz IV reform took place.

[8]UI benefits are paid for by worker and employer contributions, whereas UA benefits are funded by general revenues. The wealth threshold is not very stringent, but given the wealth distribution in Germany it is likely to be binding for part of our sample.

3.2 Data

The data for this paper is the universe of social security records in Germany. For each individual working in Germany between 1975 and 2008, the data contains day-to-day longitudinal information on every instance of employment covered by social security and every receipt of unemployment insurance benefits, as well as corresponding wages and benefit levels.[9] Compared with many other social security data sets, this data is very detailed. We observe several demographic characteristics, namely gender, education, birth date, nationality, place of residence and work, as well as detailed job characteristics, such as average daily wage, occupation, industry, and characteristics of the employer.[10]

To study the effect of extensions of potential UI Durations, we created our analysis sample by selecting all periods of non-employment in this data in the age range of 40 to 46. Given changes in the institutional framework discussed in the previous section, we consider unemployment spells starting any time between July 1987 and December 2004, while focusing mainly on the pre-1999 reform period. For each non-employment spell we created variables about the previous work history (such as job tenure, experience, wage, industry and occupation at the previous job), the duration of UI benefit receipt in days, the UI benefit level, and information about the next job held after non-employment.

Since we do not directly observe whether individuals are unemployed we follow the previous literature and use length of non-employment as a measure for unemployment durations (for example, Card, Chetty, and Weber 2007b). The duration of non-employment is measured as the time between the start of receiving UI benefits and the date of the next registered period of employment. Since some people take many years to return to registered employ-

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[9]For more information on this data source, see Bender, Haas and Klose (2000)

[10]Individual workers can be followed using a unique person identifier. Since about 80 percent of all jobs are within the social security system (the main exceptions are self-employed, students, and government employees) this situation results in nearly complete work histories for most individuals. For additional description of the data see Bender, Haas and Klose (2000). Each employment record also has a unique establishment identifier that can be used to merge establishment characteristics to individual observations. Below, we will use information on occurrences of establishment-level mass-layoffs constructed, described, and analyzed further by Schmieder, von Wachter and Bender (2009).

ment while others never do so, we cap non-employment durations at 36 months and set the duration of all longer periods at this cap. This appproach reduces the influence of outliers and avoids censoring due to the end of the observation period in 2008. Our results are very robust with regard to the exact choice of the cap.

One of the main 'treatment' variables we are interested in is the potential duration of unemployment insurance benefits for any given period of non-employment. To calculate potential UI duration for each period in our sample, we use information about the law in the relevant time periods together with information on exact birthdates and work histories. This method yields exact measures for workers who have been employed for a long continuous time and are eligible for the maximum potential benefit durations for their age groups. However, the calculation is not as clear cut for workers with intermittent periods of unemployment because of complex carry-forward provisions in the law. We thus define our core analysis sample to be all unemployment spells of workers who have been employed for at least 44 months of the last seven years and who did not receive unemployment insurance benefits during that time period. The resulting sample is of intrinsic interest, because it corresponds to workers often the focus of discussion of extensions in UI benefits in difficult economic times: mature workers with high labor force attachment who absent a layoff or a recession would have been unlikely to become unemployed. Below, we show that our results are robust to broadening our sample to include workers with weaker labor force attachment. We also show that the characteristics of our sample are comparable with those of UI recipients in the United States.

3.3   Estimation

Estimating $\frac{dD}{dP}$ and $\frac{dE[w^*]}{dP}$

The institutional structure and data allow us to estimate the causal effect of UI benefit durations on non-employment duration and other outcomes using a regression discontinuity design. In a first step, we exploit the sharp age thresholds in eligibility rules for workers with

previously high labor force attachment in Germany to estimate the effect of large extensions in UI durations on labor supply. We then replicate this approach for every year or year-by-industry in our sample, and correlate it with indicators of the business cycle.

Throughout the paper, the analysis proceeds in two steps. We follow common practice and show smoothed figures to visually examine discontinuities at the eligibility thresholds (Lee and Lemieux 2010). To obtain estimates for the main causal effects, we follow standard regression discontinuity methodology and estimate variants of the following regression model:

$$y_i = \beta + \gamma \times \Delta P \times D_{a_i \geq a^*} + f(a_i) + \epsilon_i, \tag{14}$$

where y_i is an outcome variable, such as non-employment duration, of an individual i of age a_i. $D_{a_i \geq a^*}$ is a dummy variable that indicates that an individual is above the age threshold a^*. For our main estimates, we focus on the longest period for which the UI system was stable, July 1987 - March 1999, and we use the three sharp thresholds at age 42 and 44.[11] We estimate equation 14 locally around the two cutoffs and specify $f(a_i)$ as a linear function while allowing different slopes on both sides of the cutoff. We use a relatively small bandwidth of two years on each side of the cutoff.

In order to obtain additional power we also estimate a pooled regression model, where we take the estimation samples for the age 42 and the age 44 cutoffs together. For this procedure we normalize the age for all individuals within two years of the age 42 threshold to the age relative to age 42 (i.e. the rescaled age variable is set to 0 for someone who is exactly age 42 at the time of claiming UI). We then do the same for all individuals within two years of the age 44 cutoff and merge these two samples. We estimate the following model on this pooled sample:

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[11]There is a fourth discontinuity during this period at age 54. Because early retirement becomes very common at this age, and various retirement policies interact with the UI system we focus on younger workers in this paper. Early retirement in the context of the German UI system has been analyzed for example in Fitzenberger and Wilke (2010).

$$y_i = \beta + \gamma \times \overline{\Delta P} \times D_{a_i \geq a^*} + f(a_i) + \epsilon_i,$$

where a_i is the normalized age variable and $\overline{\Delta P}$ is the average change in potential UI durations at the age threshold. This value simply depends on how many individuals are at the age 42 threshold where potential UI durations increase by six months and how many individuals are at the age 44 threshold where potential UI durations increase by four months. With this specification $\hat{\gamma}$ is a direct estimate of the rescaled marginal effect, forcing it to be equal at the two cuoffs.

Estimating $\frac{dE[w^*|t]}{dP}$ and $\frac{dh_t}{dP}$

In principle it would be possible to obtain $\frac{dE[w^*|t]}{dP}$ by estimating the following regression only for individuals exiting unemployment at time t.

$$w_i^* = \delta_t P_i + \epsilon_i \,|\, t_i = t \tag{15}$$

If $cov(\epsilon_i, P_i | t_i = t) = 0$, then esimating equation (15) via OLS will yield $\hat{\delta}_t$ as consistent estimates for $\frac{dE[w^*|t]}{dP}$.

Using the RD methodology one can write equation (15) as

$$w_i^* = \delta_t P_i + f(a_i) + \epsilon_i \,|\, t_i = t \tag{16}$$

where $P_i = \Delta P \times D_{a_i \geq a^*}$. Assuming the identification assumptions of the RD design hold, we know that $cov(\epsilon_i, P_i) = 0$ in the total RD sample (after controlling for a_i). However, the RD assumptions do not impliy that ϵ_i and P_i are uncorrelated *conditional* on the duration of unemployment t_i: While the RD guarantees that individuals on both sides of the cutoff are comparable on average, the time when people exit unemployment t_i is affected by individual

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behavior and possibly by the treatment variable P_i.

Since the RD assumptions do not directly guarantee $cov(\epsilon_i, P_i | t_i = t) = 0$, we provide two alternative arguments that make $cov(\epsilon_i, P_i | t_i = t) = 0$ plausible.

First consider the decomposition of ϵ_i in to observables x_i and unobservables ε_i: $\epsilon_i = x_i \beta + \varepsilon_i$. While we do not observe ϵ_i, we can test whether observables are correlated with potential UI durations conditional on t. If $cov(x_i, P_i | t_i = t) = 0$ for all observables, then it seems plausible that: $cov(\epsilon_i, P_i | t_i = t) = 0$ and that estimating equation (16) will yield consistent estimates of $\frac{dE[w^*|t]}{dP}$.

Second, we can also make an argument based on the theoretical restriction that $\frac{dE[w^*|t]}{dP} \geq 0$, since the reservation wage has to rise in response to an increase in P. If ϵ_i is a person fixed effect, then a $\hat{\delta}_t = 0$ is only consistent with $\frac{dE[w^*|t]}{dP} > 0$ if $cov(\epsilon_i, P_i | t_i = t) < 0$. If for all t we find $\hat{\delta}_t = 0$, then it has to be the case that $\frac{dE[w^*|t]}{dP} = 0$, since on average ϵ_i is uncorrelated with P and therefore it cannot be that for all t: $cov(\epsilon_i, P_i | t_i = t) < 0$.

The same basic issues hold for estimating the effect on the hazard rate $\frac{dh_t}{dP}$. While the RD design guarantees consistent estimates for estimating the effect of an increase in P on the average duration of nonemployment, similar arguments based on observables and theory, can be made as for the wage effects conditional on t.

Estimating $\Delta_t E[w^*|t]$

In order to estimate $\Delta_t E[w^*|t]$, and if we assume linearity, we could estimate:

$$w_i^* = \pi_1 t_i + x_i \beta + u_i \tag{17}$$

If $cov(u_i, t_i) = 0$, we have consistency of the OLS estimator $\hat{\pi}_1$ for $\Delta_t E[w|t]$. Unfortunately, $cov(u_i, t_i) = 0$ seems highly implausible, clearly individuals who find a job within a very short time are different on average than individuals who are unemployed for a very long time. The regression discontinuity, does not affect this problem at all, since this type of

selection would be going on on either side of the threshold. Controlling for observables may alleviate the selection problem somewhat, but given that observables are correlated with t_i, it seems likely that even after controlling for observables we still have that $cov(u_i, t_i) \neq 0$.

However, if the $\frac{dE[w^*|t]}{dP} = 0$ and $\frac{dh_t}{dP} < 0$, then Proposition 2 suggests an alternative way for estimating π_1 using the relationship:

$$\pi_1 = \frac{dE[w^*]}{dD} = \frac{\frac{dE[w^*]}{dP}}{\frac{dD}{dP}}$$

Essentially this is an IV estimator, where we instrument nonemployment duration with potential UI durations. Both $\frac{dE[w^*]}{dP}$ and $\frac{dD}{dP}$ can be estimated consistently using the RD design and π_1 can then be calculated by dividing these two estimates or by directly estimating equation (17) using two stage least squares, whereby we instrument for D using the variation in P using the variation at the RD cutoff.

