Evaluating Welfare Effect of Extended Unemployment Insurance Benefits: Evidence from Two Natural Experiments

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Abstract

This paper evaluates the welfare effects of extended unemployment insurance (UI) benefits using two policy changes in Taiwan: the introduction of a reemployment bonus program and a benefit extension for workers aged at least 45. The reemployment bonus counteracts the moral hazard effect of unemployment insurance without providing additional liquidity during unemployment. The benefits extension, however, increases workers' ability to smooth consumption when unemployed at the cost of distortion to search. Using the variation in the bonus offer around the time the bonus was introduced and the age discontinuity in the eligibility for extended benefits, we separately identify the moral hazard effect and the liquidity effect of UI benefits extension. Our result suggests liquidity to moral hazard ratio during UI exhaustion period is around 3.5 and it is welfare enhancing to increase the potential benefit duration.

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1 Introduction

Unemployment insurance (UI) protects individuals against the risk of earnings loss during unemployment, but it also distorts incentives to search for jobs. Chetty (2008) addresses this by showing that more generous UI decreases workers' search effort through two distinct channels: a moral hazard effect and a liquidity effect. On the one hand, UI benefits increase the opportunity cost of being employed and distorts workers' incentive to find a job (i.e. moral hazard effect). On the other hand, UI benefit increase workers' ability to smooth consumption during unemployment and allows them more time to search for a job (i.e. liquidity effect). Empirically, however, only a few papers (Chetty, 2008; Card et al., 2007; Landais, 2015) distinguish the liquidity effect from the moral hazard effect of UI because the variation in unemployment benefits confounds these two sources of variation.

In this paper, we decompose the welfare effect of extended UI benefits into a moral hazard effect and a liquidity effect by exploiting two natural experiments in Taiwan. The first natural experiment is the 2009 UI extension for older workers. Starting in 2009, workers aged 45 or more became eligible for 9 months of UI benefits instead of the 6 months for those under age 45. We use a regression discontinuity (RD) design to examine the effect of extended UI benefits on reemployment hazard by comparing the job finding rate of individuals just before and just after the 45 years-old at layoff. Using a search model with borrowing constraints, we demonstrate that the effect of extended UI benefits on job finding rate is a combination of a liquidity effect and a moral hazard effect.

Thus, we use the second natural experiment – the introduction of a reemployment bonus program in January 2003 – to estimate the moral hazard effect. Since 2003, Taiwan government introduced reemployment bonuses, in which people would be paid 50% of their remaining UI benefits after they were reemployed. We can use the estimated effect of a reemployment bonus to identify the moral hazard effect of extended benefits because the bonus only increases the opportunity cost of being unemployed but not increase workers' income during unemployment. In other words, we can recover the liquidity effect as long as we are able to provide credible estimates of the labor supply responses to the UI extension and reemployment bonus. The reemployment bonus program reached back to UI recipients who were receiving benefits when the program took effect in 2003. Therefore, it results in two kinks in the bonus offer that workers are eligible for as a function of the date UI spells started. We use the regression kink (RK) design to examine the effect of reemployment bonus on reemployment hazard of middle-age UI recipients. Plugging in the RD estimate (i.e. total effect) and the RK estimate (i.e. moral hazard effect) in the search model, we can recover the liquidity effect of extended UI benefits and carry out welfare analysis.

We obtain three key findings. Firstly, our estimates using the RD design show that a three-month increase in potential benefit duration reduced monthly reemployment hazard by 7.2 percentage points for middle-aged between 7th and 9th month of nonemployment. In addition, we find that the extension of UI benefits also affect workers' search effort before UI exhaustion—a three-month extension of UI reduces the monthly reemployment hazard in the first six months of nonemployment by around 1.4 percentage points. Secondly, our estimates using the RK design show that being eligible for three months of UI benefits as reemployment bonus increase the monthly reemployment hazard by about 1.8 percentage points in the first six months of nonemployment for UI recipients aged between 40 and 50 at job loss.

Lastly, we conduct a welfare analysis of UI extension by plugging the above reduced-form estimates into the welfare decomposition formula. Following Baily (1978) and Chetty (2008), we derive the welfare effects of the benefits extension—the optimal potential duration depends on the trade-off of expected consumption smoothing benefits at the exhaustion point and the utility loss when employed due to the increased tax payment. The expected consumption smoothing benefits are identified by the product of exhaustion rate and the ratio of liquidity effect to moral hazard effect at UI exhaustion, while the welfare cost is identified by the behavioral responses to UI extension. Our estimates suggest that for middle-aged UI recipients, the liquidity to moral hazard ratio is around 3.5 at UI exhaustion. The estimated consumption smoothing benefits of extended benefits are larger than the cost of extended benefits, suggesting it is welfare enhancing to increase potential duration. In addition, consistent with Ganong and Noel (2017), the estimated liquidity to moral hazard ratio is about 0.5 before UI exhaustion, smaller than the ratio at UI exhaustion. The larger consumption smoothing benefits at the exhaustion point is also consistent with Kolsrud et al. (2018)'s findings that consumption smoothing benefits are larger later in the unemployment spell.

Our paper stands apart from previous literature on unemployment insurance in the following ways. First, we provide one of the first evidence on the consumption smoothing benefits of UI benefits extension.[1] There have been many convincing studies estimating the effects of UI benefits on unemployment duration. But the empirical evidence on consumption smoothing benefits of UI is still scarce and most of existing studies focus on the welfare effect of UI benefit level (Gruber, 1998; Chetty, 2008; Card et al., 2007; Landais, 2015).[2] In general, these studies found that the estimates of liquidity to moral hazard ratio are around 0.88 to 1.5. However, the welfare effect of UI benefit extension could be different from one of UI benefit level because the extended UI benefits mainly affect UI recipients who run out of their benefits rather than every UI recipients. An exception is Ganong and Noel (2017) who used bank account data to investigate the consumer spending during the course of unemployment. Ganong and Noel (2017)'s estimates suggest the consumption smoothing benefits from a benefits extension is at least three times larger than that from from an increased replacement rate.[3] This paper complements Ganong and Noel (2017)'s estimates using sufficient statistic approach and labor market data.

Secondly, our paper provides the first evaluation of optimal unemployment insurance in Asian countries. Most existing studies focus on unemployment insurance in U.S. or European countries. The welfare effect of unemployment insurance in Asian countries could be distinct from existing

[1] As Schmieder et al. (2016) pointed out, although there is some evidence on the consumption smoothing benefits of increasing benefit replacement rate, empirical evidence on the consumption smoothing benefits of extended UI benefits is rare.

[2] Gruber (1998) used panel data from Panel Study of Income Dynamics and state variation in UI replacement rate. Gruber (1998)'s estimates suggests a 10% increase in the replacement rate reduce the consumption drop during unemployment by 2.8%. However, Gruber (1998)'s estimates are imprecise and the estimated consumption smoothing benefits are sensitive to risk aversion coefficient.Chetty (2008) and Landais (2015) circumvent issues with estimation of risk aversion coefficient using the sufficient statistic approach—the consumption smoothing benefits equals the ratio of the liquidity effect to the moral hazard effect of UI. Chetty (2008) estimates the liquidity effect by estimating the behavioral response to severance pay. Landais (2015) uses the difference in the behavioral response to extended benefits and an increases in benefit level to identify the moral hazard, which identifies the liquidity effect indirectly. Their estimates suggest the liquidity effects explains about half of the effect of UI on unemployment duration can be attributed to the liquidity effect, suggesting significant consumption smoothing benefits of UI benefit level.

[3] Landais (2015) estimates the welfare effects of increasing benefit level and potential duration by assuming the consumption smoothing gain of these two policies are the same.

evidence since Asian countries, such as China, Japan, Korea or Taiwan, have low unemployment rate, high saving rate and close family tie. Therefore, unemployed workers in these countries could have more financial resource from their own saving or family transfer so that might be less likely to be liquidity constrained. Based on these facts, we expect our estimates of consumption smoothing benefits of UI extension could serve a lower bound for U.S. or European countries.

Finally, this paper adds new evidence on the impact of reemployment bonuses on job search. Existing evidence on the effect of reemployment bonuses still relies on the U.S. field experiments conducted in 1980s (Woodbury and Spiegelman, 1987; Decker et al., 2001; Meyer, 1995). Evidence on the effectiveness of the reemployment bonuses from these studies are mixed and still inconclusive. Ahn (2018) provides more recent evidence on this issue using a quasi-experimental method and administrative data in Korea. He finds that increases in the reemployment bonus can significantly reduce the duration of UI spells by 0.16 to 0.42 months without affecting subsequent job match quality. We contribute to this stream of literature by offering new quasi-experimental estimates. More importantly, we reinterpret the labor supply effect of bonuses as a countervailing force to the moral hazard effect of UI and use it to recover liquidity effect for UI exhaustee.

The outline of this paper is as follows. In Section 2, we describe the Taiwanese UI system. Section 4 describes our data and estimated sample. In Section 3, we present our theoretical framework for welfare analysis. In Section 5 and Section 6, we estimate the effects of extended benefits and the effects of reemployment bonus on reemployment hazard. In Section 7, we plug the reduced-form estimates into the model to obtain the welfare effects of extended benefits. Section 8 summarizes the findings and discuss possible extension to this paper.

2 Unemployment Insurance in Taiwan

Unemployment benefits in Taiwan form one part of the overall employment insurance program, which is a mandatory national program that offers unemployment benefits, reemployment bonuses, vocational training living allowances, parental leave allowances and national health insurance 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, individuals aged 15 to 65 who lose their jobs must have at least one year of employment history in the three years prior to the job loss.[4] To receive the first month's benefits, a claimant must register with the government employment service and complete a 14-day waiting period. If the worker does not find a job by the end of the waiting period, the insured period begins (up to the maximum duration of benefits the claimant is entitled to). Since 2009, the maximum duration of benefits has been 6 months for workers aged below 45 at the time of job loss, and 9 months for those aged 45 or older when they lost their job.[5]

Unlike in the United States, where benefits are paid 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 a worker is reemployed before the end of a given 30-day interval, the amount of benefits paid in that month is prorated. The monthly UI benefits replace 60% of the average insured wage during the six months prior to job loss[6] for those without non-working dependents. For UI recipients with non-working dependents the replacement rate is increased, and can reach as high as 80% depending on the number of dependents.

Workers are required to actively search for a job while receiving benefits. Specifically, they have to list at least two job contacts for each continued claim. In general, this work search test plays the role of the stick, promoting rapid employment via undesirable consequences. The other strategy is the carrot: Taiwan's UI program offers a generous financial incentive to workers who return to work quickly. This incentive, which takes the form of a reemployment bonus, offers 50% of any

[4] Only workers losing their jobs involuntarily or due to the ending of a fixed term contract are eligible. According to Employment Insurance Act and Labor Standard Act, involuntary separation from employment refers to separation from employment because the insured unit has closed down, relocated, suspended business, dissolved, filed bankruptcy, or business cycle induced layoff and downsizing. Employment history is the number of days for which a worker has 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 included when determining eligibility.

[5] There is only one exception: UI recipients who hold disability cards are eligible for nine months of benefits regardless of their age at the time of job loss. However, few UI recipients are disability card holders; our data showed that only 0.8% of workers younger than 45 received unemployment benefits for longer than six months during our study period.

[6] This refers to the last six months for which a worker was enrolled in employment insurance prior to their job loss.

remaining unemployment benefits to UI recipients who find jobs before the end of their eligibility period, and who then accumulate at least three months of employment history after reemployment. The three months of reemployment does not have to be continuous, or with a single employer. A person who worked for multiple employers for three months after reemployment would also qualify for the bonus.[7]

3 Theoretical Framework

In this Section, we use a simple search model to show that the effect of extended benefits on the search effort can be decomposed into a liquidity effect and a moral hazard effect, which can be identified using the effect of the reemployment bonus. Then, we show the labor supply response to extended benefits and that to the bonus are sufficient to identify the welfare effect of UI extension. Here, we focus on the intuition and leave the derivations to the Appendix A.

3.1 Moral Hazard versus Liquidity Effects of Extended Benefits

The discrete-time search model with borrowing constraints we use in this paper comes from Chetty (2008). Since we focus on the effects of extended benefits rather than the effects of increasing benefit level analyzed in Chetty (2008), we also consider Landais (2015)'s extension to Chetty (2008)'s model when deriving the effects of extended benefits.

