The Causal Effect of Unemployment Duration on Wages: Evidence from Unemployment Insurance Extensions*

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Abstract

The causal effect of nonemployment durations on reemployment wages plays an important role in evaluating the costs of unemployment for workers and the economy, and in devising appropriate policy responses. However, almost no direct empirical estimates of this effect exist. This paper presents an analytical framework for analyzing the causal effect of nonemployment duration on wages and implements it using a regression discontinuity design and a large administrative dataset from Germany. Using a job search model with heterogeneous agents, endogenous search choices, and a time-varying wage offer distribution, the paper shows that correlations of nonemployment durations and wages are not only hard to interpret because of selective exit from nonemployment, but also because both search behavior (labor supply) and wage offers (labor demand) change throughout the nonemployment spell. We show that if the path of reemployment wages throughout the nonemployment spell does not shift in response to an exogenous change in workers' outside options – in our case unemployment insurance (UI) durations – this implies reservation wages do not bind. In that case the model implies that UI extensions can be used to obtain an IV estimator that estimates a weighted sum of the slopes of the wage offer distribution at different durations. We find that UI extensions occurring with exact age in Germany had small but precisely estimated effects on reemployment wages, but that they did not affect the observed path of reemployment wages. As a result, our IV estimates point to substantial negative causal effects of nonemployment durations on wages of 0.8% per month. These findings imply high costs of long-term unemployment.

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1 Introduction

Whether and how much longer unemployment durations hurt the job prospects of unemployed workers has been a longstanding question among economists, policy makers, and the public. This issue has been actively debated at least since the lasting rise in unemployment rates and unemployment durations in Europe in the 1980s. It has received renewed attention and urgency with the staggering rise in unemployment durations in the United States in the aftermath of the 2008 recession. Persistent negative effects of rising nonemployment durations bear on the well being of nonemployed job losers, but have also been cited as potentially hurting economic recovery and growth (e.g., Blanchard and Summers 1996, Pissarides 1992, Ljungqvist and Sargent 1998, 2008, Coles and Masters 2000, Ball 2009). To the least, they are thought to possibly permanently reduce the economic potential of large groups of workers (e.g. Bernanke 2012). The causal effect of nonemployment duration on wages is also an important aspect in designing labor market policies to deal with unemployment (e.g., Shimer and Werning 2006, Pavoni and Violante 2007, Pavoni 2009, Spinnewijn 2013, Pavoni, Setty, and Violante 2013). For example, if nonemployment durations lower reemployment wages, extensions in unemployment insurance (UI) durations – the largest government program geared towards job losers in recessions – instead of helping to obtain better job matches may hurt the prospects of job losers and the economy as a whole.

While an important parameter in gauging the adverse effects of large-scale and long-term unemployment and in determining the best policy response, the causal effect of nonemployment durations on wages and other job outcomes is difficult to estimate. It is believed that those workers with longer nonemployment durations also have other, potentially unobserved characteristics that make them hard to employ and lower their wages. A common approach to deal with such selection would be to find a source of manipulation of nonemployment durations that is independent of worker attributes. An additional key difficulty that has received less attention is that it is hard to find a manipulation that would not induce *both* changes in job search behavior (i.e., a labor supply response) and changes in wage offers (driven by labor demand). As a result, even estimates free of selection cannot be easily interpreted as the causal effect of nonemployment durations on a decline in wages.

To address these issues, this paper develops a new framework for estimating the causal effect of nonemployment durations on wages in presence of worker heterogeneity, endogenous job search behavior, and wage offers that change with nonemployment durations. Based on the model the paper develops a test for whether UI extensions affect workers' reservation wages, and clarifies when UI extensions and other quasi-experimental changes in the outside option can be used as instrumental variables for nonemployment durations. The paper then implements the research strategy resulting from the theory using a discontinuity in UI durations and a large administrative data set from Germany. The results indicate small but precisely estimated negative effects of UI extension on wages that translate into substantial negative causal effects of nonemployment durations on wages. Building on our new framework, these empirical findings point towards potentially large costs of pervasive increases in long nonemployment spells.

We base our analysis of the effect of changes in the outside option of workers – such as increases in UI durations – on search behavior and wages, on a partial equilibrium, nonstationary model of job search in which unemployed workers choose reservation wages and search intensity in the face of potentially declining wage offers. The model predicts that if workers value their outside option, reemployment rates fall and reservation wages rise. A key insight from the model is that if the path of observed reemployment wages at different nonemployment durations does not shift in response to a rise in UI durations, this implies that reservation wages do not bind, at least in the part of the wage offer distribution relevant for workers' employment decisions. Hence, the only effect of nonemployment durations on wages must be from a change in the wage offer distribution. In that case, we show that UI durations can be used as instrumental variable for nonemployment duration, and that the resulting estimate is the weighted average of the slope of the wage offer distribution at different nonemployment durations.

Up until recently, in Germany UI durations were a step function of exact age at benefit claiming, such that the causal effect of UI durations on job outcomes can be estimated using a regression discontinuity design. A key feature of the German environment is that we have access to the universe of social security records with information on day-to-day nonemployment spells, exact dates of birth, as well as a broad range of worker and job characteristics. We obtain three main findings. First, we find small but precisely estimated negative effects of UI extensions on wages and other job outcomes. Second, we show that the path of reemployment wages at different nonemployment durations does not shift, implying that reservation wages do not bind. As a result, reservation wages do not contribute to declining wages over the nonemployment spell, and one can use UI extensions as valid manipulation of nonemployment durations. Third, we obtain instrumental variable (IV) estimates of the causal effect of nonemployment durations. We find that for each additional month in nonemployment duration, daily wages decline by a bit less than one percent. This effect declines over time, and is statistically indistinguishable from zero after five years. This negative effect can arise from multiple sources, including skill depreciation, stigma effects, or changes in job characteristics, something we address in our empirical analysis.

Our paper contributes to several strands of literature. Foremost, it presents both a framework for obtaining causal estimates of the effect of nonemployment durations on wages, and a new set of causal estimates, neither of which is currently available in the literature. Existing estimates are typically based on correlations of nonemployment durations and wages (e.g., Addison and Portugal 1989, Gregory and Jukes 2001), or derived from structural models (e.g., Keane and Wolpin 1997).[1] Both sets of estimates indicate effects of similar order of magnitude as we find here. An alternative approach to address the problems we discuss here is to estimate the causal effect of nonemployment durations on employer behavior using an

[1]In an exception, Edin and Gustavsson (2008) document a significant negative effect of nonemployment spells on direct measures of skills in Sweden. Estimates of the earnings losses of displaced workers have also been used to infer the correlation of nonemployment duration and wages (e.g., Neal 1995).

audit study (Kroft, Lange, and Notowidigdo 2012). However, despite other advantages, it is hard to use the audit study approach to estimate the effect of nonemployment durations on actual wage offers.

Our paper also adds to the literature estimating the effect of UI benefits on nonemployment durations, wages and other job outcomes. 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; Schmieder, von Wachter, and Bender 2012a,b; Kroft and Notowidigdo 2012), and the consumption smoothing effect of UI (for example, Gruber 1997), a much smaller literature has found mixed results regarding the effects on earnings and wages 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), Card, Chetty and Weber (2007), and Centeno and Novo (2009) used regression discontinuity designs to more clearly identify the effects and find negative impacts on wages. While these results are relatively imprecisely estimated and hence not statistically significantly different from zero, confidence intervals contain possible negative and positive values that are economically meaningful.[2]

Our findings also relate to the literature examining the properties and effects of reservation wages. The bulk of that literature is based on survey evidence containing direct information on both reservation wages and accepted wages (e.g., Feldstein and Poterba 1984, Blau and Robins 1986, DellaVigna and Paserman 2005, Krueger and Mueller 2011, Mueller 2013). In contrast, here we show that it is possible to infer about the effect of and changes in reservation wages even absent such direct information when quasi-experimental variation of the outside option is available together with information on accepted wages and nonemployment durations. Two exceptions are Hornstein, Violante, and Krueger (2011), who infer about reservation wages using flow data, and Lalive, Landais, and Zweimueller (2013), who

[2]Consistent with a negative effect of nonemployment durations, Black, Smith, Berger and Noel (2003) find positive effects on reemployment and quarterly earnings of UI recipients who are randomly assigned to (but not necessarily participate in) more intensive job search services. Meyer (1995) reports imprecisely estimated positive effects on earnings for UI recipients who receive a bonus upon faster reemployment.

replicate our approach of analyzing reemployment wage paths for Austria.

Finally, our estimates can be used to calibrate models in macroeconomics or in public finance in which the causal effect of nonemployment plays an important role. In public finance, a growing theoretical literature shows how the structure of labor market depends on the degree of wage decline with nonemployment. For example, Pavoni and Violante (2007) show that this parameter plays a key role when multiple labor market policies are chosen jointly. In macroeconomics, in prominent papers the causal effect of nonemployment on wage offers is important in explaining the persistence of aggregate unemployment rates. For example, in a series of papers Ljungqvist and Sargent (1998, 2008) and den Haan, Haefke, and Ramey (2005) show how the rate of skill depreciation can be a key determinant of the effect of economic shocks and UI benefits on equilibrium unemployment.

The next section describes the nonstationary job search model we use and derives our main theoretical results. Section 3 describes the institutional setting, the data, and the empirical framework implied by the theory. In section 4 we present the main results of how potential UI durations affect average match quality. 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 summarizes an extensive sensitivity analysis. Section 7 discusses these results and concludes.

2 Theory

We analyze a discrete time, non-stationary search model, based on Mortensen (1986) and van den Berg (1990). The model features endogenous search intensity in addition to the choice of accepting or rejecting job offers, thus allowing for UI extensions to affect nonemployment durations through changes in reservation wages as well as through changes in search effort. The distribution of wage offers may shift downward throughout the unemployment spell, which may represent skill depreciation or stigmatization. We focus on three aspects: First, we show how the effect of UI extensions on reemployment wages can be decomposed into changes in reservation wages and changes in the wage offer distribution over the nonemployment spell.[3] Second, we use the model to clarify how the effect of UI extensions on the reemployment wage *path* (i.e., reemployment wages conditional on the time of exiting unemployment) is informative about the response of reservation wages to UI extensions and with it about their effect on the reemployment wage path. Third, we show under what conditions it is possible to identify the change in the wage offer distribution over the nonemployment spell – and hence the causal effect of nonemployment durations on wages – separately from the reservation wage effect using UI extensions as a form of exogenous variation.

2.1 Setup of Model

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 a workers receive UI benefits b_t and choose search intensity λ_t, which is normalized to be equal to the probability of receiving a job offer in that period. Without loss of generality we focus on the case of a two-tiered UI system, where UI benefits are at a constant level b up to the maximum potential duration of receiving UI benefits P. After benefit exhaustion, individuals receive a second tier of payments indefinitely, so that $b_t = b$ for all $t \leq P$ and $b_t = \underline{b}$ for all $t > P$. The cost of job search $\psi(\lambda_t)$ is an increasing, convex and twice differentiable function.

Jobs offer a wage w_t^* and wage offers are drawn from a distribution with cumulative distribution function $F(w^*; \mu_t)$, which may vary with the duration of unemployment t. To simplify the exposition we assume that the distribution can be summarized by its mean in period t: μ_t.[4] In this case we can write $w_t^* = \mu_t + u_t$, where $E[u_t|t] = 0$ such that u_t reflect random draws from the wage offer distribution. If a job is accepted, the worker starts working at the beginning of the next period and stays at that job forever. Optimal search

[3]For an early, informal discussion of these issues, see Addison and Portugal (1989).

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

behavior of the worker is described by a search effort path λ_t and a reservation wage path ϕ_t, so that all wage offers $w_t^* \geq \phi_t$ are accepted. In the appendix we provide details on the value functions, the first order conditions, as well as the derivations for the following results.

2.2 The Causal Effect of Unemployment Durations on Wages

The expected wage of an individual exiting unemployment in month t is $w^e(t;P) = \frac{\int_{\phi_t}^{\infty} w^* dF(w^*;\mu_t)}{1-F(\phi_t)}$, which given the above assumptions can be written as: $w^e(\phi_t, \mu_t) \equiv w^e(t;P) = \mu_t + E[u_t | u_t \geq \phi_t(P) - \mu_t]$. Note that the change in $w^e(t;P)$ over time can be either due to changes in ϕ_t or due to changes in μ_t.