3.4   Validity of RD Design

The regression discontinuity method may yield inconsistent results if factors apart from the treatment variable vary discontinuously at the threshold. In that case, estimates for β_1 may pick up the impact of these other variables rather than only the causal effect of the treatment. Such discontinuous variation in other variables at the threshold can occur when individuals have control over the forcing variable of the regression discontinuity estimator–in this case, the age of the UI claimant. Individuals having precise control over the forcing variable may lead to systematic selection, where optimizing agents choose to land on different sides of the cutoff based on other characteristics. Such a scenario may strongly violate the necessary identification assumptions for the RD design.

Two ways to manipulate the age of the UI claimant are feasible in our setting. On the one hand, employers can decide who to lay off and when. For example, if some employers care about workers' welfare or need approval by work councils, they may find it preferable

to lay off workers with higher UI eligibility. If these employers differ from those who do not base their layoff decisions on workers age or UI eligibility, there will be systematic differences in observable and unobservable characteristics on both sides of the cutoff. One the other hand, once a worker has lost a job, he or she can decide how long to wait before claiming UI. For example a worker who loses his or her job five days before his or her 42nd birthday, can claim UI benefits directly and will be eligible for a maximum of 12 months. Or he or she can wait five days before claiming UI benefits and be eligible for a maximum of 18 months. While such selection is a serious potential problem, a major advantage of the RD design is that it is straightforward to test for it.

A standard test for sorting around the threshold is to investigate density plots to locate spikes near the threshold or permanent shifts of observations at the thresholds. Figure 1 shows the number of periods of unemployment in two-week age intervals around the cutoff. The figure indicates that there is a small increase in the density right after the threshold.[12] Further investigation showed that this increase is driven not by higher layoff rates right after the age cutoffs, but by individuals who get laid off very shortly before their birthdays and postpone their claim until they are eligible for longer benefits. The magnitude of this effect, however, is very small: only about 200 instances relative to about 500,000 observations in the sample.

The second standard test is to investigate whether predetermined characteristics of individuals in the sample vary discontinuously at the threshold. Table 1 presents results estimating equation (1) and (2) using two year bandwidths around the cutoffs. The first panel shows the estimates for the age 42 threshold, where potential UI durations increase from 12 to 18 months; the second panel shows the estimates for the increase in potential UI durations from 18 to 22 months at the age 44 threshold. The last panel shows the results for pooling both thresholds. The only statistically significant change at the threshold is the fraction female in the age 42 model and the pooled model: The fraction of UI recepients who

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[12]These increases in density are statistically significant according to the McCrary (2008) test.

are female is estimated to increase by about three-quarters of a percentage point (or 0.5%
in the pooled model). All other variables show essentially no (economically or statistically)
meaningful difference at the threshold.

Both the increase in the density at the threshold as well as the increase in fraction of
women are very small relative to the average density and the overall fraction of women.
In smaller datasets, such small discontinuities and density shifts would almost certainly be
undetectable. While they point to a small violation of the RD identification assumptions,
they should have a relatively small impact on the overall results. To ensure that they
are unlikely to drive our results, our empirical section checks repeatedly for robustness.
Among other things, we show how controlling for observable characteristics affects our results,
what happens when we exclude observations close to the age discontinuity, and Manski-type
bounds of the treatment effect allowing for selection around the threshold.

4 The Average Effect of UI extensions on Job Quality

4.1 The Effects of UI extensions on Nonemployment Durations and Wages

One would expect an effect of potential UI durations on job quality only if longer potential
UI durations have a substantial effect on the actual time of nonemployment. If the increase
in nonemployment duration is very small, it is unlikely that either the additional time to
search for jobs or human capital depreciation would noticeably impact match quality. We
therefore first investigate whether potential UI durations affect the time of UI benefit receipt
and of unemployment before turning to the impact on wage outcomes and other measures
of match quality.

Figure 2 (a) shows the effect of an increase in potential UI durations on the number of
months of receiving UI benefits. Each dot represents the average length of UI benefit receipt
for individuals who began collecting UI benefits within a 2 month age window. Figure 2
shows that increasing potential UI durations from 12 to 18 months increases the actual time
of receiving UI benefits by 1.7 months. At the second threshold, the time of receiving UI

benefits increases by about 1 month–less than at the first threshold but expected given the smaller increase in potential durations. This effect is partly mechanical, since individuals who would have exhausted their benefits at 12 months or 18 months are now covered for up to 6 more months, and partly behavioral, since individuals may reduce their search effort and thus stay unemployed longer. Either way, the effect is quite large and clearly shows that the policy change is highly significant for individuals.[13]

To demonstrate the purely behavioral effect of an increase in potential UI durations, Figure 2 (b) shows the effect on nonemployment durations. At the first age threshold the increase in potential UI durations leads to an increase of nonemployment durations of almost 0.9 months. At the second threshold, nonemployment durations increase by about two weeks. Increases in potential UI durations thus have a very clear effect on nonemployment durations and substantially change behavior of unemployed individuals.

In Table 2, columns (1) and (2) confirm the visual impression. The effects on actual UI duration and nonemployment duration are very precisely estimated (for example, a t-statistic of larger than 10 for nonemployment durations in the joint model). The table also shows the marginal effect of an increase in potential UI durations by 1 month, i.e. the estimated RD coefficient rescaled by the increase in potential UI durations. For one additional month of potential UI benefits unemployed individuals receive about 0.3 months of additional benefits and remain unemployed for about 0.15 months longer. These marginal effects are similar to findings from previous research including Moffitt (1985), Katz and Meyer (1990), Meyer (1990), Hunt (1995), or Card et al. (2007a), although much more precisely estimated. Column (3) shows that this increase in nonemployment durations has long lasting effects: Individuals to the right of the threshold are more likely to permanently exit the labor force. The probability of ever again working in a social security liable job decreases by about 0.5 percentage points at the age 42 threshold or less than 0.1 percent per additional month of potential UI benefits in the pooled model (Panel C). This effect is very small relative to the

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[13]Another way to see this is that about 30 percent of recipients who are eligible for 12 months of UI exhaust their benefits

average probability of being employed again (which is close to 90 percent).

Next, we turn to the effects of increases in potential UI durations on job quality measured by the wage at the post-unemployment job. Figure 3 (a) shows the effect on the log wage at the first job after the period of unemployment. There appears to be a small decline by about 0.01 log points in the post-unemployment wage at the age 42 threshold. At the age 44 threshold, the lines (fitted quadratic polynomials) also seem to indicate a small drop in the post-unemployment wage. Figure 3 (b) shows the difference in the pre-unemployment log wage and the post-unemployment log wage. This difference is essentially a way to remove an individual fixed effect and can be viewed as a way to control for possible selection or simply to soak up variance and produce more precise estimates. The figure shows that there are large wage losses for the unemployed, about 24 percent at the age 42 threshold and close to 26 percent at the age 44 threshold. While the gain in precision is modest, Figure 3 (b) indicates that selection along the previous wage has little impact on the results and still clearly points to a negative effect of potential UI durations on post-unemployment wages.

In Table 2, columns (4) and (5) provide the corresponding regression estimates. The regression results confirm the visual impression that the increase in potential UI durations has a negative effect on post-unemployment wages. Panel A shows that the post-unemployment wage is about 1 percent lower when potential UI durations increase by six months. Similarly, the pre-post wage difference drops by about 0.01 log points, and both estimates are statistically significant on the 5 percent level. The point estimates at the age 44 threshold are very similar (especially when rescaled by the increase in potential UI durations) but too noisy to be statistically significant. Panel C shows the results from pooling both cutoffs and thus benefits from a small gain in statistical precision. Overall, the point estimates are clearly negative, and while the 95 percent confidence interval includes values very close to a zero match effect, we can rule out positive wage effects with high confidence. The estimate from the pooled model implies that an increase in potential UI durations by one month decreases post-unemployment wages by about 0.18 percent. Although this effect may not seem large,

it may add up to substantial losses if individuals remain in lower paying jobs for a long period of time, something we will return to below.

Wage is an important measure for job quality, but there are clearly many other dimensions of jobs that are important to workers. While it seems likely that lower paying jobs are also inferior along other dimensions, it is possible to think of scenarios where this may not be the case. One example might be when workers look for career potential in new jobs. Possibly, jobs that offer higher investments in human capital have lower initial wages but higher wage growth in the medium term and may be more attractive. Following another scenario, many unemployed workers live in economically weak regions with low wages. If moving is costly (for example, because of social costs), unemployed individuals might prefer staying close to their old jobs even if they could earn higher wages by moving to a different region. A worker with a relatively short potential UI duration may be more likely to be forced to search for a job outside of the region where he or she is living, but the higher wage may be nullified due to the high costs of relocation. It is therefore worthwhile to investigate additional measures of job match quality before concluding that the decline in post-unemployment wages is due to human capital depreciation.

Table 3 shows the effect of increases in potential UI durations on a number of other outcome variables. Column (1) shows the effect on the log wage five years after the start of the unemployment spell. A month increase in potential UI durations in the pooled model is associated with a 0.18 percent decrease in the long term wage. This point estimate is exactly the same as in the pooled model in Column (4) of Table 2, thus there does not appear to be an effect of potential UI durations on wage growth. Similarly Columns (2) and (3) of Table 3 show that individuals with more time to search for jobs are not more likely to be employed five years out and in fact have a slightly higher probability. This marginal effect of about 0.1 percentage points is statistically significant at the first threshold and in the pooled model, of being unemployed again five years later. Column (4) complements this finding by looking directly at whether the first job after unemployment is more stable if individuals have more

time to search for a job. Column (4) shows that there is a small but statistically significant, decrease in the duration of the post-unemployment job of about 0.012 years, indicating that there is a negative effect of increased potential UI durations on job stability.

Although there appear to be no gains in job quality when individuals have more time to search for jobs, they might still trade off better location matches for jobs that are inferior along other dimensions. Column (5) of Table 3 estimates the effect of potential UI durations on the probability of taking a job in a different region (county) than the pre-unemployment job. There is no indication that longer potential UI durations increase the probability of finding a job in the same region as the previous job.

It also seems likely that remaining in the same industry or occupation is important to job seekers. Columns (6) and (7) show longer potential UI durations increase the probability of switching to both a different industry and a different occupation by about 0.15 to 0.2 percentage points.

Overall all measures of job quality either point to negative effects of longer potential UI durations or no effect.

[[About here discussion of effect from industry changing (Table 6) and long-term effects (Table 7).]]

4.2   Robustness

In this section, we discuss the effect of changes in the specification of the RD model on our results. Our main results are all based on a two-year bandwidth around the age thresholds (that is, the sample includes workers within two years of age relative to the thresholds at the beginning of their unemployment) with linear age controls. Focusing on the model pooling both thresholds, Table 4 shows the sensitivity of our results when we allow for more flexibility in the estimation, focusing on five outcome variables: nonemployment durations, the two wage outcomes, relocation probability, and post-unemployment job duration. The first column shows the baseline estimates using a two-year bandwidth, while columns (2)

and (3) show the estimated effects when the bandwidth is reduced to 1 year and 0.5 years. Interestingly, while the sample size drops dramatically and the standard errors increase correspondingly, the point estimates all become larger in absolute terms, pointing to worse match outcomes than in the baseline estimates. This pattern is very similar when we control for age with quadratic or cubic polynomials on both sides of the cutoff (columns 4 and 5), where the point estimates are similar to the linear specification with 0.5 years of bandwidth.

