The Effects of Extended Unemployment Benefits: Evidence from a Regression Discontinuity Design (Latest version available here)

Po-Chun Huang* Tzu-Ting Yang†

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Abstract

This paper uses administrative unemployment insurance (UI) data and the unemployment benefit extension for workers age 45 or more in Taiwan to estimate the effects of extended benefits on duration and reemployment outcomes. In Taiwan, since workers who are eligible for unemployment benefits can receive 50% of their remaining entitlements as reemployment bonuses after they become employed, extending potential duration not only extends unemployment benefits but also increases the bonus amount for workers reemployed before exhausting the benefits. We show in a search model that increasing the potential duration of benefits increases the unemployment duration and reservation wage because the bonus offer partly offsets the moral hazard effect of extended benefits, while the liquidity effect of extended benefits stays intact. The model further points out that the effects of extended benefits on search effort and reservation wage are stronger for workers who tend to run out of benefits and those who are liquidity-constrained. Our estimates using the regression discontinuity design suggest that an increase in potential duration from 180 to 270 days increases insured duration by 57 days (39%) and nonemployment duration by 41 days (15%). While we do not find wage gains for overall UI recipients around 45 years old, benefit extension is estimated to increase the reemployment wage for lower-wage workers who are most likely to exhaust benefits. Our findings suggest the number of liquidity-constrained workers at exhaustion point relative to all UI recipients might play an important role on the effect of benefit extension on job match quality.

*PhD Candidate, Michigan State University, email: huangpo5@msu.edu. I thank Steve Woodbury, Carl Davidson, Steven Haider and Todd Elder for their careful guidance and support, and Gary Solon for his advice in the early stage of the project. The unemployment insurance data were provided by the Bureau of Labor Insurance. This paper represents the views of the authors and does not reflect the views of the Bureau of Labor Insurance.

†Academia Sinica, Email: ttyang@econ.sinica.edu.tw.

1 Introduction

In order to provide additional liquidity for older unemployed workers, many countries, such as Germany, France, Japan, and South Korea, extend the potential duration of unemployment benefits for older workers. While benefit extension increases workers' abilities to smooth consumption during unemployment (Rothstein and Valletta (2014); Ganong and Noel (2016)) and to find better job matches (Nekoei and Weber (2016)), it also reduces workers' willingness to search for jobs (Schmieder et al. (2012a)).[1] It is hence crucial to estimate the effects of benefit extension on labor market outcomes in order to evaluate what the optimal unemployment insurance (UI) might be. While there have been estimates of the effects of extended benefits in Europe and the United States, the evidence in Asia is still scarce.

Estimates from Europe and the United States cannot be reliably extrapolated to Asia for at least two reasons. On the one hand, based on 2014 World Bank data, the gross saving rate in East Asia and Pacific is 35%, higher than the 21% in the European Union and 18% in North America. As pointed out by Chetty (2008), the labor supply response to increasing generosity of unemployment benefits is stronger for liquidity constrained households. We might expect benefit extension to have a smaller effect on labor supply for high saving rate countries. On the other hand, the data from World Bank also shows that the unemployment rate in East Asia and Pacific in 2014 was 4.5%, lower than the 10.2% in the European Union and 6.3% in North America. Workers in a low unemployment area might react more intensely to extended benefits due to a higher return on the search for work, as suggested by Kroft and Notowidigdo (2016) who found a negative relationship between the labor supply response to an increasing benefit level and the unemployment rate.

In this paper, we fill this gap by exploiting UI benefit extension to older workers in Taiwan. Starting in 2009, workers in Taiwan aged at least 45 when losing their job are eligible for 9 months of unemployment benefits, rather than the 6 months offered to younger workers. The

[1] Several studies have offered evidence on the consumption smoothing benefits of increasing unemployment benefit levels; e.g., Gruber (1998), Kroft and Notowidigdo (2016) and Kolsrud et al. (2016). Rothstein and Valletta (2014) and Ganong and Noel (2016) estimates income or consumption losses when benefits are exhausted. The direct evidence on consumption smoothing benefits for benefit exhaustees is still scarce.

sharp discontinuity in the eligibility rule for extended benefits and administrative data on the population of UI recipients in Taiwan allows us to compare the duration and reemployment outcomes of UI recipients aged around 45 at time of job loss using a regression discontinuity (RD) design.

Taiwan's benefit extension, however, is worth investigation in its own right because of its creative program for reducing the disincentive effect of benefit extension. In Taiwan, people are paid 50% of their remaining unemployment benefits after they are reemployed. As a result, the benefit extension not only increases the potential duration, but also the bonus amount among those who are reemployed before exhausting the benefits. In this sense, Taiwan's benefits extension program is not a pure benefit extension, but a benefit extension with reemployment bonuses. A vast previous literature has shown that extended benefits increase unemployment duration (Solon, 1979; Moffitt and Nicholson, 1982; Card and Levine, 2000; Schmieder et al., 2012a; Landais, 2015). On the other hand, three random experiments conducted in the United States provide strong evidence that reemployment bonuses significantly reduce unemployment duration (Woodbury and Spiegelman, 1987; Decker et al., 2001). The countervailing effect of bonuses makes it unclear how Taiwan's benefit extension might affect the labor supply.

To help understand how Taiwan's benefit extension affects workers' search behavior, we incorporate reemployment bonuses into a search model with borrowing constraints (Lentz and Tranaes, 2005). We show that the effect of benefit extension on search effort is a combination of three effects: a liquidity effect, a substitution effect, and a bonus effect. First, the additional liquidity from extended benefits allows workers more time to search. Second, the benefit extension decreases the net wage from being employed and distorts the incentive to search. Third, the bonus effect reduces the substitution effect by 50% because the reemployment bonus program offers the unemployed 50% of their remaining entitlements when reemployed. Hence, the search model predicts that the benefit extension in Taiwan still increases the unemployment duration and reservation wage because of the bonus effect only partly offsetting the substitution effect of benefit extension.

Empirically, consistent with the theoretical prediction, our RD estimates show an increase in potential duration from 180 days to 270 days increases the insured duration of unemployment by 57 days (39%) and the nonemployment duration of unemployment by 41 days (15%). Benefit extension not only increases the duration for workers who would have exhausted benefits in the absence of benefit extension, but also for those who would have not. Specifically, we estimate that the benefit extension decreases the probability of finding a job in nine months after the initial claim by 21%, but that in six months after the initial claim also declines by 12%, suggesting workers are forward looking in their search behavior. Furthermore, the employment effect of benefit extension appears to be persistent - the probability of being reemployed in two years after the initial claim is still 3% lower for workers who are eligible for extended benefits.

Another important issue we investigate is whether the extra time spent in unemployment produced by the benefit extension helps workers find better jobs. Contrary to the prediction on the effect of benefit extension on unemployment duration, job search models do not produce sharp predictions on the effect of benefit extension on the reemployment wage. In the classic job search models from Burdett (1977) and Mortensen (1977), increasing unemployment benefits shifts reservation wage curve outward, while reservation wage declines before exhaustion point and stay constant afterwards. Alternatively, Nekoei and Weber (2016) uses a directed search model to show extended benefits cause workers to seek better jobs, while skill depreciates over the unemployment spell.[2] While we do not find any significant effect of extended benefits on wage with the first post-claim employer, we find a three month increase in potential duration decreases the wage two years after the initial claim by 2% for overall UI recipients. These results are in line with previous estimates finding either negative or

[2] Learning about the unknown wage distribution also generate declining reservation wage (Burdett and Vishwanath (1988)). Also see Gonzalez and Shi (2010) incorporating learning into a directed search model and shows that the desired wage decreases over unemployment spell.

insignificant wage effect of extended benefits (Schmieder et al., 2016; Nekoei and Weber, 2016; Barbanchon, 2016; Card et al., 2007).

The search model with borrowing constraints in this paper points out the reservation wage response of UI extension is stronger for workers who are more likely to be exhaustees and those who are more liquidity-constrained at exhaustion point. Consistent with Nekoei and Weber (2016) and Caliendo et al. (2013), benefit extension is estimated to increase the reemployment wage by 2.6% for workers whose predicted probability of exhausting benefits fell into the 4th quartile, but not for workers at lower quartiles. Using average pre-unemployment earnings as an index for liquidity constraints (Centeno and Novo, 2009), we show that the positive wage effect is disproportionately driven by the wage gains for low-wage workers. In other words, on average, benefit extension only generates match quality gains for workers who are likely to exhaust benefits and become liquidity-constrained. Our results suggest the number of liquidity-constrained workers at exhaustion point play an important role on the effect of benefit extension on job match quality.

In the next section, we describe the UI system and benefit extension program in Taiwan. Section 3 provides a short review of the literature related to our paper. Section 4 present a search model for predicting the effect of Taiwan's benefit extension program on unemployment duration. Section 5 introduces the administrative UI data and the sample for empirical analysis. In Section 6, we use the RD design to estimate the effects of benefit extension on the unemployment duration and reemployment wage. In Section 7, we discuss the liquidity effect of benefit extension. Section 8 concludes.

2 Institution Background

2.1 Unemployment Insurance in Taiwan

The unemployment benefits in Taiwan form one part of the overall benefits of employment insurance, which is a mandatory and national program that offers unemployment benefits, reemployment bonuses, vocational training living allowances, parental leave allowances and National Health Insurance (NHI) premium subsidies. It covers all Taiwanese workers excluding civil servants and the self employed. It is financed by 1% of the monthly insured wage. 20% is imposed on workers, 70% on employers, and the government pays the remaining 10%.

To be eligible for unemployment benefits, job losers aged 15 to 65 must have at least one year of employment history in the three years prior to job loss.[3] To receive the first month of benefits, a claimant must register for the governmental employment service and serve a 14 days waiting period. If workers do not find jobs by the end of the waiting period, the insured duration begins up to the maximum duration of benefits. The maximum duration of benefits is six months for workers aged below 45 at job loss, while it extends to nine months for workers aged 45 or more at job loss.[4] Unlike in the United States, where benefits are weekly, unemployed workers in Taiwan claim benefits on a monthly basis. The Bureau of Labor Insurance treats one month as a period of 30 days. If workers are reemployed before the end of a 30-day interval, the duration of benefits in the month is reduced. The monthly unemployment benefits replace 60% of the average insured wage during the six months prior to job loss for those with no non-working dependents.[5] For UI recipients having non-working dependents, the replacement rate is increased up to 80%, depending on the number of dependents. Monthly unemployment benefits are subject to a maximum. The maximum equals $26,340$ NTD (≈ 878 dollars) throughout our sample period from 2009 to 2012.

Workers are required to actively search for a job when receiving benefits. Specifically, they have to list at least two job contacts for each continued claim, which the UI agency

[3]Only workers losing their jobs involuntarily or due to the ending of a fixed term contract are eligible. Involuntary separation from employment refers to separation from employment because the insured unit has closed down, relocated, suspended business, dissolved, or filed bankruptcy. In addition, employment history is the number of days that workers have been enrolled in the employment insurance. Since part-time workers must be insured according to the Employment Insurance Act, history as a part-time worker is also included.

[4]The only exception is that UI recipients who are disability card holders are eligible for nine months of benefits regardless of age at time of job loss. However, few UI recipients hold disability cards. Our data show that only 0.8% of workers aged under 45 receive unemployment benefits for longer than six months during our sample period.

[5]The refers to the last six months for which workers were enrolled in employment insurance prior to job loss.

will confirm by contacting the employers to which workers sent applications. In general, the work search test plays the role of the stick that promotes rapid employment through punishment.[6] The other strategy is a carrot: Taiwan's UI provides generous financial incentives to encourage workers to go back to work early, in the form of a reemployment bonus. The reemployment bonus program offers 50% of remaining unemployment benefits to UI recipients who find jobs before exhausting their benefits and who accumulate at least three months of employment history after reemployment. For example, a worker eligible for six months of benefits who found a job at the end of the second month of the insured duration would be eligible to receive an additional two months of benefits as a bonus if she accumulated at least three months of employment history after reemployment. The three months reemployment period does not have to be a consecutive three months for the same employers. Someone who worked for multiple employers for three months after reemployment would also qualify for a bonus.

Unlike in many European countries, e.g. Austria and Germany, Taiwan's UI does not offer mean tested unemployment assistance after the benefits have been exhausted. However, job losers in Taiwan are eligible for 6 months of vocational training subsidies regardless of age if they register for employment service and participate in full-time vocation training. Same as unemployment benefits, the monthly subsidies equals 60% of average insured wage during the 6 months prior to job loss. Workers are not eligible for unemployment benefits when they participate in vocational training. However, it is not prohibited that workers claim unemployment benefits after completing training program and still unemployed, or participate in training after they have received benefits for a certain amount of time. During our sample period, about 6.5% of UI recipients participated in vocational training.

[6] For introduction and discussion on work search test in the United States, please see Woodbury (2015).