Using this notation, we can be explicit about what we mean by causal effect. We define the **slope of the reemployment wage path** as the total (right) derivative of the reemployment wage with respect to unemployment duration:[5]

$$\frac{dw^e(t;P)}{dt} = \frac{\partial w^e(t;P)}{\partial \phi_t}\frac{\partial \phi_t}{\partial t} + \frac{\partial w^e(t;P)}{\partial \mu_t}\frac{\partial \mu_t}{\partial t} \tag{1}$$

Furthermore we define the **causal effect of unemployment duration on wages** to be the part of the slope of the reemployment wage path that is due to changes in the wage offer distribution over time:

$$\frac{\partial w^e(\phi_t, \mu_t)}{\partial \mu_t}\frac{\partial \mu_t}{\partial t} \tag{2}$$

The causal effect of unemployment duration on wages is thus defined as the change in expected reemployment wages that would result from exogenously increasing unemployment duration by one month while holding the endogenous variables (i.e., the reservation wage) constant over time. Note that if the reservation wage is not binding at t, i.e., $F(\phi_t) = 0$,

[5] Our model is a discrete model in time, but for the following the notation will be simpler if we can work with time derivatives. In the model only the values of ϕ_t, μ_t and $w^e(t,P)$ at discrete values of time $\{0, 1, 2, ...\}$ are necessary to describe the relevant environment for an individual and the optimal search strategy. Without loss of generality we can therefore define the values of ϕ_t, μ_t and $w^e(t,P)$ for the time values between these discrete values such that they are linear between the discrete points. For example for $0 < t < 1$ let $w^e(t,P)$ be defined as: $w(0) + [w(1) - w(0)]t$. This means that ϕ_t, μ_t and $w^e(t,P)$ are piecewise linear, with kinks at the integer values. All time derivatives below are right derivatives so that by construction we have that: $\frac{df(t)}{dt} = f(t+1) - f(t)$,where $f(t)$ is any function $\phi_t, \mu_t, w^e(t,P)$.

then $w^e(\phi_t, \mu_t) = \mu_t$ and $\frac{\partial w^e(\phi_t,\mu_t)}{\partial \mu_t}\frac{\partial \mu_t}{\partial t} = \frac{\partial \mu_t}{\partial t}$, that is the causal effect of unemployment duration on the reemployment wage is simply the **change in mean offered wages over time**. We will argue below that this seems plausible in the light of our empirical results. Therefore, for simplicity, we will alternatively refer to $\frac{\partial w^e(\phi_t,\mu_t)}{\partial \mu_t}\frac{\partial \mu_t}{\partial t}$ in (2) as the causal effect of nonemployment durations on wages or as the change in the wage offer distribution in the rest of the paper.

Notice that the duration of unemployment t itself is determined by the search intensity and reservation wage of an individual. Hence, reemployment wages and unemployment durations form a system of two simultaneous equations. Regressing w on unemployment durations t using OLS will not result in a meaningful parameter for two reasons: First, since both t and w are affected by individual characteristics (such as preference, human capital, etc), the correlation between the error term of the wage equation and t leads to the standard omitted variable bias in the estimate of the slope of the reemployment wage path $\frac{dw^e(\phi_t,\mu_t)}{dt}$. Second, even if we could fully condition on individual heterogeneity - which seems highly unlikely - we would obtain an estimate of (1) but not of the causal effect of unemployment duration on wages as defined in (2).

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

To simplify the exposition we will first analyze the model under the additional assumption that the expected reemployment wage is a linear function of unemployment duration and thus can be written as: $w^e(t; P) = \xi + \frac{dw^e(t;P)}{dt}\, t$, where we assume that $\frac{dw^e(t;P)}{dt}$ is a constant. Furthermore we will ignore individual heterogeneity for now and focus on the case of homogenous workers. Below we will show that our result generalizes in a straightforward way to the nonlinear case with heterogenous workers.

Note that the expected reemployment wage of an individual (not conditioning on unemployent duration) 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 $E[w^e(t;P)] = \sum_0^\infty w^e(t;P)\, g(t)$. An extension in potential UI durations P affects the expected reemployment wage through two components:

$$\frac{dE[w^e(t;P)]}{dP} = \sum_{t=0}^{\infty}\left[\frac{\partial w^e(t,P)}{\partial P} g(t)\right] + \sum_{0}^{\infty}\left[w^e(t,P)\,\frac{\partial g(t)}{\partial P}\right] \tag{3}$$

The first term $E\left[\frac{\partial w^e(t,P)}{\partial P}\right] = \sum_{t=0}^\infty \left[\frac{\partial w^e(t,P)}{\partial P} g(t)\right]$ represents the average (weighted by the distribution of nonemployment durations) *shift* in the reemployment wage path that is caused by the benefit extension. The second term is due to the shift in the distribution of nonemployment durations *along* the reemployment wage path. Note that the expected nonemployment duration is $D = \sum_{t=0}^\infty [t\, g(t)]$ and the effect of extending UI benefits thus: $\frac{dD}{dP} = \sum_{t=0}^\infty \left[t\, \frac{dg(t)}{dP}\right]$. Given our assumption of linearity for $w^e(t;P)$, the constant ξ cancels out because the changes in the probability mass function have to sum up to 0, so that $\sum_{t=0}^\infty \frac{dg(t)}{dP} = 0$. Equation (3) can then be written as:

$$\frac{dE[w^e(t;P)]}{dP} = E\left[\frac{\partial w^e(t,P)}{\partial P}\right] + \frac{dw^e(t;P)}{dt}\,\frac{dD}{dP} \tag{4}$$

where $\frac{dD}{dP}$ is the marginal effect of an increase in P on the expected non-employment duration D. This formula holds independently from our model and shows how in general the reemployment wage effect can be decomposed into shifts of the reemployment wage path and movement along the reemployment wage path due to increases in nonemployment durations.

While the decomposition in equation (4) is mechanical, results from the search model provide key insights into *how* changes in the outside option (in this case UI durations) affect wages. Combining equations (4) and (1) it follows that the reemployment wage effect can then be written as a combination of the reservation wage effect and the change in the wage offer distribution over time:

$$\frac{dE[w^e(t;P)]}{dP} = E\left[\frac{\partial w^e(t;P)}{\partial \phi_t}\frac{\partial \phi_t}{\partial P}\right] + \left[\frac{\partial w^e w^e(t;P)}{\partial \phi_t}\frac{\partial \phi_t}{\partial t} + \frac{\partial w^e(t;P)}{\partial \mu_t}\frac{\partial \mu_t}{\partial t}\right]\frac{dD}{dP} \tag{5}$$

where $E[.]$ takes the expectation over nonemployment durations. The reservation wage response affects the reemployment wage in two ways: through a shift in the reservation wage and through movements along the reservation wage path. A key implication of equation (5) is that in order to isolate the causal effect of unemployment duration on wages $\frac{\partial w^e w^e(t;P)}{\partial \mu_t}\frac{\partial \mu_t}{\partial t}$ it is necessary to isolate it from these two reservation wage effects. Direct estimates of the effect of UI extensions (or other changes in the outside options) capture all three components. Hence, even absence a potential bias from selective reentry into the labor market, such estimates are hard to interpret.

A final point of equation (5) is that the sign of the effect of extending UI benefits on the reemployment wage is ambiguous, reflecting the contrasting hypotheses about the effect of UI mentioned in the introduction: The first component – due to an upward shift in the reservation wage – will tend to increase the reemployment wage. The second component - longer nonemployment durations leading to more job offers drawn from a different wage offer distribution with lower reservation wages - will tend to decrease the reemployment wage.

2.4 Estimating the Causal Effect of Nonemployment Durations on Wages

To obtain an estimate of the effect of nonemployment durations on the wage offer distribution, we need to infer about how reservation wages change with UI durations and over the nonemployment spell. The effect of the reservation wage on the reemployment wage conditional on exiting at time t is given as $\frac{\partial w^e(t,P)}{\partial \phi_t}$, which enters in both the second and third terms of equation (5).

Notice that in a search model the response of the reemployment wage conditional on nonemployment duration is directly dependent on shifts in the reservation wage. Furthermore the reservation wage is a function of the value of unemployment:

$$\frac{\partial w^e(t,P)}{\partial P} = \frac{\partial w^e(t,P)}{\partial \phi_t}\frac{\partial \phi_t}{\partial P} = \frac{\partial w^e(t,P)}{\partial \phi_t}\frac{dV_t^u}{dP}\rho \tag{6}$$

Rearranging this, the effect of the reservation wage on the reemployment wage can be expressed by the reemployment wage effect of UI extensions divided by the change in the value function: $\frac{\partial w^e(t,P)}{\partial \phi_t} = \frac{\partial w^e(t,P)}{\partial P} / \left(\frac{dV_t^u}{dP}\rho\right)$. This holds as long as $\frac{dV_t^u}{dP}$ is not equal to 0, i.e. the UI extension does in fact affect the outside value. The latter is easily testable, since if a change in P affects the hazard rate of exiting unemployment $\frac{dh_t}{dP}$ this implies that the outside option has in fact changed.[6] [7]

This argument provides a basic but important insight: the effect of UI extensions on reemployment wages conditional on exiting unemployment at time t is informative about the effect of reservation wages on reemployment wages. We therefore have a straightforward test for whether or not reservation wages affect reemployment wages. If the exit hazard is changing ($\frac{dh_t}{dP} < 0$) and there is no effect of UI durations on reemployment wages ($\frac{\partial w^e(t,P)}{\partial P} = 0$), then changes in the reservation wage do not affect reemployment wages.[8]

Note that if the reservation wage does not affect reemployment wages, then the first two terms of equation (5) drop out. In other words, an increase in UI durations does neither affect reemployment wages through a shift in reservation wages (the first term) nor through a rise in nonemployment durations and hence shifts along the, likely declining, reservation wage path (the second term), since $\frac{\partial w^e(\phi_t,\mu_t)}{\partial \phi_t}\frac{\partial \phi_t}{\partial t} = 0$.

Thus, in this case the *causal effect of nonemployment durations on wages* is simply the

[6] In the appendix we show that: $\frac{dh_t}{dP} = -\frac{dV_{t+1}^u}{dP}\left[\frac{(1-F_t(\phi_t))^2}{\psi''(\lambda_t)} + \rho\lambda_t f(\phi_t)\right]$, where the part in the brackets is positive.

[7] Note that we have implicitly assumed that there is no direct of UI extensions on the reemployment wage path itself, i.e., $\partial\mu_t/\partial P = 0$. This would fail for example if firms set wages taking a worker's outside option into account, in which case $\partial\mu_t/\partial P > 0$. As long as wages are assumed to be responding positively to an increase in the outside option, our approach is robust. This is because given the theory implies $\frac{\partial w^e(t;P)}{\partial \phi_t}\frac{\partial \phi_t}{\partial P} \geq 0$ and we have $\partial w^e(t;P)/\partial\mu_t > 0$, a finding of $\partial w^e(t,P)/\partial P = 0$ also implies that $\partial\mu_t/\partial P = 0$.

[8] If the effect of UI durations on reemployment wages is not equal to zero, the equation (6) allows one to infer about the sign in the effect of reservation wages on accepted wages, which is equal to that of the wage response. The magnitude depends on specific model parameters.

ratio between the effect of UI extensions on the average wage and the effect of UI extensions on nonemployment durations. Since both numerator and denominator of this expression can be estimated in the data, this suggests a procedure to estimate the causal effect of nonemployment. This is the same formula as the standard IV estimator and thus suggests that if the conditions on the reemployment hazard and the path of reemployment wages hold the change in the wage offer distribution - the causal effect of nonemployment durations on wages - can be estimated by regressing wages on nonemployment durations using two-stage least squares with UI extensions as an instrument.

2.5 Heterogeneity and Nonlinearity

The model above can be extended to allow for observed and unobserved heterogeneity across individuals, as well as relaxing the assumption of linearity in the reemployment wage path. We use i subscripts to make the individual heterogeneity explicit and allow individuals to be different in terms of the model parameters (such as the cost of job search, the wage offer distribution, preferences, etc.).

In the appendix we show that equation (4) can be generalized to:[9]

$$\frac{dE[w_i^e(t_i, P)]}{dP} = E\left[\frac{\partial w_i^e(t,P)}{\partial P}\right] + \int_0^\infty E_\zeta \left[\frac{\partial w_i^e(t)}{\partial t}\middle| \frac{\partial S_i(t)}{\partial P} > 0\right] \frac{\frac{\partial S(t)}{\partial P}}{\frac{dD}{dP}} dt \, \frac{dD}{dP} \tag{7}$$

where ζ denotes the vector of model parameters and E_ζ is the expectation taken over ζ. Thus in the heterogeneous, nonlinear case, the basic intuition still holds, that the average effect of extending UI benefits on wages can be decomposed into the shift of reemployment wages conditional on unemployment durations and movement along the reemployment wage path. Equation (7) shows that these movements along the reemployment wage path, can still be expressed in terms of what is now a weighted average of the individual slopes of the reemployment wage path $\frac{\partial w_i^e(t)}{\partial t}$, multiplied by the overall increase in nonemployent durations $\frac{dD}{dP}$. The weighted average is such that, only the individual slopes in the reemployment wage

[9] The proof is based on Angrist, Graddy, and Imbens (2000).

path receive positive weight of individuals for whom unemployment durations are infact changing. Furthermore the average slopes at each month t receive weight proporational to the overall change in the survivor function in that month.

As in the linear, homogenous case, the first part of equation (7) will depend on the change in reservation wages at each point, while the slopes of the reemployment wage path depend both on the reservation wage and the change in the wage offer distribution throughout the unemployment spell. As in the base case, if the reemployment wage path is again not affected by changes in potential UI durations, then we can infer that the reservation wage does not affect reemployment wages. Thus, the second term in equation (7) would be a weighted average of of causal effects of nonemployment duration on wages for different individuals at different durations, $\frac{\partial w_i^e(\phi_{it},\mu_{it})}{\partial \mu_{it}} \frac{\partial \mu_{it}}{\partial t}$. In that case we can derive an IV estimator of the causal effect of nonemployment durations on wages. The following proposition states the exact interpretation of this IV estimator explicit for the case that potential UI durations P take on discrete values (as in our empirical part):

Proposition 1. *Suppose the reservation wage is not binding for all individuals for whom the duration of unemployment is responding to changes in UI durations. If potential UI durations P take on exactly two values (P, P'), then the IV estimand, defined as the ratio of the difference in average wage at two values of the durations instrument, to the difference in average durations at the same two values of the durations instrument,*

$$\beta^* = \frac{E[w_i(t, P')] - E[w_i(t, P)]}{D(P') - D(P)}$$

equals the following weighted average of the derivative of the wage function:

$$\beta^* = \int_0^\infty E\left[\frac{\partial w_i^e(\phi_{it}, \mu_{it})}{\partial \mu_{it}} \frac{\partial \mu_{it}}{\partial t} \middle| t_i^e(P') > t > t_i^e(P)\right] \omega^*(t) dt$$

where the weights

$$\omega^*(t) = \frac{Pr(t < t_i^e(P')) - Pr(t < t_i^e(P))}{\int_0^\infty Pr(t < t_i^e(P')) - Pr(t < t_i^e(P))dt} = \frac{S(t;P') - S(t;P)}{D(P') - D(P)}$$

are nonnegative and integrate to one.