There is always a tradeoff between precision and bias in an RD design. And while smaller bandwidth and higher order polynomials should reduce the bias, they may do so at the cost of increases in noise and overfitting of the age controls (Lee and Lemieux 2010). Our impression from investigating the relevant figures is that these smaller bandwidths and higher order polynomials lead to such overfitting, and we prefer the more precise estimates from our baseline specification. Nevertheless, it is reassuring that these alternative specifications support the overall conclusion of negative match quality effects.

In section 3, we showed that there is a slight increase in density just to right of the two age thresholds. Furthermore, we found a small increase in the fraction of female UI recipients at the threshold. Here, we provide several methods to investigate whether this increase will affect our results. The first column of Table 5 shows the results from estimating the marginal effect of potential benefit durations on employment outcomes using our RD design pooling both thresholds, when we exclude all observations within one month of the age threshold. These restrictions reduce the sample size by about 30,000 observations. Nevertheless, we get similar effects for increases in potential benefit durations on nonemployment duration, the log post wage, and the probability of moving to a new region. The negative effect for the log wage differences is reduced from about 0.14 percent to 0.096 percent and loses statistical significance. The effect on the duration of the post-unemployment job is reduced from about 0.3 percent to about 0.14 percent but is still statistically significant from zero. Overall, while excluding the observations close to the cutoff reduces statistical power somewhat, it does not affect our overall conclusions. Column (2) of Table 5 shows how the estimates change when we

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control for a rich set of observables, including year, state, and industry fixed effects, as well as human capital and experience measures. The effects on nonemployment durations, the post-unemployment wage, and the duration of the post-unemployment job are slightly reduced but still clearly imply negative match effects and remain strongly statistically significant except for the log wage difference and post-unemployment job duration.

The increase in density just to the right of the cutoff might be driven by individuals who lose their jobs shortly before their birthdays and decide to wait before claiming UI benefits. This phenomenon would be particularly concerning if the people most likely to wait before claiming UI benefits expected to be unemployed the longest. The effect might explain higher nonemployment durations to the right of the threshold and, if these workers were also negatively selected along other dimensions, may be responsible for worse math quality measures. We employ a bounds analysis to gauge the possible magnitude of such an effect (See Manski 1990). For this purpose, we compute the excess mass of observations within a two-week window on the right side of the threshold. Once we calculated the number N of excess observations, we move the N observations with the highest outcome variable (for example, the longest nonemployment durations) to the left side of the threshold. This step provides a lower bound of the effect of potential UI durations on the outcome variable. Alternatively, we move the N observations with the lowest outcome variable to the left of the threshold to obtain an upper bound of the effect. These upper and lower bounds are very conservative in that they assume that the excess mass is due to the worst possible selection.

The lower bound on the nonemployment effect is 0.11 months, while the upper bound is 0.17 months. The bounds are thus quite wide, but they remain informative. For the wage effects (both the log post wage and the log wage difference), the lower bound is nearly twice as large (in absolute terms) as in the baseline specification and highly statistically significant. The upper bound is positive but quite close to zero and not statistically significant. The results are very similar for the two other outcomes. Thus, while the bounds analysis yields bounds that cover positive match quality effects, they are very small and statistically not

significant, while the bounds on the negative match quality side are highly significant. Given the nature of this method, we find it reassuring that even the very conservative bounds rule out larger positive match quality effects.

Column (5) shows another method robustness check to limit the effect of selective waiting before claiming UI, where we limit the sample to individuals who claim UI within two weeks of losing their job. These effects are quite similar to our baseline results.

5 The Dynamic Effects of UI Extensions on Reemployment Wages

5.1 Selection Throughout the Nonemployment Spell

In this section we provide result as to how observable characteristics and reemployment wages conditional on the duration of non-employment change as a result of the increase in potential UI durations. Throughout this section we focus on the age 42 threshold where potential UI durations increase from 12 to 18 months and we have sufficient power to observe even relatively small effects.

Figures 4 and 5 present estimates of how observables of individuals exiting unemployment change at the age discontinuity conditional on the duration of nonemployment. Vertical bars indicate that the point estimates at time t are statistically significant at the 5 percent level. Overall the figures show that there is some correlation between observables and nonemployment durations, e.g. years of schooling or fraction female is positively correlated with nonemployment duration. However there is little indication that observables are changing at the age threshold, conditional on t. While there are a few statistically significant point estimates in each figure, given that each figure is created from 24 separate point estimates, it is expected that about one to two of the estimates are statistically significant on the 5 percent level purely because of sampling variation. The one exception to this appears to be the spikes at the exhaustion point for fraction female. Individuals who are exiting from unemployment at the exhaustion points, are significantly more likely to be female. This is consistent with larger labor supply effects of UI benefits for women. The fact that the spikes

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in fraction women cancel each other out, seems to indicate that some women are simply waiting until their benefits expire before going back to work.

Figure 6 shows pre-unemployment wage and the predicted reemployment wage as summary measures for observables that are relevant to the labor market. In both of these figures the pre-unemployment wage paths and the predicted reemployment wage path are essentially unaffected by changes in potential UI durations. These figures therefore strongly support the notion that obervables are essentially uncorrelated with potential UI durations conditional on t and that therefore unobservables are also unlikely to be correlated with potential UI durations: $cov(\epsilon_i, P_i | t_i = t) = 0$. Given this, estimating equation (16) is likely going to yield consistent estimates of $\frac{dE[w^*|t]}{dP}$.

5.2 Esimates of $\frac{dE[w^*|t]}{dP}$ and $\frac{dh_t}{dP}$

Figure 7 shows estimates of the shift in the hazard rate at the age 42 discontinuity. We clearly see that the hazard rate shifts downward in response to increasing P for all $t \leq P$. This is statistically significant for nearly all point estimates, even in the first period $t = 0$, so individuals are clearly forward looking and responding to the increase in P a long time before they are running out of benefits.

Figure 8 Panel (a) shows the effect of changes in P on the reemployment wage conditional on t. Note that despite the clear shift in the hazar rate for all $t < P$, we do not observe a change in the reemployment wage $\frac{dE[w^*|t]}{dP} = 0$ for all $t < P$. The only statistically significant changes in the reemployment wages are at the exhaustion points for the two groups. Right in the period when individuals exhaust their UI benefits, reemployment wages go down relative to the other group. It is noteworthy that the two downward spikes are of very similar magnitude and essentially cancel each other out. Figure 8 Panel (b) shows an almost unchanged pattern when we control for individual heterogeneity by plotting the difference in post and pre unemployment log wage. Wages still decline by about 25 percent within the first year. Extending UI benefits does not shift the reemployment wage upwards for $t < P$

(in fact the one ststictically significant difference at $t = 3$ goes in the opposite direction). There are negative spikes at the exhaustion points. Given that the two spikes go in the same direction this still points towards selection (though selection that is not picked up by the pre unemployment wage). In fact these differences disappear once we look at women and men separately, indicating that the negative wage spikes are indeed driven by more women exiting at the exhaustion points.

We thus have evidence that $\frac{dE[w^*|t]}{dP} = 0$ for all $t < P$, and $\frac{dh_t}{dP} < 0$. Thus the conditions for Proposition 1 apply and we thus have that the decline in the reemployment wage path is entirely due to the decline in the wage offer distribution. Note however that we do not have a good estimate of the decline in the reemployment wage path yet, since the observed change of the reemployment wage with the duration of nonemployment may still be due to selection. We can now however use the result in Proposition 3, to estimate the slope of the reemployment wage path using the IV estimator: $\frac{dE[w^*|t]}{d\mu_t}\Delta_t\mu_t = \frac{\frac{dE[w^*]}{dP}}{\frac{dD}{dP}}$.

Table 8 shows 2SLS results for the effect of nonemployment durations on reemployment wages using extensions in potential UI durations as an instrument. The first stage shows the first stage of the regression (the effect of nonemployment durations on wages), which easily passes weak instrument concerns. The second column shows the reduced form that effects, which correspond to the baseline estimates above. Column 3 reports the 2SLS estimate of nonemployment durations on wages. We find that $\Delta_t E[w^*|t] = -1.0\%$. One additional month of nonemployment lowers reemployment wages by about 1 percent.

Furthermore if the support of wage offers is convex and given Proposition 3, we know that $\frac{dE[w^*|t]}{dP} = 0$ can only be the case if the reservation wage is not binding, i.e. all wage offers are above the reservation wage of unemployed individuals.

5.3 Bounding the change in the wage offer distribution if the reemployment wage path shifts

Our analysis so far was based on the empirical result that $\frac{dE[w^*|t]}{dP}$ is approximately zero. It is difficult to empirically establish that a moment is indeed exactly zero. Fortunately we can still derive meaningful results about the change in the wage offer distribution for the case that $\frac{dE[w^*|t]}{dP} \neq 0$.

Suppose we want to know what we can learn about the change in the wage offer distribution if we don't assume that the empirical part shows that $\frac{dE[w^*|t]}{dP} = 0$. For example suppose we have an estimate of the $\frac{dE[w^*|t]}{dP}$ with a standard error and we want to make inference on $\pi_1 = \frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t$ under the assumption that the true $\frac{dE[w^*|t]}{dP}$ is at the upper bound of the estimated confidence interval.

To estimate this we run the following regression:

$$w_i^* = \delta P_i + \sum_{t=1}^{T} \theta_t + F(age_i) + \epsilon_i \tag{18}$$

where w_i* is the observed reemployment wage and θ_t are time dummies for the duration of non-employment. We are still estimating the regression in the RD setting and therefore controlling for age at the time of entering unemployment. Thus we get an estimate $\hat{\delta} = \frac{dE[\hat{w}^*|t]}{dP}$.