The only difference between our model and the models in Chetty (2008) and Landais (2015) is that we incorporate Taiwan's reemployment bonus into the model. A worker reemployed before running out of benefits receives a reemployment bonus, r_t, equal to θ percent of the remaining benefits; otherwise, $r_t = 0$. Formally,

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

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

$$\frac{\partial s_t}{\partial P} = b\frac{\partial s_t}{\partial A_t} - b(1-\theta)S_{t+1}(P)\frac{\partial s_t}{\partial w_t}; \forall t \leq P \qquad (1)$$

[7] The three months reemployment period does not include recalls (the work experience in the firm prior to layoff).

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 $t + 1$. The effect of extended benefits on search intensity is a combination of the liquidity effect ($\frac{\partial s_t}{\partial A_t}$) and the moral hazard effect ($-\frac{\partial s_t}{\partial w_t}$). On the one hand, the liquidity effect of UI increases workers' ability to smooth consumption during unemployment, allowing them more time to search for a job, so the liquidity effect is welfare increasing. On the other hand, the moral hazard effect is distortionary because it decreases workers' net wages, which decreases the incentive to search. Therefore, the relative importance of the liquidity effect and the moral hazard effect identifies the consumption smoothing benefits of a benefit extension.[8]

Empirically, to separate the effect of the liquidity effect from the moral hazard effect, we have to estimate at least any two of $\frac{\partial s_t}{\partial P}$ (i.e. total effect of UI extension), $\frac{\partial s_t}{\partial w_t}$ (i.e. moral hazard effect), and $\frac{\partial s_t}{\partial A_t}$ (i.e. liquidity effect). We can identify the total effect of extended benefits ($\frac{\partial s_t}{\partial P}$) by exploiting the age discontinuity in the eligibility rule, but we have neither information on the asset amount nor exogenous variation in wage offer.

To address this, we recognize that the variation in the wage rate and that in the bonus offer affect the search intensity in the same way, so that $\frac{\partial s_t}{\partial w_t} = \frac{\partial s_t}{\partial r_t}$. Since the introduction of the bonus program provides credible exogenous variation for identifying $\frac{\partial s_t}{\partial r_t}$ (i.e. moral hazard effect), we can recover $\frac{\partial s_t}{\partial A_t}$ (i.e. liquidity effect) indirectly.

In section 7.2, we show that the consumption smoothing benefits of UI extension depends on the probability of exhausting six months of benefits and the consumption smoothing benefits at UI exhaustion. To estimate the consumption smoothing benefits at UI exhaustion, we need to estimate the total effect and moral hazard effect of extended benefits at the exhaustion point ($\frac{\partial s_P}{\partial P}$ and $\frac{\partial s_P}{\partial r_P}$). However, there is no variation in reemployment bonuses between 6th and 9th month of nonemployment spell because workers lose the eligibility for bonuses after exhausting their benefits.

[8]The formula also shows the reemployment bonus counteracts the moral hazard effect by offering θ remaining benefits for workers reemployed before exhaustion point, suggesting the benefit extension in the UI with the reemployment bonus will not increases unemployment duration as much as extending potential without the bonus. This prediction is consistent with Davidson and Woodbury (1991)'s findings that bonus reduce insured duration more for workers eligible for longer potential duration.

To circumvent this issue, we assume the hazard rate response to reemployment bonus before UI exhaustion is comparable to that at UI exhaustion. Note that we will underestimate the consumption smoothing benefits at UI exhaustion if we assume the moral hazard effect does not vary over the spell. Kolsrud et al. (2018) estimate that the duration response to unemployment benefits decrease over the spell, implying that the moral hazard effect declines over the spell. Therefore, our estimated consumption smoothing benefits are likely to be a lower bound for true consumption smoothing benefits.

4 Data and Sample

4.1 Data

We use two sources of administrative data from the Taiwanese Bureau of Labor Insurance: the unemployment benefits files and the employment insurance enrollee file dating from January 1999 to December 2013. Each entry in the unemployment benefits file represents one beneficiary case on a monthly basis (i.e. 30 days), and contains each UI recipient's date of birth, date of job loss, starting date of UI benefits, average previous insured earnings in the six months prior to layoff (hereafter, previous earnings), an individual identifier, and some demographic information, including gender, place of birth, and a four-digit code indicating the recipient's previous industry. We use a recipient's birthday and the date of job loss to precisely measure our key variable—a recipient's age at the time of a job loss.

In employment insurance enrollee file, each entry represents a change in the employment record: (1) new enrollments in employment insurance (job change/entry); (2) cancellations of employment insurance (job separation), or wage changes, and contains the date of change, an enrollee's insured wage, and an individual identifier. We use the individual identifiers to merge unemployment benefits files with employment insurance enrollee file and use the date of new enrollments in employment insurance after job loss to represent the date of reemployment.

4.2 Sample

To examine the effect of UI benefits extension, since the benefits extension took effect on May 1, 2009, we use the recipients who lost their job during May 2009 and January 2012 as post-reform sample (i.e. 2009-2012 sample) and focus on the individuals losing their jobs when they were age 43 to 47 (i.e. two years before and after age 45). The first two columns of Table 1 reports the summary statistics of selected characteristics for control group (column (1), age 43-45) and treatment group (column (2), age 45-47), respectively.

We find that treatment group has lower reemployment hazard in 7^{th} to 9^{th} month of a UI spell than control group. In addition, both groups have similar individual characteristics except that treatment group has higher share of recipients working in manufacturing sector than control group. In our main analysis, we further control other confounding factors, that might affect our outcome of interest, by including a pre-reform sample (i.e. 2006-2008 sample): the individuals who lost their job during May 2006 and November 2008. The last two columns of Table 1 reports the summary statistics of selected characteristics for pre-reform sample. We find that the difference in reemployment hazard between individuals losing their job before age 45 (column (3), age 43-45) and those losing their job after age 45 (column (4), age 45-47) is quite small before the benefits extension reform. Similar to post-reform sample, the only difference in selected characteristics is that older group also has higher share of recipients working in manufacturing sector than younger group.

To investigate the impact of reemployment bonus, we utilize the recipients who lost their job around 6 months (i.e. 180 days) before and after July 1, 2002 as reform sample (i.e. 2002-2003 sample). Since the relationship between eligibility for reemployment bonus and staring date of a UI spell change for the recipeients starting their UI spell after July 1, 2002. Thus, we define this cohort as treatment group and those starting their UI spell before July 1, 2002 as a control group. The first two columns of Table 2 suggests treatment group (column (2)) has higher reemployment hazard than control group (column (1)). For other characteristics, both groups are quite similar. Moreover, in order to control any seasonality of employment opportunities, we also include sample for placebo test (2001-2002 sample), namely, the recipients who lost their jobs between 6 months (i.e. 180 days) before and after July 1, 2001 in our main analysis (see column (3) and (4) of Table 2).

5 Effects of Extended Benefits

5.1 Regression Discontinuity Design

Our identification strategy is similar to that of other recent studies using an "age discontinuity" to identify the UI effect on labor market outcomes (Schmieder et al., 2016). To quantify the effects of UI benefits extension, we estimate the following regression:

$$y_{im} = \alpha_m + \beta Age45_i + f(a_i) + v_{im} \tag{2}$$

where y_{im} represent our outcome variable: reemployment hazard, which is equal to 1 if UI recipient i in month m finds a job in month $m+1$. α_m represents monthly baseline hazards. The variable a is individual i's age at layoff and is measured in days. The variable $Age45_i$ is a treatment dummy indicating an individual is eligible for a three-month extended UI benefits (i.e. from six to nine months), being equal to 1 if individual i's age at the time of job loss is greater than 45. $f(a_i)$ is a smooth function of age at job loss that controls the age profile of reemployment hazard. In our main analysis, we specify $f(a_i)$ as a linear function that allows for different slopes below and above age cutoff.[9] v_{im} is an error term that reflects all of the other factors that affect the outcome of interest.

Our primary interest is in β, which measures any deviation from the continuous relation between the age at layoff and the reemployment hazard if an individual has an involuntary job loss after age 45 (i.e., when the treatment variable switches from 0 to 1). The key identification assumption is that all factors except the eligibility for extended UI benefits vary continuously around the individual's 45^{th} birthday at layoff, so that β can be interpreted as the causal effect of a three-month extended UI benefits on the reemployment hazard.

In order to control for potential discontinuity around the age cutoff prior to extended benefits reform, we then extend our empirical specification to RD-DID model by exploiting the policy

[9] Specifically, we use the following linear function: $f(a_i) = \gamma_1(a - c_{45}) + \gamma_2 Age45_i(a - c_{45})$, where c_{45} is the age cutoff of interest (i.e. age 45 at layoff). Note that c_{45} is also measured in day.

changes. Specifically, we include the individuals who lost their job before the reform as the additional control group so that we can account for other time-invariant jumps taking place at the age 45. The RD-DID specification can substantially increase the credibility of our research design and is estimated as follows:

$$y_{im} = \alpha_m + \kappa_1 Age45_i + \kappa_2 PostEB_i + \kappa_3 PostEB_i \times Age45_i + g(a_i) + \varepsilon_{im} \tag{3}$$

where $PostEB_i$ indicates an individual i becomes unemployed after extended benefits reform (i.e. $PostEB_i = 1$). Here, $Age45_i$ is still an indicator for being layoff after age 45 so that κ_1 can represent any change in reemployment hazard for those who lose their jobs around age 45, which is not induced by the extended UI benefits, during pre-reform period. Again, we use a smooth function $g(a_i)$ to control the age profile of reemployment hazard and specify it as a linear function of age at job loss that fully interacts with $Age45_i$ and $PostEB_i$.[10] The key variable in this specification is $PostEB_i \times Age45_i$ indicating the individuals losing their job after age 45 in the post-reform period. Its coefficient κ_3 can capture the causal effect of the extended UI benefits on the hazard of transition to employment. Following Schmieder et al. (2016), we estimate equation (2) and (3) locally within a bandwidth of two years (i.e. 730 days) before and after the age 45 at layoff. In the later section, we examine whether our main results are sensitive to different bandwidth choices and specifications.

5.2 Estimation Results

In this section, we examine the effect of UI extension on reemployment hazard. We first discuss the estimates for UI exhaustees and then investigates the impact of UI extension on UI non-exhaustees. Figure 3 displays how the monthly reemployment hazard during 7^{th} to 9^{th} months of an UI spell (i.e. exhausted period) varies with an individual's age at job loss.[11] As shown in Figure 3a, for those who lose job after age 45, the average reemployment hazard in exhausted period shows a discernible drop at the cutoff by about 7 percentage points. To examine any confounding factors affecting our

[10] Specifically, we use the following linear function: $g(a_i) = \gamma_1(a - c_{45}) + \gamma_2 Age3_a(a - c_{45}) + \gamma_3 PostEB_i \times (a - c_{45}) + \gamma_4 PostEB_i \times Age45_a(a - c_{45})$

[11] We plot monthly reemployment hazard within the 10 years before and after the age 45 and group them into 90-day bins. For example, we group the first 90 days after the age 45 to construct the first bin after the cutoff.

estimates, we repeat the above analysis using pre-reform data (i.e. 2006-2008 sample) as a placebo test. Since worker above the age of 45 at job loss were not eligible for the extended UI benefit during this period, we should not observe any discontinuity in our outcomes if the discontinuity at the cutoff in Figure 3 are mainly driven by extended UI benefit. In sharp contrast to the Figure 3a, we find no visible discontinuity in reemployment hazard at age 45 when using pre-reform data. This provides clear evidence that the change in monthly reemployment hazard at the age 45 is driven exclusively by extension of UI benefits.

Table 3 reports our main estimates for the effect of extended benefits on the reemployment hazard within the 7^{th} to 9^{th} months of an UI spell. The first four columns display the results based on RD models and the last column shows the estimate of RD-DID model. Column (1) displays a basic RD estimate using a linear function to control age profile of reemployment hazard. The result suggests a three-month increase in potential benefit duration significantly reduces the monthly reemployment hazard for UI recipients exhausting their benefits by 7.4 percentage points. Column (2) further includes covariates, such as gender, birth place, and previous job industry. We find the estimate is quite similar, suggesting the estimated effect is not driven by the observed difference between eligible and ineligible workers. In Column (3), we use a quadratic function on either side of the cutoff to control age profile of reemployment hazard and find the estimate decreases slightly to 6.9 percentage points. Column (4) reports a bias corrected estimate using a local linear regression with a triangular kernel and the optimal bandwidth according to algorithm proposed by Calonico et al. (2014). In addition, its standard error is adjusted for bias correction. We find the estimate is robust to this specification.