2.2 Benefit Extension for Older Workers

As shown in Figure 1, the UI extension went into effect on May, 1, 2009 before the reemployment bonus program took effect in 2003. As a result, Taiwan's version of UI extension not only extended the unemployment benefits, but also increased the bonuses for which workers were potentially eligible. Taking the two UI recipients in Figure 2 as an example, UI recipient 1 (eligible for six months of unemployment benefits) finds a job at the end of the second month of insured duration and is eligible for an additional two months of benefits as a bonus, while UI recipient 2 (eligible for nine months of unemployment benefits) finds a job at the end of the second month of insured duration and is eligible for an additional three and a half months of benefits as a bonus. Therefore, the reemployment bonus creates a countervailing force that alleviates the moral hazard effect of benefits extension. In Section 4, we use a search model to predict the effect of benefits extension in a UI system with the reemployment bonus.

3 Previous Literature

Our studies are connected to the literature on labor supply and wage effect of extended unemployment benefits. We summarize previous findings and make a simple comparison with closely related studies.

Early research correlating the duration and reemployment outcomes to potential duration conditional on worker characteristics and unemployment rate suffers from omitted variable bias and reverse causality since the potential duration of benefits is a function of previous earnings and (insured) unemployment rate, which are determinants for labor supply.[7] Researchers have exploited the differing timing and degrees of benefit extension across states to identify the effect of extended benefits on the labor supply (Farber and Valletta, 2015; Farber et al., 2015). However, it is hard to completely eliminate the possibilities that the increase

[7] For studies before the 2000s, please see Krueger and Meyer (2002). Hamermesh (1977) provides a concise survey of the literature from the 1970s.

in potential duration may be confounded with the change in labor market conditions when using the identification strategy. Recent evidence from Europe, including Schmieder et al. (2012a) in Germany, Barbanchon (2016) in France, (Card et al., 2007), Lalive (2008) and Nekoei and Weber (2016) in Austria, and Centeno and Novo (2009) in Portugal, addresses this concern using age or tenure based discontinuities in the eligibility for extended benefits to identify the effects of benefit extension on duration outcomes.

The estimated nonemployment duration elasticities in Europe and the United States range from 0.1 to 1, with a median around 0.4, and the insured duration elasticities range from 0.52 to 1.35, with a median of 0.58. Our estimated nonemployment duration elasticity of 0.3 and the insured duration elasticity of 0.78 are in the range of previous estimates.[8] Cross country comparison, however, is a difficult task because the estimates from different countries might differ for many reasons, including difference in institutional background and labor market conditions, difference in sample years, difference in affected workers, and difference in measure of duration. For simplicity, we make comparison with Schmieder et al. (2012a) and Nekoei and Weber (2016), who study benefit extension for older workers in Germany and Austria, respectively. Our estimated nonemployment duration elasticity of 0.3 and the insured duration elasticity of 0.78 are both larger than theirs. We offer two possible explanations. First, the potential benefit duration in Taiwan is shorter compared to Germany and Austria. The shorter potential duration increases the exhaustion rate. Since exhaustees are more affected by benefit extension, the duration elasticities tend to be larger for countries have shorter potential duration. Second, Taiwan does not have Unemployment Assistance (UA) program as Germany and Austria. The UA program offers a second tier benefits with a lower replacement rate for exhaustees. The existence of the UA mitigates the income loss after exhausting benefits, and decreases the net replacement rate of extended benefits, making the effect of the effect of benefit extension smaller.

On the other hand, most of studies obtain insignificant estimates for the effect of extended

[8] Please see Schmieder and von Wachter (2016) for excellent survey.

benefits on the reemployment wage for overall UI recipients (Card et al., 2007; Lalive, 2007; van Ours and Vodopivec, 2008; Barbanchon, 2016) except Schmieder et al. (2016) and Nekoei and Weber (2016). Schmieder et al. (2016) estimates a six months increase in potential duration precisely lowers the reemployment wage by less than 1%, whereas Nekoei and Weber (2016)'s estimates show a nine weeks increase in potential duration increases reemployment wage by less than 0.5%. Nekoei and Weber (2016) argues that this can be attributed to the smaller duration response to extended benefits in Austria. He reconciles their findings with previous estimates by showing, both theoretically and empirically, a negative relationship between the effect of benefit extension on duration and the effect of benefit extension on the reemployment wage. Given the fact that our estimated nonemployment duration elasticity is larger than Schmieder et al. (2016), it is not surprising that we do not find wage gain for overall UI recipients.

Several studies have found heterogeneous effects of benefit extension or related UI programs on job match quality. Nekoei and Weber (2016) and Caliendo et al. (2013) find that benefit extension causes a larger wage gain for benefit exhaustees. Centeno and Novo (2009) estimates that benefit extension in Portugal increases the reemployment wage for lower-wage workers. Lachowska et al. (2016), who examine the effect of work search test on long-term employment outcomes using Washington Alternative Work Search experiment, provide evidence that work test increases job tenure with the first post-claim employer for permanent job losers, and it tends to select lower-wage workers into reemployment. They conclude work test may be an important policy for improving reemployment prospects for lower-wage, permanent job losers. Our paper links to this line of research by showing the UI extension in Taiwan generates significant wage gain for lower-wage, potential exhaustees.

4 Theorectical Discussions

For our discussion on empirical analysis on the effect of Taiwan's benefit extension on duration and reemployment wage, we incorporate Taiwan's version of a reemployment bonus into a discrete time search model with borrowing constraints from Lentz and Tranaes (2005), Chetty (2008) and Landais (2015).

An unemployed worker becomes unemployed at time 0 and lives for T periods. She determines the probability of finding a job in period t by varying search intensity, s_t, at a cost of $g(s_t)$, which is strictly increasing and convex. If she is unemployed at time t, she receives an unemployment benefit, b for at most P periods. If she is employed at time t, she earns a wage rate w_t, pays a tax rate, τ, and keeps the job forever. In period t, the worker holds asset A_t and faces a borrowing constraint: $A_{t+1} \geq L$. A worker reemployed before running out of benefits receives a reemployment bonus, r_t, equal to θ percent of her remaining benefits; otherwise, $r_t = 0$.[9] Formally,

$$r_t = \theta \cdot \sum_{k=t}^{P-1} b_k, 0 < \theta < 1$$

The worker's consumption at time t equals the difference between income and saving. The income depends on her employment status, while the change in asset, $A_{t+1} - A_t$ reflects her saving. When employed, she earns wage rate, w_t, a bonus, r_t, and pays a tax, τ. The flow utility when employed at time t equals $u(c_t^e) = u(A_t - A_{t+1} + w_t + r_t - \tau)$, where c_t^e indicates the consumption when employed at time t. The value of being employed in period t is

$$V_t = \max_{A_{t+1}} u(A_t - A_{t+1} + w_t + r_t - \tau) + \beta V_{t+1}(A_{t+1})$$

If an unemployed worker cannot find a job at time t, her flow utility is equal to $u(c_t^u) =$

[9] In the case of Taiwan's reemployment bonus program, $\theta = 0.5$. Here, we assume an arbitrary θ between 0 and 1.

$u(A_t - A_{t+1} + b_t)$. The value of being unemployed in period t is

$$U_t = \max_{A_{t+1}} u(A_t - A_{t+1} + b_t) + \beta J_{t+1}(A_{t+1}),$$

where $J_{t+1}(A_{t+1})$ is the value of entering period $t+1$ unemployed with asset A_{t+1}. The worker without a job at the beginning of period t maximizes

$$J_t(A_t) = \max_{s_t} s_t V_t(A_t) + (1 - s_t)U_t(A_t) - g(s_t),$$

For $t \leq P$, the effect of extended benefits on the search intensity at time t can be written

$$\begin{aligned} \frac{\partial s_t}{\partial P} &= b\frac{u'(c_t^e) - u'(c_t^u)}{g''(s_t)} - b(1-\theta)S_{t+1}(P)\frac{u'(c_t^e)}{g''(s_t)} && (1) \\ &= b\frac{\partial s_t}{\partial A_t} - b(1-\theta)S_{t+1}(P)\frac{\partial s_t}{\partial w_t} && (2) \end{aligned}$$

where $S_{t+1}(P) = (1 - s_{t+1})..(1 - s_P)$ is the survival rate in period P conditional on being unemployed in period $t+1$. Equation 1 is similar in spirit to the formulas from Chetty (2008) and Landais (2015) with an adjustment on the moral hazard effect. On the one hand, an increase in potential duration, P, increases workers' ability to smooth consumption during unemployment, giving more time to search for a job ($\frac{\partial s_t}{\partial A_t}$). On the other hand, it decreases workers' net wages by one dollar, which distorts the incentive to search ($\frac{\partial s_t}{\partial w_t}$). The reemployment bonus counteracts the moral hazard effect by offering θ remaining benefits for workers reemployed before the exhaustion point. Therefore, extending the potential duration in a UI system offering reemployment bonuses increases the unemployment duration by a smaller magnitude than in the absence of bonuses, because reemployment bonuses countervail the moral hazard effect of extended benefits while leaving the liquidity effect of extended benefits intact.

To guide our discussion on the effect of benefit extension on the reemployment wage, in the appendix, we include reservation wage choice, R_t, into the model. For $t \leq P$, the effect of extended benefits on the reservation wage at time t is

$$\begin{aligned} \frac{\partial R_t}{\partial P} &= bS_{t+1}(P)\frac{u'(c_P^u) - u'(c_P^e)}{T-t} + b(1-\theta)\frac{u'(c_P^e)}{T-t} && (3) \\ &= bS_{t+1}(P)\frac{\partial R_t}{\partial A_P} - b(1-\theta)\frac{\partial R_t}{\partial r_P}. && (4) \end{aligned}$$

Equation 3 shows that benefit extension increases workers selectivity for jobs not only because it increases the opportunity cost of being employed at time P but also because workers suffers consumption drop when exhausting benefits. Combining equations 1 and 3 yield an important predictions on heterogeneous effects of benefit extension: workers who are more likely to exhaust their benefits (higher $S_{t+1}(P)$) and experience larger consumption drop at exhaustion point (larger $u'(c_P^u) - u'(c_P^e)$) should have a larger increase in unemployment duration and reservation wage. We provide evidence on the heterogeneous effect of benefit extension in our empirical section.

5 Data and Sample

This section describes the administrative data from Taiwan. The day-to-day information on the universe of UI benefit receipt, and the corresponding employment and wage records allows us to precisely calculate our key variables, including worker's age at job loss, insured duration, nonemployment duration, and the reemployment wage.

The data we use come from two sources: the administrative unemployment benefits file from January 1999 to December 2013 and the corresponding employment insurance enrollee file from the Bureau of Labor Insurance. First, each observation in the unemployment benefits file represents one beneficiary case, and records each UI recipient's date of birth, date of job loss, first and last date of benefit receipt, average previous insured wage in the six months prior to layoff, individual identifier and some demographics, including gender, number of dependants, place of birth, four-digit previous occupation, etc. Using the unemployment benefits file alone, we can create a dataset in which each observation represents one

UI spell, containing information on the UI recipient's exact age at job loss and the insured duration of each unemployment spell, which is the total number of days for which workers received unemployment benefits during an unemployment spell.

Second, we match the dataset to the employment insurance enrollee file on each UI spell to construct reemployment outcomes. In the employment insurance enrollee file, each observation represents a change in the employment record, including any new enrollments in employment insurance, cancellations of Employment Insurance (job separation) or wage changes. Using the matched dataset, we define the nonemployment duration as the total number of days from the start of receiving unemployment benefits to the next registered date employment (Schmieder et al., 2012a). For UI recipients not observed to have been reemployed during our samples, indicating they failed to find a job, became self-employed or dropped out of the labor force. We cap nonemployment duration at 730 days.

We consider two outcomes for the reemployment wage. On the one hand, we consider the wage with first post-claim employer, which is the monthly insured wage applying to the first spell of registered employment after job loss. On the other hand, we also consider the reemployment wage two years after the initial claim. Both of the reemployment wage outcomes are conditional on employment. Monthly insured wage is categorized into one of 20 buckets, ranging from a minimum to a maximum. The minimum insured wage is the minimum wage. During our sample period from 2009 to 2012, the minimum wage was $17,280$ NTD in 2009, $17,880$ NTD in 2011, and $18,780$ NTD in 2012. The maximum insured wage equals $43,900$ NTD (≈ 1463 dollars) throughout the sample period.[10]

We impose three sample restrictions on our data. First, since the benefit extension took effect on May 1, 2009, we drop any UI spell starting before that date. Second, we exclude UI spells starting after January 1, 2012 because the nonemployment duration for those UI spells are capped at a duration shorter than 730 days. Third, we focus on the sample around the age 45 cutoff. Column 1 of Table 1 reports the summary statistics for the UI recipients aged

[10]For example, in 2012, there are 20 categories from $18,780$, $19,200$,.., $42,000$ to $43,900$.

25-65 during the sample period. Our baseline RD sample are workers aged 43-46 at job loss in Columns 2. The exhaustion rate for workers around 45 is high, and the nonemployment duration is on average longer than 180 days, suggesting a substantial amount of workers do not find employment after exhausting benefits. The remaining columns show the statistics for workers from the other age groups. Older workers tend to have longer insured duration and nonemployment duration, and to earn higher wages.