One can write the decomposition of the average wage effect as:

$$
\begin{aligned}
\frac{dE[w^*]}{dP} &= \sum_{t=0}^{\infty} \left[\frac{dE[w^*|t]}{dP} g(t) \right] + \left[\frac{dE[w^*|t]}{dP} \left(\frac{dV_t^u}{dP} \rho \right)^{-1} \Delta_t \phi_t + \frac{dE[w^*|t]}{d\mu_t} \Delta_t \mu_t \right] \frac{dD}{dP} \\
&= \delta + \left[\delta \frac{\Delta_t \phi_t}{d\phi_t/dP} + \pi_1 \right] \frac{dD}{dP} \\
&= \delta + \left[\delta \frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}} + \pi_1 \right] \frac{dD}{dP}
\end{aligned}
$$

and some rearranging yields the slope of the wage offer distribution as a function of the

IV estimator from above plus a term that depends on δ and the ratio of the change in the value of unemployment over time, relative to the change in the value of unemployment when potential UI benefits are extended by one month:

$$\pi_1 = \frac{\frac{dE[w^*]}{dP}}{\frac{dD}{dP}} - \delta \left[\frac{1}{\frac{dD}{dP}} + \frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}} \right]$$

So given that we can estimate δ, $\frac{dE[w^*]}{dP}$ and $\frac{dD}{dP}$, we have to somehow say something about $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}} < 0$ in order to calculate π_1. To get a sense for $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}}$, notice that if everything were stationary except for the benefits expiring at time P, then $\frac{dV_t^u}{dt} = -\frac{dV_t^u}{dP}$ for $t < P$, since reducing benefits by one month has the same effect on the value function as moving forward one more month (in unemployment). In practice it is probably true that $\frac{dV_t^u}{dt} < -\frac{dV_t^u}{dP}$, that is the value function declines faster than what one would expect simply from moving one month closer to the exhaustion date. This would be because skills are depreciating, people run out of savings, the cost of job search may increase (or job search becomes less effective), etc. We therefore can be pretty confident that $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}} < -1$

Note that $\frac{1}{\frac{dD}{dP}} \approx 7$. Therefore the IV estimate provides an upper bound (but it's a negative number, so a lower bound of the absolute value) of the change in the wage offer distribution π_1, as long as $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}} > -6$. Only for a very strong decline in the value function $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}} < -6$ would the IV estimate be biased downward (or the absolute value be biased upward).

Table 9 shows different estimates for δ. The first column shows results for estimating equation (18). Column (1) shows the results controlling for a linear effect of nonemployment duration. This reduces statistically significant negative wage effect of an additional month of potential UI from -0.13% for the 12 to 18 month discontinuity to very close to zero (point estimate 0.012% with a standard error of 0.048%). If we control more flexibly for the nonemployment duration effect (Columns 2 and 3), the point estimate is even closer to 0. For the 12 to 18 month discontinuity the point estimate becomes in fact negative -0.0039% [SE 0.048%], while for the pooled sample it is positive but extremely small: 0.0038% [SE

0.042%] and statistically indistinguishable from zero despite very precise estimates. Taking the pooled sample point estimate as the most conservative one and taking the range of the confidence interval, we get that the true δ likely lies somewhere between -0.038% and 0.046%.

Table 10 shows the implied π_1 for various values of $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}}$ and δ. Essentially as long as δ is small (i.e. close to the estimated range in Table 9 which is always clearly less then 0.1%) or $\frac{\frac{dV_t^u}{dt}}{\frac{dV_t^u}{dP}}$ is not too high (between -2 and -8 or so) we get values for the change in the slope of the wage offer distribution that are quite close to the IV estimate or even smaller. For example for the upper bound of the confidence interval for $\hat{\delta} = 0.046$ the range of slopes for the wage offer distribution is between -1.4% (actually even smaller than the IV estimate) to -1.0%, just slightly larger than the IV estimate of -1.1% decline in wage offers per month.

6 Discussion

Overall, all measures imply that longer potential UI durations negatively affect reemployment wages through a decline of either the actual productivity (skill depreciation) or perceived productivity (stigmatization) of workers. Our results indicate that the reservation wage is not binding and thus has no impact on post-unemployment wages. The IV estimate of the decline in offered wages is quite large (though only half as large as the naiv OLS estimate): one month out of work reduces productivity by about 0.9 percent, i.e. a worker who is unemployed for 1 year will earn 11 percent lower wages due to the decline in the offered wages. This points to very high costs of long-term unemployment.

While our results may seem surprising and do indeed imply large decreases in (actual or perceived) productivity due to time spend out of work, they are broadly consistent with the previous literature. Other papers that have estimated the wage effect of increases in potential UI durations have found similar point estimates (though generally with less precision) as we do. For example Card, Chetty and Weber (2007) found a negative point estimate, quite comparable when rescaled to a marginal effect. Similarly van Ours and Vodopivec (2008) and Centeno and Novo (2009) find negative effects of similar magnitude.

One important contribution relative to these papers is that we have enough statistical power to reject positive wage effects, and that we provide the dynamic results that allow us to separate the wage offer and the reservation wage effect. Our results are also consistent with structural estimates in van den Berg (1990) who found that most job offers are indeed accepted and that unemployed workers do not seem to reject many jobs based on wages. The result is also in line with DellaVigna and Paserman (2005) who calibrate a model similar to our model here and find that very few wage offers fall below the reservation wage. Finally our results are consistent with the recent paper by Krueger and Muller (2011) using time use data, where self reported reservation wages stay remarkably constant over time, while the decrease in the hazard rate throughout the unemployment spell is explained by unemployed workers lowering their search efforts dramatically.

The possibility that UI benefits increase job matches is often brought forward in support of existing UI sytem. The possibility that UI benefits increase job matches is often brought forward in support of existing UI systems. Do our results imply that longer potential UI durations are costly to society because of lower match quality? Not necessarily: If individuals reap all the benefits from job matches and thus get all the surplus from higher match quality, then the effects of potential UI durations on match quality can be ignored from a social welfare perspective. The reason is essentially that optimizing individuals choose a reservation wage and search intensity that is optimal, which implies, by the first order conditions, that marginal changes in their behavior have no impact on their welfare. If the workers reap all the surplus from a job match, increasing this surplus through changes in search behavior can therefore not affect social welfare and only the effects on nonemployment are relevant.[14] This situation is different, however, if workers do not reap all the benefits of better matches—for example, because the surplus is shared with the employer or because the government receives taxes. To our knowledge, literature has not explored the extent to which negative match quality effects would affect the optimal design of the UI system, as, for example, in the

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[14]This is essentially an application of the envelope theorem. See Chetty (2008) and Schmieder et al. (2011) for details.

typical Baily-Chetty type welfare formula (See Baily 1978, Kiley 2003, Shimer and Werning 2007, Chetty 2008, Sanchez 2008, Kroft and Notowidigdo 2010 or Schmieder, von Wachter and Bender 2011). Therefore, this area is an important avenue for future research.

Even if the implications for optimal UI policy design are not obvious, the results imply very high costs of long-term unemployment and therefore suggest the importance of policies that avoid long unemployment spells to avoid the loss in human capital, such as job search assistance programs, job training programs or macroeconomic stabilization policies.

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Table 1: Smoothness of Predetermined Variables around Age Thresholds

	(1) Years of Education	(2) Female	(3) Foreign Citizen	(4) Tenure Last Job	(5) Experience Last Job	(6) Pre Wage
Increase in Potential UI Dur. from 12 to 18 Months						
D(Age above Cutoff)	0.030	0.0086	0.0038	0.0039	-0.038	0.12
	[0.014]*	[0.0028]**	[0.0020]	[0.025]	[0.049]	[0.18]
Observations	510955	510955	510955	510955	510955	480724
Mean of Dep. Var.	11.0	0.36	0.10	2.69	10.8	70.8
Increase in Potential UI Dur. from 18 to 22 Months						
D(Age above Cutoff)	-0.0049	0.0036	-0.0020	-0.019	0.0044	0.053
	[0.013]	[0.0027]	[0.0022]	[0.027]	[0.047]	[0.19]
Observations	501282	501282	501282	501282	501282	469627
Mean of Dep. Var.	10.9	0.37	0.11	2.69	10.8	70.8
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)						
D(Age above Cutoff)	0.013	0.0061	0.00092	-0.0074	-0.017	0.089
	[0.0091]	[0.0020]**	[0.0017]	[0.018]	[0.033]	[0.13]
Observations	1012237	1012237	1012237	1012237	1012237	950351
Mean of Dep. Var.	10.9	0.37	0.10	2.69	10.8	70.8

Notes: Standard errors clustered on day relative to cutoff level (* P<.05, ** P<.01)).
The sample are individuals who started receiving unemployment insurance between 1987 and 1999 within 2 years from the age thresholds. Each coefficient is from a separate regression discontinuity model with the dependent variable given in the column heading. The first panel shows the increase at the discontinuity at the age 42 threshold (where potential UI durations increase from 12 to 18 months). The second panel shows the increase at the age 44 threshold (where potential UI durations increase from 18 to 22 months). The third panel pools both thresholds. The models control for linear splines in age with different slopes on each side of the cutoff.

Table 2: The Effect of Potential UI Durations on Non-employment Duration and the Post Unemployment Wage

	(1) ALG Duration	(2) Non-Emp Duration	(3) Ever emp. again	(4) Log Post Wage	(5) Log Wage Difference
Increase in Potential UI Dur. from 12 to 18 Months					
D(Age above Cutoff)	1.77	0.88	-0.0094	-0.0078	-0.0070
	[0.034]**	[0.080]**	[0.0020]**	[0.0030]**	[0.0029]*
$\frac{dy}{dP}$	0.29	0.15	-0.0016	-0.0013	-0.0012
	[0.0057]**	[0.013]**	[0.00033]**	[0.00050]**	[0.00049]*
Observations	510955	510955	510955	437182	420311
Mean of Dep. Var.	7.57	14.8	0.86	4.01	-0.14
Increase in Potential UI Dur. from 18 to 22 Months					
D(Age above Cutoff)	1.01	0.42	-0.0056	-0.0032	-0.0047
	[0.045]**	[0.084]**	[0.0022]*	[0.0030]	[0.0028]
$\frac{dy}{dP}$	0.28	0.12	-0.0016	-0.00089	-0.0013
	[0.013]**	[0.023]**	[0.00061]*	[0.00084]	[0.00079]
Observations	501282	501282	501282	417324	401778
Mean of Dep. Var.	9.21	15.7	0.83	4.01	-0.15
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)					
D(Age above Cutoff)	1.39	0.65	-0.0075	-0.0055	-0.0059
	[0.028]**	[0.059]**	[0.0015]**	[0.0021]**	[0.0021]**
$\frac{dy}{dP}$	0.30	0.14	-0.0016	-0.0012	-0.0012
	[0.0060]**	[0.012]**	[0.00031]**	[0.00044]**	[0.00044]**
Observations	1012237	1012237	1012237	854506	822089
Mean of Dep. Var.	8.38	15.2	0.85	4.01	-0.14

Notes: Standard errors clustered on day relative to cutoff level (* P<.05, ** P<.01)). The sample are individuals who started receiving unemployment insurance between 1987 and 1999 within 2 years from the age thresholds. Each coefficient is from a separate regression discontinuity model with the dependent variable given in the column heading. The first panel shows the increase at the discontinuity at the age 42 threshold (where potential UI durations increase from 12 to 18 months). The second panel shows the increase at the age 44 threshold (where potential UI durations increase from 18 to 22 months). The third panel pools both thresholds. The models control for linear splines in age with different slopes on each side of the cutoff.