The estimates in column (1) to (4) are based on RD model, which assumes that the effect of age is continuous around age 45 before the extended benefits reform. To examine this assumption, column (5) includes the workers were unemployed before the reform (i.e. those who lost their job during 2006-2008) since worker above the age of 45 were not eligible for the extended UI benefit during this period. In practice, we implement RD-DID model (i.e. equation 3) by subtracting from our RD estimates any changes in reemployment hazard at age 45 during periods when age-based extended benefits policy was not introduced. The estimated coefficient on $PostEB_i \times Age45_i$ suggest a three-month increase in potential benefit duration reduces the monthly reemployment hazard for UI exhaustees by 7.2 percentage points. Compared to baseline mean (around 12%), it represents 60% reduction in average monthly reemployment hazard. In the later analysis, we will use this estimate as our main result. Overall, the point estimates are quite similar across different specifications.

The extension of UI benefits might also affect the non-exhaustees' reemployment hazard if UI recipients are forward looking. Therefore, we replicate the above analysis for UI non-exhaustees (i.e. reemployment hazard during non-exhausted period). The results in Figure 2 and Panel B of Table 3 suggest the extension of UI benefits has smaller effect on UI recipients' search effort when they have not run out of their benefits. Column (5) in Panel B of Table 3 (i.e. RD-DID estimate) suggests a three-month increase in potential benefit duration significantly reduces the monthly reemployment hazard during non-exhausted period by only 1.4 percentage points. Consistent with the findings in previous literature (Card et al., 2007; Nekoei and Weber, 2016; Caliendo et al., 2013), our result shows that at least some UI recipients can anticipate the longer duration of UI benefits and reduce their effort of finding a job before the benefit extension takes effect.

5.3 Robustness Checks

In this section, we examine the robustness of our main estimate (i.e. column (5) in Table 3) for the effect of extended UI benefits. Table 4 display the results using different specifications and sample criteria. In column (1)-(2), we examine whether our main results are robust to different sample selection criteria. First, we restrict our sample to those who continue claiming UI benefits.[12] The sample size decreases substantially by 70% due to this restriction. We find the estimated effect of extended UI benefits decrease slightly to 6.8 percentage points, which is not significantly different from our main estimate in Table 3. Following Card et al. (2007) and Nekoei and Weber (2016), column (2) displays the estimate based on the sample that excludes temporary layoffs. That is, we eliminate individuals recalled to their prior firms. This restriction reduces sample size

[12] We define such sample by using the recipients whose maximum gap between two UI claims are within 3 days.

by 16% and increase the estimated effect to 8.5 percentage points. However, this estimate is not significantly different from our main estimate in Table 3. Instead of clustering the standard errors by unemployment spell, in column (3), we adjust for potential correlation in errors within age group by clustering the standard errors by age at job loss. We find the standard error of estimated coefficient on $PostEB_i \times Age45_i$ changes slightly due to this setting.

Next, we examine the robustness of our main estimate to bandwidth choices. Figure C1 of Online Appendix C displays the RDD-DID estimates over various bandwidth choices (from 100 to 2,000 days before and after age 45 cutoff). The result in Figure C1 suggests our estimates are quite robust to bandwidth choices. Most estimates are in the range of 6 to 8 percentage points.

Finally, we investigate the validity of our research design, which depends on whether the UI recipients around the age-45 cutoff have similar characteristics except for their eligibility for extended UI benefits. Under Taiwan's UI program, it is unlikely that workers will be able to manipulate the eligibility rule for extended benefits because it is based on their age at the time of job loss rather than their age when claiming benefits.[13] It seems possible, however, that some firms might be willing to delay laying off workers for a certain period of time, so that they would qualify for extended benefits. If many firms were doing this, we would likely see a larger-than-expected number of workers just above age 45 claiming UI benefits. Furthermore, if these workers or employers fell into certain types or industries, then this sorting would not be random, and would need to be addressed. As Imbens and Lemieux (2008) and Lee and Lemieux (2010) suggest, we investigate the validity of our RD design by examining the frequency of UI recipients over different ages.

Figure C2 of Online Appendix C shows the number of UI recipients for each age at layoff (from age 40 to 50). Each age bin represents the total number of new claimants in a 90-day interval. Around the age 45 cutoff, there are roughly 1,200 to 1,400 new claimants within each age interval, and the number of new claimants decreases with age at job loss. Similar to Schmieder et al. (2012), we find that there are about 130 more workers losing their jobs within the first three months (i.e.

[13] The eligibility rules for extended benefits in Germany and Austria are based on an applicant's age when claiming unemployment benefits. Schmieder et al. (2012) found a slight increase in the number of new claimants on the right of each age cutoff, and addressed this concern using a variety of methods, including adding covariates, a donut RD, and bounding.

90 days) after age 45 than just before that cutoff, and that the number of new claimants within a few months past age 45 is still slightly higher than that just before age 45. This increase in the number of UI 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 only 1% of workers aged 43-46 at the time of job loss, and is thus unlikely to invalidate our main estimate.

In order to understand potential bias resulted from this small discontinuity in number of UI recipients around the cutoff, we conduct the following robustness checks. First, we check for the possibility of non-random sorting. In Table C2 of Online Appendix C, we use the pre-determined observables as dependent variables in RDD-DID model (i.e. equation 3). Most estimates are either insignificant or small. The only exception is that UI recipients who lost their jobs right after age 45 tend to have higher previous pay than those right before age 45 (i.e. around 3% higher). It is possible that high-wage employees can negotiate better "severance pay" (i.e. a three-month extension of UI benefits). In addition, such workers are also more likely to find a new job, which might make us underestimate the effect of extended benefits (in absolute value). Moreover, following Card et al. (2007), we predict the a worker's reemployment hazard based on his/her observable characteristics. Then, we examine the effect of extended UI benefits on this predicted reemployment hazard using equation 3. In contrast to our main result, we find the estimated coefficient on $PostEB_i \times Age45_i$ is very small and insignificantly different from zero.

Second, to understand the lower (upper) bound effect of benefits extension on reemployment hazard, following Schmieder et al. (2012), we select a number of UI recipients equal to the bunching (i.e. excess mass) in the density above the cutoff and intentionally reallocate the UI recipients with the longest (shortest) nonemployment durations from the right to the left of the cutoff. Column (4) in Panel A of Table 4 displays the lower bound estimate and suggests a three-month benefits extension reduces monthly reemployment hazard by 5.5 percentage points, which is not significantly different from our main estimate. In addition, the upper bound estimate (see Column (5) in Panel A) is quite similar to the main result.

Third, we follow Card et al. (2007) and replicate our findings with new UI spells that originate from employers experiencing multiple layoffs who hence have less scope for selectively laying off workers. Column (6) in Panel A of Table 4 suggests our main estimate is similar to the estimate based on the sample with less selective layoffs.

Finally, in the Table C1 of Online Appendix C we implement the donut RDD-DID model suggested by Barreca et al. (2016). We exclude observations within 30 to 210 days around the cutoff to examine how selective layoff around the cutoff affects the results. The estimates in Table C1 suggest removal of observations around the cutoff does not overturn our results. To sum up, the above results suggest the impact of manipulating age at layoffs on our estimates could be limited.

6 Effects of Reemployment Bonuses

6.1 Regression Kink Design

In this section, we investigate the effect of reemployment bonus on an individual's search effort to get moral hazard effect of UI. The reemployment bonus program in Taiwan offers 50% of remaining benefits to UI recipients reemployed before exhaustion point and holding new jobs for at least three months. The program was announced by the government on May 15, 2002, before it officially began on January 1, 2003. Importantly, it not only applies to workers starting their UI spells after January 1, 2003, but also workers having UI spells span across January 1, 2003. Therefore, depending on when a recipient starts to receive UI benefits, his/her potential bonus increased as the starting date of UI spell approaches Januray 1, 2003.

Figure 4 displays the relationship between individuals' potential reemployment bonus measured by duration of UI benefits and their starting date of UI spell. There are three segments distinguished by two cutoffs. The first cutoff is July 1, 2002: the recipients starting their spells before this date (i.e. six months before Januray 1, 2003) would run out of their UI benefits so that they were not eligible for bonus. The second cutoff is Januray 1, 2003 because the recipients who start to receive benefits after this date are potentially eligible for full reemployment bonus – 3 months (90 days) of UI benefits as bonuses. Finally, those who start their UI benefits between July 1, 2002 (i.e. the first cutoff) and Januray 1, 2003 (i.e. the second cutoff) are potentially eligible for partial reemployment bonus and their bonus increased linearly as the starting date of UI spell approaches Januray 1, 2003, namely, one day increase in staring date would lead to 0.5 day increase in duration of UI benefits as bonus. Thus, the first kink located at July 1, 2002 where the slope of bonus offer with respect to UI starting date changes from 0 to 0.5. The second kink located at Januray 1, 2003 where the slope changes from 0.5 to 0.

Based on the above observation, to estimate effects of reemployment bonuses, we look for induced kinks in the relationship between the date UI spells start and reemployment outcomes around the cutoffs, and compare the magnitude of the kinks at the cutoffs in the outcome to that of the potential bonus amount. The intuition is that we can attribute the slope change in the outcome (i.e. reemployment hazard) to that in the treatment (i.e. reemployment bonus) if workers are similar around the kinks. To formalize this idea, we implement a regression kink design (Nielsen et al., 2010; Card et al., 2015).

$$E(\frac{\partial y}{\partial RB(t)}|t=c_k)=\frac{\lim\limits_{t\to c_k^+}\frac{dE(y|t)}{dt}-\lim\limits_{t\to c_k^-}\frac{dE(y|t)}{dt}}{\lim\limits_{t\to c_k^+}\frac{dRB(t)}{dt}-\lim\limits_{t\to c_k^-}\frac{dRB(t)}{dt}} \tag{4}$$

where t represents the starting date of UI spell. $E(\frac{\partial y}{\partial RB(t)}|t=c_k)$ is the causal effect of interest: the effect of reemployment bonus $RB(t)$ on the conditional expectation of y (i.e. reemployment hazard) around cutoff date c_k (i.e. c_1 is July 1, 2002 and c_2 is January 1, 2003). We can express it as the slope change in $E(y|t)$ with respect to UI starting date t divided by the slope change in potential reemployment bonus $RB(t)$. In this case, the denominator is straightforward to calculate, since the slope change in $RB(t)$ at these two kinks are deterministic. Specifically, the slope change is 0.5 for the first kink and -0.5 for the second one. To get estimated numerator of equation (4), we can estimate the following model:

$$\mathrm{E}[\mathrm{y}_{im}|t]=\mu_m+\sum_{p=1}^{P}\gamma_p Kink_i(t-c_k)^p+\sum_{p=1}^{P}\delta_p(t-c_k)^p \tag{5}$$

where the variable $Kink_i$ is a dummy indicating an individual starting his/her UI spell after July 1, 2002 (i.e. c_1 cutoff date of the first kink) or January 1, 2003 (i.e. c_2 cutoff date of the second kink).

μ_m represents monthly baseline hazards. In our main analysis, we would focus on the first kink since UI recipients should have limited scope for manipulating their starting date of UI benefits right after policy announcement. We also use the results based on the second kink as a robustness check. $t - c_k$ measures the difference between UI starting date and cutoff date. γ_p and δ_p represent the coefficients on the polynomial terms. We use a linear model (i.e. $P = 1$) as our main specification and conduct a robustness check using a quadratic model (i.e. $P = 2$). The slope change in $E(y|t)$ with respect to UI starting date t (i.e. the estimated numerator of equation (4)) can be measured by γ_1. Combining the the slope change in potential reemployment bonus $RB(t)$ (i.e. denominator of equation (4)), the effect of being eligible for reemployment bonus equivalent to one-day UI benefits on the reemployment hazard can be represented by $\frac{\gamma_1}{0.5}$. To match the estimated effect of three-month (90 days) UI benefit extension, in the following analysis, we will multiply $\frac{\gamma_1}{0.5}$ by 90 (i.e. $\gamma_1 \times 180$) to give effects of eligibility for reemployment bonus that is equivalent to three-month UI benefits. In our main specification, we estimate equation (5) locally within a bandwidth of 180 days before and after the cutoff dates (i.e., c_k).