6 Regression Discontinuity Design

Using the fact that UI recipients in Taiwan aged 45 or older at time of job loss are eligible for nine months of unemployment benefits, while those younger than 45 at time job loss are only eligible for six months of benefits, the RD design isolates the variation in potential duration as long as workers around the age 45 cutoff are similar to each other except for the eligibility for the UI extension. Consider

$$y_i = \alpha + \beta_{EB} Age45_i + f(a_i) + v_i, \tag{5}$$

where y_i is an outcome variable, including insured duration, nonemployment duration and difference in log wage between two periods of employment. $Age45_i$ is an indicator for a UI recipient's age at time of job loss being at least 45, and a_i is the worker's age at job loss. β_{EB} is the coefficient of interest, capturing the effects of a three-month increase in potential duration. The key identification assumption here is that the outcome variables should evolve smoothly over the cutoff in the absence of extended benefits. Note that $Age45_i$ depends solely on a_i. If the effects of age at job loss are adequately controlled by $f(a_i)$ such that $E(v_i|a_i) = 0$, β_{EB} will identify the effects of the extended benefits. For our baseline results, we estimate the equation using the sample of workers age 43 to 46 at time of job loss, and consider $f(a_i)$ to be a linear function interacted with the extended benefits dummy. Specifically,

$$f(a_i) = (1 - Age45_i)[\pi_0(a_i - 45)] + Age45_i[\pi_1(a_i - 45)]. \tag{6}$$

Since the choice of the estimation sample as UI recipients age 43 to 46 at time of layoff is arbitrary, we also report RD estimates using the optimal bandwidth and robust confidence interval proposed by Calonico et al. (2014), and test for the sensitivity of the bandwidth choice.

6.1 Identification Assumptions

The validity of RD design depends on whether the UI recipients age around 45 cutoff are identical on average except for the eligibility for their eligibility for extended benefits. In Taiwan's UI, it is unlikely that workers will be able to manipulate the eligibility rule for extended benefits because of it being based on their age at time of job loss rather than age when claiming benefits.[11] It seems possible, however, that firms might be willing to wait to lay off workers for a certain period of time, until they were eligible for extended benefits. We would expect an extra mass of workers just above 45 years old if firms did so. Furthermore, if these workers or employers fell into certain types, then this sorting would not be random and that would need to be addressed. We investigate the validity of our RD design by examining the frequency of UI recipients over different ages, and the means of the observables around the cutoff as Lee and Lemieux (2010) suggest.

Figure 3 shows the number of UI recipients aged 40 to 50 at layoff within each age bin. Each age bin represents the total number of new claimants in a 30 days interval. Below the age-45 cutoff, there are roughly 450 new claimants within each age interval, and the number of new claimants decreases with age at job loss. Consistent with Schmieder et al. (2012a), we find that there are about 150 more workers losing their jobs within the 30 days above age 45 than just below it, and that the number of new claimants within a few months above age 45 is slightly higher than that just below age 45. This increase in the number of UI

[11] The eligibility rule for extended benefits in Germany is based on age when claiming unemployment benefits. Schmieder et al. (2012a) finds a slight increase in the number of new claimants on the right of each age cutoff, and address this concern using a variety of methods, including adding covariates, donut RD and bounding.

recipients at and just above the cutoff is significant at the 5% level using the density test proposed by Cattaneo et al. (2016). However, it accounts for less than 1% of workers aged 43 to 46 at time of job loss, and is unlikely to invalidate our RD design. To alleviate the concern that this small discontinuity in bin size at the cutoff might bias our RD estimates, we implement the donut RD strategy suggested by Barreca et al. (2016). We intentionally exclude observations within 180 days around the cutoff to examine how selective laying off around the cutoff affects the results. Our results are robust to this removal of observations.

To check whether the sorting behavior is non-random, we look for any discontinuities in the means of workers' characteristics around the cutoff. Figure 4 plots the number of days between losing one's job and claiming benefits, whether workers are female, whether workers previously worked in the manufacturing industry, the number of dependants and the average log wage in the six months prior to layoff. The means either evolves smoothly or shows economically small discontinuities around the cutoff. To make sure this small degree of discontinuities does not invalidate our RD design, we estimate the average reemployment hazard in the first six months conditional on available observables, excluding the treatment indicator as suggested by Card et al. (2007). Figure 4 shows that the predicted nonemployment duration before workers exit UI is smooth around the age-45 cutoff. In Table 2, we estimate a local linear regression using the observables as dependent variables. The estimates are either insignificant or small. In particular, the estimated effect on the predicted nonemployment duration is insignificant different from zero, showing no evidence of manipulation of the running variable.

6.2 Results on Duration Outcomes

The eligibility rule for extended benefits in Taiwan generates clear discontinuities in the relationship between age and duration outcomes, with no discernible change in wage growth at the age cutoff. Figure 5 plots our average outcomes against age at job loss using a [40, 50] window. Each bin indicates the conditional means within a width of 30 days. As shown in Figure 5 (a) and (b), the average number of days of receiving benefits shifts up by about 60 days to the right of the cutoff, while the average nonemployment duration shifts up by roughly 40 days. In Figure 5 (c), we plot the probability of having nonemployment duration lower than 180 days. The probability shows a sharp decline at the cutoff, suggesting workers are not myopic in terms of their labor supply. Importantly, there are no comparable discontinuities at other age points and the trend relationship between age and duration outcomes are approximately linear. Overall, the RD graphs of the duration outcomes imply that extending the potential duration lowers the search intensity in a UI system with a variable reemployment bonus.

Column 1 of Table 3 reports our baseline estimates for the effect of extended benefits on insured duration, nonemployment duration and the probability of finding a job within 180 days after the initial claim. A increase in potential benefits duration from 180 days to 270 days is estimated to increase the insured duration of unemployment by about 57.96 days, a 39% increase in the average insured duration and an elasticity with respect to potential duration equal to 0.78. The estimated effect of extended benefits on nonemployment duration is about 41.14 days, a 15% increase in nonemployment duration and an elasticity of 0.3.[12] The nonemployment duration elasticity is smaller than the insured duration elasticity because the effect of extended benefits on benefit exhaustees is partly mechanical. Benefit extension also increases the nonemployment duration for workers who would have not exhausted benefits. In particular, the likelihood of having nonemployment duration lower than 180 days is estimated to decline about six percentage points, a 12% declines in the baseline mean. Including covariates barely affects the estimates and precision. Columns 3 to 6 use the optimal bandwidth algorithm proposed by Calonico et al. (2014) and Calonico et al. (2016). The optimal bandwidth algorithm chooses a somewhat larger bandwidth than two

[12]The insured duration elasticity equals the percentage change in insured duration divided by the percentage change in potential duration, which is $\frac{57.96/147.32}{(9-6)/6}$. Similarly, the nonemployment duration elasticity is $\frac{41.14/276.39}{(9-6)/6}$.

years, but in the range of three to six years. In general, our estimates are robust to using the optimal bandwidths, and the robust standard errors are also similar to the conventional ones.

In order to investigate how benefit extension changes the distribution of nonemployment duration. We show the cumulative distribution of nonemployment in Figure 6 and the density distribution in Figure 7. Specifically, we plot $P(\text{nonemplyment duration} \leq k)$ and $P(k-30 < \text{nonemplyment duration} \leq k)$, where k ranges from 30 to 730, for workers aged 43-44 and 45-46. The cumulative distribution of nonemployment duration makes three points. First, the cumulative probability is significantly smaller for workers eligible for extended benefits even when k is smaller than 180, indicating workers respond to benefit extension before exhausting regular benefits. Second, the difference in the cumulative probabilities is particularly large between $k = 180$ and $k = 270$. This supports our theoretical prediction that workers who would otherwise exhaust benefits are most affected by benefit extension. Third, the difference begins to shrink around the exhaustion point of extended benefits, but this difference does not disappear even in the two years after the initial claim, providing evidence that benefit extension might have medium-term effects on labor supply. On the other hand, the density distribution shows the density for each age group declines over nonemployment duration, while it exhibits a clear spike right after the maximum duration of benefits. These spikes are consistent to Schmieder et al. (2012a) and Nekoei and Weber (2016), finding spikes in reemployment hazard around exhaustion points.

Table 4 estimates the effect of benefit extension on the probability of finding a job in k days using equation 5. We use the optimal bandwidths and specify a linear function on either side of the cutoff. Column 1 shows the effect of benefit extension on the probability of finding a job within 90 days is statistically indistinguishable from zero. As shown in Columns 2 and 3, we estimate the probability of finding a job within 180 days declines by 12%, and the that within 270 days by 21%. The magnitude of the decline in cumulative probability decreases after the exhaustion point of extended benefits, but the declines are still substantial: 8% and 3% in terms of the probability of being reemployed in 360 and 730 days.

We are especially interested in the declines in the probability of being reemployed around exhaustion point because the main policy goal is to improves these workers' ability to maintain assumption upon benefit exhaustion. It might because extended benefits provide additional liquidity for exhaustees, or because workers decrease their job search effort in order to claim extended benefits. Rothstein and Valletta (2014) and Ganong and Noel (2016) show that workers experience significant consumption drop upon exhausting their benefits, suggesting that part of the decline might be attributed to workers' increased ability to smooth consumption. We return to this issue in Section 7.

6.3 Results on Reemployment Wage

A natural follow-up question is whether workers invest the extra time on searching for better employment or it is used as leisure. We examine this issue by directly estimating the effect of benefit extension on the reemployment wage. Figure 5 (e) plots the difference between the log pre-unemployment wage and the log wage with first post-claim employer conditional on age at job loss, where the solid lines are fitted values from a local linear regression on either side of the cutoff. By differencing, we control for individual heterogeneity and alleviate the concern of selection into treatment (Hotz, 2006; Schmieder et al., 2012b). While job losers around 45 suffer from about 23% wage declines upon reemployment, the graph shows no discontinuity at the cutoff, indicating extended benefits have little impact on the reemployment wage with first post-claim employer for overall UI recipients age around 45. The RD estimates in Table 5 also shows no evidence on wage gains in any of the specifications.

It is puzzling that benefit extension is not estimated to have match quality gains. As equation 3 shows, benefit extension increases workers selectivity conditional on the timing of the unemployment spell, so we should expect higher selectivity reflects on higher reemployment wages. This argument ignores the role of the declining reservation wage over an unemployment spell. On the one hand, the model with liquidity constraints (Mortensen, 1986; Lentz and Tranaes, 2005) points out that workers lower their reservation wage or increases their search effort closer to the constraint. On the other hand, previous literature has discussed possible mechanism on declining reservation wage, including time-varying benefit schedule (Mortensen, 1977), learning (Burdett and Vishwanath, 1988), and skill depreciation (Nekoei and Weber, 2016). The insignificant wage effect implies the increased selectivity is offset by declining reservation wage.

One possible explanation for not finding increasing potential benefit duration increases the wage with first post-claim employer is that the match gain of an increase in potential duration reflects on other dimensions of the job quality (Card et al., 2007). For example, benefit extension might help workers find jobs with higher wage growth after reemployment, which is not captured by the wage with first employer. In Panel C of Table 5, we report the estimated effect of increasing potential duration from 180 days to 270 days on the the difference between the log previous wage and the log wage two years after the initial claim. Contrary to the hypothesis that benefit extension helps workers find jobs with higher wage growth, the estimates show a three months increase in potential duration decrease the wage two years after the initial claim by about 2%. Overall, we do not find natch quality gain for UI recipients around 45 years old.

6.4 Heterogeneity

Equation 3 points out the two sources of heterogeneity in the effect of UI extension on duration and wage: the survival rate at benefit exhaustion and the gap in marginal utilities between the unemployed and employed. To be specific, workers who exhaust their benefits and are liquidity-constrained at exhaustion point are predicted to be most affected. We have shown in Figure 6 and Table 4 that the labor supply response is the largest around exhaustion point. If labor supply response to extended benefits mostly occurred around exhaustion, the duration and reservation wage response should be the largest for those who are most likely to exhaust benefits.

To explore how the effects of benefit extension on duration and the reemployment wage vary across the likelihood of exhausting benefits, we first estimate the probability that workers are going to exhaust benefits, that is receiving 6 months (180 days) of unemployment benefits similar to Nekoei and Weber (2016). Specifically, using all unemployment spells before the reform, we estimate a linear probability model of a dummy variable denoting whether workers exhaust unemployment benefits on a set of predictors, including previous wage, square of previous wage, previous industry, gender, birthplace, number of job loss, month when unemployed, number of days between job loss and initial claim, and whether workers recalled to previous employers. We then divide our sample into four categories by quartiles of predicted probability of exhausting benefits, and estimate the effects of benefit extension for each group using the RD design.