Table 3: The Effect of Potential UI Durations on Other Match Quality Measures

	(1) Log wage 5 years after start of UI	(2) Employed 5 years after start of UI	(3) Receiving UI 5 years after start of UI	(4) Duration of post unemp job in years	(5) Move to different county to take up job after unemp	(6) Post unemp job is different industry	(7) Post unemp job is different occupation
Increase in Potential UI Dur. from 12 to 18 Months							
$\frac{dy}{dP}$	-0.00015	-0.00075	0.00092	-0.012	0.00011	0.0012	0.0018
	[0.00061]	[0.00046]	[0.00032]**	[0.0049]*	[0.00049]	[0.00050]*	[0.00051]**
Observations	266147	510955	510955	437899	437690	425131	437899
Mean of Dep. Var.	4.13	0.52	0.15	2.91	0.42	0.69	0.61
Increase in Potential UI Dur. from 18 to 22 Months							
$\frac{dy}{dP}$	0.00072	-0.0020	0.0013	-0.012	-0.00028	0.00087	0.0018
	[0.0011]	[0.00079]*	[0.00059]*	[0.0081]	[0.00089]	[0.00083]	[0.00086]*
Observations	249436	501282	501282	418041	417849	405748	418041
Mean of Dep. Var.	4.11	0.50	0.17	2.93	0.41	0.69	0.61
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)							
$\frac{dy}{dP}$	0.00017	-0.0012	0.0011	-0.012	-0.000036	0.0011	0.0018
	[0.00056]	[0.00042]**	[0.00031]**	[0.0043]**	[0.00047]	[0.00045]*	[0.00047]**
Observations	515583	1012237	1012237	855940	855539	830879	855940
Mean of Dep. Var.	4.12	0.51	0.16	2.92	0.41	0.69	0.61

Notes: Standard errors clustered on day relative to cutoff level (* P<.05, ** P<.01)).
The sample are individuals who started receiving unemployment insurance between 1987 and 1999 within 2 years from the age thresholds. Each coefficient is from a separate regression discontinuity model with the dependent variable given in the column heading. The first panel shows the increase at the discontinuity at the age 42 threshold (where potential UI durations increase from 12 to 18 months). The second panel shows the increase at the age 44 threshold (where potential UI durations increase from 18 to 22 months). The third panel pools both thresholds. The models control for linear splines in age with different slopes on each side of the cutoff.

Table 4: The Effect of Potential UI Durations on Nonemployment and Match Quality: Varying Bandwidth and Polynomial Order

	(1) Bandwidth: 2 Years	(2) Bandwidth: 1 Year	(3) Bandwidth: 0.5 Years	(4) Quadratic Age Control	(5) Cubic Age Control
Non-employment duration					
$\frac{dy}{dP}$	0.14	0.18	0.20	0.16	0.21
	[0.012]**	[0.018]**	[0.026]**	[0.019]**	[0.026]**
Observations	1012237	506070	252984	1012237	1012237
Mean of Dep. Var.	15.2	15.2	15.2	15.2	15.2
Log post wage					
$\frac{dy}{dP}$	-0.0012	-0.0021	-0.0022	-0.0015	-0.0028
	[0.00044]**	[0.00063]**	[0.00096]*	[0.00068]*	[0.00095]**
Observations	854506	427846	214043	854506	854506
Mean of Dep. Var.	4.01	4.01	4.01	4.01	4.01
Log wage difference					
$\frac{dy}{dP}$	-0.0012	-0.0021	-0.0029	-0.0016	-0.0035
	[0.00044]**	[0.00063]**	[0.00092]**	[0.00068]*	[0.00091]**
Observations	822089	411714	205934	822089	822089
Mean of Dep. Var.	-0.14	-0.14	-0.14	-0.14	-0.14
Moved to different county to takeup job after unemployment					
$\frac{dy}{dP}$	-0.000036	0.00045	0.00014	-0.00028	0.00036
	[0.00047]	[0.00067]	[0.00096]	[0.00071]	[0.00096]
Observations	855539	428368	214287	855539	855539
Mean of Dep. Var.	0.41	0.41	0.41	0.41	0.41
Duration of post unemployment job					
$\frac{dy}{dP}$	-0.012	-0.030	-0.025	-0.019	-0.037
	[0.0043]**	[0.0063]**	[0.0089]**	[0.0067]**	[0.0090]**
Observations	855940	428564	214388	855940	855940
Mean of Dep. Var.	2.92	2.92	2.91	2.92	2.92

Notes: Standard errors clustered on day relative to cutoff level (* P<.05, ** P<.01)).
The sample are individuals who started receiving unemployment insurance between 1987 and 1999. Each panel shows the increase at the age threshold of the dependent variable (given in the panel title) rescaled by the average increase in potential UI durations at the thresholds. The columns refer to different estimating the RD model with different bandwidths and controlling for different polynomials in age.

Table 5: The Effect of Potential UI Durations on Nonemployment and Match Quality: Robustness Checks

	(1) Excluding Obs within 1 month of threshold	(2) Controlling for observable characteristics	(3) Lower bound estimates	(4) Upper bound estimates	(5) Sample restricted to UI takeup within 15 days of job end
Non-employment duration					
$\frac{dy}{dP}$	0.13	0.11	0.11	0.17	0.15
	[0.013]**	[0.012]**	[0.014]**	[0.014]**	[0.013]**
Observations	969810	893505	1012237	1012237	874684
Mean of Dep. Var.	15.2	15.2	15.2	15.2	14.9
Log post wage					
$\frac{dy}{dP}$	-0.0011	-0.00090	-0.0028	0.0012	-0.0014
	[0.00045]*	[0.00038]*	[0.00057]**	[0.00061]*	[0.00046]**
Observations	818526	771197	854506	854506	745167
Mean of Dep. Var.	4.01	4.01	4.01	4.01	4.02
Log wage difference					
$\frac{dy}{dP}$	-0.00073	-0.00076	-0.0029	0.0011	-0.0014
	[0.00047]	[0.00040]	[0.00059]**	[0.00057]	[0.00044]**
Observations	787532	771197	822089	822089	724136
Mean of Dep. Var.	-0.14	-0.14	-0.14	-0.14	-0.14
Moved to different county to takeup job after unemployment					
$\frac{dy}{dP}$	0.00018	-0.00043	-0.0012	0.00076	0.000015
	[0.00051]	[0.00049]	[0.00054]*	[0.00050]	[0.00049]
Observations	819529	771827	855539	855539	745956
Mean of Dep. Var.	0.41	0.41	0.41	0.41	0.41
Duration of post unemployment job					
$\frac{dy}{dP}$	-0.0096	-0.0074	-0.037	-0.0070	-0.012
	[0.0047]*	[0.0046]	[0.0068]**	[0.0045]	[0.0048]*
Observations	819910	772129	855940	855940	746315
Mean of Dep. Var.	2.92	2.92	2.92	2.92	3.04

Notes: Standard errors clustered on day relative to cutoff level (* P<.05, ** P<.01)).
The sample are individuals who started receiving unemployment insurance between 1987 and 1999 within 2 years from the age thresholds. Each panel shows the increase at the age threshold of the dependent variable (given in the panel title) rescaled by the average increase in potential UI durations at the thresholds. The columns refere to different sample restrictions and model specifications (see text for details).

Table 6: The Effect of Potential UI Durations on Other Match Quality Measures, Pre-1999 only, Duruptake<84 days

	(1) Baseline Spec.	(2) Control for Occupation changes	(3) Control for Industry Changes
Increase in Potential UI Dur. from 12 to 18 Months			
D(Age above Cutoff)	-0.0078	-0.0064	-0.0059
	[0.0030]**	[0.0029]*	[0.0028]*
$\frac{dy}{dP}$	-0.0013	-0.0011	-0.00099
	[0.00050]**	[0.00048]*	[0.00047]*
Effect relative to mean	-0.0019	-0.0016	-0.0015
Observations	437182	398829	398829
Mean of Dep. Var.	4.01	4.01	4.01
Increase in Potential UI Dur. from 18 to 22 Months			
D(Age above Cutoff)	-0.0032	-0.0013	-0.0010
	[0.0030]	[0.0028]	[0.0029]
$\frac{dy}{dP}$	-0.00089	-0.00037	-0.00028
	[0.00084]	[0.00080]	[0.00080]
Effect relative to mean	-0.00080	-0.00033	-0.00025
Observations	417324	380021	380021
Mean of Dep. Var.	4.01	4.01	4.01
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)			
D(Age above Cutoff)	-0.0055	-0.0039	-0.0035
	[0.0021]**	[0.0019]*	[0.0019]
$\frac{dy}{dP}$	-0.0012	-0.00082	-0.00073
	[0.00044]**	[0.00041]*	[0.00041]
Effect relative to mean	-0.0014	-0.00096	-0.00086
Observations	854506	778850	778850
Mean of Dep. Var.	4.01	4.01	4.01

Notes: Coefficients from RD regressions. Local linear regressions (different slopes) on each side of cutoff. Standard errors clustered on day level (* P<.05, ** P<.01)).