It is possible that the RKD estimates might confound with the seasonality of job opportunities. In order to control for any seasonal factors affecting reemployment hazard prior to reemployment bonus reform, we then extend our empirical specification to RK-DID model by exploiting the policy changes. Specifically, besides original sample (i.e. 2002-2003 sample), we include the individuals who started their UI spell in January 2001 to June 2002 as the additional control group (i.e. 2001-2002 sample) so that we can account for other time-invariant kink taking place on July 1 or January 1. The RK-DID specification can substantially increase the credibility of our research design and is estimated as follows:

$$\begin{aligned}\mathrm{E}[\mathrm{y}_{im}|t] \quad = \quad & \mu_m + \kappa PostRB_i + \sum_{p=1}^{P} \lambda_p Kink_i (t - c_k)^p + \sum_{p=1}^{P} \theta_p PostRB_i \cdot Kink_i (t - c_k)^p \\ + \quad & \sum_{p=1}^{P} \delta_p (t - c_k)^p + \sum_{p=1}^{P} \pi_p PostRB_i (t - c_k)^p \qquad (6)\end{aligned}$$

where $PostRB_i$ indicates an individual i belong to reform sample: those who started their UI spell between January 2002 to June 2003 (i.e. $PostRB_i = 1$). The coefficient λ_1 on $Kink_i(t - c_k)$

can measure any slope change in relationship, not due to reemployment bonus, between UI starting date and reemployment hazard for those who start their UI benefits around July 1. The parameter of interest in RK-DID model is θ_1, which can capture the causal effect of the reemployment bonus on the hazard of transition to employment. Again, in order to represent the effect of a reemployment bonus equivalent to three-months (90-days) of UI benefits on monthly reemployment hazard, we first divide θ_1 by 0.5 and then multiply it by 90 (i.e. $\theta_1 \times 180$).

6.2 Estimation Results

Figure 5a presents the relationship between average monthly reemeployment hazard within the 1^{st} to 6^{th} months of an UI spell and starting date of UI benefits for the benefit recipients who were age 40 to 50 during January 2002 to June 2003. Each bin represents the total number of UI spells starting within 20 days interval. We find that the monthly reemeployment hazard of the UI recipients who started their spells between July 1, 2002 and Januray 1, 2003 (i.e. partially eligible for reemployment bonus) increased as their UI starting date approaches Januray 1, 2003. On average, the monthly reemeployment hazard increased substantially from 0.04 (those who started UI around July 1, 2002) to 0.07 (those who started UI around Januray 1, 2003).

For those who started their UI spells before July 1, 2002 (i.e. ineligible for reemployment bonus) or after Januray 1, 2003 (i.e. fully eligible for reemployment bonus), we find that their monthly reemeployment hazard change relatively little with their UI starting date. In other words, there are two changes in slope of average reemployment hazard against workers' UI starting date, which is consistent with the relationship of potential bonus offer and UI starting date depicted in Figure 4. Figure 5b displays the relationship between average monthly reemeployment hazard within the 1^{st} to 6^{th} months of an UI spell and starting date of UI benefits for the UI recipients who were age 40 to 50 during January 2001 to June 2002 (i.e. 2001-2002 sample). In sharp contrast to Figure 5a, figure 5b displays the relationship between average monthly reemeployment hazard within the 1^{st} to 6^{th} months of an UI spell and starting date of UI benefits using placebo sample. We find that the monthly reemeployment hazard of 2001-2002 sample (i.e. individuals unaffected by reemployment bonus reform) was constantly around 0.04.

Table 5 reports our main estimates for the effect of reemployment bonus on the reemployment hazard within the 1^{st} to 6^{th} months of an UI spell. The first four columns display the results based on RK models and the last column shows the estimate of RK-DID model. Column (1) displays a basic RK model using a linear function to control the relationship between starting date of UI spell and reemployment hazard. The result suggests being eligible for reemployment bonus equivalent to three-months (90-days) of UI benefits can increase the monthly reemployment hazard by 1.2 percentage points. Column (2) further includes covariates, such as gender, birth place, and previous job industry. We find the estimate is quite similar, suggesting the estimated effect is not driven by the observed difference between eligible and ineligible workers. In Column (3), we use a quadratic function on either side of the cutoff to control age profile of reemployment hazard and find the estimate increase substantially to 6.4 percentage points. However, all quadratic terms are not significantly different from zero and model selection criteria (i.e. Bayesian information criterion) suggests quadratic specification is dominated by linear specification. Column (4) reports a bias corrected estimate using a local linear regression with a triangular kernel and the optimal bandwidth according to algorithm proposed by Calonico et al. (2014). In addition, its standard error is adjusted for bias correction. We find the estimate is robust to this specification.

The estimates in column (1) to (4) are based on RK model, which assumes that the there is no kinked relationship between UI starting date and reemployment hazard before introducing reemployment bonus. To examine this assumption, column (5) includes the workers were unemployed before the reform (i.e. those who started their UI spell during Januray 2001 to June 2002) since no worker were eligible for reemployment bonus during this period. In practice, we implement RK-DID model (i.e. equation 6) by subtracting from the RK estimates that capture kinked relationship in UI starting date and reemployment hazard before the reform. The estimated coefficient on $PostRB_i \times Kink_i(t - c_k)$ suggest a reemployment bonus which is equivalent to three-months (90-days) of UI benefits can increase monthly reemployment hazard by 1.8 percentage points. Compared to baseline mean (around 4.1%), it represents 43% increase in average monthly reemployment hazard. In the later analysis, we will use this estimate as our main result.

6.3 Robustness Checks

In this section, we examine the robustness of our main estimate for the effect of reemployment bonus. Table 6 display the results using different specifications and sample criteria. In column (1)-(2), we examine whether our main results are robust to different sample selection criteria. First, we restrict our sample to those who continue claiming UI benefits.[14] The estimated effect of reemployment bonus based on this sample increases slightly to 2.8 percentage points. Again, following Card et al. (2007) and Nekoei and Weber (2016), column (2) displays the estimate using the sample that excludes temporary layoffs. We find the estimate is quite close to our main result. Instead of clustering the standard errors by unemployment spell, in column (3), we adjust for potential correlation in errors for those with the same UI starting date by clustering the standard errors at UI starting date. We find the standard error of estimated coefficient on $PostRB_i \times Kink_i(t - c_k)$ changes slightly due to this setting. Column (4) display the estimates for the second kink. We find the estimate is similar to the first kink, suggesting being eligible for reemployment bonus equivalent to three-months (90-days) of UI benefits can increase the monthly reemployment hazard by 1.6 percentage points. In the later analysis, we would combine main estimate for the effect of reemployment bonus with the estimate of extended benefits effect to conduct a welfare analysis of UI extension. An important caveat to our main estimate is that we use early 2000s sample to estimate reemployment bonus effect. Following Schmieder et al. (2012), we use the typical procedure to reweight reemployment bonus sample to match the observable characteristics of extended benefits sample. Column (5) reports reweighting estimate, which is similar to our main result.

Next, we examine the robustness of our main estimate to bandwidth choices. Figure D1 of Online Appendix D displays the RK-DID estimates over various bandwidth choices (from 90 to 240 days before and after July 1). The result in Figure D1 suggests our estimates are quite robust to bandwidth choices. Most estimates are in the range of 1 to 3 percentage points.

Finally, we investigate the validity of our research design. That is, UI recipients cannot manip-

[14] Similar to the analysis for extended UI benefits, we define such sample by using the recipients whose maximum gap between two UI claims are within 3 days.

ulate their UI starting date to be eligible for reemployment bonus. Figure D2 in Online Appendix D shows that there is little evidence that workers manipulate their UI starting date around cutoff dates (i.e. July 1, 2002 and January 1, 2003). Consistent with the above findings, in Table D1 of Online Appendix D, we use the observable characteristics as dependent variables in RK-DID model (i.e. equation (6)) and find most RK-DID estimates are either insignificant or small, suggesting the composition of UI recipients are quite similar around cutoff date.

7 Welfare Implications

7.1 Estimated Liquidity Effect

To back out the liquidity effect, we first define each period as an interval of three months such that the regular six months of potential duration is equal to two periods. With this timing definition, increasing the potential duration from six months to nine months is equivalent to a one-period increase in potential duration.[15] According to Table 3, a three-month increase in potential benefit duration $\frac{\partial s_t}{\partial P}$ is estimated to decrease the monthly reemployment hazard between the 7th and 9th month of nonemployment by 7.3 percentage points. On the other hand, based on Table 5, the eligibility for three-month benefit as reemployment bonus $b\frac{\partial s_t}{\partial r_t}$ is estimated to increase the monthly reemployment hazard between the 1th and 6th month of nonemployment by 1.8 percentage points for workers aged between 40 and 50. Plugging the reduced form estimates into equation (1) yields

$$\begin{aligned} b\frac{\partial s_P}{\partial A_P} &= \frac{\partial s_P}{\partial P} + 0.5b\frac{\partial s_P}{\partial r_P} \\ &= -0.072 + 0.5 \cdot 0.018 \\ &= -0.063. \end{aligned}$$

[15] This definition is similar to that of Card et al. (2007). They define ten weeks in a UI spell as one period. Therefore, extending unemployment benefits from 20 weeks to 30 weeks is equivalent to a one-period increase in potential duration under their timing definition.

The above result suggests the ratio of the liquidity effect to the moral hazard effect at UI exhaustion is

$$
\begin{aligned}
R_P &= -\frac{\partial s_P}{\partial A_P} / \frac{\partial s_P}{\partial r_P} \\
&= 0.063/0.018 \\
&= 3.5.
\end{aligned}
$$

The estimated liquidity to moral hazard ratio at UI exhaustion point (3.5) is about four times larger than the exiting estimates from Chetty (2008) and Landais (2015) on the liquidity to moral hazard ratio of increasing benefit level (0.88-1.5).

We can investigate the liquidity to moral hazard ratio before exhaustion using the estimated effect of extended benefits on the monthly reemployment hazard between 1st and 6th month instead of that between 7th and 9th month.

$$
\begin{aligned}
b\frac{\partial s_t}{\partial A_t} &= \frac{\partial s_t}{\partial P} + 0.5bS_{t+1}(P)\frac{\partial s_t}{\partial r_t} \\
&= -0.014 + 0.5 \cdot 0.6 \cdot 0.018 \\
&= -0.009.
\end{aligned}
$$

Therefore, we estimate that the liquidity to moral hazard ratio before exhaustion equals

$$
\begin{aligned}
R_t &= -\frac{\partial s_t}{\partial A_t} / \frac{\partial s_t}{\partial r_t} \\
&= 0.009/0.018 \\
&= 0.5
\end{aligned}
$$

The larger liquidity to moral hazard ratio at UI exhaustion is consistent with Kolsrud et al. (2018)'s findings that consumption smoothing benefits are larger later in the unemployment spell.

7.2 Welfare Effect of Extended Benefits

Next, we analyze the welfare effect of UI extension. Suppose a social planner chooses the potential benefit duration (P) to maximize the expected utility of a job loser at the beginning of period

0 subject to the worker's optimal search behavior and government budget constraint, $Bb + (P - B)\theta b = (T - D)\tau$, where B and D are insured duration and nonemployment duration, respectively. The welfare effect of a balanced-budget increase in unemployment benefits at time P is

$$\frac{dW_0^*}{db_P} = \frac{dW_0}{db_P}/u'(c_P^e) = S_0(P)R_P - \{(1-\theta)\sum_{t=0}^{P-1}\frac{dS_0(t)}{dP} + \tau \cdot \frac{dD}{dP}\} \tag{7}$$

where $\sum_{t=0}^{P-1} \frac{dS_0(t)}{dP}$ is the increase in duration due to reduced search effort before the exhaustion point and $R_P = -\frac{\partial s_P}{\partial A_P} / \frac{\partial s_P}{\partial r_P}$ is the ratio of liquidity effect to moral hazard effect at time P. The welfare effect of UI extension balances the utility gain from the increased ability to maintain consumption during unemployment, and the utility loss due to higher taxes levied on employment. On the one hand, the benefits of extended benefits are determined by the product of the exhaustion rate and the marginal utility gain of redistributing income from when workers are employed to income when unemployed. This product will be large for individuals who are liquidity constrained at the exhaustion point and for individuals whose unemployment duration is long. On the other hand, the tax rate rises because the UI extension increases the insured and nonemployment durations of unemployment, implying a shorter employment duration with which to finance increased benefits.