Table 6 reports the baseline means and the RD estimates for each quartile. The more likely workers are going to be exhaustees, the longer the insured duration and nonemployment duration. Workers also experience a larger wage drop if they are more likely to exhaust benefits. Consistent with the predictions of equation 3, the RD estimates show the benefit extension has the strongest effect on the duration outcomes and the reemployment wage for workers most likely to exhaust benefits. To be specific, the estimates in Column 4 suggests that, for workers whose predicted probability falls into the 4th quartile, a 90-day increase in potential duration raises insured duration by 64 days and nonemployment duration by 50 days, larger than the points estimates at lower quartiles. Meanwhile, the reemployment wage foe these workers increases by 2.7%, but not for others. Hence, the UI extension appears to be more beneficial for workers who are more likely to exhaust benefits.

Recent evidence from Ganong and Noel (2016) and Rothstein and Valletta (2014) demonstrate that unemployed workers experience a significant income or consumption loss after exhausting unemployment benefits. To examine the role of liquidity constraints, we split the workers into four groups by whether workers' likelihood of exhausting benefits falls into the 4th quartile and the median of average previous wage (Centeno and Novo, 2009). Columns 1 and 2 of Table 7 show that the wage gains for workers most likely to be exhaustees are driven by workers whose average previous wage are below the median. On the contrary, for workers at lower quartiles of predicted probability of exhausting benefits, neither lower-wage nor higher-wage workers are estimated to experience wage gain sue to benefit extension. Therefore, this exercise demonstrate both the exhaustion rate and the degree of liquidity constraints are critical determinants for the effect of benefit extension on reemployment wage.

6.5 Robustness

We have demonstrated our results are robust to a variety of specifications. This section provides additional robustness analysis by checking for the sensitivity of bandwidth choice, using donut RD design, and conducting a placebo test. We also examine whether the presence of training program plays a role in our RD estimates.

To test for the sensitivity to bandwidth choice, Figure 8 plots out estimated effects of extended benefits and their 95% confidence intervals using our baseline specification and a bandwidth from 40 to 2000 days. When the bandwidth is smaller than about 400 days, the estimates are relatively volatile and imprecise. As the bandwidth size increases, the confidence intervals become narrow and the estimates stabilize at a bandwidth around 600 days. The robustness of our RD estimates to a variety of specifications and varying bandwidth assures that the estimates do not pick up nonlinear relationship between workers' age and the outcomes.

To alleviate the concern that the small discontinuity at the cutoff might bias our RD estimates, we intentionally exclude observations close to the cutoff suggested by Barreca et al. (2016) and estimate a linear regression on either side of the cutoff using a two years bandwidth. Table 8 reports the estimates removing observations within one month, two months,.., six months of the cutoff. The estimates are close to our baseline estimates, suggesting the small extra mass of UI recipients just on the right of the cutoff affects our results little. As a placebo test for our RD estimates, we plot the average outcomes conditional on age at layoff before the reform in Figure 9. There are no discernible discontinuities at the cutoff. The estimates in Table 9 also suggest there are no permanent difference between workers laid off on either side of the age threshold. The smooth relationship of workers' age at layoff with observed covariates after the reform and outcomes before the reform raises our confidence in the validity of the RD design.

Finally, the fact that job losers in Taiwan are not only eligible for unemployment benefits but also the vocational training subsidies arouses the concern that whether benefit extension induces any interaction of the two programs. On the one hand, extended benefits might be a substitute for the training subsidies. In particular, claiming extended benefits might be less costly than participating in a full-time training class. On the other hand, since extended benefits increase the duration of unemployment, a longer unemployment might increase the chance of participating in training program. Table 10 examines whether an increase in potential duration changes the probability of participating in training and the number of days workers receiving training subsidies. We find that the use of training program is not affected by benefit extension, mitigating the concern that our RD estimates might be affected by the presence of training subsidies. The results, however, are not surprising given the fact that only 8% of UI recipients age 43-46 participated in the training program.

7 Discussion on Liquidity Effect of Benefit Extension

The positive wage effect of UI extension for potential exhaustees (predicted probability at 4th quartile) who earned below the median of income distribution in Section 6.3 provides suggestive evidence on the existence of the liquidity effect of benefit extension. Can we gauge the magnitude of liquidity effect of UI extension using the estimates in Table 7?

To do this, we rewrite equation 3

$$\frac{\partial logR_t}{\partial P} = bS_{t+1}(P)\frac{\partial logR_t}{\partial A_P} - b(1-\theta)\frac{\partial logR_t}{\partial r_P}. \tag{7}$$

The difference in the wage responses between low-wage and high-wage individuals ($0.34 - 0.21$) is simply the difference in the liquidity effect between two groups if the effect of reemployment bonus does not vary over wage distribution. If we further assume the liquidity effect for high income individuals is zero, we can get the ratio of the liquidity effect to the total effect of UI extension:

$$\frac{0.34 - 0.21}{0.34} = 0.38. \tag{8}$$

This means the liquidity effect of benefit extension explains 38% of the effect of benefit extension on the reemployment wage for potential exhaustees whose previous wage were below median. This calculation might not be satisfactory for at least three reasons. First, it will underestimate the liquidity effect if the liquidity effect also exists for high-wage workers. Second, it will overestimate the liquidity effect if the marginal utility of consumption when employed is lower (smaller moral hazard) for high income individuals. Third, the most pressing concern of this calculation is that the reservation wage response to benefit extension is not necessarily the same as the mean reemployment wage response. How relaxation of these assumptions affects the above result requires future research.

We are currently investigating a different way to disentangle the liquidity effect form the total effect of extended benefits. We use the variation brought about by the reach back provision of the reemployment bonus program to estimate the effects of reeemployment bonus as well as the moral hazard effect of extended benefits. This is perhaps the most important extension of this paper because it allows us to recover the liquidity effect and estimate the welfare effect of benefit extension.

8 Conclusion

In this article, we study the unemployment benefit extension for workers aged at least 45 in Taiwan. Taiwan's benefit extension is different from those in other countries due to the reemployment bonus program. Since workers can receive half of their remaining benefits as a bonus after they become reemployed, the benefit extension not only increases the potential duration of unemployment benefits but also the bonuses for workers reemployed before exhausting benefits. Using a search model with borrowing constraints, we find that the bonus offer reduces the moral hazard effect of benefit extension by 50%, while it does not change the liquidity effect of extended benefits. Since both the liquidity and moral hazard effects increases unemployment duration and reservation wage, the model predicts that Taiwan's version of benefit extension still lengthen unemployment duration and raise reservation wage. Furthermore, the model generates an testable prediction that workers who are more likely to run out of benefits and become liquidity-constrained at exhaustion point are more responsive to benefit extension.

Empirically, we use the administrative records for the population of UI recipients and the RD design to estimate the effects of extended benefits on unemployment duration and reemployment wage. Our RD design yields three sets of findings for overall UI recipients around 45 years old. First, an increase in potential benefit duration from 180 days to 270 days increases the nonemployment duration by 39% and insured duration by 15%, in the range of previous estimates from the United States and Europe. Second, the three months increase in potential duration decreases the probability of being reemployed within 180 days after the initial claim by 12%, and that within 270 days by 21%. Moreover, the probability of finding a job within 730 days is still lower for workers eligible for extended benefits. Third, we do not find match quality gain due to benefit extension for overall UI recipients aged around 45. Specifically, our estimates show insignificant effect of benefit extension on wage with the first post-claim employer, and a significant 2% decrease in the wage two years after the initial claim.

Guided by the theoretical predictions of the search model, we examine how the duration and wage effects of UI extension vary across the probability of exhausting benefits as well as the degree of liquidity constraint at the exhaustion point. Our estimates show workers most likely to run out of benefits increases duration more, and experience a significant wage gain by 2.6%, but not for others. The positive wage effect for potential exhaustees is significant at 3.5% for lower-wage workers, but it is insignificant for higher-wage workers. The wage effect for overall UI recipients is small and insignificant might be because the wage gains are averaged out by insignificant wage effect for potential non-exhaustees or workers who can sustain consumption after exhausting benefits. These results suggest the ratio of exhaustees who are liquidity constrained at exhaustion point could be an important determinant in the wage effect of extended benefits. It will be a helpful future work to conduct a meta analysis by correlating the exhaustion rate and some measure of liquidity constraints to the estimated wage effect of benefit extension.

Finally, we still have limited knowledge on the ratio of liquidity effect to the moral hazard effect of UI extension. We propose a approach that compares the wage response to benefit extension for high-wage and -low-wage workers. Assuming the mean wage response is the same as the reservation wage response and the moral hazard effect is constant over wage distribution, we estimate the liquidity effect accounts for 38% of the total effect of benefit extension on the reemployment wage for potential exhaustees who earned below median prior to job loss. The robustness of this result requires further investigation.

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9 Tables

Table 1: Descriptive Statistics

	25-65 (1)	43-46 (2)	43-44 (3)	45-46 (4)	25-45 (6)	45-65 (7)
age (years)	37.79	45.00	43.99	45.97	33.96	50.24
female	0.52	0.49	0.50	0.48	0.54	0.46
number of dependants	0.67	1.14	1.17	1.10	0.64	0.80
previous wage (NTD)	29,810	30,853	30,757	30,947	29,401	31,134
worked in manufacturing	0.30	0.33	0.31	0.34	0.30	0.32
participate in training	0.07	0.08	0.08	0.08	0.07	0.07
duration of training (days)	6.45	7.94	7.98	7.91	6.36	6.73
insured duration (days)	145.74	175.04	144.22	204.90	124.91	213.30
nonemployment duration (days)	255.95	294.97	272.41	316.83	224.23	358.83
right censored at 730 days	0.12	0.15	0.13	0.15	0.09	0.21
exhaustion rate	0.53	0.65	0.61	0.69	046	0.73
recall rate	0.13	0.12	0.12	0.13	0.12	0.18
reemployment wage (NTD)	25,554	25,907	25,703	26,104	25,303	26,367
observations	187,450	20,893	10,283	10,610	143,284	44,166

Note: This table shows the means of our main variables from the extended benefits sample. The sample in Column 1 consists of all UI recipients starting UI spells between May, 1, 2009 and Jan. 1, 2012. Columns 2-6 report the results for UI recipients in the same sample period for workers from five different age groups. Nonemployment duration is censored at 730 days. We define exhaustion rate as the ratio of workers whose insured duration is 180 days or longer.

Table 2: Estimates of Smoothness of Predetermined Covariates

	(1) Delay Days	(2) Female	(3) Manu. Sector	(4) # of Dependents	(5) Log Previous Wage	(6) Predicted Nonemp. Dur.
β_{EB}	-0.70	-0.00	0.018*	0.01	0.013*	1.34
	(2.06)	(0.10)	(0.010)	(0.01)	(0.007)	(1.31)
Sample size	46,916	43,035	42,036	37,961	50,903	50,706
Poly. model	linear	linear	linear	linear	linear	linear
Bandwidth (days)	CCT	CCT	CCT	CCT	CCT	CCT

Note: This table checks for smoothness of mean predetermined variables by estimating a local linear regression using the optimal bandwidth by Calonico et al. (2014) and triangular kernel. The sample are workers aged within the bandwidth and starting UI spells between May 1, 2009 and Jan. 1, 2012. The predictors for nonemployment duration are previous wage, squared previous wage, previous industry, gender, place of birth, number of dependants, month/year at job loss and the number of days between job loss and initial claim. Standard errors in parentheses are clustered by age in days. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table 3: The Effect of Extended Benefits on Insured Duration and Nonemployment Duration

	(1)	(2)	(3)	(4)	(5)	(6)
Insured Duration						
β_{EB}	57.96***	58.29***	56.55***	56.12***	57.09***	57.23***
	(1.97)	(1.95)	(1.50)	(1.74)	(2.25)	(2.25)
Baseline mean			147.32			
Sample size	20,906	20,893	40,507	40,507	37,785	50,680
Nonemployment Duration						
β_{EB}	41.14***	43.02***	36.23***	37.76***	40.41***	41.86***
	(6.90)	(6.90)	(5.18)	(6.01)	(7.96)	(7.95)
Baseline mean			276.39			
Sample size	20,906	20,893	40,987	40,987	36,589	48,906
Nonemployment Duration < 180 days						
β_{EB}	-0.06***	-0.06***	-0.06***	-0.06***	-0.06***	-0.07***
	(0.01)	(0.01)	(0.01)	(0.01)	(0.02)	(0.02)
Baseline mean			0.46			
Sample size	20,906	20,893	31,887	31,887	34,770	34,382
Bias-corrected	–	–	–	Yes	Yes	Yes
Covariates	–	Yes	–	–	–	Yes
Poly. model	linear	linear	linear	linear	quadratic	quadratic
Bandwidth (days)	730	730	CCT	CCT	CCT	CCT