Table 7: The Effect of Potential UI Durations on Other Match Quality Measures, Pre-1999 only, Duruptake<84 days

	(1) Log Post Wage	(2) Mean wage at post-UI estab.	(3) Size (# emp) of post-UI estab.	(4) Log wage 1 year after reemployment	(5) Log wage 2 years after reemployment	(6) Log wage 3 years after reemployment	(7) Log wage 5 years after reemployment
Increase in Potential UI Dur. from 12 to 18 Months							
D(Age above Cutoff)	-0.0078	0.20	6.23	-0.0086	-0.0049	-0.0056	-0.0054
	[0.0030]**	[0.12]	[7.21]	[0.0035]*	[0.0038]	[0.0040]	[0.0043]
$\frac{dy}{dP}$	-0.0013	0.033	1.04	-0.0014	-0.00082	-0.00093	-0.00089
	[0.00050]**	[0.020]	[1.20]	[0.00059]*	[0.00064]	[0.00066]	[0.00071]
Effect relative to mean	-0.0019	0.0031	0.027	-0.0022	-0.0012	-0.0014	-0.0013
Observations	437182	410127	415648	382089	361787	345073	311833
Mean of Dep. Var.	4.01	63.7	233.3	3.95	3.94	3.95	3.97
Increase in Potential UI Dur. from 18 to 22 Months							
D(Age above Cutoff)	-0.0032	0.18	-5.58	0.00047	-0.0030	0.00046	-0.0048
	[0.0030]	[0.13]	[7.26]	[0.0037]	[0.0040]	[0.0042]	[0.0045]
$\frac{dy}{dP}$	-0.00089	0.051	-1.56	0.00013	-0.00084	0.00013	-0.0013
	[0.00084]	[0.035]	[2.03]	[0.0010]	[0.0011]	[0.0012]	[0.0012]
Effect relative to mean	-0.00080	0.0029	-0.025	0.00012	-0.00077	0.00012	-0.0012
Observations	417324	391431	396776	365900	345703	329278	296141
Mean of Dep. Var.	4.01	63.4	222.9	3.93	3.92	3.93	3.94
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)							
D(Age above Cutoff)	-0.0055	0.19	0.40	-0.0041	-0.0040	-0.0025	-0.0051
	[0.0021]**	[0.088]*	[5.12]	[0.0024]	[0.0028]	[0.0028]	[0.0030]
$\frac{dy}{dP}$	-0.0012	0.040	0.085	-0.00088	-0.00084	-0.00054	-0.0011
	[0.00044]**	[0.019]*	[1.09]	[0.00052]	[0.00059]	[0.00059]	[0.00064]
Effect relative to mean	-0.0014	0.0030	0.0018	-0.0011	-0.0010	-0.00065	-0.0013
Observations	854506	801558	812424	747989	707490	674351	607974
Mean of Dep. Var.	4.01	63.5	228.2	3.94	3.93	3.94	3.96

Notes: Coefficients from RD regressions. Local linear regressions (different slopes) on each side of cutoff. Standard errors clustered on day level (* P<.05, ** P<.01)).

Table 8: The Effect of Time Out of Work on Reemployment Wages, Pre-1999

	(1) First Stage Nonemp Dur	(2) Reduced Form Reemp Wage	(3) OLS Reemp Wage	(4) 2SLS Reemp Wage
Increase in Potential UI Dur. from 12 to 18 Months				
Potential UI Dur in months	0.13	-0.0013	.	.
	[0.012]**	[0.00050]**	.	.
Nonemp Dur in months	.	.	-0.011	-0.010
	.	.	[0.000070]**	[0.0038]**
Observations	437182	437182	437182	437182
Mean of Dep. Var.	11.3	4.01	4.01	4.01
Increase in Potential UI Dur. from 18 to 22 Months				
Potential UI Dur in months	0.095	-0.00089	.	.
	[0.021]**	[0.00084]	.	.
Nonemp Dur in months	.	.	-0.011	-0.0094
	.	.	[0.000069]**	[0.0086]
Observations	417324	417324	417324	417324
Mean of Dep. Var.	11.7	4.01	4.01	4.01
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)				
Potential UI Dur in months	0.12	-0.0012	.	.
	[0.011]**	[0.00044]**	.	.
Nonemp Dur in months	.	.	-0.011	-0.010
	.	.	[0.000049]**	[0.0036]**
Observations	854506	854506	854506	854506
Mean of Dep. Var.	11.5	4.01	4.01	4.01

Notes: Coefficients from RD regressions. Local linear regressions (different slopes) on each side of cutoff. Standard errors clustered on day level (* P<.05, ** P<.01)).

Table 9: The Effect of Potential UI Durations on Reemployment Wages Conditional on Nonemployment Durations, Pre-1999

	(1) Reemp Log Wage Contrl. for Nonemp Dur	(2) Reemp Log Wage Contrl. for Nonemp Dur Polyn.	(3) Reemp Log Wage Contrl. for Nonemp Dur Dummies
Increase in Potential UI Dur. from 12 to 18 Months			
D(Age above Cutoff)	0.00070	-0.000082	-0.00024
	[0.0029]	[0.0029]	[0.0029]
Potential UI Dur	0.00012	-0.000014	-0.000039
	[0.00048]	[0.00048]	[0.00048]
Observations	437182	437182	437182
Mean of Dep. Var.	4.01	4.01	4.01
Increase in Potential UI Dur. from 18 to 22 Months			
D(Age above Cutoff)	0.00069	0.00032	0.00029
	[0.0029]	[0.0029]	[0.0029]
Potential UI Dur	0.00019	0.000090	0.000082
	[0.00082]	[0.00082]	[0.00081]
Observations	417324	417324	417324
Mean of Dep. Var.	4.01	4.01	4.01
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)			
D(Age above Cutoff)	0.00072	0.00018	0.00018
	[0.0020]	[0.0020]	[0.0020]
Potential UI Dur	0.00015	0.000038	0.000038
	[0.00042]	[0.00042]	[0.00042]
Observations	854506	854506	854506
Mean of Dep. Var.	4.01	4.01	4.01

Notes: Coefficients from RD regressions. Local linear regressions (different slopes) on each side of cutoff. Standard errors clustered on day level (* P<.05, ** P<.01)).
Table was generated on 11 Jun 2012 at 15:35:23.

Table 10: Slope of Mean Wage Offers as Function of (dVu/dt)/(dVu/dP) and the effect of UI extensions conditional on duration of nonemployment $dE[w|t]/dP$

$dE[w|t]/dP$ in percent	$\frac{dVu/dt}{dVu/dP}$								
	-1	-2	-3	-4	-5	-6	-7	-8	-9
0	-0.011	-0.011	-0.011	-0.011	-0.011	-0.011	-0.011	-0.011	-0.011
0.045	-0.014	-0.013	-0.013	-0.012	-0.012	-0.012	-0.011	-0.011	-0.010
0.1	-0.017	-0.016	-0.015	-0.014	-0.013	-0.012	-0.011	-0.010	-0.009
0.2	-0.023	-0.021	-0.019	-0.017	-0.015	-0.013	-0.011	-0.009	-0.007
0.3	-0.029	-0.026	-0.023	-0.020	-0.017	-0.014	-0.011	-0.008	-0.005
0.4	-0.036	-0.032	-0.028	-0.024	-0.020	-0.016	-0.012	-0.008	-0.004
0.5	-0.042	-0.037	-0.032	-0.027	-0.022	-0.017	-0.012	-0.007	-0.002
0.6	-0.048	-0.042	-0.036	-0.030	-0.024	-0.018	-0.012	-0.006	0.000
0.7	-0.054	-0.047	-0.040	-0.033	-0.026	-0.019	-0.012	-0.005	0.002
0.8	-0.060	-0.052	-0.044	-0.036	-0.028	-0.020	-0.012	-0.004	0.004
0.9	-0.066	-0.057	-0.048	-0.039	-0.030	-0.021	-0.012	-0.003	0.006
1.1	-0.072	-0.062	-0.052	-0.042	-0.032	-0.022	-0.012	-0.002	0.008

Notes: The table shows the implied slope of the mean wage offer distribution if the effect of potential UI durations on reemployment wages conditional on nonemployment durations is not equal to zero $dE[w|t]/dP$. Rows show the implied slope for different values of $dE[w|t]/dP$ and columns for different values of $\frac{dVu/dt}{dVu/dP}$. The preferred point Estimate for $dE[w|t]/dP$ is 0.0038% (from last column and bottom panel of Table 10).
The upper bound of the 95% confidence interval for $dE[w|t]/dP$ is 0.045%.

Figure 1: Potential Unemployment Insurance Durations by Period for Workers with High Prior Labor Force Attachment

Notes: The figure shows density of spells by age at the start of receiving unemployment insurance (i.e. the number of spells in 2 week interval age bins). The vertical lines mark age cutoffs for increases in potential UI durations at age 42 (12 to 18 months), 44 (18 to 22 months). The sample are unemployed worker claiming UI between July 1987 and March 1999 who had worked for at least 44 months in the last 7 years without intermittent UI spell.

Figure 2: The Effect of Extended Potential UI Durations on Benefit and Nonempoyment Durations
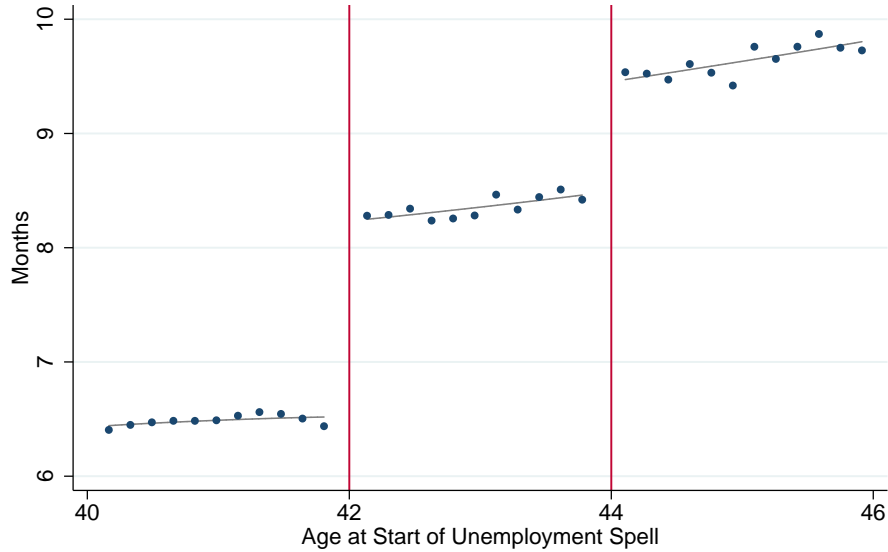
(a) Months of receiving UI benefits
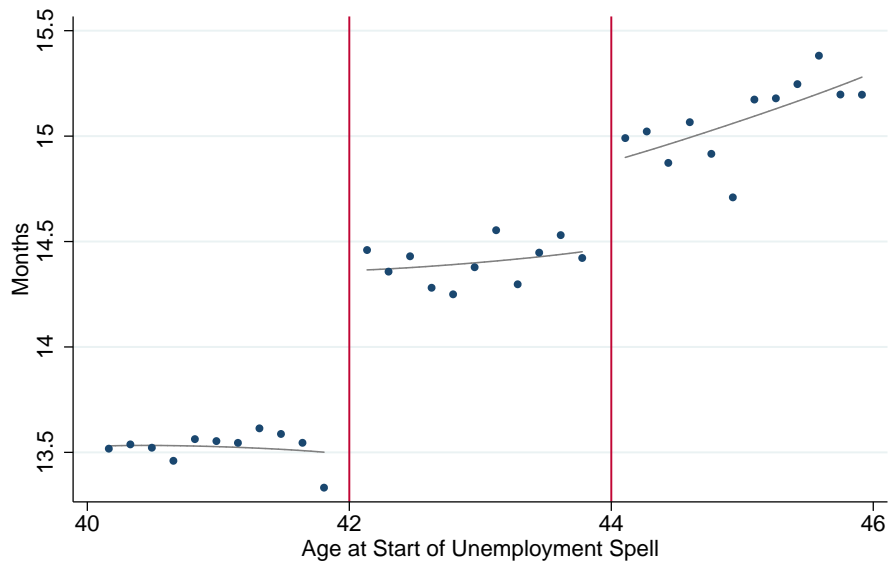
(b) Months of Nonemployment

Notes: The top figure shows average durations of receiving UI benefits by age at the start of unemployment insurance receipt. The bottom figure shows average nonemployment durations for these workers, where nonemployment duration is measured as the time until return to a job and is capped at 36 months. Each dot corresponds to an average over 120 days. The continuous lines represent quadratic polynomials fitted separately within the respective age range. The vertical lines mark age cutoffs for increases in potential UI durations at age 42 (12 to 18 months), 44 (18 to 22 months).