Plugging our reduced form estimates into the welfare formula 7, we get

$$\begin{aligned}
\frac{dW_0^*}{db_P} &= S_0(P)R_P - \{(1-\theta)\sum_{t=0}^{P-1}\frac{dS_0(t)}{dP} + \frac{(1-\theta)B + \theta P}{T - D} \cdot \frac{dD}{dP}\} \\
&= 0.61 \cdot 3.5 - \{0.5 \cdot (\frac{58.29}{90} - 0.61) + (0.5 \cdot 0.15 + 1.38 \cdot 0.5 \cdot 0.15) \cdot \frac{43.02}{90}\} \\
&> 0
\end{aligned}$$

This means increasing potential duration is estimated to be welfare enhancing for workers aged around 45 because the expected consumption smoothing benefits at exhaustion point are larger than the increased welfare cost arising from lengthened insured and nonemployment durations.

8 Conclusion

This paper has exploited the introduction of a reemployment bonus program and a UI extension to older workers to evaluate optimal potential duration of unemployment benefits. The duration response to the reemployment bonus is linked to the moral hazard effect of UI, because the bonus offer increases workers' incentive to search, while it does not increase workers' ability to maintain consumption during unemployment. On the other hand, the duration response to extended benefits is composed of a liquidity effect and a moral hazard effect, since it not only increases income during unemployment but also distorts the incentive to search. Using a search model with borrowing constraint and the RD (RK) design, we separately identify the liquidity and moral hazard effects of UI. Our estimates show that, for workers aged around 45, the liquidity effect is 3.5 times more than the moral hazard effect of extended benefits, and increasing potential duration is welfare enhancing for them.

Note that our welfare calculation is based on at least two assumptions. First, our model assume a flat labor demand and every unemployed worker is eligible for unemployment benefits. However, in the search model including reservation wage, when the government increase the generosity of unemployment benefits, workers will raise their selectivity, so the firms might be less willing to open vacancies. Second, since workers eligible for more generous benefits will decrease the search effort, those ineligible will have better chance to be employed and become more willing to exert search effort. Recent evidence from Lalive et al. (2015) suggests the latter dominates the former.

We have also assumed the effects of reemployment bonuses on hazard rate in each period before exhaustion point is constant. However, it is possible that workers in different timing of unemployment spell respond differently for reemployment bonuses. Future studies on how the ratio of a liquidity and a moral hazard effect varies over unemployment duration will provide insights on optimal time profile of unemployment benefits (Kolsrud et al. (2018)).

Finally, this paper can be extended to study design optimal age-dependent UI. Intuitively, optimal UI should consider heterogeneity in consumption smoothing benefits and heterogeneity in moral hazard costs, offering a longer potential duration or a higher replacement rate for workers suffering greater losses and having smaller duration responses. However, in the literature, it is unclear how the liquidity effect and moral hazard effect vary over age. On the one hand, in the early years of the life cycle, workers are likely to be more liquidity constrained due to low income, and want jobs that build up a high return of human capital (Michelacci and Ruffo (2015)). On the other hand, young workers might return to live at home when unemployed, operating as an alternative form of insurance (Kaplan (2012)). Estimating how the effects of extended benefits and reemployment bonuses vary over ages will be useful for optimal UI design over life-cycle.

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Tables

Table 1: Descriptive Statistics for Extended Benfits Sample

	2009-2012 Sample		2006-2008 Sample	
	Before Age 45	After Age 45	Before Age 45	After Age 45
Reemployment hazard	0.13	0.06	0.11	0.10
	(0.33)	(0.24)	(0.32)	(0.30)
Age	44.01	45.97	44.02	46.00
	(0.57)	(0.58)	(0.58)	(0.58)
Male	0.51	0.49	0.57	0.56
	(0.50)	(0.50)	(0.50)	(0.50)
Born in Taipei	0.13	0.13	0.14	0.13
	(0.34)	(0.34)	(0.35)	(0.34)
Work in manufacturing previously	0.31	0.35	0.39	0.44
	(0.46)	(0.48)	(0.49)	(0.50)
Insured wage	30,079.27	30,746.21	30,581.57	30,775.61
	(10,472.44)	(10,499.84)	(10,257.87)	(10,206.44)
Temporary layoff	0.18	0.17	0.12	0.13
	(0.38)	(0.38)	(0.33)	(0.33)
Number of previous UI spell	1.20	1.20	1.17	1.16
	(0.45)	(0.46)	(0.40)	(0.39)
Number of recipients	5,409	6,398	4,122	4,231
Number of observations	14,020	18,189	10,809	11,256

Notes: Data are from 2006-2012 unemployment benefits files and the employment insurance enrollee file. We focus on the UI recipients who lost their job two years (730 days) before and after age 45. This table displays the means and standard deviations of our outcome variable and related individual characteristics for the extended benefits sample (2009-2012 sample) and corresponding placebo sample (2006-2008). Reemployment hazard is the average monthly hazard rate in the 7^{th} to 9^{th} of UI spells. Standard deviations are in parentheses.

Table 2: Descriptive Statistics for Reemployment Bonus Sample

	2002-2003 Sample		2001-2002 Sample	
	Before July 1	After July 1	Before July 1	After July 1
Reemployment hazard	0.04	0.06	0.04	0.04
	(0.20)	(0.23)	(0.19)	(0.21)
Age	44.92	44.88	45.13	44.84
	(2.87)	(2.86)	(2.87)	(2.86)
Male	0.50	0.53	0.53	0.52
	(0.50)	(0.50)	(0.50)	(0.50)
Born in Taipei	0.11	0.12	0.09	0.09
	(0.32)	(0.33)	(0.28)	(0.28)
Work in manufacturing previously	0.50	0.49	0.63	0.62
	(0.50)	(0.50)	(0.48)	(0.48)
Insured wage	28,805.42	28,035.32	27,360.93	27,840.58
	(10,228.38)	(9,977.84)	(9,816.49)	(9,859.18)
Temporary layoff	0.15	0.13	0.14	0.13
	(0.36)	(0.34)	(0.35)	(0.34)
Number of previous UI spell	1.01	1.02	1.01	1.01
	(0.11)	(0.15)	(0.09)	(0.11)
Number of recipients	8,942	6,650	6,893	10,411
Number of observations	25,943	18,992	20,098	30,324

Notes: Data are from 2001-2004 unemployment benefits files and the employment insurance enrollee file. We focus on the UI recipients who lost their job 180 days before and after July 1. This table displays the means and standard deviations of our outcome variable and related individual characteristics for the reemployment bonus sample (2002-2003 sample) and corresponding placebo sample (2001-2002 sample). Reemployment hazard is the average monthly hazard rate in the 4^{th} to 6^{th} of UI spells. Standard deviations are in parentheses.

Table 3: The Effect of Extended Benefits on Monthly Reemployment Hazard

	(1)	(2)	(3)	(4)	(5)
Panel A: *UI exhaustees (7^{th} to 9^{th} month)*					
$Age45$	-0.074***	-0.075***	-0.067***	-0.064***	
	(0.007)	(0.007)	(0.010)	(0.005)	
$PostEB \times Age45$					-0.072***
					(0.011)
Baseline mean			0.131		
Sample size	32,209	32,184	32,184	255,827	54,249
Panel B: *UI non-exhaustees (1^{st} to 6^{th} month)*					
$Age45$	-0.015***	-0.016***	-0.017***	-0.015***	
	(0.004)	(0.004)	(0.006)	(0.004)	
$PostEB \times Age45$					-0.014**
					(0.006)
Baseline mean			0.131		
Sample size	96,573	96,505	96,505	861,013	164,292
RDD	Yes	Yes	Yes	Yes	–
RDD+DID	–	–	–	–	Yes
Covariates	–	Yes	Yes	–	Yes
Poly. model	linear	linear	quadratic	linear	linear
Bandwidth (days)	730	730	730	CCT	730

Notes: The first four columns display the estimated coeeficients on $Age45$ using equation (2). The outcome variable is monthly reemployment hazard in the 7^{th} to 9^{th} of a UI spell (Panel A) or that in the 4^{th} to 6^{th} of a UI spell (Panel B). Column (1) displays a basic RD estimate using a linear function to control age profile of reemployment hazard. Column (2) further includes covariates, such as gender, birth place, and previous job industry. Column (3) uses a quadratic function on either side of the cutoff to control age profile of reemployment hazard. The bandwidth choice in Column (1) to (3) is 730 days. Column (4) reports a bias corrected estimate using a local linear regression with a triangular kernel and the optimal bandwidth according to algorithm proposed by Calonico et al. (2014). In addition, its standard error is adjusted for bias correction. Column (5) displays the estimated coefficients on $PostEB \times Age45$ using equation (3). The bandwidth choice in Column (5) is 730 days. Standard errors are in parentheses. Except column (4), all standard errors are clustered at UI spell. *** 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 Monthly Reemployment Hazard: Robustness Checks

	(1) Continuous UI	(2) Non-temporary layoff	(3) Clustered at Age	(4) Lower bound	(5) Upper bound	(6) Mass layoff
Panel A: *UI exhaustees (7^{th} to 9^{th} month)*						
$PostEB \times Age45$	-0.068***	-0.085***	-0.072***	-0.055***	-0.072***	-0.071***
	(0.021)	(0.013)	(0.010)	(0.011)	(0.011)	(0.022)
Baseline mean			0.131			
Sample size	15,747	45,225	54,249	54,249	54,249	12,723
Panel B: *UI non-exhaustees (1^{st} to 6^{th} month)*						
$PostEB \times Age45$	-0.011	-0.014**	-0.014**	-0.003	-0.031***	-0.013
	(0.012)	(0.012)	(0.006)	(0.006)	(0.006)	(0.006)
Baseline mean			0.131			
Sample size	53,745	143,939	164,292	164,292	39,525	164,292

Notes: This table displays the estimated coefficients on $PostEB \times Age45$ in equation (3) using different sample criteria and specifications. The outcome variable is monthly reemployment hazard in the 7^{th} to 9^{th} of a UI spell (Panel A) or that in the 4^{th} to 6^{th} of a UI spell (Panel B). All columns use a linear function to control age profile of reemployment hazard and the bandwidth choice is 730 days. Column (1) displays estimate based on estimated sample who continue claiming UI benefits – the recipients whose maximum gap between two UI claims are within 3 days. The estimate in column (2) is based on the sample that excludes temporary layoffs. Column (3) adjusts for potential correlation in errors within age group by clustering the standard errors by age at job loss. The estimate in column (4) represents lower bound of our main estimate by selecting a number of UI recipients equal to the bunching (i.e. excess mass) in the density above the cutoff and intentionally reallocating the UI recipients with the longest nonemployment durations from the right to the left of the cutoff. Similarly, the estimate in column (5) represents upper bound of our main estimate by intentionally reallocating the excess UI recipients with the short nonemployment durations from the right to the left of the cutoff. Column (6) displays the estimate using the sample from firms experiencing multiple layoffs. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table 5: The Effect of Reemployment Bonus on Monthly Reemployment Hazard

	(1)	(2)	(3)	(4)	(5)
$Kink$	0.012**	0.014***	0.064***	0.046**	
	(0.005)	(0.005)	(0.021)	(0.019)	
$PostRB \times Kink$					0.018**
					(0.007)
Baseline mean			0.106		
Sample size	96,876	96,876	96,876	333,591	204,746
RKD	Yes	Yes	Yes	Yes	–
RKD+DID	–	–	–	–	Yes
Covariates	–	Yes	Yes	–	Yes
Poly. model	linear	linear	quadratic	linear	linear
Bandwidth (days)	180	180	180	CCT	180