Note: This table shows the estimates of the effect of increasing potential duration from 6 to 9 months on insured duration, nonemployment duration and the probability that nonemployment duration is less than 180 days. Column 1 estimates a linear regression on either side of the cutoff using sample from workers aged 43-46 at job loss, and starting UI spells between May 1, 2009 and Jan. 1, 2012. Column 2 includes the following covariates: previous wage, squared previous wage, previous industry, gender, place of birth, number of dependants, month/year at job loss and number of job loss. Columns 3 reports the estimates using optimal bandwidth algorithm from Calonico et al. (2014). The optimal bandwidths vary with the outcome variables, in the range of 3 to 6 years. The bias correction estimates and the corresponding robust standard errors are presented in the Column 4. In Column 5, we report the bias correction estimates and robust standard error using a local quadratic regression. Column 6 reports the bias correction estimates and robust standard error using a local quadratic regression with covariates (Calonico et al. (2016)). Standard errors in parentheses are all clustered by age in days. Columns 1 and 2 use rectangular kernel. Columns 3-6 use triangular kernel. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table 4: The Effect of Extended Benefits on the Probability of Being Reemployed in k Days

	(1) $k=90$	(2) $k=180$	(3) $k=270$	(4) $k=360$	(5) $k=720$
β_{EB}	-0.013	-0.059***	-0.139***	-0.062***	-0.030***
	(0.010)	(0.010)	(0.011)	(0.008)	(0.008)
Baseline mean	0.309	0.478	0.652	0.724	0.824
Sample size	38,376	36,463	30,801	46,588	50,551
Poly. model	linear	linear	linear	linear	linear
Bandwidth (days)	CCT	CCT	CCT	CCT	CCT

Note: This table shows the estimates of the effect of increasing potential duration from 6 to 9 months on the probability of finding a job in k days. We estimate equation 5 using the optimal bandwidth by Calonico et al. (2014) and triangular kernel. The sample are workers aged within the bandwidth, and starting UI spells between May 1, 2009 and Jan. 1, 2012. Standard errors in parentheses are clustered by age in days. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table 5: The Effect of Extended Benefits on Reemployment Wage

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A:						
Log Reemp. Wage with 1st Post-Claim Employer						
β_{EB}	0.014	0.006	0.013	0.012	0.011	0.007
	(0.011)	(0.010)	(0.007)	(0.009)	(0.012)	(0.012)
Baseline mean			10.03			
Sample size	17,845	17,835	49,721	49,721	51,361	40,857
Panel B:						
Diff. in Log Wage with 1st Post-Claim Employer						
β_{EB}	-0.005	0.006	-0.001	-0.004	-0.004	0.006
	(0.012)	(0.010)	(0.009)	(0.009)	(0.013)	(0.013)
Baseline mean			-0.234			
Sample size	17,845	17,835	28,645	28,645	26,426	41,249
Panel C:						
Diff. in Log Wage 2 years After Initial Claim						
β_{EB}	-0.021*	-0.011	-0.019*	-0.023**	-0.025*	-0.019*
	(0.011)	(0.009)	(0.010)	(0.012)	(0.013)	(0.011)
Baseline mean			-0.234			
Sample size	17,845	17,835	25,631	25,631	26,426	44,866
Bias-corrected	–	–	–	Yes	Yes	Yes
Covariates	–	Yes	–	–	–	Yes
Poly. model	linear	linear	linear	linear	quadratic	quadratic
Bandwidth (days)	730	730	CCT	CCT	CCT	CCT

Note: This table shows the estimates of the effect of increasing potential duration from 6 to 9 months on insured duration, nonemployment duration and daily reemployment hazard in the first 6 months. Column 1 estimates a linear regression on either side of the cutoff using sample from workers aged 43-46 at job loss, and starting UI spells between May 1, 2009 and Jan. 1, 2012. Column 2 includes the following covariates: previous wage, squared previous wage, previous industry, gender, place of birth, number of dependants, month/year at job loss and number of job loss. Columns 3 reports the estimates using optimal bandwidth algorithm from Calonico et al. (2014). The optimal bandwidths vary with the outcome variables, in the range of 3 to 6 years. The bias correction estimates and the corresponding robust standard errors are presented in the Column 4. In Column 5, we report the bias correction estimates and robust standard error using a local quadratic regression. Standard errors in parentheses are all clustered by age in days. Column 6 reports the bias correction estimates and robust standard error using a local quadratic regression with covariates (Calonico et al. (2016)). Columns 1 and 2 use rectangular kernel. Columns 3-6 use triangular kernel. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table 6: The Effect of Extended Benefits on Unemployment Duration and Reemployment Wage: By Quartiles of Predicted Probaility of Exhausting Benefits

	(1) 1st	(2) 2nd	(3) 3rd	(4) 4th
Insured Duration				
β_{EB}	41.61***	52.79***	60.45***	64.04***
	(3.85)	(2.97)	(2.96)	(2.44)
Baseline mean	139.94	140.93	147.28	156.93
Sample size	7,777	12,512	11,049	14,757
Nonemployment Duration				
β_{EB}	28.56**	24.94***	33.54***	50.81***
	(11.04)	(9.22)	(9.88)	(10.54)
Baseline mean	211.63	239.33	276.43	346.25
Sample size	7,773	11,412	12,120	13,155
Diff. in Log Wage with 1st Post-Claim Employer				
β_{EB}	-0.009	0.005	0.000	0.027**
	(0.018)	(0.013)	(0.015)	(0.014)
Baseline mean	-0.184	-0.197	-0.250	-0.289
Sample size	7,038	12,307	10,045	12,604
Covariates	Yes	Yes	Yes	Yes
Poly. model	linear	linear	linear	linear
Bandwidth (days)	CCT	CCT	CCT	CCT

Note: This table shows the estimates of the effect of increasing potential duration from 6 to 9 months on insured duration, nonemployment duration, and the difference in log wage between post- and pre- unemployment for four different groups divided by quartiles of predicted probability of exhausting benefits. Please see section ?? for estimation on the probability of exhausting benefits. We estimate a local linear regression using the optimal bandwidth by Calonico et al. (2014) and triangular kernel. The sample are workers aged within the optimal bandwidth, and starting UI spells between May 1, 2009 and Jan. 1, 2012. Standard errors in parentheses are clustered by age in days. All columns include the following covariates: previous wage, previous industry, gender, place of birth, number of dependants, number of job loss, number of days between job loss and initial claim, and month/year at job loss. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

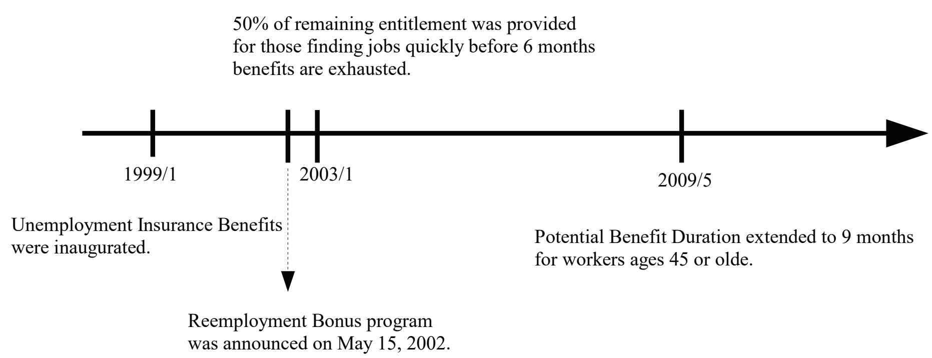
Table 7: The Effect of Extended Benefits on Unemployment Duration and Reemployment Wage: By Quartiles of Predicted Probaility of Exhausting Benefits and Previous Wage

	(1)	(2)	(3)	(4)
	4th Quartile of Exhaustion Rate		1-3 Quartiles of Exhaustion Rate	
	Below 30,300	Above 30,300	Below 30,300	Above 30,300
Insured Duration				
β_{EB}	61.70***	66.24***	53.11***	52.23**
	(3.34)	(2.77)	(2.68)	(2.77)
Baseline mean	155.02	158.33	147.24	139.32
Sample size	8,039	9,545	14,383	13,995
Nonemployment Duration				
β_{EB}	46.54***	44.41***	19.76**	37.14***
	(14.35)	(13.24)	(8.26)	(8.69)
Baseline mean	339.90	350.74	256.75	237.83
Sample size	6,841	8,478	16,091	14,284
Diff. in Log Wage with 1st Post-Claim Employer				
β_{EB}	0.035**	0.021	-0.002	0.002
	(0.017)	(0.024)	(0.011)	(0.015)
Baseline mean	-0.10	-0.46	-0.06	-0.37
Sample size	3,544	6,172	10,772	14,104
Covariates	Yes	Yes	Yes	Yes
Poly. model	linear	linear	linear	linear
Bandwidth (days)	CCT	CCT	CCT	CCT

Note: This table shows the estimates of the effect of increasing potential duration from 6 to 9 months on insured duration, nonemployment duration, and the difference in log wage by quartiles of predicted probability of exhausting benefits and average previous wage. We estimate a local linear regression using the optimal bandwidth by Calonico et al. (2014) and triangular kernel. The sample are workers aged within the bandwidth, and starting UI spells between May 1, 2009 and Jan. 1, 2012. Standard errors in parentheses are clustered by age in days. All columns include the following covariates: previous wage, previous industry, gender, place of birth, number of dependants, number of job loss, number of days between job loss and initial claim, and month/year at job loss. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