Figure 3: The Effect of Extended Potential UI Durations on Post Unemployment Wages
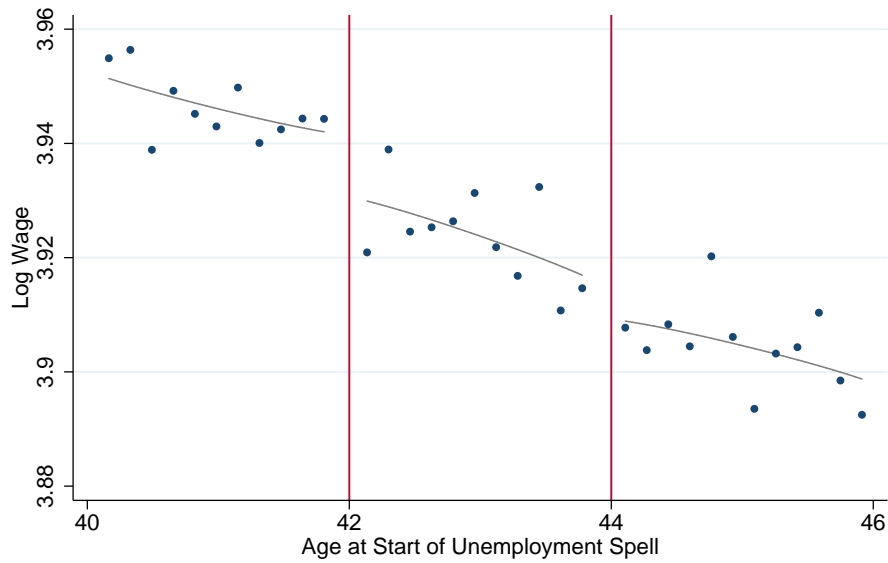
(a) Log post unemployment wage
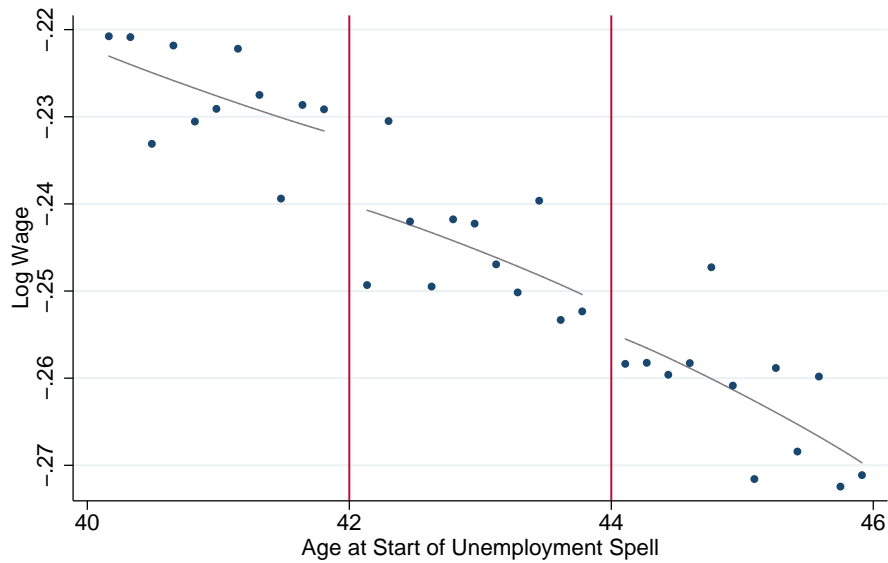
(b) Log wage difference (pre unemployment minus post unemployment)

Notes: The top figure shows average post unemployment log wages by age at the start of unemployment insurance receipt. The bottom figure shows average difference in the pre and post unemployment log wage for these workers. Each dot corresponds to an average over 120 days. The continuous lines represent quadratic polynomials fitted separately within the respective age range. The vertical lines mark age cutoffs for increases in potential UI durations at age 42 (12 to 18 months), 44 (18 to 22 months).

Figure 4: The Effects of Extended Potential UI Durations on Selection throughout the Spell of Non-employment
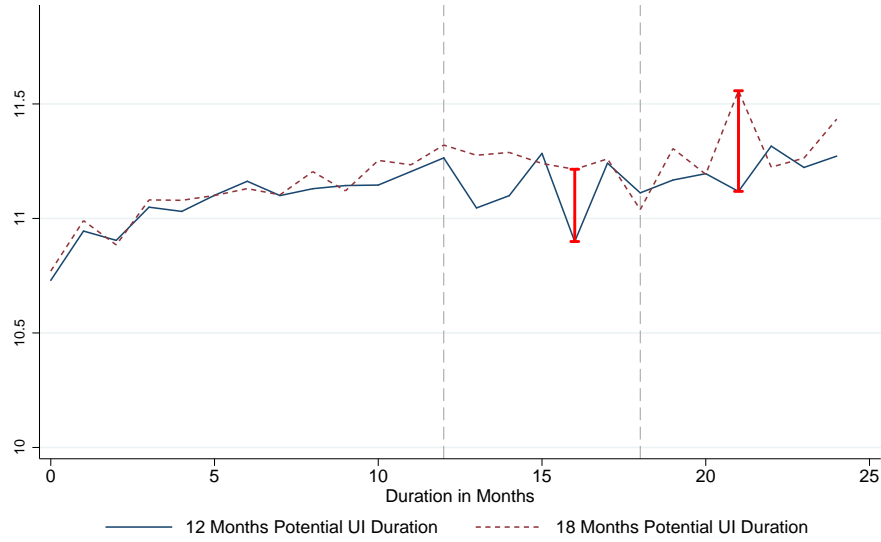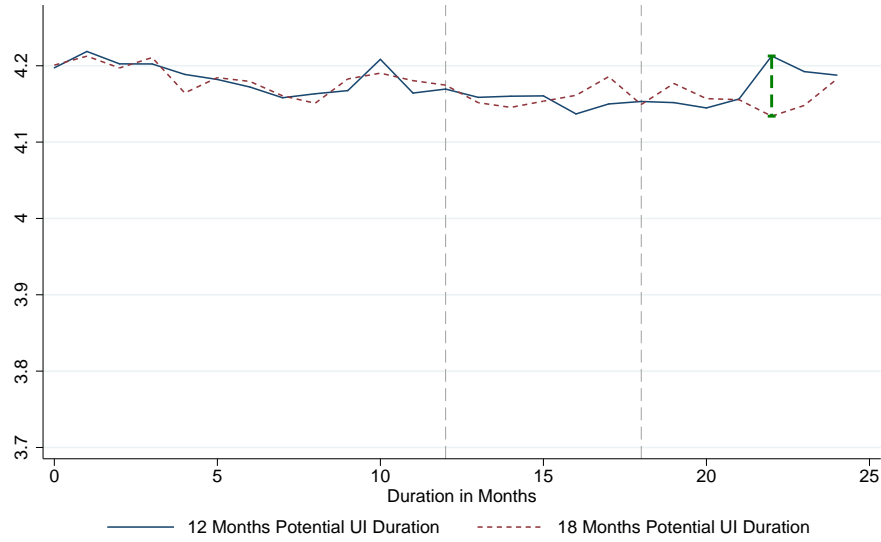

(a) Years of Schooling
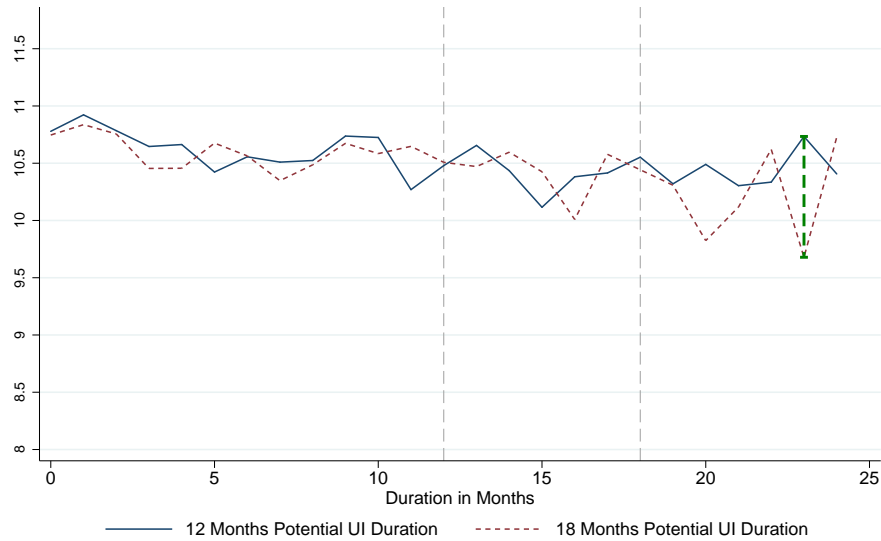
(b) Pre-unemployment Experience

Notes: The difference between the lines is estimated pointwise at each point of support using regression discontinuity estimation. Vertical bars indicate that the differences are statistically significant from each other at the five percent level. The sample are unemployed worker claiming UI between July 1987 and March 1999 who had worked for at least 36 months in the last 7 years without intermittent UI spell. For details see text.