Notes: The first four columns display the estimated coefficients on $Kink$ in equation (5) and multiply them by 180 to give effects of eligibility for reemployment bonus that is equivalent to three-month UI benefits. The outcome variable is monthly reemployment hazard in the in the 4^{th} to 6^{th} of a UI spell. Column (1) displays a basic RK estimate using a linear function to control age profile of reemployment hazard. Column (2) further includes covariates, such as gender, birth place, and previous job industry. Column (3) uses a quadratic function on either side of the cutoff to control age profile of reemployment hazard. The bandwidth choice in Column (1) to (3) is 180 days. Column (4) reports a bias corrected estimate using a local linear regression with a triangular kernel and the optimal bandwidth according to algorithm proposed by Calonico et al. (2014). In addition, its standard error is adjusted for bias correction. Column (5) displays the estimated coefficients on $PostRB \times Kink$ in equation (6) and multiply it by 180 to give effects of eligibility for reemployment bonus that is equivalent to three-month UI benefits. The bandwidth choice in Column (5) is 180 days. Standard errors are in parentheses. Except column (4), all standard errors are clustered at UI spell. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

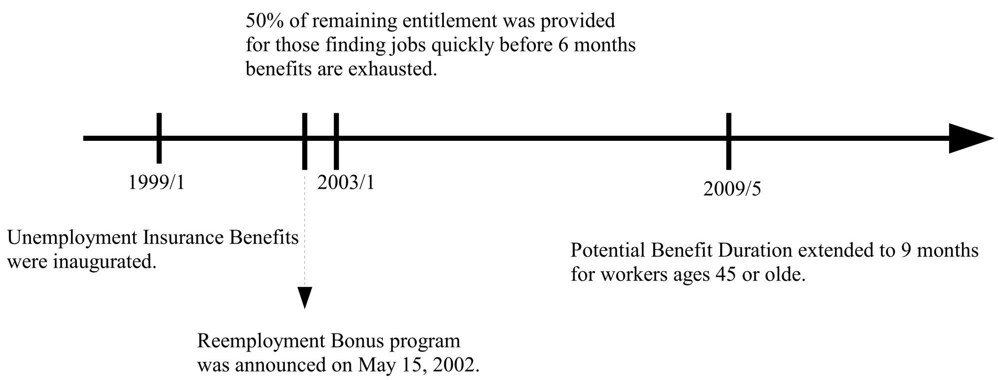
Table 6: The Effect of Reemployment Bonus on Monthly Reemployment Hazard: Robustness Checks

	(1) Continuous UI	(2) Non-temporary layoff	(3) Clustered at UI starting date	(4) Kink 2	(5) Reweighting
$PostRB \times Kink$	0.028**	0.015**	0.018**	-0.016**	0.020*
	(0.010)	(0.008)	(0.009)	(0.008)	(0.011)
Baseline mean			0.106		
Sample size	114,363	176,504	204,746	199,040	204,742
RKD+DID	Yes	Yes	Yes	Yes	Yes
Covariates	Yes	Yes	Yes	Yes	Yes
Poly. model	linear	linear	linear	linear	linear
Bandwidth (days)	180	180	180	180	180

Notes: This table displays the estimated coefficients on $PostRB \times Kink$ in equation (6) and multiply them by 180 to give effects of eligibility for reemployment bonus that is equivalent to three-month UI benefits. The outcome variable is monthly reemployment hazard in the in the 4^{th} to 6^{th} of a UI spell. All columns use a linear function to control age profile of reemployment hazard and the bandwidth choice is 180 days. Column (1) displays estimate based on estimated sample who continue claiming UI benefits – the recipients whose maximum gap between two UI claims are within 3 days. The estimate in column (2) is based on the sample that excludes temporary layoffs. Column (3) adjusts for potential correlation in errors within age group by clustering the standard errors by UI starting date. Column (4) display the estimate for the second kink. Column (5) reports reweighting estimate based on the reweighting sample that match the observable characteristics of extended benefits sample. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Figures

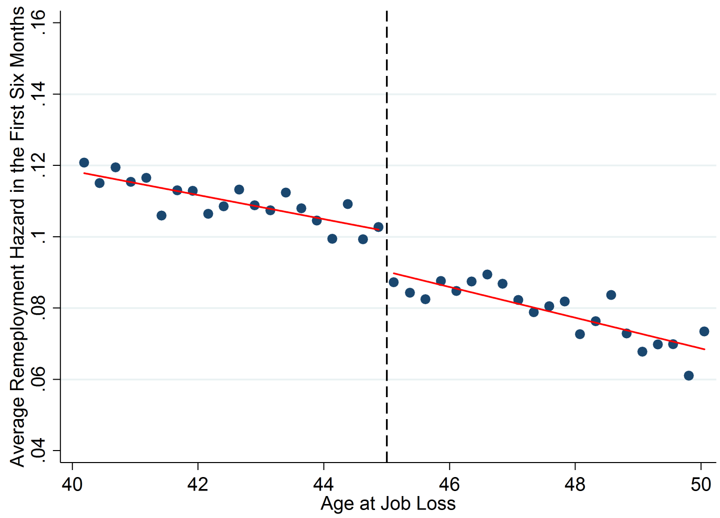
Figure 1: UI Timeline



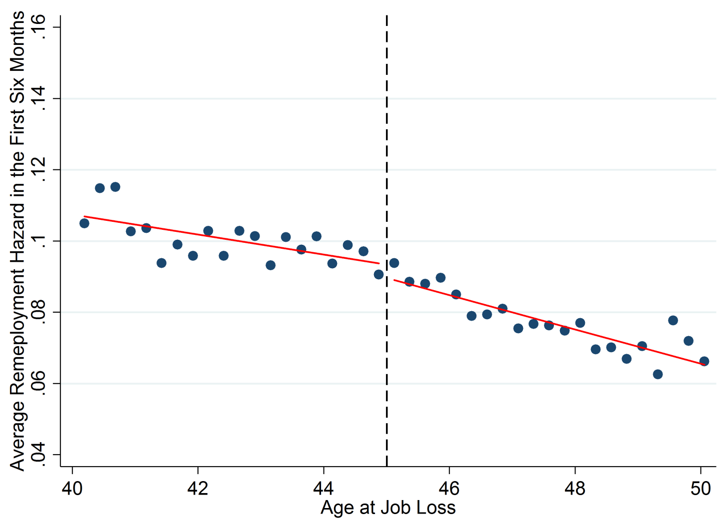
Notes: This figure summarizes the evolution of Taiwan's UI. UI in Taiwan was inaugurated in Jan 1999. On May 15, 2002, the reemployment bonus program was announced. On January 1, 2003, a bonus, equal to 50% of remaining benefits, began to offer for UI recipients who find jobs before exhausting benefits.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

Figure 2: Effects of Extended UI Benefits

(a) Monthly Reemployment Hazard in the 1^{st} to 6^{th} of a UI spell: 2009-2012



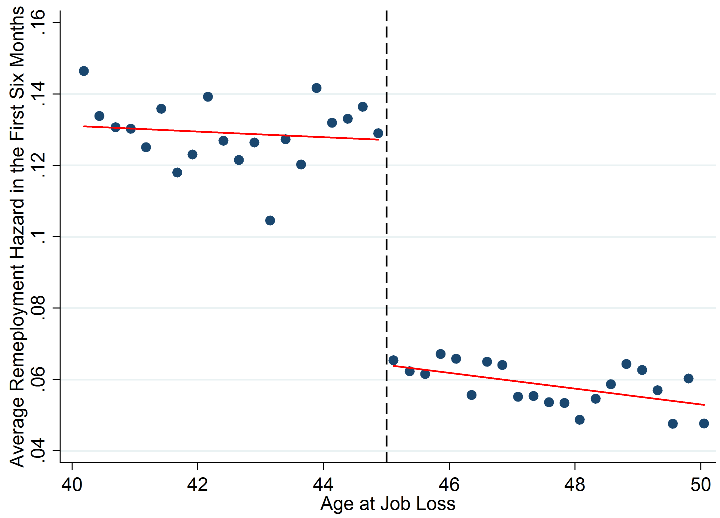
(b) Monthly Reemployment Hazard in the 1^{st} to 6^{th} of a UI spell: 2006-2008



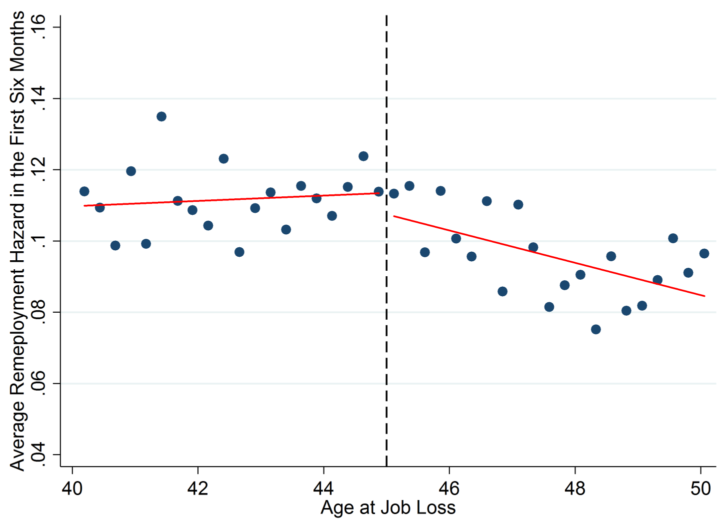
Notes: Data are from 2006-2012 unemployment benefits files and the employment insurance enrollee file. Figure 2a displays result for 2009-2012 sample and figure 2b is for 2006-2008 sample. Both figure plot the average monthly reemployment hazard in the in the 1^{st} to 6^{th} of a UI spell for UI recipients aged 40 to 50 at job loss. Each bin represents the average monthly reemployment hazard within 90 days (3 months) interval. The solid lines are fitted values from a linear regression on either side of the cutoff.

Figure 3: Effects of Extended UI Benefits

(a) Monthly Reemployment Hazard in the 7^{th} to 9^{th} of a UI spell: 2009-2012

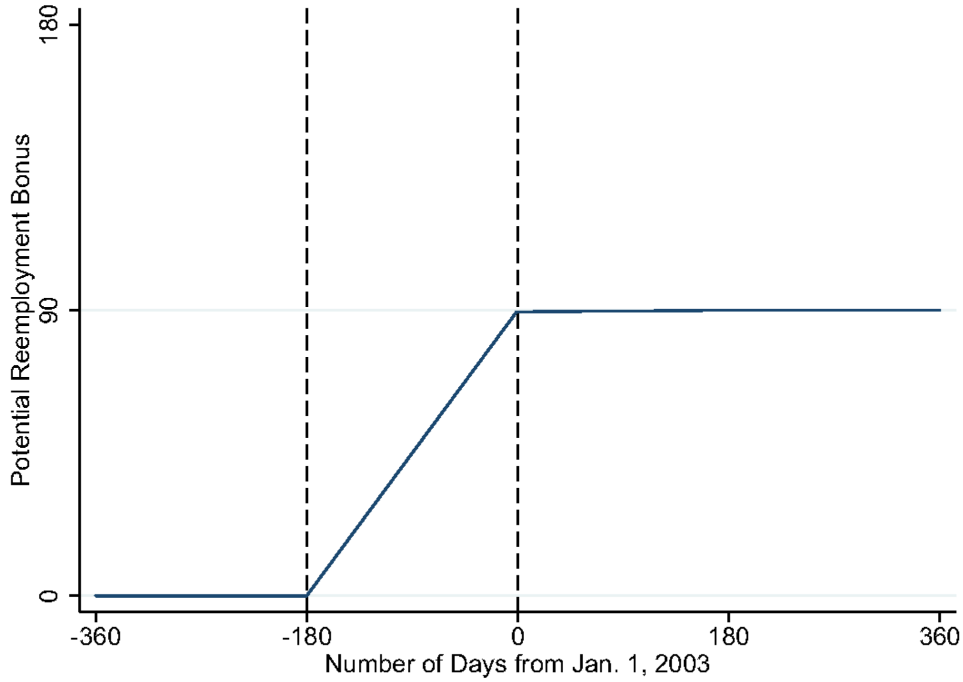


(b) Monthly Reemployment Hazard in the 7^{th} to 9^{th} of a UI spell: 2006-2008



Notes: Data are from 2006-2012 unemployment benefits files and the employment insurance enrollee file. Figure 3a displays result for 2009-2012 sample and figure 3b is for 2006-2008 sample. This figure plots the average monthly reemployment hazard in the in the 7^{th} to 9^{th} of a UI spell for UI recipients aged 40 to 50 at job loss. Each bin represents the average monthly reemployment hazard within 90 days (3 months) interval. The solid lines are fitted values from a linear regression on either side of the cutoff.