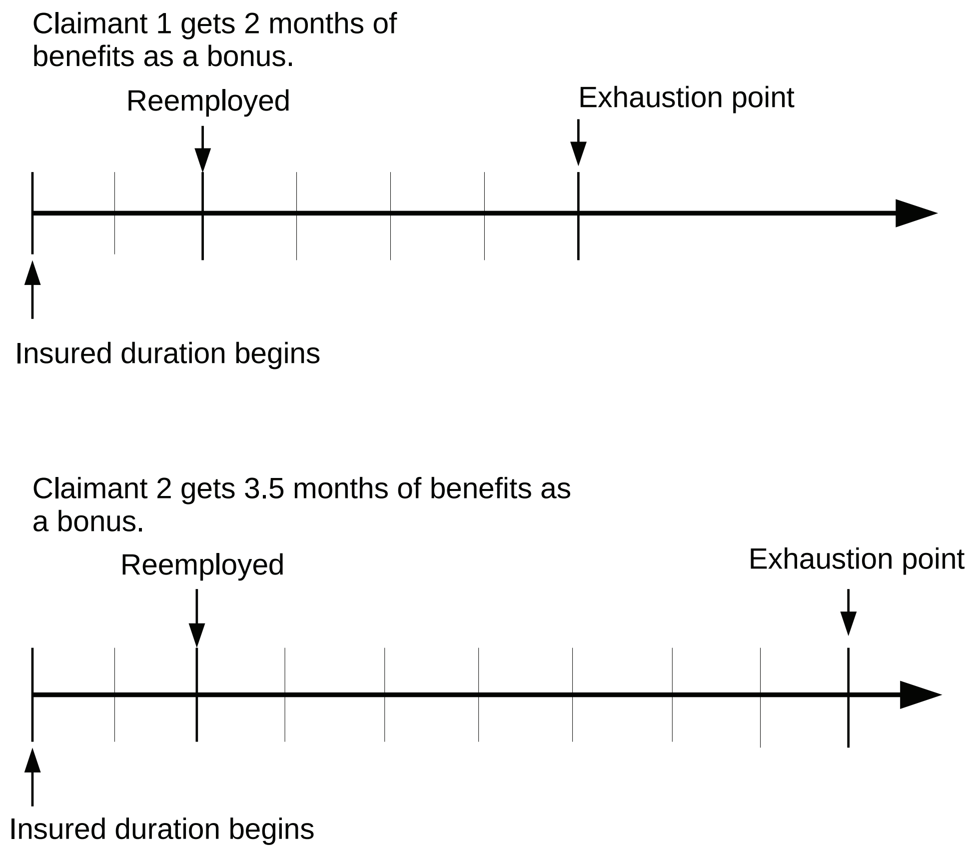
10 Figures

Figure 1: Timeline of UI Reforms



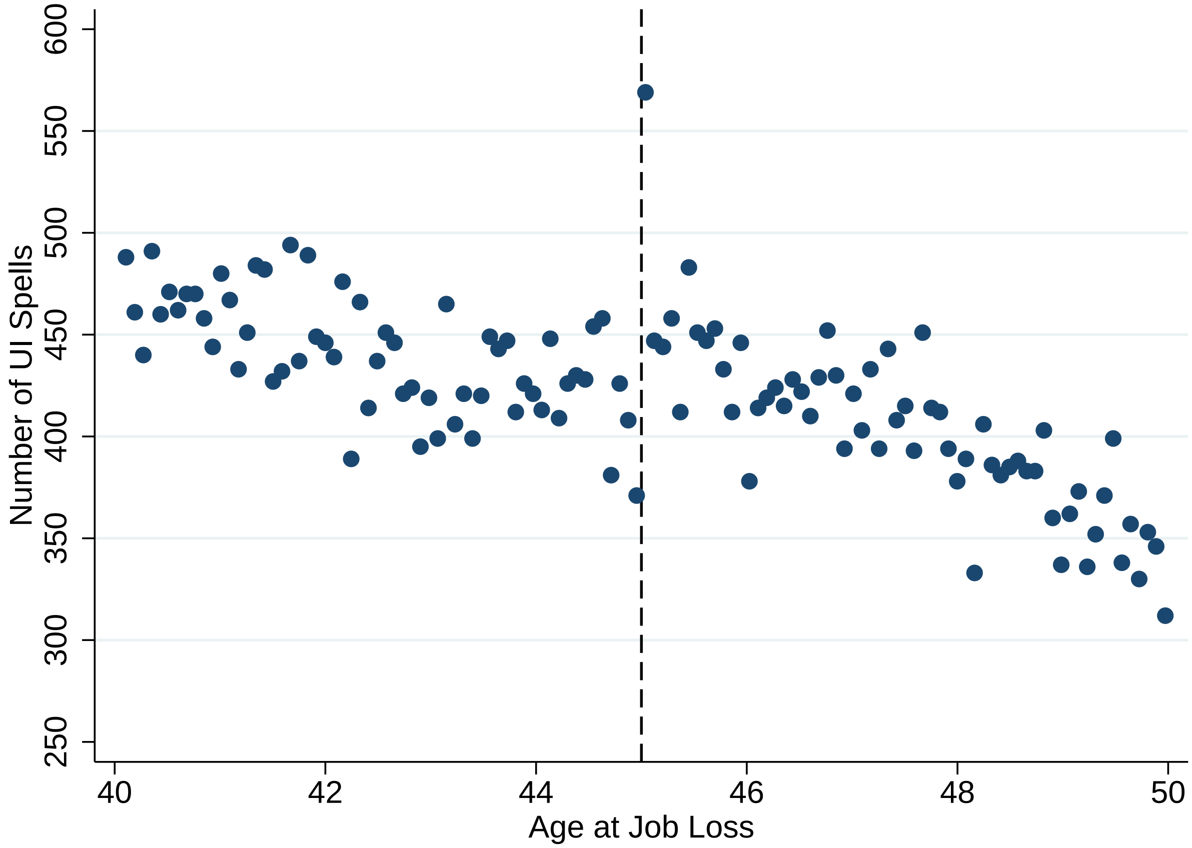
Notes: This figure summarizes the evolution of Taiwan's UI. UI in Taiwan was inaugurated in Jan 1999. On May 15, 2002, the reemployment bonus program was announced. On Jan. 1, 2003, a bonus, equal to 50% of remaining benefits, began to offer for UI recipients who find jobs before exhausting benefits. The potential duration for the worker aged 45 or older has extended from 6 months to 9 months since May 1, 2009.

Figure 2: Examples of Benefit Extension in Taiwan



Notes: The figure provides two examples for Taiwan's version of benefit extension. Claimant 1 age below 45 at job loss is eligible for 6 months of benefits. If he finds a job in the end of the second month of the UI spell, he will be eligible additional two months of benefits as a bonus. Claimant 2 aged above 45 at job loss is eligible for 9 months of benefits. He also finds a job in the end of second month of the UI spell, but he can receive additional 3.5 months of benefits as a bonus after reemployment.

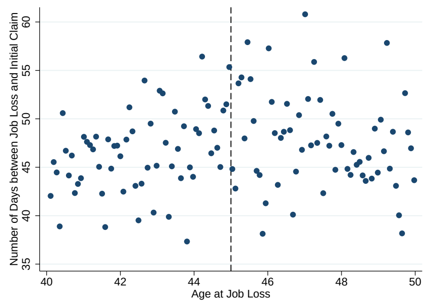
Figure 3: Validity of RDD: Density Test



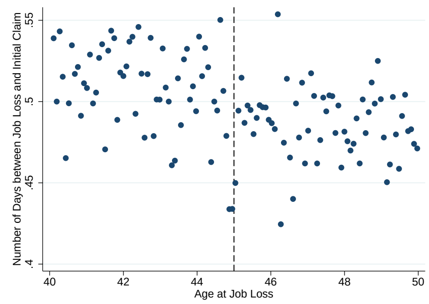
Notes: This figure plots the number of workers starting UI spells between May 1, 2009 and Jan. 1, 2012, conditional on age at job loss. Each bin corresponds to the total number of workers starting UI spells within a 30 days interval.

Figure 4: RD: Smoothness of Covariates

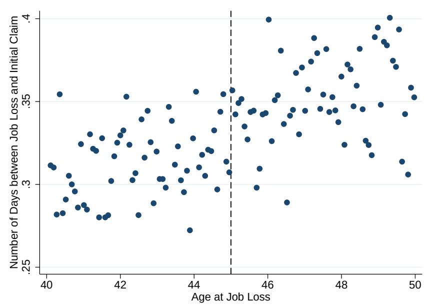
(a) Number of Days Between Job Loss and Initial Claim



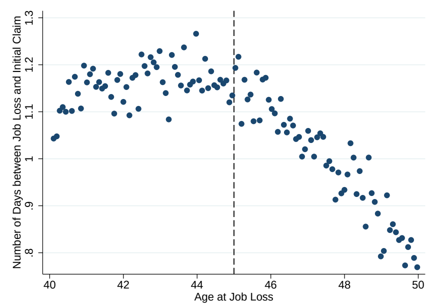
(b) Female



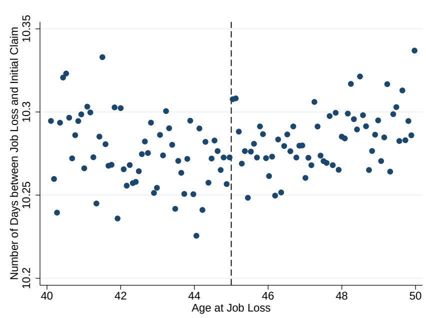
(c) Worked in Manufacturing Sector (Last Job)



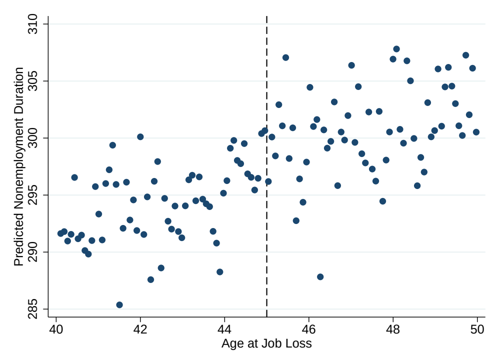
(d) Number of Dependents



(e) Log Average Monthly Wage Prior to Layoff

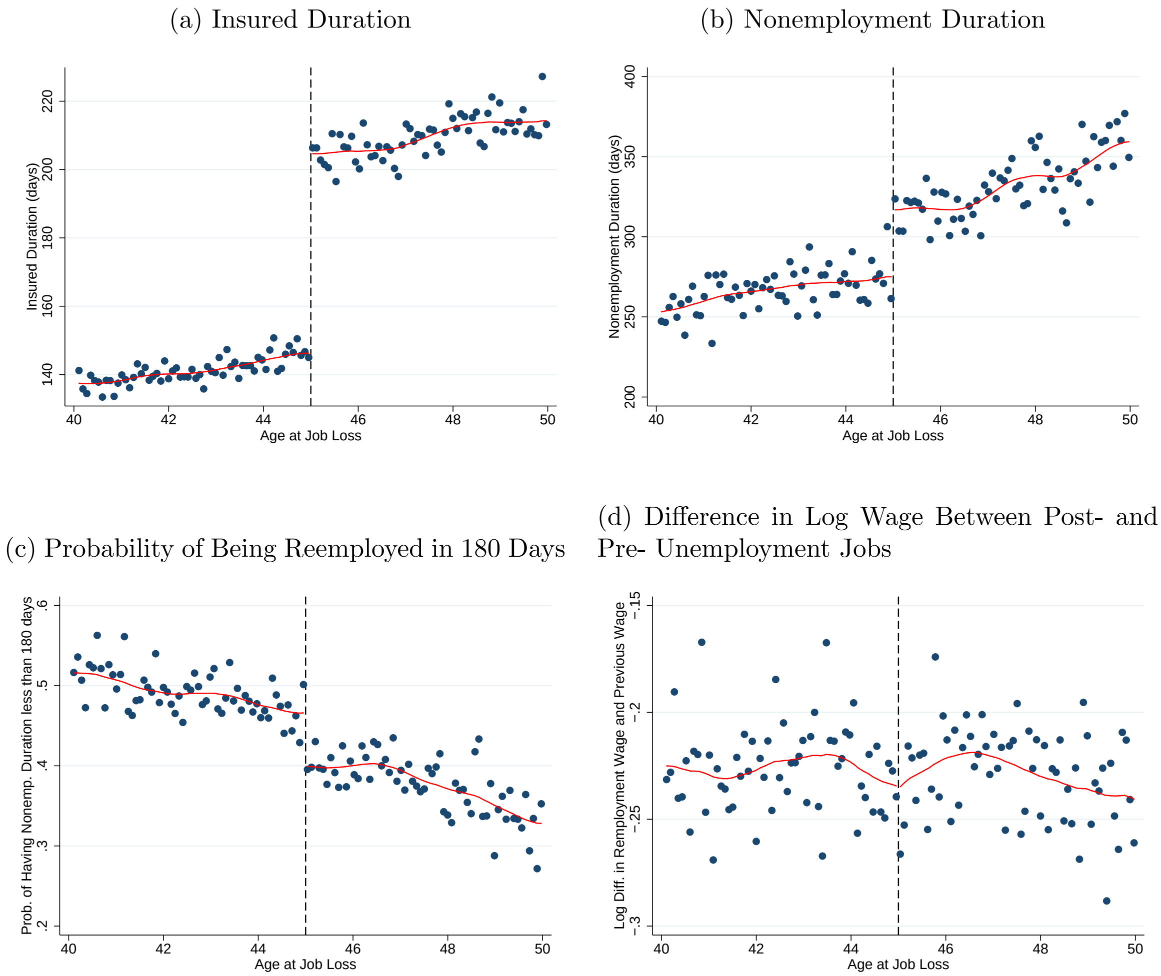


(f) Predicted Nonemployment Duration



Notes: This figure plots the averages workers' characteristics for those starting UI spells between May 1, 2009 and Jan. 1, 2012, conditional on age at job loss. The predictors for average reemployment hazard before exiting UI are previous wage, squared previous wage, previous industry, gender, place of birth, number of dependants, month/year at job loss and the number of days between job loss and claiming benefits. Each bin represents the average number of UI recipients within 20 days interval.