Proposition 1 states that IV estimator from a regression of wages on nonemployment durations using UI extnsions as an instrument has an interpretation of a local average treatment effect (LATE) of unemployment durations on wages. The weighting function $\omega^*(t)$ is proportional to the differences in survivor functions. The IV estimator puts more weight on those individuals whose nonemployment durations respond more strongly to the instrument (i.e., whose survival functions are shifting). This is akin to the standard result in linear models with heterogeneous parameters (Angrist, Imbens, and Rubin 1996), but is here derived for the general case in which wages may be a nonlinear function of nonemployment durations (Angrist, Graddy, and Imbens 2000). Hence, as in the more standard linear case, the weighting function can be estimated from the data. In the empirical section, we discuss how the IV estimator is affected if the underlying conditions fail, and we present bounds for the case in which there are small shifts in the reservation wage path.

2.6 Empirical Content of Model

The key new insight of Propositions 1 to 4 is to show that estimating whether reemployment wages conditional on unemployment durations are affected by changes in the UI benefit path (or other factors affecting the value of nonemployment), provides a test for the importance of the outside option of unemployed workers in the wage determination process. If reemployment wages conditional on unemployment duration do not respond to changes in the outside option, then the decline of reemployment wages over the unemployment spell can not be due to a response to the the outside option throughout the unemployment spell. Instead, it must be due to a decline of the wage offer distribution over the nonemployment spell. For this to be meaningful, individuals must value the outside option, as implied by a change in hazard rates.

While we illustrated this insight in a model of wage posting, a symmetric intuition applies in wage bargaining models, where wages should in principle also be affected by the outside option of the unemployed worker. If they are not, then changes in the value of the outside option throughout the unemployment spell should also not have an effect on reemployment wages and thus cannot explain the observed decline in reemployment wages. A similar intuition would hold in a directed search model where workers choose to search for jobs in a segment of the labor market. In such a model wages are affected by the choice of the labor market and the reservation wage when searching in a market. If the wage conditional on unemployment duration does not respond to UI benefit changes, then again workers are not responding to changes in the outside option and the outside option cannot explain the decline in wages over the unemployment spell.

The theory suggests a straightforward strategy for the empirical work. The first step is to estimate the effect of UI durations on average reemployment wages. According to Proposition 1, by itself this just obtains the expression in equation (5). The second step then is to estimate the response of the hazard rate and reemployment wages at each nonemployment duration. If these estimates satisfy the conditions of Proposition 2, in a final step UI extensions can be used as an instrument for nonemployment durations to obtain the causal effect of nonemployment durations on reemployment wages. A key step in the empirical implementation is to address the potential role of changes in the distribution of observable and unobservable characteristics throughout the nonemployment spell. This will be discussed in our methods 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, workers losing a job through no fault of their own in Germany are eligible for UI benefits that provide a fixed replacement rate of 63 percent for an individual without children.[10] This paper focuses on the time period between 1987 and 1999, which is the longest period for which the UI system was stable, and during which the maximum duration of benefits was tied to the exact age of the start of benefit receipt and to prior 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.[11] For workers age 42 to 43 maximum potential UI duration increased to 18 months and for workers age 44 to 48, the maximum duration further rose to 22 months.[12] 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.

In Germany individuals who exhaust regular UI benefits are eligible for means tested unemployment assistance benefits (UA), which do 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, which may explain why only about 50% of UI exhaustees take up UA benefits. In Schmieder, von Wachter and Bender (2012a) we provide an in-depth

[10]For individuals with children the replacement rate is 68 percent. There is a cap on earnings insured, but it affects only 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. Sanctions for not taking suitable jobs exist but appear to be rarely enforced (Wilke 2005).

[11]The age cutoffs were changed in 1999, a period we analyze in Schmieder, von Wachter, and Bender (2012a). The system was reformed substantially and most age cutoffs were abolished in 2004. For an investigation of the stepwise introduction of these age cutoffs between 1983 and 1987 see Hunt (1995).

[12]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 both the sample sizes and the proportional increases in UI durations are smaller at these thresholds the wage estimates are quite noisy and not very informative; therefore, we do not present results on them here. Furthermore in particular 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.

assessment of the role of UA.

3.2 Data

For this paper we have obtained access to the universe of social security records in Germany from 1975 to 2008. The data covers day-to-day 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. 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.[13]

For our analysis sample, we extracted all unemployment insurance spells where the claimant was between age 40 and age 46 on the claim date. For reasons discussed above, we consider unemployment spells starting any time between July 1987 and April 1999. For each UI spell we created variables about the previous work history (such as job tenure, labor market 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, in addition to duration of UI benefit receipt, we 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. Our analysis period assures that we can follow individuals for at least 9 years after the start of the UI spell.

[13]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.

The core part of our identification strategy is to use variation in potential UI durations at the age thresholds for any given UI claim spell. We calculate each individual's potential UI duration at the beginning of the UI spell, using information about the law together with information on exact birth dates 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.[14] In our companion paper, we also show that the characteristics of our sample are comparable with those of UI recipients in the United States.

3.3 Estimation

The institutional structure and data allow us to estimate the causal effect of UI durations on wages, and – once the conditions on the path of reemployment wages described in Section 2 are satisfied – to obtain estimates of the causal effect of nonemployment duration on wages. Our empirical strategy follows three consecutive steps.

Estimating the Causal Effect of UI Durations on Employment and Wages. The institutional structure and data allow us to estimate the causal effect of large extensions in UI benefit durations on non-employment duration, reemployment wages and other outcomes for workers with previously stable employment using a regression discontinuity design. We follow common practice and first show smoothed figures to visually examine discontinuities at the eligibility thresholds (e.g., Lee and Lemieux 2010). To obtain estimates for the main causal effects, we follow standard regression discontinuity methodology and estimate variants

[14] 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.

of the following regression model:

$$y_i = \beta + \gamma \times \Delta P \times D_{a_i \geq a^*} + f(a_i) + \epsilon_i, \tag{8}$$

where y_i is an outcome variable, such as non-employment duration (D) or reemployment wages (w), 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^*. In the notation from Section 2, we obtain estimates for $\frac{dD}{dP}$ and $\frac{dE[w]}{dP}$.

For our main estimates, we focus on the period from July 1987 - March 1999, and we use the sharp threshold at age 42. We estimate equation (8) 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, and summarize our extensive sensitivity analysis below.

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.[15] For this procedure we normalize the age for all individuals within two years of the age 42 (44) threshold to the age relative to age 42 (44) (i.e. the rescaled age variable is set to 0 for someone who is exactly age 42 (44) at the time of claiming UI). We estimate the following model on the pooled sample:

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

where a_i is the normalized age variable and $\overline{\Delta P}$ is the average change in potential UI durations at the age threshold. With this specification $\hat{\gamma}$ is a direct estimate of the rescaled marginal effect, forcing it to be equal at the two cutoffs.

[15]We also estimated all results at the age 44 cutoff separately. The point estimates are very similar but lack precision.

Estimating the Shift in the Path of Reemployment Wages and Hazards. The main goal of the paper is to estimate the causal effect of nonemployment durations on wages. As derived in Section 2, the first step in obtaining such an estimate is to assess whether the path of reemployment wages and the reemployment hazard shift in response to the UI extensions. If the probability of exiting nonemployment (the hazard rate) declines at each duration, this implies that individuals value future increases in UI durations. If the path of reemployment wages does not shift as well, this means the outside option does not bind, and we can use UI duration as an instrumental variable for nonemployment durations.

To analyze how the reemployment wage path responds to UI durations we begin by a graphical examination. We estimate the following regression separately for each nonemployment exit month t:

$$w_i^* = \delta_t P_i + f(a_i) + \epsilon_i | \, t_i = t, \tag{9}$$

where $P_i = \Delta P \times D_{a_i \geq a^*}$ captures the effect of a change in UI durations for those reaching the age of eligibility. If $cov(\epsilon_i, P_i | t_i = t) = 0$, then estimating equation (9) via OLS will yield $\hat{\delta}_t$ as consistent estimates for the effect of UI durations on reemployment wages at each nonemployment duration t, $E\left[\frac{\partial w_i^e(t,P)}{\partial P} | t_i = t\right]$. A corresponding analysis and condition yields consistent estimates for the effect of UI durations on reemployment hazards at each nonemployment duration t.

However, while the identification assumptions of the RD design guarantee that individuals on both sides of the cutoff are comparable *on average* (i.e., $cov(\epsilon_i, P_i) = 0$) in the total RD sample close to the age cutoff, they do not imply that ϵ_i and P_i are uncorrelated *conditional* on the duration of unemployment t_i. The time when people exit unemployment t_i is affected by individual behavior and possibly by the treatment variable P_i.

We provide two alternative arguments that make $cov(\epsilon_i, P_i | t_i = t) = 0$ plausible in our context. First, 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 (9) will yield consistent estimates of, $\frac{\partial w^e(t,P)}{\partial P}$, the effect of UI durations on the path of reemployment wages.

Second, we can also make an argument based on the theoretical restriction that the reservation wage has to rise in response to an increase in P (i.e., $\frac{\partial w^e(t,P)}{\partial P} \geq 0$ for all t). If ϵ_i is a person fixed effect, then an estimate of $\hat{\delta}_t = 0$ for any given nonemployment duration t is only consistent with $\frac{\partial w^e(t,P)}{\partial P} > 0$ if $cov(\epsilon_i, P_i | t_i = t) < 0$. If we find that $\hat{\delta}_t = 0$ at all nonemployment durations t, then it has to be the case that $\frac{\partial w^e(t,P)}{\partial P} = 0$, since it cannot be that for all t: $cov(\epsilon_i, P_i | t_i = t) < 0$ given the RD assumtions. Similar arguments based on observables and theory can be made as for estimates of the reemployment hazard.[16]

Following the same intution, we also directly estimate the average shift in the reemployment wage path and test whether it is equal to zero. To do so, we estimate the following regression on the entire sample

$$w_i^* = \delta P_i + \sum_{t=1}^{T} \theta_t + f(a_i) + \epsilon_i \qquad (10)$$

where w_i^* is the observed reemployment wage and θ_t are time dummies for the duration of non-employment. The parameter δ captures the average shift in the reemployment wage path, such that $\delta = E\left[\frac{\partial w_i^e(t,P)}{\partial P}\right]$. We are still estimating the regression in the RD setting and hence control for age at the time of entering unemployment. Given the RD assumptions, the resulting estimate for δ is consistent; i.e., while individuals exiting at each nonemployment duration may be different in terms of unobservable characteristics in the high and low UI duration regimes, again on average this must cancel out close to the cutoff. In other words, differential selection over the nonemployment spell cannot explain a shift in the entire reemployment wage path. Hence, if we cannot reject that the estimated $\hat{\delta}$ is equal to zero, we can conclude that the reemployment wage path and hence reservation wages have not shifted.[17]

[16] Most of the literature presents estimates of the effect of UI durations on reemployment hazards without specifically addressing this selection issue.

[17] Selection could, in principle, lead to a rotation of the observed reemployment wage paths. Yet, following the intuition outlined above we again use the prediction from the theory that $\frac{\partial w^e(t,P)}{\partial P} \geq 0$ since the

Estimating the Causal Effect of Nonemployment Durations on Wages. If the hazard rate declines ($\frac{dh_t}{dP} < 0$) and there is no change in reemployment wages ($E\left[\frac{\partial w_i^e(t,P)}{\partial P}\right]$=0) then the relevant conditions in Proposition 3 hold, and the effect of UI durations on reemployment wages are driven by higher nonemployment durations. The final step then is to directly estimate the causal effect of nonemployment duration on wages using an instrumental variables strategy. 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 an estimate for π can then be calculated by dividing these two estimates. Alternatively, one can directly estimate the causal effect of nonemployment durations 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

A key aspect in all three steps of our empirical strategy is the validity of the RD design. The regression discontinuity method only yields consistent results if factors apart from the treatment variable do not vary discontinuously at the threshold. If individuals have control over the forcing variable of the regression discontinuity estimator, in this case the age of claiming UI benefits, then the estimates resulting from the RD can be biased. In our setting, both potential UI claimants and their employers face potential incentives to manipulate the age of claiming. We have examined this issue at length in our related paper (Schmieder, von Wachter, and Bender 2012a) and its Web Appendix, and conclude that sorting around the threshold is not a concern in this case. We only summarize the main findings here and refer the interested reader to our precursor paper for a more detailed discussion.