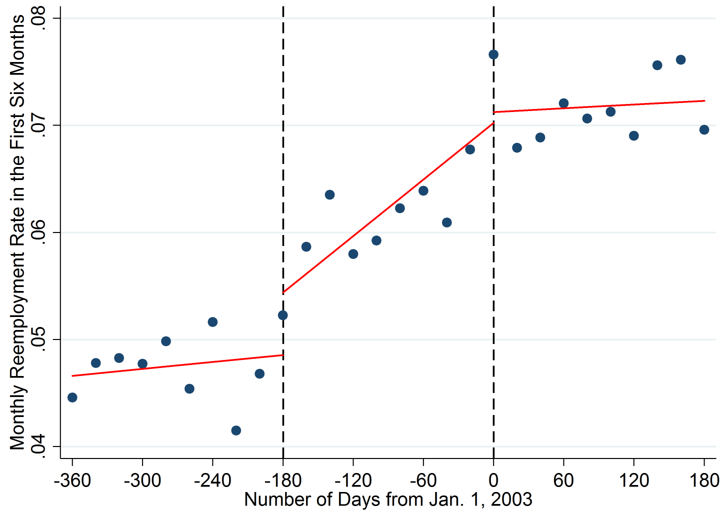
Figure 4: RKD's First Stage



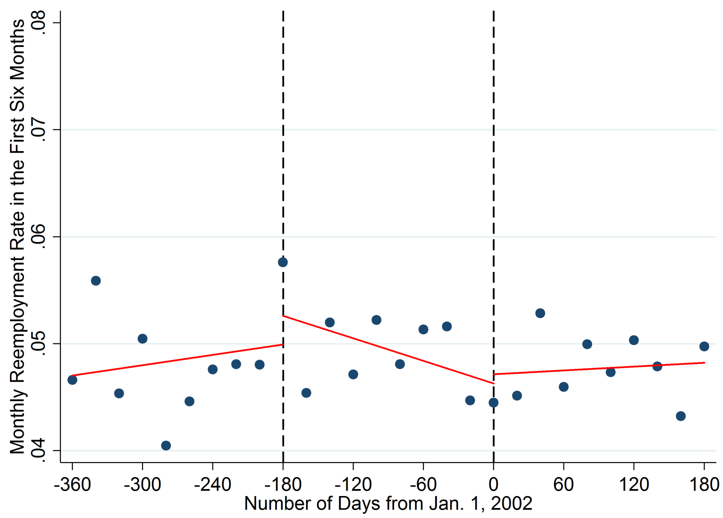
Notes: This figure demonstrates the relationship between the length of qualification period and the date UI spells started. UI recipients starting receiving benefits before July 1, 2002 are not eligible for any reemployment bonus. As the program phased in, UI recipients are potentially eligible for a more generous bonus offer, while the potential reemployment bonus is constant for UI recipients start receiving benefits after Januray 1, 2003.

Figure 5: Effects of Reemployment Bonus

(a) Average Monthly Reemployment Hazard in the 1^{st} to 6^{th} of a UI spell: 2002-2003



(b) Average Monthly Reemployment Hazard in the 1^{st} to 6^{th} of a UI spell: 2001-2002



Notes: Figure 5a plots the average monthly reemployment hazard over the number of days between January 1, 2003 and the date UI spells started. The sample include every UI spell started within 360 days from January 1, 2003. Each bin represents the average monthly reemployment hazard within 20 days interval. The first dash line indicates July 1, 2002, 6 months before the bonus program began. The second line indicates January 1, 2003, the date bonus program began. Figure 5b is for a placebo test. It plots the average monthly reemployment hazard over the number of days between January 1, 2002 and the date UI spells started. The sample include every UI spell started within 360 days from January 1, 2002. The first dash line indicates July 1, 2001, 6 months before the bonus program began. The second line indicates January 1, 2002.

Online Appendix: For Online Publication

A Decomposition of the Effect of Extended Benefits

Consider a discrete time search model based on 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,

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

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

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

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

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

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 in the beginning of period t maximizes

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

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:

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

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

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

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

$$u'(c_t^e) = u'(c_P^e);$$

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

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.

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

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

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

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.

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

B Welfare Effects of Extending Potential Duration

Given the level of benefits, b, and the generosity of bonuses, θ, the social planner chooses the potential duration P to maximize the agents expected utility in the beginning of period 0 subject to the agent's optimization and the government's budget constraint.

$$W_0 = \max_{s_0} s_0 V(A_0) + (1 - s_0)U(A_0) - g(s_0)$$

$$s.t. Bb + (P - B)\theta b = (T - D)\tau;$$

Differentiating W_0 with respect to P yields

$$\begin{aligned}
\frac{dW_0}{db_P} &= (1 - s_0)[\frac{\partial U_0}{\partial b_P} - \frac{\partial U_0}{\partial w}\frac{d\tau}{db_P}] + s_0[\frac{\partial V_0}{\partial b_P} - \frac{\partial V_0}{\partial w}\frac{d\tau}{db_P}] \\
&= (1 - s_0)\frac{\partial U_0}{\partial b_P} + s_0\frac{\partial V_0}{\partial b_P} - [(1 - s_0)\frac{\partial U_0}{\partial w} + s_0\frac{\partial V_0}{\partial w}]\frac{d\tau}{db_P}
\end{aligned}$$

Using envelope theorem and Euler equations,

$$\begin{aligned}
\frac{\partial U_0}{\partial b_P} &= S_1(P)u'(c_p^u) + s_1\theta u'(c_1^e) + .. + (1 - s_1)..(1 - s_{P-1})s_P\theta u'(c_P^e) \\
&= S_1(P)u'(c_p^u) + [1 - S_1(P)]\theta u'(c_P^e)
\end{aligned}$$

$$\frac{\partial V_0}{\partial b_P} = \theta u'(c_0^e)$$

$$\frac{\partial U_0}{\partial w} = \sum_{t=1}^{T-1}[\prod_{i=1}^{t-1}(1 - s_i)]s_t(T - t)u'(c_t^e)$$

$$\frac{\partial V_0}{\partial w} = Tu'(c_0^e)$$

Define $\mathbb{E}_{0,T-1}u'(c_t^e)$ as the average marginal utility of consumption when employed, that is

$$(T - D)\mathbb{E}_{0,T-1}u'(c_t^e) = (1 - s_0)\sum_{t=1}^{T-1}[\prod_{i=1}^{t-1}(1 - s_i)]s_t(T - t)u'(c_t^e) + s_0 Tu'(c_0^e)$$

Plug in into $\frac{dW_0}{db_P}$,

$$\begin{aligned}
\frac{dW_0}{db_P} &= S_0(P)u'(c_p^u) + (1 - s_0)[1 - S_1(P)]\theta u'(c_P^e) + s_0\theta u'(c_0^e) - (T - D)\mathbb{E}_{0,T-1}u'(c_t^e)\frac{d\tau}{db_P} \\
&= S_0(P)u'(c_p^u) + (1 - s_0)[1 - S_1(P)]\theta u'(c_p^e) + s_0\theta u'(c_p^e) - (T - D)u'(c_p^e)\frac{d\tau}{db_P} \\
&= S_0(P)u'(c_p^u) + [1 - S_0(P)]\theta u'(c_p^e) - (T - D)u'(c_p^e)\frac{d\tau}{db_P}
\end{aligned}$$

An increase in benefits in period P increases the tax rate by

$$\frac{d\tau}{db_P} = \frac{d\tau}{dP}\frac{1}{b} = \frac{1}{T-D}[(1-\theta)\frac{dB}{dP} + \theta + \frac{dD}{dP}],$$

where

$$\frac{dB}{dP} = S_0(P) + \sum_{t=0}^{P-1}\frac{dS_0(t)}{dP}$$

In other words, the tax rate, τ, increases because a longer potential duration increases the benefits and bonus payment in a shorter period of employment. Using the decomposition of $\frac{dB}{dP}$ and $=$ $\frac{(1-\theta)B+\theta P}{T-D}$, $\frac{dW_0}{db_P}$ can be written

$$\begin{aligned}
\frac{dW_0}{db_P} &= S_0(P)u'(c_p^u) + [1-S_0(P)]\theta u'(c_p^e) - u'(c_p^e)[(1-\theta)\frac{dB}{dP} + \theta + \frac{(1-\theta)B+\theta P}{T-D}\frac{dD}{dP}] \\
&= S_0(P)[u'(c_p^u) - u'(c_P^e)] - u'(c_p^e)\{(1-\theta)\sum_{t=0}^{P-1}\frac{dS_0(t)}{dP} + \frac{(1-\theta)B+\theta P}{T-D}\frac{dD}{dP}\}
\end{aligned}$$

$$\frac{dW_0}{db_P}/u'(c_P^e) = S_0(P)R_P - \{(1-\theta)\sum_{t=0}^{P-1}\frac{dS_0(t)}{dP} + \frac{(1-\theta)B+\theta P}{T-D}\frac{dD}{dP}\},$$

where we define $R_P = -\frac{\partial s_p/\partial A_p}{\partial s_p/\partial w_p}$.

Finally, the ratio of liquidity to moral hazard effect at time P can be transformed to the ratio of liquidity to moral hazard effect at time P using

$$\begin{aligned}
R_t &= \frac{u'(c_t^u) - u'(c_t^e)}{u'(c_t^e)} \\
&= \frac{S_{t+1}(P)u'(c_p^u) + [1-S_{t+1}(P)]u'(c_p^e) - u'(c_t^e)}{u'(c_t^e)} \\
&= S_{t+1}(P)R_p.
\end{aligned}$$

This implies the ratio of liquidity to moral hazard effect increase over the UI spell.

C Robustness Checks for Regression Discontinuity Design

Table C1: The Effect of Extended Benefits on Monthly Reemployment Hazard: Donut RD Analysis

	Monthly Reemployment Hazard							
Size of Donut around age 45 (days)	0	15	30	45	60	75	90	105
Panel A: *UI exhaustees (7^{th} to 9^{th} month)*								
$PostEB \times Age45$	-0.073***	-0.070***	-0.072***	-0.078***	-0.081***	-0.079***	-0.078***	-0.072***
	(0.011)	(0.011)	(0.012)	(0.012)	(0.013)	(0.014)	(0.014)	(0.015)
Sample size	54,249	53,082	51,878	50,764	49,611	48,492	47,404	46,260
Panel B: *UI non-exhaustees (1^{st} to 6^{th} month)*								
$PostEB \times Age45$	-0.014**	-0.013**	-0.013*	-0.013*	-0.012*	-0.012	-0.011	-0.013
	(0.006)	(0.006)	(0.007)	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
Sample size	164,292	160,790	157,176	153,793	150,356	146,912	143,599	140,180