Figure 5: Effects of Extended UI Benefits



(a) Insured Duration

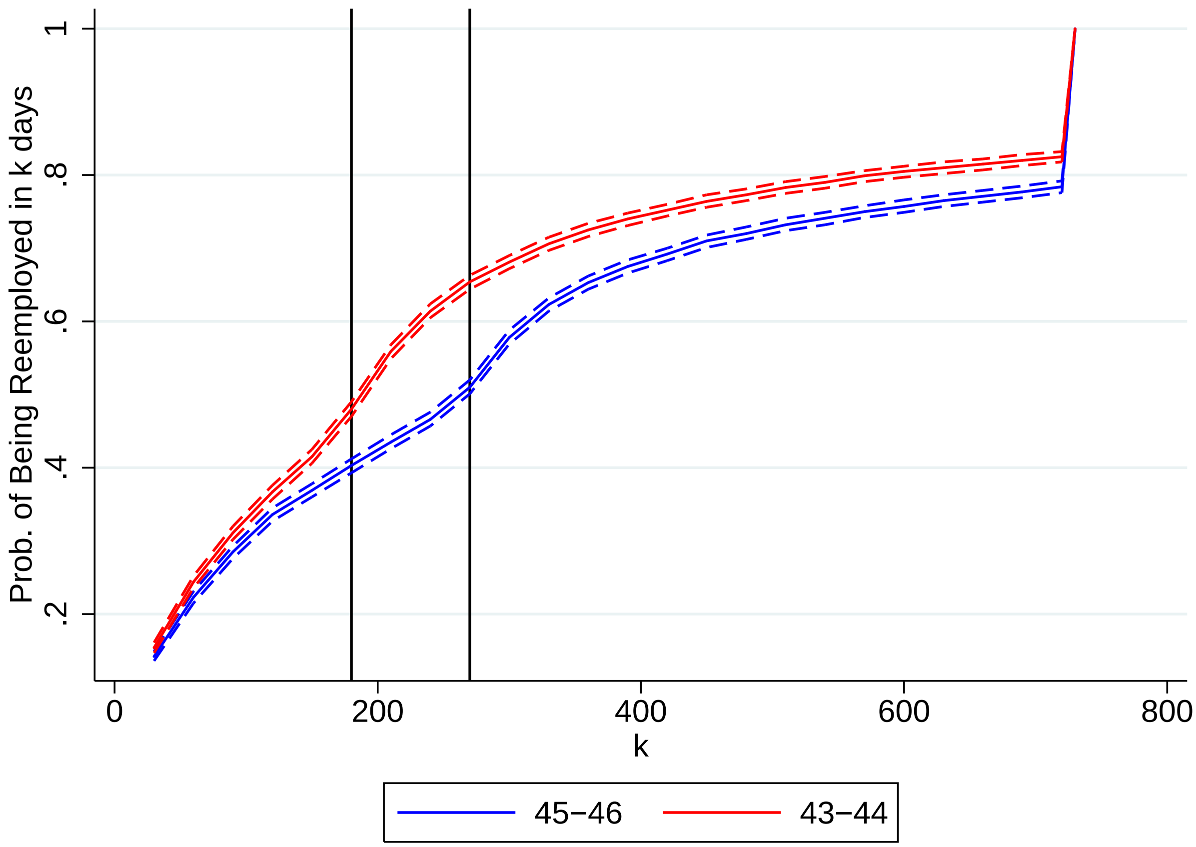
(b) Nonemployment Duration

(c) Probability of Being Reemployed in 180 Days

(d) Difference in Log Wage Between Post- and Pre- Unemployment Jobs

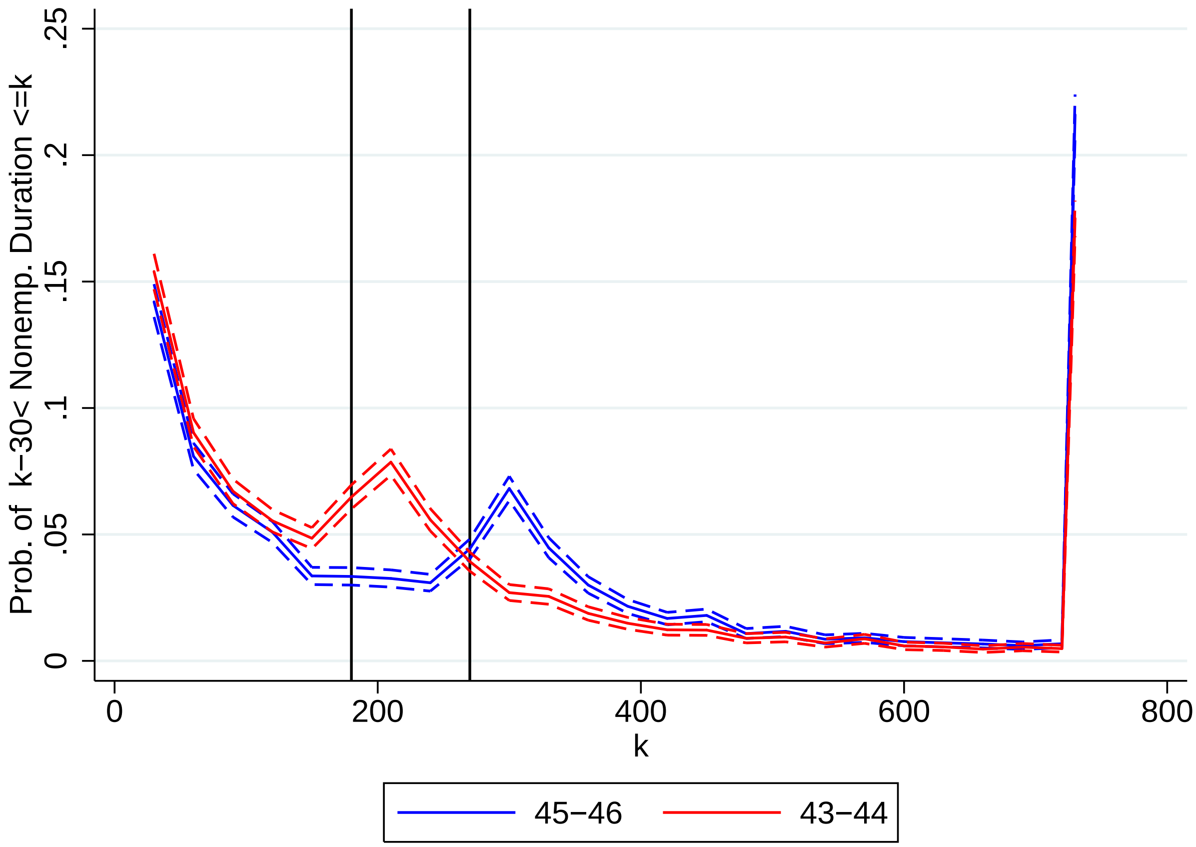
Notes: This figure plots the average outcomes for UI recipients aged 40 to 50 at job loss and starting UI spells between May 1, 2009 and Jan. 1, 2012, conditional on age at job loss. Each bin represents the average number of UI recipients within 30 days interval. The solid lines are fitted values from a local linear regression on either side of the cutoff using edge kernel, with a bandwidth of one year.

Figure 6: Cumulative Distribution of Nonemployment Duration



Notes: This figure plots the probability of being reemployed in k days for workers aged 45-46 and 43-44 at job loss, respectively. We plot the probability from $k = 30$, $k = 60$, .., to $k = 720$, and $k = 730$. The dash lines are 95% confidence interval. The solid vertical lines indicate 180th day and 270th day of the nonemployment spell.

Figure 7: Density Distribution of Nonemployment Duration



Notes: This figure plots $P(k - 30 < \text{nonemployment duration} \leq k)$ for workers aged 45-46 and 43-44 at job loss, respectively. We plot the probability from $k = 30$, $k = 60$, .., to $k = 720$, and $k = 730$. The dash lines are 95% confidence interval. The solid vertical lines indicate 180th day and 270th day of the nonemployment spell.

11 Appendix

11.1 Decomposition of the Effect of Extended Benefits

Consider a discrete time search model based on Lentz and Tranaes (2005); Chetty (2008) and Landais (2015). An unemployed worker becomes unemployed at time 0 and holds an initial asset A_0. She lives for T periods and determines the probability of finding a job in period t by varying search intensity, s_t, at a cost of $g(s_t)$, which is strictly increasing and convex. If she is unemployed at time t, she receives an unemployment benefit, b_t, with a potential duration, P, that is

$$b_t = \begin{array}{ll} b, & \text{if } 0 \leq t < P \\ 0, & \text{if } t \geq P \end{array}$$

If she is employed at time t, she earns a wage rate w_t, pays a tax rate, , and keeps the job forever. Moreover, if she is reemployed before running out of benefits, she receives a reemployment bonus, r_t, equal to θ percent of remaining benefits; otherwise, $r_t = 0$. Formally,

$$r_t = \theta \cdot \sum_{k=t}^{P-1} b_k, 0 < \theta < 1$$

The worker's consumption at time t equals the difference in income and saving. The income depends on her employment status, while the change in asset, $A_{t+1} - A_t$ reflects her saving. When employed, she earns wage rate, w_t, bonus, r_t, and pays a tax, τ. The flow utility when employed at time t equals $u(c_t^e) = u(A_t - A_{t+1} + w_t + r_t - \tau)$, where c_t^e indicates the consumption when employed at time t. Assuming the interest rate and the time discount rate are zero, the value of being employed in period t is

$$V_t = \max_{A_{t+1}} u(A_t - A_{t+1} + w_t + r_t - \tau) + \beta V_{t+1}(A_{t+1})$$

If an unemployed worker cannot find a job at time t, her flow utility is equal to $u(c_t^u) = u(A_t - A_{t+1} + b_t)$. The value of being unemployed in period t is

$$U_t = \max_{A_{t+1}} u(A_t - A_{t+1} + b_t) + \beta J_{t+1}(A_{t+1}),$$

where $J_{t+1}(A_{t+1})$ is the value of entering period $t+1$ unemployed with asset A_{t+1}. The value in the beginning of period t without a job is

$$J_t(A_t) = \max_{s_t} s_t V_t(A_t) + (1 - s_t) U_t(A_t) - g(s_t),$$

In the model, since workers face no uncertainty after they are employed, the marginal utility of consumption when workers are employed at time t equals the marginal utility of consumption at time $t+1$ when employed if the liquidity constraint does not bind. Otherwise, workers set consumption at time t equal to after tax wage rate. Formally, we can write the intertemporal first order condition when employed as follows:

$$u'(c_t^e) = \begin{array}{c} \beta u'(c_{t+1}^e);\ \text{if } A_t > L \\ u'(w-\tau);\ \text{if } A_t = L \end{array}$$

Similarly, if workers are unemployed at time t, they smooth consumption such that the marginal utility of consumption when unemployed at time t equals the expected marginal utility of consumption at time $t+1$. That is, the intertemporal first order condition when unemployed is

$$u'(c_t^u) = \begin{array}{c} \beta[s_{t+1} u'(c_{t+1}^e) + (1 - s_{t+1}) u'(c_{t+1}^u)]; \text{if } A_t > L \\ u'(b_t);\ \text{if } A_t = L \end{array}$$

If liquidity constraint is not binding yet at exhaustion point, $P-1$,

$$u'(c_t^e) = \beta_{P-t} u'(c_P^e);$$

$$u'(c_t^u) = \beta^{P-t}[1 - S_{t+1}(P)] u'(c_P^e) + S_{t+1}(P) u'(c_P^u).$$

The intratemporal first order condition balances the marginal cost of search and the difference between the value of being employed and unemployed at time t.

$$g'(s_t) = V_t(A_t) - U_t(A_t)$$

The effect of one dollar increase in unemployment benefits in period P on search intensity in period t is dependent on the effect on the value of employment in period t and the value of unemployment in period t, respectively.

$$g''(s_t)\frac{\partial s_t}{\partial b_P} = \frac{\partial V_t(A_t)}{\partial b_P} - \frac{\partial U_t(A_t)}{\partial b_P}$$

For $t \leq P$, one dollar increase in b_P raises r_t by θ dollar, and increases the value of employment in period t by $\theta u'(c_t^e)$.

$$\frac{\partial V_t(A_t)}{\partial b_P} = \theta u'(c_t^e)$$

One dollar increase in b_P increases the value of unemployment in period t through two channels. On one hand, it increases the value of unemployment in period t because it increases the utility of being unemployed in period P. On the other hand, it also increases the utility of finding a job in any period before exhaustion point.

$$\begin{aligned}
\frac{\partial U_t(A_t)}{\partial b_P} &= \beta^{P-t}(1-s_{t+1})..(1-s_P)u'(c_P^u) + s_{t+1}\beta\theta u'(c_{t+1}^e) + .. + (1-s_{t+1})..s_P\beta_{P-t}\theta u'(c_P^e) \\
&= \beta^{P-t}S_{t+1}(P)u'(c_P^u) + [1-S_{t+1}(P)]\theta u'(c_t^e)
\end{aligned}$$

Hence, we can write the effect of one dollar increase in unemployment benefits at time P on search effort at time t as below

$$\begin{aligned}
\frac{\partial s_t}{\partial b_P} &= \frac{\theta u'(c_t^e) - \{\beta^{P-t}S_{t+1}(P)u'(c_P^u) + \theta u'(c_t^e)[1-S_{t+1}(P)]\}}{g''(s_t)} \\
&= \frac{-\beta^{P-t}S_{t+1}(P)u'(c_P^u) + \theta u'(c_t^e)S_{t+1}(P)}{g''(s_t)}; \forall t \leq P
\end{aligned}$$

The liquidity effect and the moral hazard effect of one dollar increase in unemployment benefits at time P on search effort at time t is captured by $\frac{\partial s_t}{\partial A_t}$ and $\frac{\partial s_t}{\partial w_t}$, respectively.

$$\frac{\partial s_t}{\partial A_t} = \frac{u'(c_t^e) - u'(c_t^u)}{g''(s_t)};$$

$$\frac{\partial s_t}{\partial w_t} = \frac{u'(c_t^e)}{g''(s_t)}.$$

Using intertemporal first order conditions and assuming liquidity constraint is not yet binding at time $P-1$, we decompose the effect of an increases in b_P on search intensity into a liquidity and a moral hazard effect.

$$\begin{aligned}
\frac{\partial s_t}{\partial b_P} &= \frac{[1-S_{t+1}(P)]u'(c_t^e)-u'(c_t^u)+\theta S_{t+1}(P)u'(c_t^e)}{g''(s_t)} \\
&= \frac{u'(c_t^e)-u'(c_t^u)-(1-\theta)S_{t+1}(P)u'(c_t^e)}{g''(s_t)} \\
&= \frac{\partial s_t}{\partial A_t}-(1-\theta)S_{t+1}(P)\frac{\partial s_t}{\partial w_t}; \forall t \leq P
\end{aligned}$$

11.2 Reservation Wage Choice

To include reservation wage choice into the model, we assume the unemployed worker samples one i.i.d. offer each period from a known wage distribution $F(w)$. The value of entering period t without a job is

$$J_t(A_t) = \max_{s_t, R_t} s_t P(w \leq R_t) V_t + [1 - s_t P(w \leq R_t)] U_t$$