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 unemployment spells in two-week age intervals around the cutoff. The

reservation wage has to rise or stay constant. In this case, if $\delta = E\left[\frac{\partial w_i^e(t,P)}{\partial P}\right] = 0$, then it must be that $\frac{\partial w^e(t,P)}{\partial P} = 0$ at all nonemployment durations t. Below, we will use the confidence interval for the estimate $\hat{\delta}$ to derive bounds for our causal estimates for small shifts in reservation wages.

figure indicates that there is a small increase in the density right after the threshold.[18] Further investigation showed that this increase is not driven by individuals who postpone their claim until they are eligible for longer benefits. Instead, if at all the incidence of lay off rises slightly at the eligibility age. The magnitude of this effect, however, is very small: only about 200 instances relative to about 500,000 observations in the sample close to the age cutoff.

A 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 pooling the age 42 and the age 44 cutoff (where potential UI durations increase from 18 to 22 months. There is a statistically significant change at the threshold is the fraction female in the age 42 model and the pooled model: The fraction of UI recipients who are female is estimated to increase by about 0.8 percentage points (or 0.5% in the pooled model). Furthermore there is a tiny difference in the years of education variable at the first threshold of about 0.03 years (or 10 days) of education. 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 minor discontinuities and density shifts would almost certainly not be detectable. While these findings point to a small violation of the RD identification assumptions, these should have a relatively small impact on the overall results. In fact, neither trimming observations close to the eligibility thresholds nor directly controlling for observable characteristics affects our results. To ensure that our results are not affected by sorting around the threshold and by particular implementation choices of the RD estimator,

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

we performed multiple robustness checks summarized in the sensitivity section (Section 6).

4 The Average Effect of UI extensions on Job Quality

4.1 The Effects of UI extensions on Nonemployment Durations

In our precursor paper we have shown that the UI extensions studied here lead to significant increases in nonemployment duration (Schmieder, von Wachter, and Bender 2012a). To show the strength of our research design, to obtain nonemployment estimates used later on, and because the sample in this paper differs slightly from the previous paper, Table 2 and Figure 2 replicate the analysis of the effect of UI extensions on benefit and nonemployment durations.

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 at the two age thresholds lead to substantial increases in UI 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. The purely behavioral effect of an increase in potential UI durations is demonstrated in Figure 2 (b), which shows the effect on nonemployment durations. Increases in potential UI durations have a very clear effect on nonemployment durations and hence 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. 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.

Finally, as further discussed in von Wachter, Bender, and Schmieder (2012b), columns (3) to (5) show that these increases in nonemployment durations have small long lasting effects. Shown in Column (3), the probability of ever again working in a social security liable job decreases by about 0.5 (0.1) percentage points per additional month of potential UI benefits at the age 42 threshold (pooled). Columns (4) and (5) show that the incidence of employment is lower and the probability of receiving UI is higher even five years after employment. However, these effects are all very small, less than 5% relative to the mean. In the sensitivity section, we will assess whether these findings could imply a potential bias from sample selection in our main wage estimates.

4.2 The Effect of UI Extensions on Reemployment Wages

Longer nonemployment durations in response to higher potential UI durations could either raise wages as individuals have more time to search for a better job, or lower wages if the negative effect from longer nonemployment durations dominates. 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 a way to remove an individual fixed effect and hence can be viewed as a way to both control for possible selection and to obtain more precise estimates. The figure shows that the average wage loss for the unemployed in our sample is substantial, ranging from 13% to 16%. While the gain in precision is modest, Figure 3 (b) indicates that selection along the previous wage has little impact on the results, and again clearly points to a negative effect of a rise in potential UI durations on post-unemployment wages.

The corresponding regression estimates in Table 3 columns (1) and (2) show that increases in potential UI durations lead to precisely estimated negative effects on post-unemployment wages. Panel A shows that the post-unemployment wage is about 1 percent lower in both levels and first differences when potential UI durations increase by six months. Panel B shows the results from pooling both cutoffs and reveals similar estimates with a small gain in statistical precision. While precisely estimated, the effects are small. 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. However, although per se this effect may not seem large, below we show that it can imply substantial negative effects of nonemployment durations on wages. Moreover, losses can add up to more substantial effects if individuals remain in lower paying jobs for a long period of time.

Columns (4) to (6) of Table 3 shows the effect on the log wage one, three, and five years after the start of the new employment spell. A one-month increase in potential UI durations in the pooled model is associated with a 0.089 percent decrease in the wage five years after start of the employment spell. This point estimate is smaller than the effect in years 1 and 3 after start of employment, confirming the result in column (3) of Table 3 that there is a small positive (yet insignificant) effect of potential UI durations on wage growth. Although the longer-term effects are not estimated precisely, these findings are suggestive of potentially substantial cumulated wage losses. We will return to the implications for the total wage loss and individual behavior in the conclusion.

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 (2007a) found a negative point estimate of UI durations on wages, 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 of UI extensions. As further discussed in Section 5, an additional value added with respect to these papers is that we provide a framework and dynamic results that allow us to separate the wage offer and the reservation wage effect.

4.3 The Effect of UI Extensions on Other Job Outcomes

It could be that individuals accept lower wages in return for other desirable job characteristics. For example, relative to those with shorter UI durations, individuals with longer UI durations may have access to jobs that are more stable, jobs that have potentially higher wage growth, or jobs that require a lower commute.[19] Similarly, they may be able to trade off lower wages against the need to change occupation or industry. In this section we show that this is not the case – i.e., jobs tend to be worse among all the dimensions we can measure here. The analysis of other job outcomes is of interest in its own right. In addition, it provides insights into the potential channels underlying the reduction in wages and the role of nonemployment durations, which we further discuss below.

Table 4 shows the effect of increases in potential UI durations on a number of job-related outcome variables. The first outcome is the completed job tenure at the post-unemployment job, which is often used as an indicator of the quality of the job match. Column (1) shows that there is a small decrease in the duration of the post-unemployment job of about 0.0081 (0.0099) years in Panel A (B), which is statistically significant for the full sample. This confirms findings in Table 2 that higher potential UI durations reduce job stability. Hence, it does not appear individuals with longer UI durations trade lower wages for more stable jobs or jobs that appear to represent better matches.

We analyzed several additional indicators of job quality. There is no indication that longer potential UI durations increase the probability of finding a job in the same region as the previous job (column (2)).[20] An important finding of the literature on displaced workers is that those switching to another industry or occupation experience much larger declines in

[19]Taking one example, 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.

[20]We also analyzed changes in firm size as proxy for employer quality and found no significant change.

earnings (e.g., Neal 1995, Addison and Portugal 1989). Hence, one would expect that longer UI durations may help individuals to find jobs in their previous line of work. Columns (4) and (5) of Table 4 show that this is not the case. Longer potential UI durations increase the probability of switching to a different industry and a different occupation by about 0.12 to 0.18 percentage points, respectively.

Overall, all measures of job quality available in our data either point to negative effects of longer potential UI durations or no effect. Hence, at least based on this limited set of job characteristics, it does not appear that workers with longer UI durations accept lower wages in return to better job outcomes along other dimensions. Instead, longer UI durations appear to lead to worse job outcomes among multiple dimensions studied here.

5 The Causal Effect of Nonemployment Durations on Reemployment Wages

In this section we assess to what extent the causal effects of UI durations on wages we document in Section 4 result from longer nonemployment durations. As discussed in the theory section (Section 2) and method section (Section 3.3), a crucial step in this analysis is to assess whether the hazard rate and the path of reemployment wages over the nonemployment spell shift in response to UI durations. As a preliminary step, to interpret our findings on hazard rates and reemployment wages, we analyze how observable characteristics 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.

5.1 Selection Throughout the Nonemployment Spell

As summary measures for observable characteristics that are relevant to the labor market Figure 5 shows the mean of pre-unemployment wages and the mean of predicted reemployment wages (based on a broad range of pre-determined characteristics) by month of nonemployment duration. Vertical bars indicate that the point estimates at time t are statistically significant at the 5 percent level. The figures show two key findings. First, as expected, the figures show that there is some correlation between pre-determined characteristics and nonemployment duration, though the gradient is not very strong. For example, mean pre-unemployment wages fall by about 5% and mean predicted wages fall by about 7% in the first year of nonemployment duration. In separate analysis, we found that years of schooling or fraction female is positively correlated with nonemployment duration. Second, 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. While there are a few statistically significant point estimates in each figure (and in the figures of single characteristics not shown here), 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.[21] Overall, these figures therefore support the notion that observables are essentially uncorrelated with potential UI durations conditional on t and that therefore unobservables, captured by the person effect ϵ_i in equation (9),are also unlikely to be correlated with potential UI durations: $cov(\epsilon_i, P_i|t_i = t) = 0$. Given this, we are confident that shifts in the reemployment hazard or mean reemployment wages conditional on unemployment duration reflect true behavioral changes and not shifts in worker composition.

5.2 Estimates of the Shift of Reemployment Hazards and Wages

Figure 4 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 nonemployment durations t smaller than the maximum potential UI duration P. This is statistically signifi-

[21] The one exception 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 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. To address this aspect, we show in the sensitivity section that our results hold within gender groups.

cant 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. A similar pattern has been observed in other studies of the effect of UI extensions on nonemployment duration (e.g., Card, Chetty, and Weber 2007b).

Figure 6 Panel (a) shows the effect of changes in P on the reemployment wage conditional on t. On average, wages decline by about 25 percent within the first year. However, note that despite the clear shift in the hazard rate for all nonemployment durations $t < P$, we do not observe a corresponding change in the path of reemployment wages over the nonemployment spell. In the notation of the model, it appears that indeed $\frac{\partial w_i^e(t,P)}{\partial P} = 0$ for all nonemployment durations $t < P$. Figure 6 Panel (b) shows an almost unchanged pattern when we control for individual heterogeneity by plotting the difference in post and pre unemployment log wage. Extending UI benefits does not shift the reemployment wage upwards.

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. As we discuss in the robustness section, these differences are reduced significantly once we look at women and men separately, indicating that the negative wage spikes are partly driven by more women exiting at the exhaustion points.

Figure 6 provides visual and statistical evidence that $\frac{\partial w_i^e(t,P)}{\partial P} = 0$ at each nonemployment duration t lower than the exhaustion point P. The validity of this evidence relies on the additional assumption that similarity in observable characteristics over the nonemployment spell for individuals with different potential UI durations documented in Section 5.1 implies similarity in unobserved characteristics as well. To dispense with this additional assumption, we move to our second strategy outlined in Section 3.3 that only relies on the RD assumptions. Table 5 presents estimates of the *average shift* in the reemployment wage path, $E\left[\frac{\partial w_i^e(t,P)}{\partial P}\right]$, obtained from implementing equation (10). Column (1) of Table 5 shows the results controlling for a linear effect of nonemployment duration. This yields an estimate for δ for the 12 to 18 month discontinuity very close to zero (point estimate -0.016% with a standard error of 0.048%). If we control more flexibly for the nonemployment duration effect (Columns 2 to 3), the point estimate is even closer to 0. For the pooled sample it is positive but extremely small and statistically indistinguishable from zero despite very precise estimates (point estimate 0.00015 log points, or 0.015%).

Recall that given there is a decline in the hazard rate in response to UI durations, reservation wages are predicted rise. In that case, it has to be that $\frac{\partial w_i^e(t,P)}{\partial P} \geq 0$ for all nonemployment durations t. Hence, given that the estimates in Table 5 are very close to zero, and that theory excludes cases for which $\frac{\partial w_i^e(t,P)}{\partial P} < 0$, this confirms the visual impression of Figure 6 that $\frac{\partial w_i^e(t,P)}{\partial P} = 0$ for all nonemployment durations $t < P$. This important finding implies that the effect of UI durations on wages found in Section 4 arises due to a rise in nonemployment durations, not due to a shift in the reservation wage schedule.

These results also imply that reservation wages do not appear to bind in our sample. This is consistent with related findings in the literature. For example, DellaVigna and Paserman (2005) who calibarate a model similar to ours find that very few wage offers fall below the reservation wage. Using time use data, Krueger and Muller (2011) show that self-reported reservation wages stay remarkably constant over time and do not appear to respond to UI durations. The decrease in the hazard rate throughout the unemployment spell is explained by unemployed workers lowering their search efforts dramatically. 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. Similarly, consistent with our findings, Hornstein, Krusell, and Violante (2011) show that in broad classes of search models, the value of non market time – and hence the reservation wage – has to be low to be able to reconcile why despite high wage dispersion, and hence a high option value of searching, workers in practice accept jobs quickly. Following our approach, Lalive, Landais, and Zweimueller (2013) also find that the path of reemployment wages does not respond to increases in UI durations in Austria.[22]

5.3 Estimates of the Causal Effect of Nonemployment Durations on Wages

The results on the hazard and reemployment wage path imply that reservation wages do not bind. As a result, the observed decline in the reemployment wage path is entirely due to the decline in the wage offer distribution (equation 1), and the effect of UI durations on wages is only due to a rise in nonemployment durations (equation 5). Proposition 3 then allows us to obtain a valid estimate of the average decline in the reemployment wage path. In the final step of our empirical analysis, we hence estimate the slope of the reemployment wage path using UI durations as instrumental variables for nonemployment durations: $\pi = \frac{\frac{dE[w]}{dP}}{\frac{dD}{dP}}$.