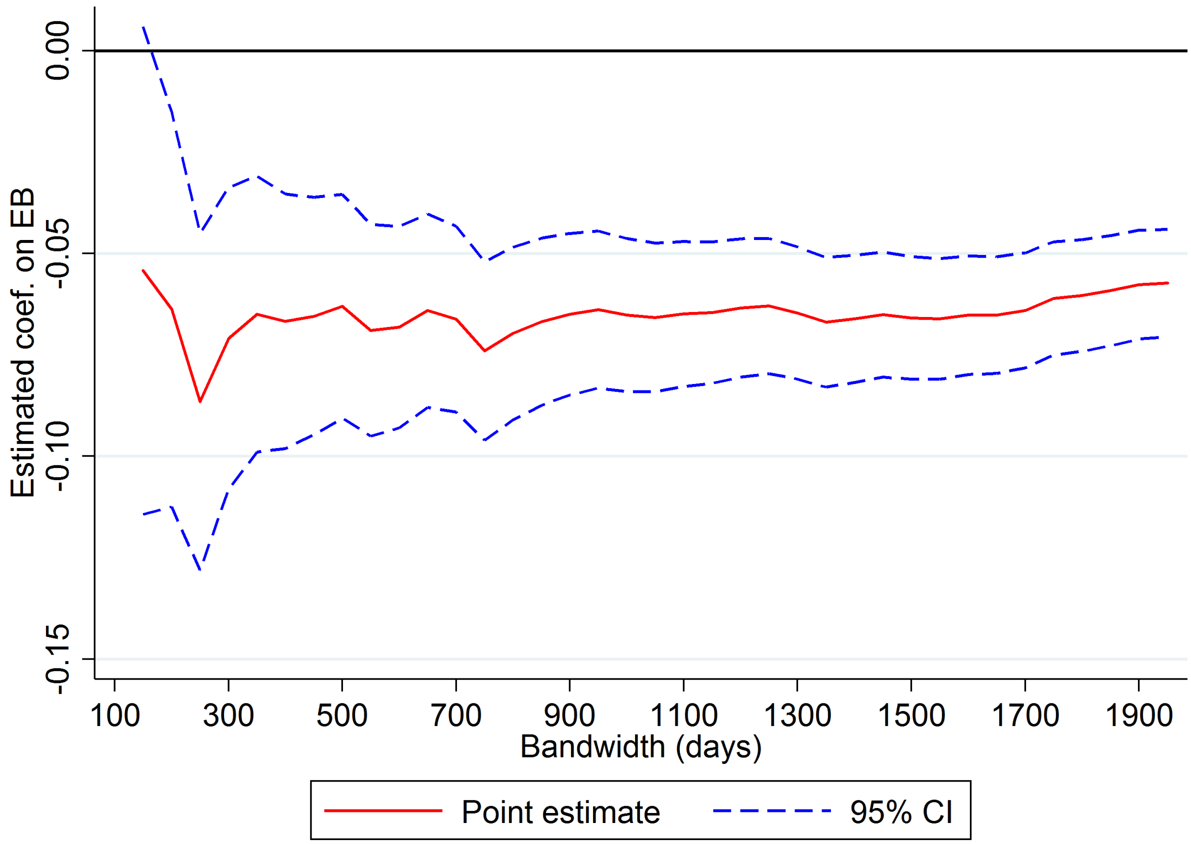
Notes: This table displays the estimated coefficients on $PostEB \times Age45$ in equation (3) by excluding different number of days around age 45. The outcome variable is monthly reemployment hazard in the 7^{th} to 9^{th} of a UI spell (Panel A) or that in the 4^{th} to 6^{th} of a UI spell (Panel B). All columns use a linear function to control age profile of reemployment hazard. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table C2: Examining Smoothness of Predetermined Covariates for Extended Benefits Sample

	(1) Female	(2) Born in Taipei	(3) Manu. Sector	(4) Previous Wage	(5) Recall Job	(6) Number of Unemployment	(7) Predicted Reemp. hazard
$PostEB \times Age45$	-0.017	-0.001	-0.014	876.3**	-0.015	0.025	0.001
	(0.02)	(0.015)	(0.021)	(443.9)	(0.012)	(0.02)	(0.001)
Sample size	35,550	35,563	35,563	35,563	35,563	35,563	54,249

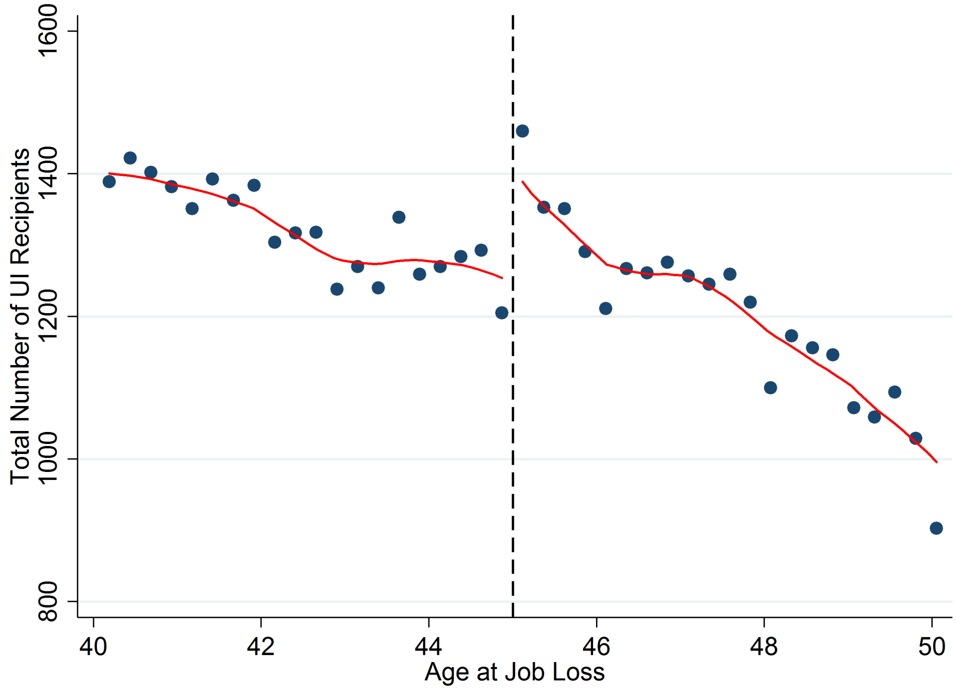
Notes: This table displays the estimated coeeficients on $PostEB \times Age45$ in equation (3) using different covariates as outcomes. All columns use a linear function to control age profile of reemployment hazard and the bandwidth choice is 730 days. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Figure C1: RD Estimates with Varying bandwidths



Notes: This figure the estimated coeeficients on $PostEB \times Age45$ in equation (3) using different bandwidth choices ranging from 100 to 2000 days. The solid line indicates the point estimates, and the dash lines are corresponding 95% confidence intervals.

Figure C2: 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 90 days interval.

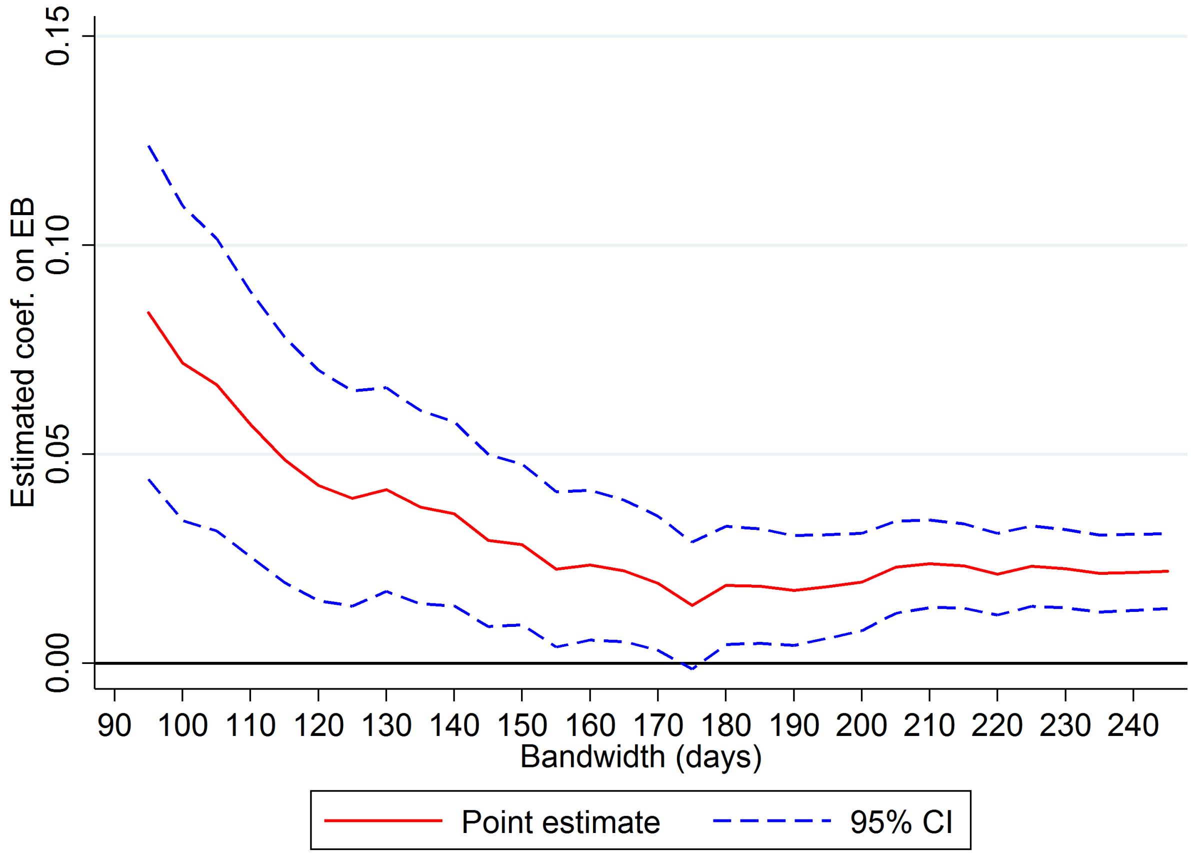
D Robustness Checks for Regression Kink Design

Table D1: Examining Smoothness of Predetermined Covariates for Reemployment Bonus Sample

	(1) Female	(2) Born in Taipei	(3) Manu. Sector	(4) Previous Wage	(5) Recall Job	(6) Number of Unemployment	(7) Predicted Reemp. hazard
$PostRB \times Kink$	0.000	-0.000	0.001***	3.444	0.000	-0.0000	-0.000
	(0.000)	(0.000)	(0.000)	(3.915)	(0.000)	(0.0000)	0.001
Sample size	39,079	39,079	39,079	39,079	39,079	39,079	95,357

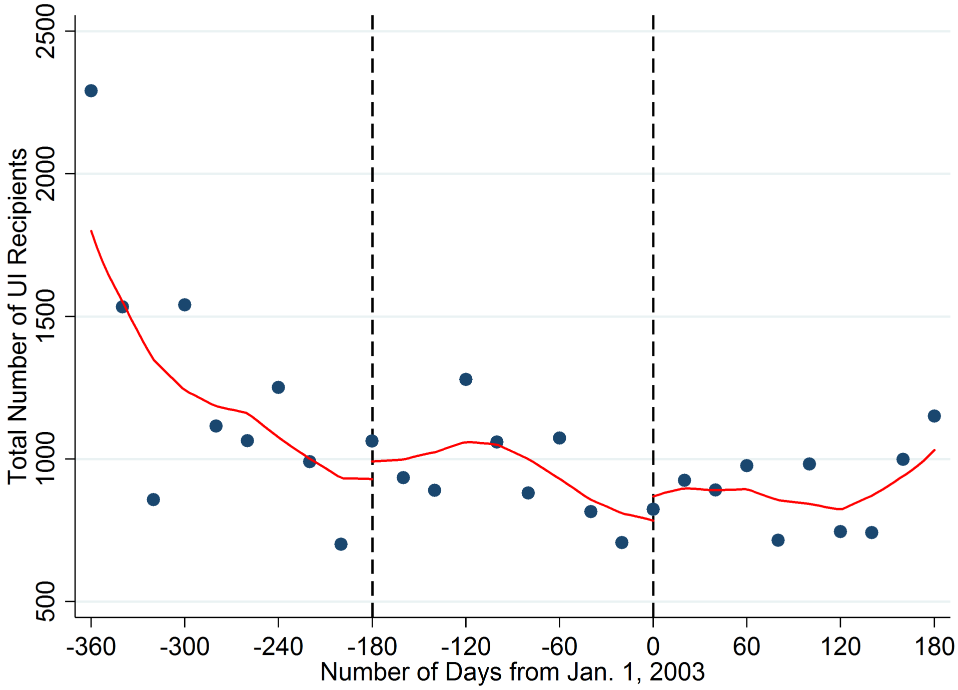
Notes: This table displays the estimated coefficients on $PostRB \times Kink$ in equation (6) using different covariates as outcomes. All columns use a linear function to control age profile of reemployment hazard and the bandwidth choice is 180 days. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Figure D1: RK Estimates with Varying Bandwidths Using Kink 1



Notes: This figure the estimated coeeficients on $PostRB \times Kink$ in equation (6) using different bandwidth choices ranging from 100 to 240 days. The solid line indicates the point estimates, and the dash lines are corresponding 95% confidence intervals.

Figure D2: Validity of RKD: Density Test



Notes: This figure plots the number of workers starting UI spells between January 1, 2002 and June 1, 2003, conditional on age at job loss. Each bin corresponds to the total number of workers starting UI spells within a 20 days interval.