The optimal reservation wage is the wage offer such that the worker is indifferent between accepting or rejecting the offer, that is

$$U_t = V_t(R_t).$$

To derive the derivative of the reservation wage to one dollar increase in unemployment benefits at time P, using

$$\frac{\partial U_t}{\partial b_P} = S_{t+1}(P) u'(c_P^u) + [1 - S_{t+1}(P)] u'(c_P^e),$$

and

$$\frac{\partial V_t}{\partial b_P} = \theta u'(c_t^e) + \frac{\partial R_t}{\partial b_P}(T - t),$$

we can write

$$\begin{aligned} \frac{\partial R_t}{\partial b_P} &= \frac{S_{t+1}(P) u'(c_P^u) + [1 - S_{t+1}(P)] u'(c_P^e) - \theta u'(c_P^e)}{T - t} \\ &= S_{t+1}(P) \frac{u'(c_P^u) - u'(c_P^e)}{T - t} + (1 - \theta) \frac{u'(c_P^e)}{T - t}. \end{aligned}$$

To decompose the effect of extended benefits on reservation wage, we use

$$\frac{\partial U_t}{\partial A_t} = u'(c_t^u)$$

$$\frac{\partial V_t}{\partial A_t} = u'(c_t^e) + \frac{\partial R_t}{\partial A_t}(T - t)$$

$$\frac{\partial R_t}{\partial A_t} = \frac{u'(c_t^u) - u'(c_t^e)}{T - t} > 0$$

and

$$\frac{\partial U_t}{\partial r_t} = 0$$

$$\frac{\partial V_t}{\partial r_t} = u'(c_t^e) + \frac{\partial R_t}{\partial r_t}(T-t)$$

$$\frac{\partial R_t}{\partial r_t} = -\frac{u'(c_t^e)}{T-t} < 0$$

Therefore, we get the following decomposition formula:

$$\begin{aligned}
\frac{\partial R_t}{\partial b_P} &= S_{t+1}(P)\frac{u'(c_P^u) - u'(c_P^e)}{T-t} + (1-\theta)\frac{u'(c_P^e)}{T-t} \\
&= S_{t+1}(P)\frac{\partial R_t}{\partial A_P} - (1-\theta)\frac{\partial R_t}{\partial r_P}.
\end{aligned}$$

11.3 Tables

Table 8: Estimates of the Effects of Extended Benefits – Excluding Observations Within k Days of the Cutoff

	Insured duration (1)	Nonemp. duration (2)	Log Reemp. Wage (3)
$k = 30$	55.83***	32.07***	0.004
	(1.60)	(5.21)	(0.009)
Sample Size	42,591	49,329	36,360
$k = 60$	55.51***	36.69***	0.008
	(1.72)	(5.57)	(0.009)
Sample Size	40,444	46,199	40,616
$k = 90$	55.27***	38.43***	0.008
	(1.81)	(6.01)	(0.009)
Sample Size	39,947	42,416	39,859
$k = 120$	56.00***	36.69***	0.006
	(1.82)	(6.16)	(0.009)
Sample Size	42,867	43,203	41,519
$k = 150$	56.68***	36.20***	0.006
	(1.89)	(6.28)	(0.009)
Sample Size	42,822	43,131	41,771
$k = 180$	56.26***	33.24***	0.006
	(1.90)	(5.68)	(0.010)
Sample Size	42,524	51,829	40,678
Poly. model	linear	linear	linear
Bandwidth (days)	CCT	CCT	CCT

Note: This table shows the estimates of the effect of increasing potential duration from 6 months to 9 months on insured duration, nonemployment duration and the difference in log wage between employment. We estimate a local linear regression using the optimal bandwidth by Calonico et al. (2014) and triangular kernel. The sample are workers starting UI spells between May 1, 2009 and Jan. 1, 2012, and aged within the bandwidth excluding those aged within k days of the cutoff. Standard errors in parentheses are clustered by age in days.

Table 9: Placebo Test for RD Design

	(1)	(2)	(3)	(4)	(5)
Panel A: Insured Duration					
β_{EB}	1.49	1.34	2.09	1.48	1.19
	(1.88)	(1.87)	(1.68)	(1.93)	(2.07)
Sample size	14,023	14,023	18,531	18,531	26,690
Panel B: Nonemployment Duration					
β_{EB}	2.28	2.18	4.50	2.63	0.08
	(9.18)	(9.15)	(7.85)	(9.30)	(9.40)
Sample size	14,023	14,023	18,487	18,487	32,081
Panel C: Nonemployment Duration < 180 days					
β_{EB}	-0.01	-0.01	-0.01	-0.01	-0.01
	(0.02)	(0.02)	(0.02)	(0.01)	(0.02)
Sample size	14,023	14,023	28,935	28,935	27,002
Bias-corrected	–	–	–	Yes	Yes
Covariates	–	Yes	–	–	–
Poly. model	linear	linear	linear	linear	quadratic
Bandwidth (days)	730	730	CCT	CCT	CCT

Note: This table conducts a placebo test using sample before UI extension. Column 1 estimate a linear regression on either side of the cutoff using sample from workers age 43-46 at job loss, and starting UI spells before Nov. 1, 2008. Column 2 includes the following covariates: previous wage, squared previous wage, previous industry, gender, place of birth, number of dependants, month/year at job loss, number of job loss and number of days between job loss and initial claim. Columns 3 reports the estimates using optimal bandwidth algorithm from Calonico et al. (2014). The optimal bandwidths vary with the outcome variables, in the range of 4 to 6 years. The bias correction estimates and the corresponding robust standard errors are presented in the Column 4. In Column 5, we report the bias correction estimates and robust standard error using local quadratic regression. Standard errors in parentheses are all clustered by age in days. Column 1 and 2 use rectangular kernel. Columns 3-5 use triangular kernel. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table 10: The Effect of Extended Benefits on Particpation in Vocational Training

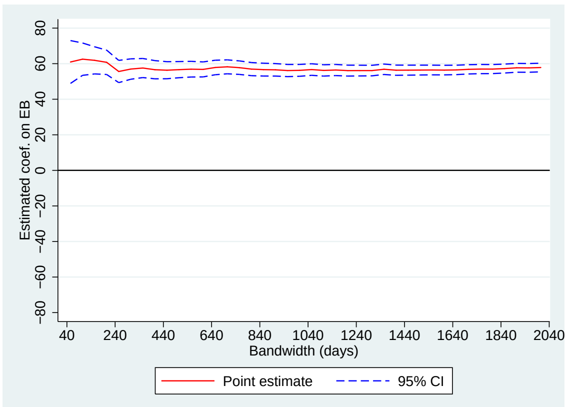
	(1) Participation in Training	(2) # of Days Receiving Training Subsidies
β_{EB}	-0.003	-0.105
	(0.005)	(0.668)
Baseline Mean	0.080	8.282
Sample size	39,996	40,352
Poly. model	linear	linear
Bandwidth (days)	CCT	CCT

Note: This table examines whether an increase in potential duration form 6 to 9 months on the participation in vocational training and duration of training. We estimate a local linear regression using the optimal bandwidth by Calonico et al. (2014) and triangular kernel. The sample are workers aged within the bandwidth and starting UI spells between May 1, 2009 and Jan. 1, 2012. Standard errors in parentheses are clustered by age in days. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

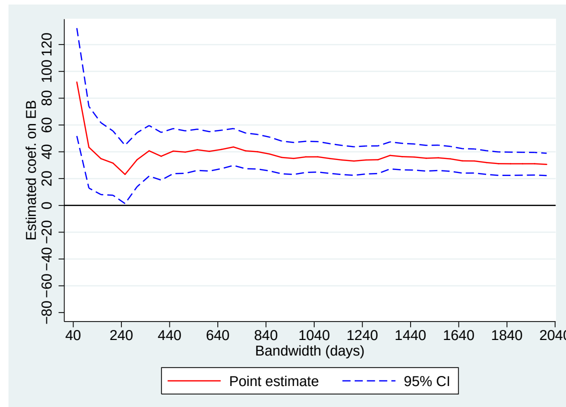
11.4 Figures

Figure 8: RD Estimates with Varying bandwidths

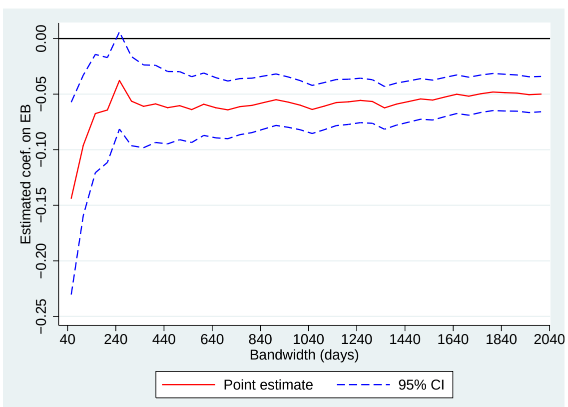
(a) Insured Duration



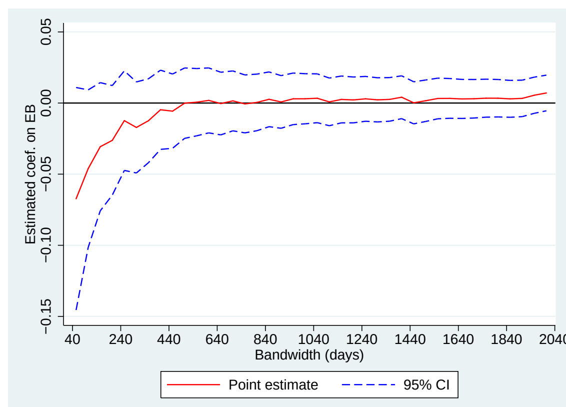
(b) Nonemployment Duration



(c) Probability of Being Reemployed in 180 Days



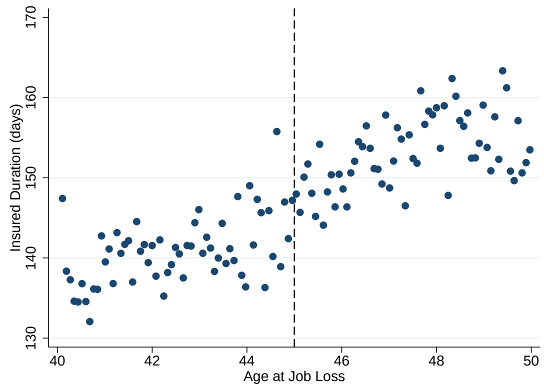
(d) Difference in Log Reemployment Wage and Log Previous Wage



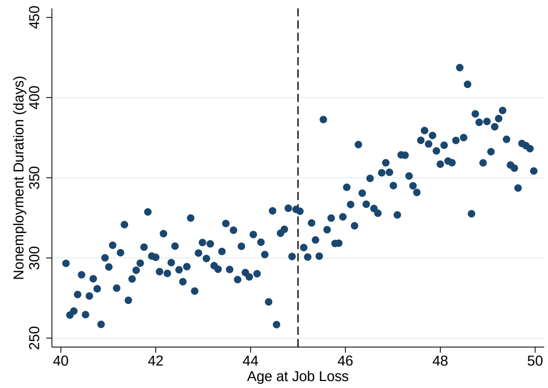
Notes: This figure test for the sensitivity to bandwidth choice for our RD estimates. We estimate a local linear regression using a bandwidth ranging from 40 to 2000 days. The solid line indicates the point estimates, and the dash lines are corresponding confidence intervals.

Figure 9: Test RDD Assumption: Placebo Test

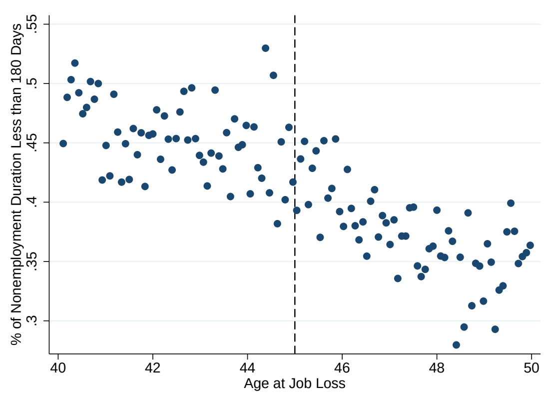
(a) Insured Duration Prior to UI Extension



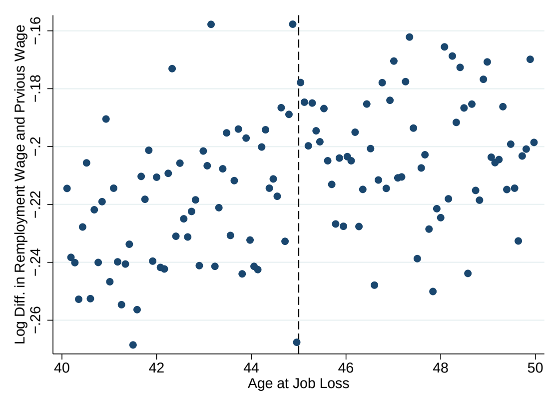
(b) Nonemployment Duration Prior to UI Extension



(c) Probability of Being Reemployed in 180 Days



(d) Difference in Log Reemployment Wage and Log Previous Wage



Notes: This Figure plots the average outcomes for UI recipients age 40 to 50 at job loss and start UI spells before Nov. 1, 2008, conditional on age at job loss. Each bin represents the average number of UI recipients within 30 days interval.