Table 6 shows 2SLS results for the effect of nonemployment durations on reemployment wages using extensions in potential UI durations as an instrument. The table first shows the first stage regression (i.e., the effect of UI extension on nonemployment durations), which easily passes weak instrument concerns. The second column shows the 'reduced form' of the IV estimator, which correspond to the baseline estimates of the effect of potential UI durations on wages of Section 4.2. Column (3) reports the resulting 2SLS estimate of nonemployment durations on wages. We find that $\Delta_t E[w|t] = -0.78\%$, which is precisely estimated. Thus, our main finding is that an additional month of nonemployment lowers reemployment wages by about 0.8 percent.

To better understand the nature of our IV estimator, recall that Proposition 3 states that in the presence of heterogeneity and nonlinearity in the slope of the reemployment wage function, the IV estimator obtains the local average treatment effect of wage declines for individuals whose nonemployment durations are most affected by the instrument. As such, it is weighted towards the treatment effects of compliers to the UI extensions that underlie

[22] In contrast, Mueller (2013) finds that jobs with wages above the self-reported reservation wage are more likely to be accepted, although a substantial fraction of jobs paying below the reservation wage as well. Taken together with the findings in Krueger and Mueller (2011), in our notation these results imply $\frac{\partial w_i^e(t,P)}{\partial \phi_{it}} > 0$ but $\frac{\partial \phi_{it}}{\partial P} = \frac{\partial \phi_{it}}{\partial t} = 0$. Hence, while in contrast to our findings on reemployment wages, the main assumptions needed to use UI durations as instruments would still hold.

our regression discontinuity (RD) estimates. As seen from the survival functions in Panel (b) of Figure 4 the compliers come from the entire range of nonemployment durations, with the largest weight being between 12 and 18 months. Hence, the IV estimator is not weighted towards particular durations, but estimates an average of the effect of nonemployment duration on wages over a broad spell of nonemployment durations.

The interpretation of our IV estimates as causal effect of nonemployment duration on wages relies on our empirical result that $\frac{\partial w_i^e(t,P)}{\partial P}$ is approximately zero. Since it is difficult to empirically establish that a moment is indeed exactly zero, we derived bounds for the effect of nonemployment durations on wage offers under the assumption that the true change in the reemployment wage path, $\frac{\partial w_i^e(t,P)}{\partial P}$, is at the upper bound of the estimated confidence interval. These bounds are quite narrow and confirm the overall magnitude of our main causal effect. The expression for the bias is shown in Section 2.5. In words, the deviation of the IV estimate from the true slope of the wage offer distribution is higher the larger the response in reservation wages (measured by δ), the smaller the first stage coefficient, and the larger 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. The first term measures the direct effect of reservation wages on the reemployment wage path (akin to a direct effect of an instrument on the outcome). The second term captures the fact that if reservation wages matter, longer nonemployment durations also induce wage changes due to changes in reservation wages.[23]

Given that we can estimate δ, $\frac{dE[w]}{dP}$ and $\frac{dD}{dP}$, to obtain bounds for the causal effect π we have to say something about $\frac{\tilde{E}\left[\frac{dV_t^u}{dt}\right]}{\tilde{E}\left[\frac{dV_t^u}{dP}\right]} < 0$. 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 as long as $\frac{\tilde{E}\left[\frac{dV_t^u}{dt}\right]}{\tilde{E}\left[\frac{dV_t^u}{dP}\right]} > -6$. Only for a very strong

[23]Recall that from equation 6 that $\frac{dE[w|t]}{dP}$ is determined by the change in the reservation wage in response to a rise in potential UI durations times the effect of reservation wages on actual wages. The intuition of the various terms is hard to see from the final equation, but clear from the derivation in the Web Appendix, to which we defer the interested reader.

decline in the value function $\frac{\tilde{E}\left[\frac{dV_t^u}{dt}\right]}{\tilde{E}\left[\frac{dV_t^u}{dP}\right]} < -6$ would the IV estimate be biased downward (or the absolute value be biased upward). In our Web Appendix, we calculate the implied π given various values of $\frac{\tilde{E}\left[\frac{dV_t^u}{dt}\right]}{\tilde{E}\left[\frac{dV_t^u}{dP}\right]}$ and δ. Essentially as long as δ is small (i.e. close to the estimated range in Table 5 which is always clearly less then 0.1%) or $\frac{\tilde{E}\left[\frac{dV_t^u}{dt}\right]}{\tilde{E}\left[\frac{dV_t^u}{dP}\right]}$ is not too high (between -2 and -8) 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 the pooled estimate in Column (3) of Table 5, $\hat{\delta}= 0.095\%$, the range of slopes for the wage offer distribution is between -1.3% (actually even smaller than the IV estimate) to -0.7%, just slightly larger than the IV estimate of -0.78% decline in mean wage offers per month.

As mentioned in Section 2.5, another source of bias in the IV estimator arises if $\frac{dh_t}{dP} = 0$. Examining the reemployment hazards shown in Panel A of Figure 4 , this occurs for nonemployment durations greater than two years, when benefit durations are exhausted on both sides of the age threshold. Thus, even though $\frac{\partial w_i^e(t,P)}{\partial P} = 0$, for $t > 24$ it may be that reservation wages affect accepted wages (i.e., the second term in equation (5) is not zero). Using the survivor functions and the definition of the weighting function of the IV estimator shown in Proposition 3, we can calculate the weight in the IV estimator that nonemployment durations beyond 24 months receive. Integrating over the difference in the two survivor curves in Panel B of Figure 4 for $t > 24$, we obtain that the sum of the weight is 0.35. The size of the bias then depends on the effect of reservation wages on accepted wages. Given our results imply no response in reservation wages for $t < 24$, and given the findings in the literature on the role of reservation wages, we find it safe to assume that the reservation wage effect at $t > 24$ is likely to be small. Hence, given our overall findings the bias of IV is likely to be minor.

5.4 Interpretation of IV Estimates

Based on our methodology and empirical analysis, the causal effect of nonemployment durations on wage offers is about 0.8% per month. For small but realistic changes in the outside options the effect ranges between -0.7% and -1.3% . As discussed in the theory section, this interpretation is independent of the particular economic model of job search we used, and holds for a broad range of how wages and outside options are determined. In this section, we further discuss the economic magnitude, potential channels, and implications of these findings.

As causal estimate of nonemployment durations, the parameter we estimate has many potential applications. Furthermost, it can be used for predicting the effect of an increase of nonemployment durations on offered wages at the individual level. In a partial equilibrium context, it can also be used to extrapolate the effect of increases in average nonemployment durations among certain groups of workers or in the labor market as a whole. In addition, an increasing number of theoretical studies in public economics show that the parameter we estimate is a key input into determining the optimal policy response to long-term unemployment. Similarly, the parameter is a key input in macroeconomic models aiming to explain the persistence of unemployment rates.

A natural question is whether the IV estimator is large or small in economic terms. For example, at 6 (12) months additional nonemployment duration, our point estimates imply a loss in daily wages of 4.8% (9.6%). Based on Figure 6, this represents about 40% of the average wage loss at 6 and 12 months, respectively. Thus, long nonemployment spells can lead to considerable declines in daily wages. Shorter nonemployment spells will also reduce reemployment wages, but obviously the predicted effect is smaller. For example, the effect of a 4 week rise in nonemployment duration (the typical response to a 6 month UI extension estimated in the literature, see Table 2, Column 2) leads to a decline in wages of about 0.8% (Table 3, Column 1). Given these effects are small, it is perhaps not surprising that previous studies using smaller sample sizes, a less precise research design, and smaller UI extensions have not detected statistically significant declines in wages. We are not aware of papers estimating the causal effect of nonemployment durations based on a quasi-experimental research design. Estimates based on the correlation of nonemployment duration of displaced workers with reemployment wages suggest effects somewhat bigger than ours (e.g., Addison and Portugal 1989, Gregory and Jukes 2001), but as explained above may be affected by selection.[24] Estimates of the rate of depreciation of human capital during unemployment based on structural models show results of similar order of magnitude as our findings (e.g., Keane and Wolpin 1997).

Effects of the magnitude we find are plausible once one realizes that they can derive from multiple channels possibly operating at the same time. One reason why long nonemployment durations may have a causal effect on wages is because they lead to a depreciation of skills.[25] This may be general skills, or industry- or occupation-specific skills. If this was the main explanation, long average nonemployment durations could have a lasting negative effect on the stock of human capital in the economy. Another reason why wages may fall with nonemployment duration is that employers may use the duration as indicator of workers' unobserved skill. If this was true and employers correctly update their beliefs, the effect should be smaller in periods of high unemployment because the long term unemployed are less likely to be negatively selected than in good times (e.g., Kroft, Lange, and Notowidigdo 2012).[26]

Per se it is hard to know what magnitude to expect from these different channels. In our RD analysis we found an effect of UI extensions on several potential channels but not others. For example, we found a significant rise in the incidence of part-time jobs, a rise in the probability of switching industry or occupation, and a small decline in completed job

[24]Absent quasi-experimental evidence or detailed worker characteristics, Addison and Portugal (1989) address selection using a Heckman correction term, and Gregory and Jukes (2001) report within-worker differences.

[25]Using longitudinal data on explicit skill measures from Sweden, Edison and Gustavsson (2008) report that one year of unemployment duration reduces skills by an equivalent of 0.7 years of schooling.

[26]Note that changes in job characteristics by themselves can not be a channel of what we find here. As discussed in Section 2, our empirical findings also exclude changes in reservation job quality – and hence an effect from directed search – as potential channel underlying the negative wage effects we find.

tenure. We did not find significant effects on switching county of employment or measures of firm quality, though results were partly imprecise. To assess the potential impact of these effects on reemployment wages, even though they are potentially endogenous, we included these outcomes as additional explanatory variables in our main RD estimates (not shown). Controlling for an indicator capturing industry and occupation changes leads to a slight drop in the effect of UI extensions on reemployment wages of 20-25%. Controlling separately for a part-time indicator and completed tenure at the new job leads to a bit larger decline of 30-40%. Including proxies for employer quality made no difference.

Overall, while such regressions have to be interpreted with caution, as expected no additional job characteristic is able to account for the entire wage effect. However, we do see some prima facie evidence of role of industry and occupation changes, which have been associated with losses in (industry or occupation) specific skills in the literature. Similarly, the rise in part-time employment could reflect a decline in job quality. Clearly, some of these outcomes could reflect several mechanisms; for example, a decline in human capital could increase part-time work and lower job quality. Hence, while interesting, we do not stress any particular interpretation here.

We have also analyzed differences in the effect of UI extensions on reemployment wages over the business cycle. On the one hand, if employers correctly update their priors, the rise in expected mean quality of job applicants during recessions should lead to lower stigma of nonemployment duration. On the other hand, it is plausible that the effect of nonemployment duration on wage offers is stronger in recessions. For example, there is ample evidence of both a decline in job quality and of a reduction of wages within jobs in recessions, which could hurt in particular workers with longer unemployment spells. When we compare our findings in periods with high and with low unemployment rates, the results are very robust in recessions, but imprecise and ambiguous in expansions (not shown). Overall, while we find no prima facie evidence in favor of a stigma effect, this could also be due to changes in the distribution of offered jobs over the cycle.[27]

6 Summary of Robustness Analysis

6.1 Robustness of Regression Discontinuity Estimates

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 7 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. 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.

[27] The ideal test would hold the distribution of job types constant (comparable to what is done in audit studies), but is not feasible in our quasi-experimental setting because of small sample sizes and endogeneity problems.

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. Column (6) of Table 7 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. Overall, while excluding the observations close to the cutoff reduces statistical power somewhat, it does not affect our overall conclusions. Column (7) of Table 7 shows how the estimates change when we 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. Column (8) of Table 7 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.

6.2 Robustness of Estimated Effects on Wages

We implemented numerous robustness checks regarding our findings on wages as well. Here, we report results addressing the aspects of selective return to employment, differences in effects across groups, and changes in unemployment rates throughout the nonemployment spell.

In Section 4, we have shown that increases in UI extensions lead to precisely estimated declines in the incidence of employment (Table 3). While these effects were very small relative to the mean (less than 1%), we investigated the potential effect of such selection on our estimates of the effect of UI durations on wages by analyzing differences in the quantiles of the distribution of outcomes on the two sides of the age cutoffs (For space reasons, these findings are contained in our Web Appendix). This standard procedure yields consistent estimates if selection depends monotonously on a single index of underlying characteristics. The analysis of differences in quantiles is also interesting in its own right, since it gives an indication of how the distribution of wages is changing due to UI extensions. The findings suggest that the decline in median wages is larger than the mean effect. Moreover, the lower percentiles of the wage distribution decline more strongly than the upper percentiles, suggesting that UI extensions not only shifts the distribution of wages downward, but also increases its skewness. These findings indicate that if at all the mean wage effects that constitute our main findings underestimate the effect of UI extensions on wages. Moreover, the results imply that labor market success becomes more unequal, with the majority of wage declines occurring in the lower part of the distribution.

The concern of selection also arises when studying reemployment wages at each point in the nonemployment spell. We have replicated a similar analysis of differences in quantiles at each point of the nonemployment spell (shown in our Web Appendix). Again, this is interesting in its own right, since if reservation wages were to matter, this would mainly affect the lower percentiles of the distribution. Interestingly, as for the mean effect, the differences in percentiles is zero at all nonemployment durations for all deciles – again with exception for a positive effect at month 12 (the exhaustion point before age 42) and a negative effect at month 18 (the exhaustion point from age 42 to 44). As discussed in Section 5, we believe this is likely due to selective exit shifting from one exhaustion point to the other other.