Figure 5: The Effects of Extended Potential UI Durations on Selection throughout the Spell of Non-employment
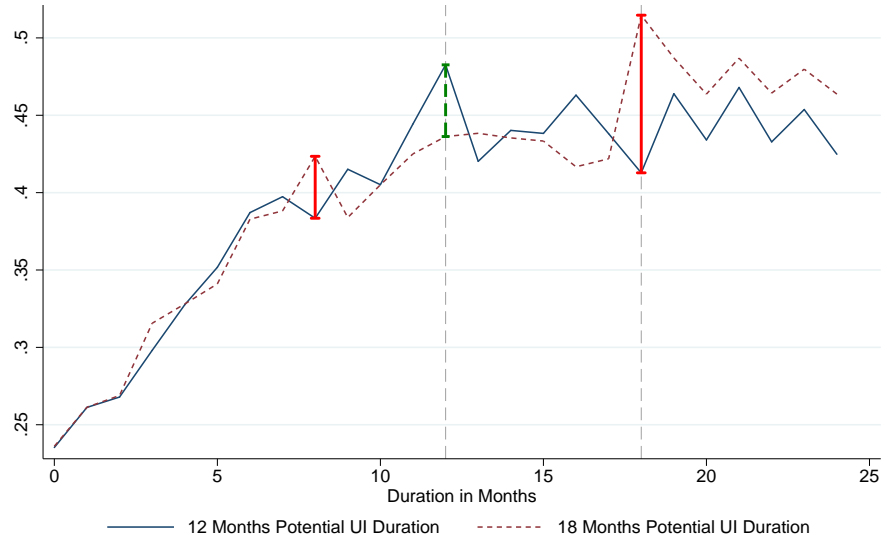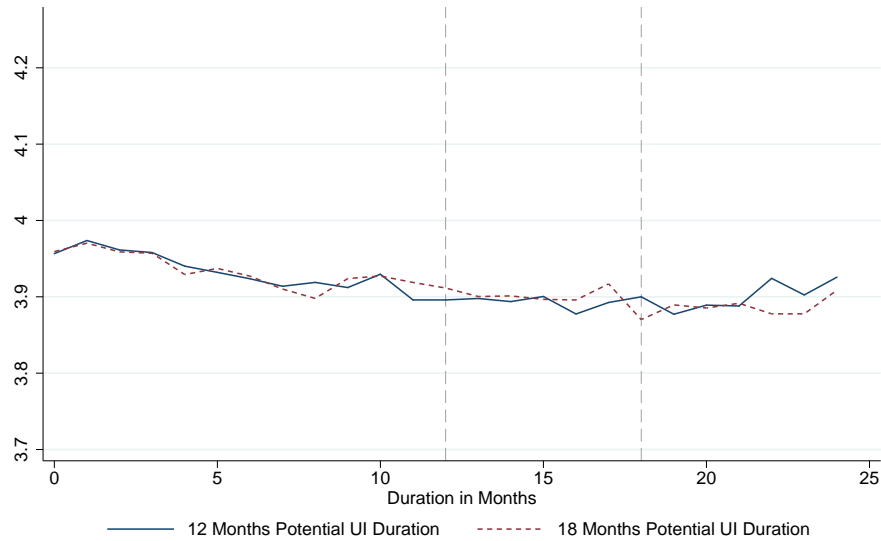

(a) Female
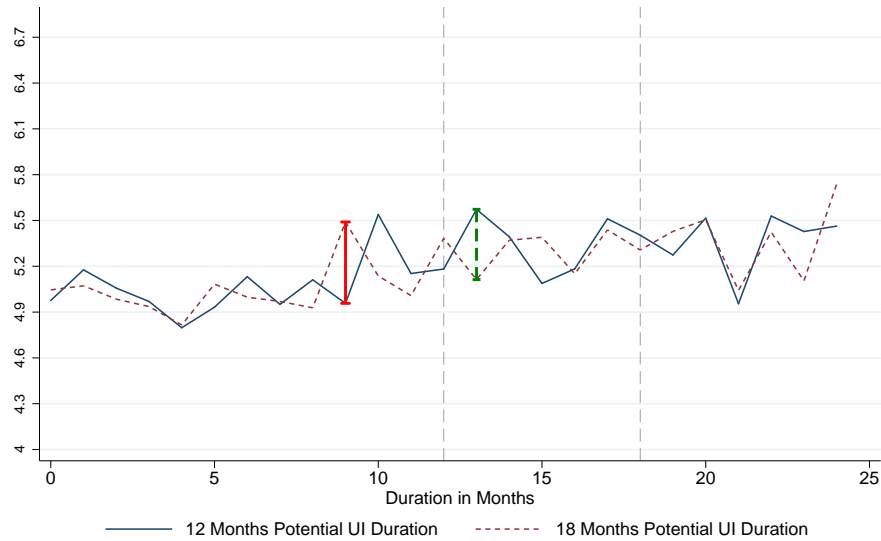
(b) Pre-unemployment Tenure

Notes: The difference between the lines is estimated pointwise at each point of support using regression discontinuity estimation. Vertical bars indicate that the differences are statistically significant from each other at the five percent level. The sample are unemployed worker claiming UI between July 1987 and March 1999 who had worked for at least 36 months in the last 7 years without intermittent UI spell. For details see text.

Figure 6: The Effects of Extended Potential UI Durations on Selection throughout the Spell of Non-employment
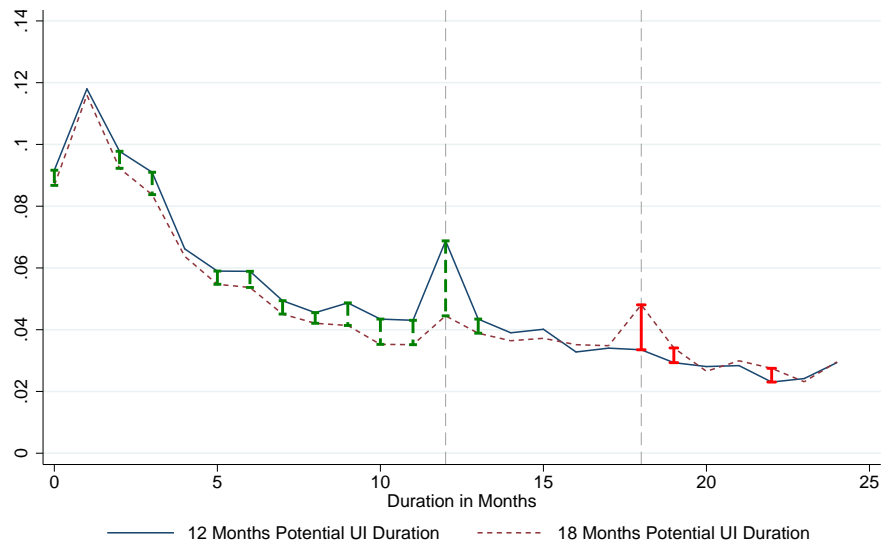

(a) Pre-unemployment log wage by time of non-emp exit

(b) Predicted reemployment log wage by time of non-emp exit

Notes: The difference between the lines is estimated pointwise at each point of support using regression discontinuity estimation. Vertical bars indicate that the differences are statistically significant from each other at the five percent level. The sample are unemployed worker claiming UI between July 1987 and March 1999 who had worked for at least 36 months in the last 7 years without intermittent UI spell. For details see text.

Figure 7: Effect of Increasing Potential Unemployment Insurance (UI) Durations from 12 to 18 Months on the Hazard Functions - Regression Discontinuity Estimate at Age 42 Discontinuity
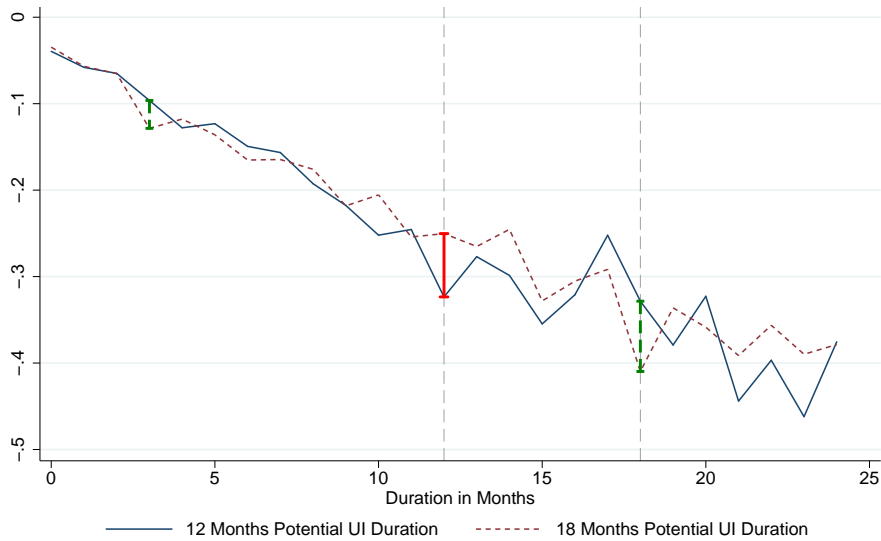

Notes: The difference between the hazard functions is estimated pointwise at each point of support using regression discontinuity estimation. Vertical bars indicate that the hazard rates are statistically significant from each other at the five percent level. The sample are unemployed worker claiming UI between July 1987 and March 1999 who had worked for at least 36 months in the last 7 years without intermittent UI spell. For details see text.

Figure 8: The Effects of Extended Potential UI Durations on Reemployment Wages throughout the Spell of Non-employment
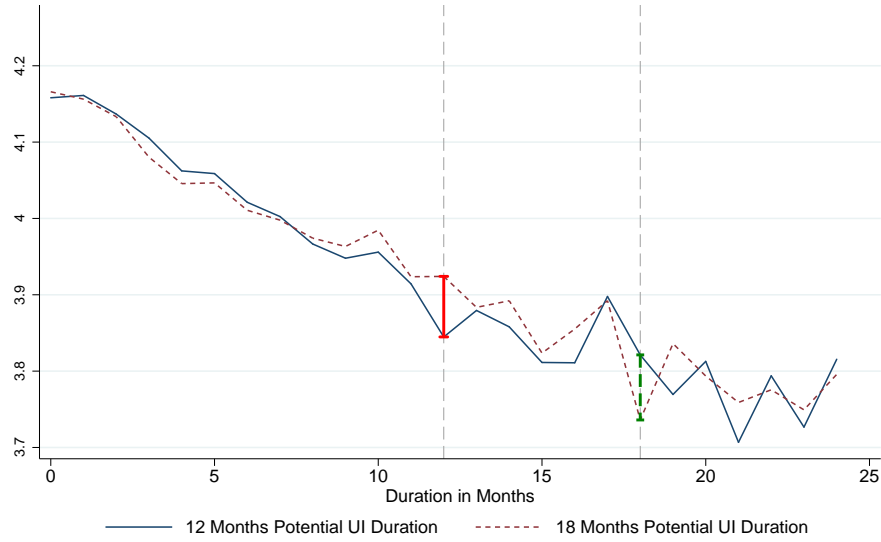
(a) Post-unemployment log wage

(b) Log wage difference (post - pre unemployment)

Notes: The difference between the reemployment wage paths is estimated pointwise at each point of support using regression discontinuity estimation. Vertical bars indicate that the differences in the reemployment wages are statistically significant from each other at the five percent level. The sample are unemployed worker claiming UI between July 1987 and March 1999 who had worked for at least 36 months in the last 7 years without intermittent UI spell. For details see text.