In Section 5, we had said that part of the effect on reemployment wages at the exhaustion points month 12 and month 18 are likely to be due by a change in sample composition. In particular, there is a rise in the fraction of women to the right of the RD cutoffs, and a rise in the fraction of women exiting at the exhaustion points. To address this point, we have replicated our main RD analysis and our analysis of reemployment wages by gender. While women's nonemployment durations clearly respond more strongly to UI extensions (Schmieder et al. 2012a), there is no precisely estimated difference in the effect on reemployment wages by gender. As a result, the implied IV estimate of the effect of nonemployment durations on wages is somewhat smaller for women. As mentioned in Section 5, when considering the necessary condition for interpretation of the IV estimate by gender, the reemployment wage path now exhibits no statistically significant differences even at the exhaustion points. Hence, our main findings are robust for the small degree of selection of women into nonemployment and UI exhaustion we find. We also considered the effect of UI durations and nonemployment durations on wages for other subgroups, but statistical precision was low and hence did not pursue this further.[28]

Another concern we addressed is the role of changes in the aggregate and local unemployment rate throughout the nonemployment spell. If labor market conditions evolve through the spell, then the wage offer distribution may shift for reasons other than nonemployment durations. Despite the fact that individuals above and below the age cutoffs face the same unemployment conditions at the start of their nonemployment spells, we found in Column (6) of Table 4 that UI extensions seem to have a small negative effect on aveage unemployment rates at reemployment. Although precisely estimated, at 0.5-1% relative to the mean the effect is again very small. To nevertheless assess its potential for confounding our estimates of the causal effect of nonemployment durations on wages, we included the unemployment rate at reemployment in the main wage RD regression as explanatory variable. Although, as discussed in our discussion section, such regressions are hard to interpret the inclusion has only small affects on our main coefficients. We thus do not believe such shifts in market conditions throughout the spell is an important factor in explaining our main findings.

[28] For example, while lower educated workers had substantially larger responses in employment duration, the effect of UI durations on wage changes appeared only slightly larger for the lower educated, implying a smaller (but not precisely estimated) causal effect of nonemployment durations on wages.

7 Conclusion

The question whether and to what extent longer unemployment durations hurt wages and other job outcomes of unemployed workers has been a longstanding question in economics. The paper makes three contributions to answering this question First, the paper derives conditions under which UI extensions and other quasi-experimental changes in workers' outside option can be used as an instrument for nonemployment durations and shows that the analysis of the evolution of reemployment wages throughout the nonemployment spell provides a test for these conditions.

The paper's second contribution is to estimate the effect of UI durations on wages and other outcomes using a discontinuous changes in UI durations by exact age in Germany. We find that an extension in UI durations leads to small but precisely estimated negative effects on wages and other job outcomes. Relative to previous work, these estimates are somewhat larger and more precisely estimated, and suggest that UI extensions tend to lower, not raise job outcomes.

Finally, the third contribution is to present new estimates for the causal effect of nonemployment duration on wages. The paper finds that the path of reemployment wages throughout the nonemployment spell does not shift in response to UI extensions. The theoretical results show that this implies reservation wages do not bind and that as a consequence, extensions in UI durations affect change wages only because rising nonemployment durations lower wage offers. The IV estimates imply that each additional month of nonemployment duration lead to a statistically significant and substantial effect on wages of 0.8%. The theory shows this estimate is a weighted average of the slope of the wage offer distribution at all nonemployment durations. Given UI durations lead to a decline in reemployment probabilities throughout the nonemployment spell, it is relevant for a broad group of unemployed workers.

The paper's findings have implications going beyond the causal effect of nonemployment and its potential mechanisms. The findings speak to the potential welfare consequences of UI extensions. If individuals get all the surplus from higher match quality, then they will have internalized the effect of their search behavior on match quality, and the effects of potential UI durations on match quality can be ignored from a social welfare perspective.[29] This situation is different, 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.

The results are also potentially informative about the effects of prolonged nonemployment spells on the aggregate economy. Our findings suggest that these can lead to persistent and substantial declines in wages that are possibly larger in recessions. In so far as workers may be receiving lower job matches or have lower productivity, this could imply a significant cost to society going beyond the direct cost of unemployment itself. However, by construction our regression discontinuity analysis is partial equilibrium in nature, and a full evaluation of the implications of causal effects of nonemployment we document here would require specifying a macroeconomic model.

Our results are also related to the value of leisure. Rational individuals incur the costs of additional wage reductions above and beyond foregone earnings during nonemployment in favor of additional leisure. A back-of-the-envelope calculation suggests that the present discounted value of the cost from lower wages due to higher nonemployment durations is about half a month of average earnings per additional month of nonemployment duration. This may be indeed rational, in so far fixed costs of working or fixed costs of leaving a job puts a wedge between the value of leisure and foregone earnings. While it also could be that individuals do not fully foresee the wage penalty they incur, without additional information and estimating structural parameters we cannot say more here.

Finally, by the nature of our regression discontinuity design and institutional framework, our estimates are based on middle age workers with stable labor force attachment. While this is the core constituency of unemployment insurance in Germany, the United States and

[29]This is essentially an application of the envelope theorem. See Chetty (2008) and Schmieder et al. (2012a) for details.

other countries, it does not speak to the potential effects of UI durations and nonemployment durations for a broader population. Studies with data and research designs encompassing broader groups of workers will help to obtain additional information on how the effects we measure here differ in the population, what the likely effect on the macro economy is, and what the underlying channels may be. Similarly, while we showed in our precursor paper that the characteristics of our sample is comparable to similar aged UI recipients in the United States and that the effect of UI durations on employment are comparable (Schmieder, von Wachter, and Bender 2012a), one has to be careful in generalizing from our results based on Germany to other countries. Beyond the specific empirical findings, our paper has presented a general framework for the analysis of the causal effects of nonemployment durations on wages and other job outcomes that will be useful for studying similar patterns in other countries.

References

Addison, John T. and McKinley L. Blackburn, "The effects of unemployment insurance on postunemployment earnings," *Labour Economics*, *7* (1), (2000), 21–53.

– and Pedro Portugal, "Job Displacement, Relative Wage Changes, and Duration of Unemployment," *Journal of Labor Economics*, *7* (3), (1989), 281.

Angrist, Joshua D., Guido W. Imbens, and Donald Rubin, "Identification of Causal Effects Using Instrumental Variables," *Journal of the American Statistical Association*, *91* (1996), 444-472.

– , Kathryn Graddy, and Guido W. Imbens, "The Interpretation of Instrumental Variables Estimators in Simultaneous Equations Models with an Application to the Demand for Fish," *Review of Economic Studies*, *67* (2000), 499-527.

Ball, Laurence M., "Hysteresis in Unemployment: Old and New Evidence," *NBER Working Papers No. 14818*, (2009).

Bender, Stefan, Anette Haas, and Christoph Klose, "IAB Employment Subsample 1975-1995 Opportunities for Analysis Provided by the Anonymised Subsample," *IZA Discussion Paper*, *117* (2000).

Bernanke, Ben S., "Recent Developments in the Labor Market," *Speech at National Association for Business Economics Annual Conference (http://www.federalreserve.gov/newsevents/speech/bernanke20120326a.pdf)*, (2012).

Black, Dan A., Jeffrey A. Smith, Mark C. Berger, and Brett J. Noel, "Is the Threat of Reemployment Services More Effective than the Services Themselves? Evidence from Random Assignment in the UI.," *American Economic Review*, *93* (4), (2003), 1313–1327.

Blanchard, Olivier and Lawrence Summers, "Hysteresis and the European Unemployment Problem," *NBER Macro Annual*, (1986).

Blau, David M and Philip K Robins, "Job search, wage offers, and unemployment insurance," *Journal of Public Economics*, *29* (2), (1986), 173–197.

Card, David, Raj Chetty, and Andrea Weber, "Cash-On-Hand and Competing Models of Intertemporal Behavior: New Evidence from the Labor Market," *Quarterly Journal of Economics*, *122* (4), (2007a), 1511–1560.

– , – , and – , "The spike at benefit exhaustion: Leaving the unemployment system or starting a new job?," *American Economic Review Papers and Proceedings*, *97* (2), (2007b), 113–118.

Centeno, M. and Á.A. Novo, "Reemployment wages and UI liquidity effect: a regression discontinuity approach," *Portuguese Economic Journal*, *8* (1), (2009), 45–52.

Chetty, Raj, "Moral hazard versus liquidity and optimal unemployment insurance," *Journal of political Economy*, *116* (2), (2008), 173–234.

Coles, Melvyn G. and Adrian Masters, "Retraining and long-term unemployment in a model of unlearning by not doing.," *European Economic Review*, *44* (9), (2000), 1801–1822.

DellaVigna, S. and M.D. Paserman, "Job search and impatience," *Journal of Labor Economics*, *23* (3), (2005), 527.

den Haan, Wouter J., Christian Haefke, and Garey Ramey, "Turbulence and Unemployment in a Job Matching Model.," *Journal of the European Economic Association*, *3* (6), (2005), 1360–1385.

Edin, Per-Andres and Magnus Gustavsson, "Time Out of Work and Skill Depreciation.," *Industrial Labor Relations Review*, *61* (2), (2008), 163–180.

Feldstein, Martin and James Poterba, "Unemployment insurance and reservation wages," *Journal of Public Economics*, *23* (1-2), (1984), 141–167.

Gregory, Mary and Robert Jukes, "Unemployment and Subsequent Earnings: Estimating Scarring Among British Men 1984-94.," *Economic Journal*, *111* (474), (2001), 607–625.

Gruber, Jonathan, "The consumption smoothing benefits of unemployment insurance," *The American Economic Review*, *87* (1), (1997), 192–205.

Hornstein, Andreas, Per Krusell, and Giovanni L. Violante, "Frictional Wage Dispersion in Search Models: A Quantitative Assessment," *American Economic Review*, *101* (7), (2011), 2873–2898.

Hunt, Jennifer, "The Effect of Unemployment Compensation on Unemployment Duration in Germany," *Journal of Labor Economics*, *13* (1), (1995), 88–120.

Katz, Lawrence and Bruce D. Meyer, "The Impact of Potential Duration of Unemployment Benefits on the Duration of Unemployment Outcomes," *Journal of Public Economics*, (1990), 45–71.

Keane, Michael P. and Kenneth I. Wolpin, "The Career Decisions of Young Men," *Journal of Political Economy*, *105* (3), (1997), 473-522.

Kroft, Kory and Matt J. Notowidigdo, "Should Unemployment Insurance Vary With the Local Unemployment Rate? Theory and Evidence," *mimeo*, (2010).

– , Fabian Lange, and Matthew Notowidigdo, "Duration Dependence and Labor Market Conditions: Evidence from a Field Experiment.," *Quarterly Journal of Economics*, *128* (3), (2012), 1123-1167.

Krueger, A.B. and A. Mueller, "Job Search, Emotional Well-Being and Job Finding in a Period of Mass Unemployment: Evidence from High-Frequency Longitudinal Data," (2011).

Lalive, Rafael, "Unemployment benefits, unemployment duration, and post-unemployment jobs: A regression discontinuity approach," *American Economic Review*, *97* (2), (2007), 108–112.

– , Camille Landais, and Josef Zweimueller, "Market Externalities of Large Unemployment Insurance Extension Programs.," *Mimeo*, (2013).

Lee, David S. and Thomas Lemieux, "Regression discontinuity designs in economics," *Journal of Economic Literature*, *48* (2), (2010), 281–355.

Ljungqvist, Lars and Thomas J. Sargent, "The European Unemployment Dilemma," *Journal of Political Economy*, *106* (3), (1998), 514–550.

– and – , "Two Questions about European Unemployment," *Econometrica*, *76* (1), (2008), 1–29.

McCrary, Justin, "Manipulation of the running variable in the regression discontinuity design: A density test," *Journal of Econometrics*, *142* (2), (2008), 698–714.

Meyer, Bruce D., "Unemployment Duration and Unemployment Spells," *Econometrica*, *58* (1990), 757–82.

– , "Lessons from the US Unemployment Insurance Experiments.," *Journal of Economic Literature*, *33* (1), (1995), 91–131.

– , "Unemployment and workers compensation programmes: rationale, design, labour supply and income support," *Fiscal Studies*, *23* (1), (2002), 1–49.

Moffitt, Robert, "Unemployment Insurance and the Distribution of Unemployment Spells," *Journal of Econometrics*, *28* (1), (1985), 85–101.

Mortensen, Dale T., "Job Search and Labor Market Analysis," *Handbook of Labor Economics*, *2* (1986), 849–919.

Mueller, Andreas, "A Contribution to the Empirics of Reservation Wages.," *Mimeo*, (2013).

Neal, Derek A., "Industry-Specific Human Capital: Evidence from Displaced Workers," *Journal of Labor Economics*, *13* (4), (1995), 653-677.

Pavoni, Nicola, "Optimal Unemployment Insurance, with Human Capital Depreciation, and Duration Dependence.," *International Economic Review*, *50* (2), (2009), 323–362.

– and Giovanni L. Violante, "Optimal Welfare-to-Work Programs," *Review of Economic Studies*, *74* (1), (2007), 283–318.

– , Ofer Setty, and Giovanni L. Violante, "Search and Work in Optimal Welfare Programs," *NBER Working Papers No. 18666*, (2013).

Pissarides, Christopher A., "Loss of Skill During Unemployment and the Persistence of Employment Shocks.," *Quarterly Journal of Economics*, *107* (4), (1992), 1371–1391.

Schmieder, Johannes F, Till Von Wachter, and Stefan Bender, "The Effects of Extended Unemployment Insurance over the Business Cycle: Evidence from Regression Discontinuity Estimates over 20 years," *The Quarterly Journal of Economics*, *127* (2), (2012a), 701–752.

– , Till von Wachter, and Stefan Bender, "The Long-Term Effects of UI Extensions on Employment," *The American Economic Review*, *102* (3), (2012b), 514–519.

Shimer, Robert and Ivan Werning, "On the Optimal Timing of Benefits with Heterogeneous Workers and Human Capital Depreciation," *NBER Working Paper No. 12230*, (2006).

Solon, Gary, "Labor Supply Effects of Extended Unemployment Benefits," *Journal of Human Resources*, *14* (1979), 247–255.

Spinnewijn, Johannes, "Training and Search during Unemployment.," *Journal of Public Economics*, *99* (2013), 49–65.

van den Berg, Gerard J., "Nonstationarity in job search theory," *The Review of Economic Studies*, *57* (1990), 255–277.

van Ours, Jan and Milan Vodopivec, "Does reducing unemployment insurance generosity reduce job match quality?," *Journal of Public Economics*, *92* (3-4), (2008), 684–695.

Wilke, Ralf A., "New Estimates of the Duration and Risk of Unemployment for West-Germany," *Journal of Applied Social Science Studies*, *128* (2), (2005), 207–237.

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	(7) UR at start of unemp	(8) County UR at start of unemp
Increase in Potential UI Dur. from 12 to 18 Months								
D(Age above Cutoff)	0.030	0.0086	0.0038	0.044	-0.046	0.12	0.0016	0.035
	[0.014]*	[0.0028]**	[0.0020]	[0.028]	[0.031]	[0.18]	[0.0087]	[0.025]
Effect relative to mean	0.0027	0.024	0.037	0.0082	-0.0041	0.0017	0.00017	0.0033
Observations	510955	510955	510955	510955	510955	480724	510955	441907
Mean of Dep. Var.	11.0	0.36	0.10	5.35	11.1	70.8	9.29	10.4
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)								
D(Age above Cutoff)	0.015	0.0054	0.0017	0.041	-0.034	0.12	-0.0095	0.017
	[0.0094]	[0.0020]**	[0.0017]	[0.023]	[0.024]	[0.13]	[0.0066]	[0.019]
Effect relative to mean	0.0014	0.015	0.016	0.0072	-0.0030	0.0016	-0.0010	0.0016
Observations	947068	947068	947068	947068	947068	888293	947068	829669
Mean of Dep. Var.	10.9	0.36	0.11	5.69	11.6	71.6	9.31	10.4

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) UI Benefit Duration	(2) Non-Emp Duration	(3) Ever emp. again	(4) Employed 5 years after start of UI	(5) Receiving UI 5 years after start of UI
Increase in Potential UI Dur. from 12 to 18 Months					
D(Age above Cutoff)	1.77	0.95	-0.0094	-0.0045	0.0055
	[0.034]**	[0.14]**	[0.0020]**	[0.0028]	[0.0019]**
$\frac{dy}{dP}$	0.29	0.16	-0.0016	-0.00075	0.00092
	[0.0057]**	[0.023]**	[0.00033]**	[0.00046]	[0.00032]**
Effect relative to mean	0.23	0.065	-0.011	-0.0086	0.036
Observations	510955	437899	510955	510955	510955
Mean of Dep. Var.	7.57	14.7	0.86	0.52	0.15
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)					
D(Age above Cutoff)	1.44	0.72	-0.0082	-0.0059	0.0049
	[0.029]**	[0.10]**	[0.0015]**	[0.0020]**	[0.0015]**
$\frac{dy}{dP}$	0.29	0.14	-0.0016	-0.0012	0.00099
	[0.0058]**	[0.021]**	[0.00030]**	[0.00041]**	[0.00030]**
Effect relative to mean	0.17	0.048	-0.0097	-0.012	0.031
Observations	947068	799105	947068	947068	947068
Mean of Dep. Var.	8.33	15.8	0.84	0.51	0.16

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 Post Unemployment Wages

	(1) Log Post Wage	(2) Log Wage Difference	(3) Log Wage Growth 5 Years	(4) Log wage 1 year after reemployment	(5) Log wage 3 years after reemployment	(6) Log wage 5 years after reemployment
Increase in Potential UI Dur. from 12 to 18 Months						
D(Age above Cutoff)	-0.0078	-0.0070	0.0016	-0.0086	-0.0056	-0.0054
	[0.0030]**	[0.0029]*	[0.0038]	[0.0035]*	[0.0040]	[0.0043]
$\frac{dy}{dP}$	-0.0013	-0.0012	0.00026	-0.0014	-0.00093	-0.00089
	[0.00050]**	[0.00049]*	[0.00064]	[0.00059]*	[0.00066]	[0.00071]
Effect relative to mean	-0.0019	0.050	-0.019	-0.0022	-0.0014	-0.0013
Observations	437182	420311	311568	382089	345073	311833
Mean of Dep. Var.	4.01	-0.14	-0.084	3.95	3.95	3.97
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)						
D(Age above Cutoff)	-0.0051	-0.0055	-0.00065	-0.0035	-0.0031	-0.0043
	[0.0021]*	[0.0022]*	[0.0028]	[0.0025]	[0.0028]	[0.0031]
$\frac{dy}{dP}$	-0.0010	-0.0011	-0.00013	-0.00070	-0.00063	-0.00086
	[0.00042]*	[0.00044]*	[0.00056]	[0.00050]	[0.00057]	[0.00062]
Effect relative to mean	-0.0013	0.038	0.0069	-0.00088	-0.00079	-0.0011
Observations	797752	767161	568540	699057	630261	569024
Mean of Dep. Var.	4.00	-0.15	-0.094	3.95	3.95	3.97

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 Other Match Quality Measures

	(1) Duration of post unemp job in years	(2) Move to different county to take up job after unemp	(3) Post unemp job is different industry	(4) Post unemp job is different occupation	(5) National UR at re-emp.
Increase in Potential UI Dur. from 12 to 18 Months					
D(Age above Cutoff)	-0.048	0.00063	0.0072	0.011	0.033
	[0.035]	[0.0030]	[0.0030]*	[0.0031]**	[0.0084]**
$\frac{dy}{dP}$	-0.0081	0.00011	0.0012	0.0018	0.0055
	[0.0058]	[0.00049]	[0.00050]*	[0.00051]**	[0.0014]**
Effect relative to mean	-0.012	0.0015	0.010	0.017	0.0036
Observations	390142	437690	425131	437899	437899
Mean of Dep. Var.	4.10	0.42	0.69	0.61	9.26
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)					
D(Age above Cutoff)	-0.049	-0.00020	0.0058	0.0088	0.018
	[0.026]	[0.0022]	[0.0022]**	[0.0023]**	[0.0062]**
$\frac{dy}{dP}$	-0.0099	-0.000040	0.0012	0.0018	0.0035
	[0.0053]	[0.00045]	[0.00044]**	[0.00046]**	[0.0012]**
Effect relative to mean	-0.012	-0.00048	0.0084	0.014	0.0019
Observations	712660	798726	775879	799105	799105
Mean of Dep. Var.	4.20	0.41	0.69	0.61	9.27

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 5: The Effect of Potential UI Durations on Reemployment Wages Conditional on Nonemployment Duration

	(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)	-0.00098	0.00056	-0.00025
	[0.0029]	[0.0028]	[0.0028]
Potential UI Dur	-0.00016	0.000093	-0.000042
	[0.00048]	[0.00047]	[0.00047]
Observations	429258	429258	429258
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)	0.00029	0.0011	0.00076
	[0.0020]	[0.0020]	[0.0020]
Potential UI Dur	0.000059	0.00021	0.00015
	[0.00040]	[0.00040]	[0.00040]
Observations	783587	783587	783587
Mean of Dep. Var.	4.02	4.02	4.02

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 6: The Effect of Time Out of Work on Reemployment Wages, OLS and IV Estimates

	(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	0.16	-0.0013	.	.
	[0.023]**	[0.00050]**	.	.
Nonemp Dur	.	.	-0.0067	-0.0078
	.	.	[0.000053]**	[0.0033]*
Observations	437182	437182	429258	429258
Mean of Dep. Var.	14.7	4.01	4.01	4.01
Pooling both Thresholds (12 to 18 Months and 18 to 22 Months)				
Potential UI Dur	0.15	-0.0010	.	.
	[0.021]**	[0.00042]*	.	.
Nonemp Dur	.	.	-0.0069	-0.0064
	.	.	[0.000039]**	[0.0031]*
Observations	797752	797752	783587	783587
Mean of Dep. Var.	14.8	4.02	4.02	4.02

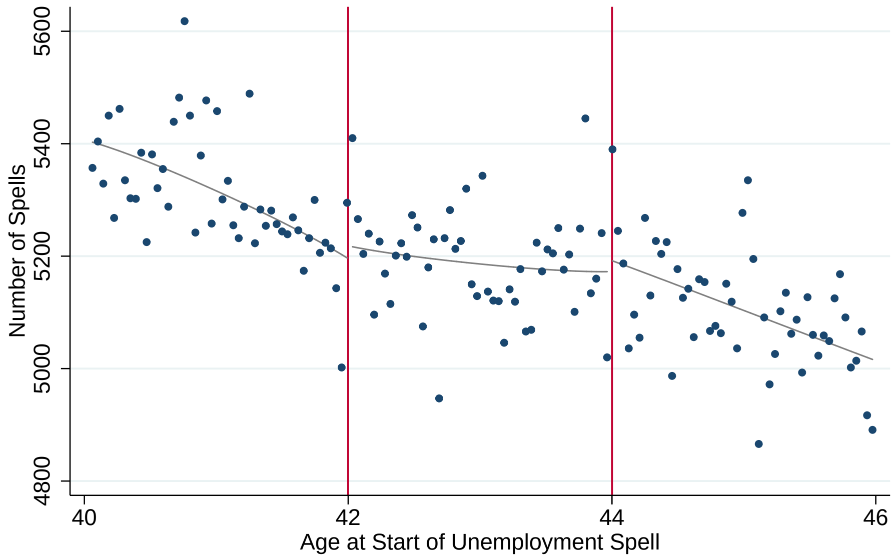
Notes: Coefficients from RD regressions. Local linear regressions (different slopes) on each side of cutoff. Standard errors clustered on day level (* P<.05, ** P<.01)).

Table 7: Sensitivity Analysis

	(1) Bandwidth: 2 Years	(2) Bandwidth: 1 Year	(3) Bandwidth: 0.5 Years	(4) Quadratic Age Control	(5) Cubic Age Control	(6) Excluding Obs within 1 month of threshold	(7) Controlling for observable characteristics	(8) Sample restricted to UI takeup within 15 days of job end
Non-employment duration								
$\frac{dy}{dP}$	0.14	0.19	0.23	0.17	0.24	0.12	0.11	0.15
	[0.021]**	[0.029]**	[0.043]**	[0.031]**	[0.043]**	[0.022]**	[0.012]**	[0.021]**
Observations	799105	399918	199889	799105	799105	765540	893505	696777
Mean of Dep. Var.	14.8	14.8	14.8	14.8	14.8	14.8	15.2	14.4
Log post wage								
$\frac{dy}{dP}$	-0.0010	-0.0020	-0.0019	-0.0017	-0.0025	-0.00090	-0.00090	-0.0012
	[0.00042]*	[0.00061]**	[0.00093]*	[0.00066]*	[0.00091]**	[0.00044]*	[0.00038]*	[0.00044]**
Observations	797752	399245	199570	797752	797752	764232	771197	695689
Mean of Dep. Var.	4.02	4.02	4.02	4.02	4.02	4.02	4.01	4.03
Log wage difference								
$\frac{dy}{dP}$	-0.0011	-0.0020	-0.0025	-0.0015	-0.0033	-0.00057	-0.00076	-0.0013
	[0.00044]*	[0.00063]**	[0.00093]**	[0.00068]*	[0.00093]**	[0.00046]	[0.00040]	[0.00043]**
Observations	767161	384054	191913	767161	767161	734989	771197	675826
Mean of Dep. Var.	-0.15	-0.15	-0.15	-0.15	-0.15	-0.15	-0.14	-0.14
Moved to different county to takeup job after unemployment								
$\frac{dy}{dP}$	-0.000040	0.00062	0.00026	-0.00013	0.00074	0.00010	-0.00043	0.000011
	[0.00045]	[0.00064]	[0.00090]	[0.00068]	[0.00091]	[0.00049]	[0.00049]	[0.00047]
Observations	798726	399737	199796	798726	798726	765180	771827	696437
Mean of Dep. Var.	0.41	0.41	0.41	0.41	0.41	0.41	0.41	0.41
Duration of post unemployment job								
$\frac{dy}{dP}$	-0.0099	-0.029	-0.027	-0.017	-0.041	-0.0066	-0.0074	-0.0093
	[0.0053]	[0.0077]**	[0.011]*	[0.0082]*	[0.011]**	[0.0056]	[0.0046]	[0.0058]
Observations	712660	356808	178324	712660	712660	682711	772129	622283
Mean of Dep. Var.	4.20	4.20	4.19	4.20	4.20	4.20	2.92	4.34

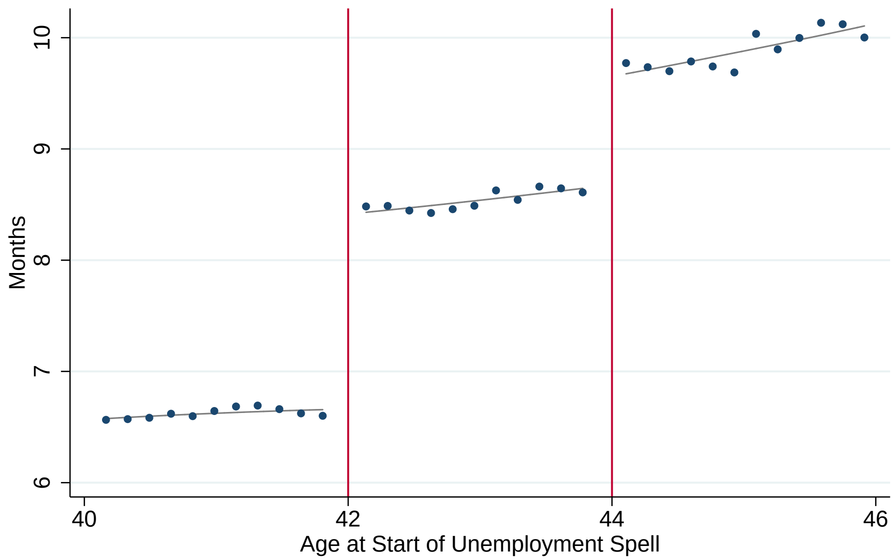
Notes: Standard errors clustered on day relative to cutoff level (* P<.05, ** P<.01)).
The sample are individuals who started receiving unemployment insurance between 1987 and 1999. Each panel shows the increase at the age threshold of the dependent variable (given in the panel title) rescaled by the 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.

Figure 1: Frequency of Observations Around Age Cutoffs for Potential Unemployment Insurance



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) and age 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

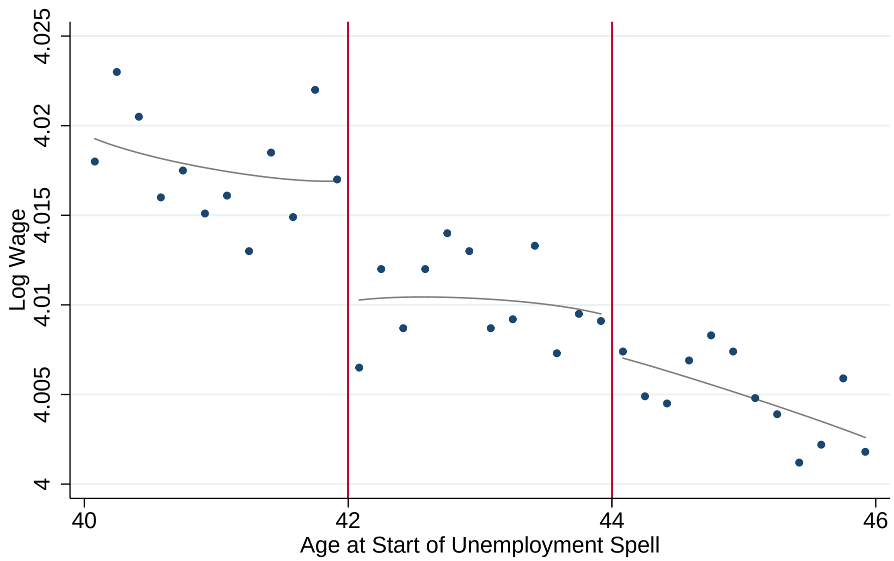


(a) Months of receiving UI benefits

(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

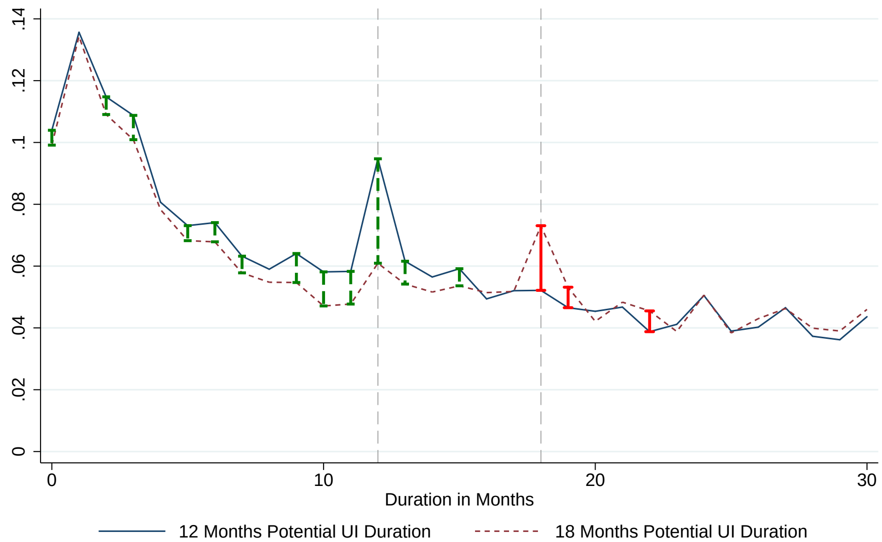


(a) Log post unemployment wage

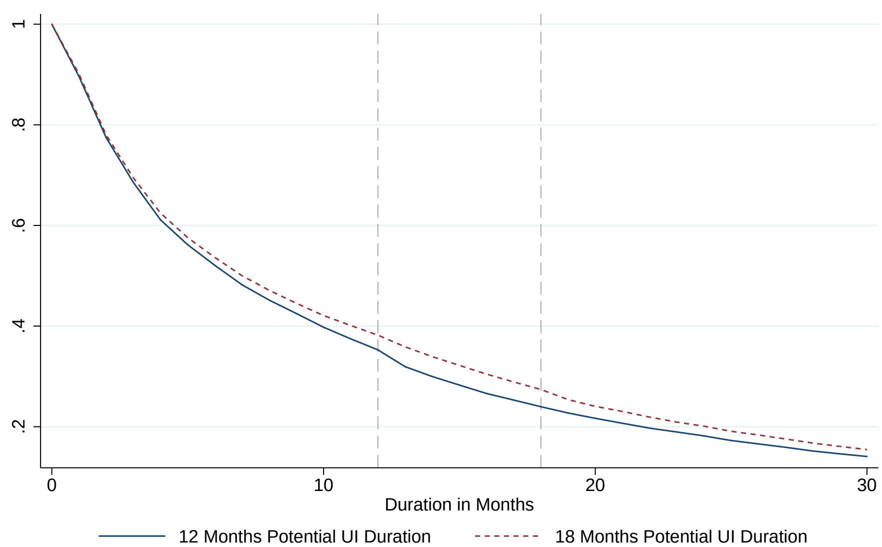
(b) Log wage difference (pre unemployment minus post unemployment)

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



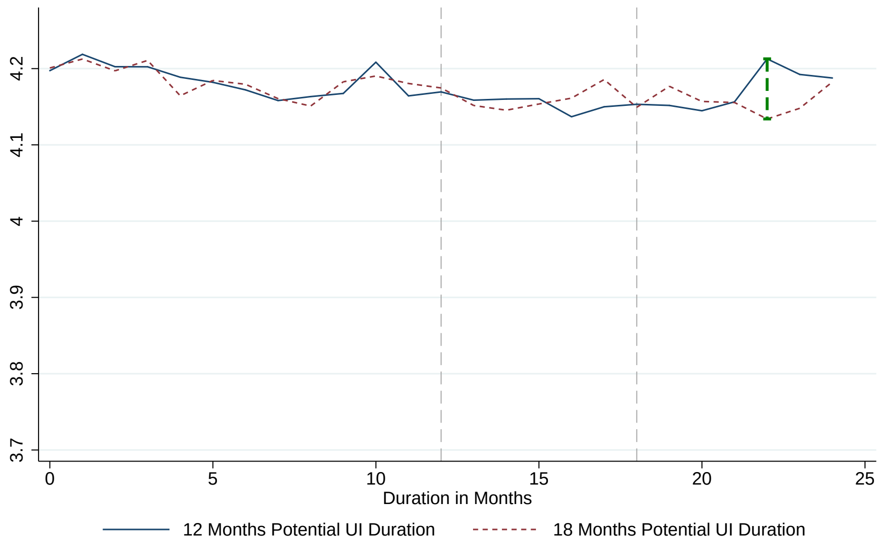
(a) Unemployment Exit Hazard



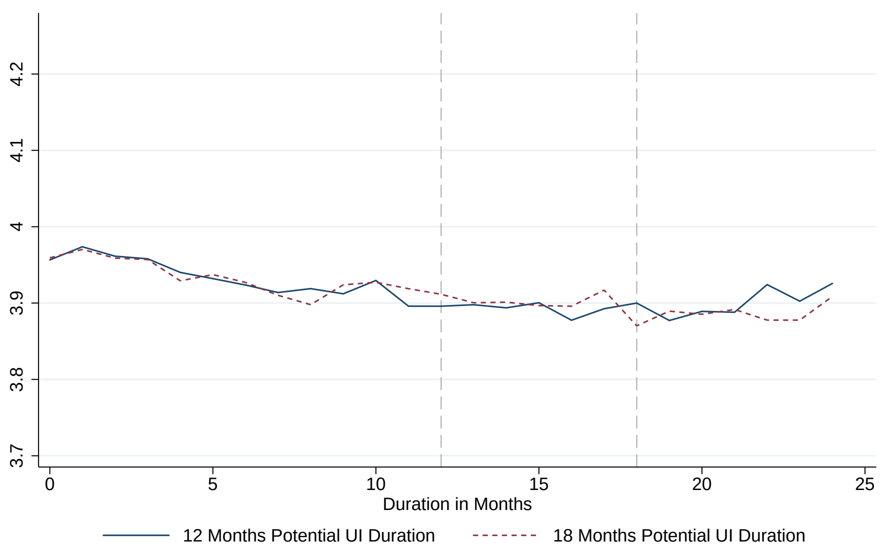
(b) Survival Functions

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



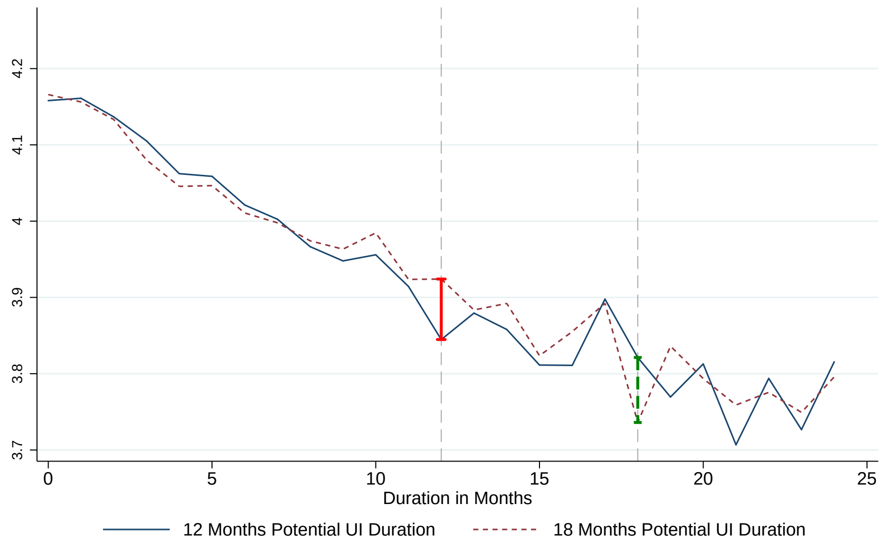
(a) Pre-unemployment log wage by time of non-emp exit



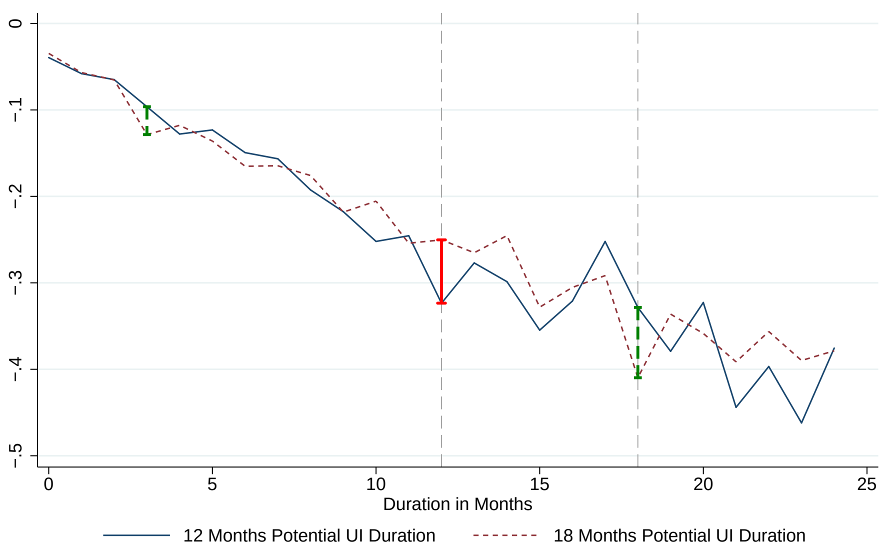
(b) Predicted reemployment log wage by time of non-emp exit

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 6: 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.