Evaluating the Welfare Effects of Extending Unemployment Insurance Benefits: Evidence from Two Natural Experiments

Po-Chun Huang* Tzu-Ting Yang

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Abstract

This paper estimates the welfare gain (i.e. ratio of liquidity to moral hazard effect) of extending UI benefits, using two natural experiments in Taiwan: a three-month benefits extension for middle-aged job losers and the introduction of a re-employment bonus. Our strategy exploits the fact that a re-employment bonus affects an individual's search efforts only through the moral hazard effect. Therefore, we recover liquidity effects by estimating the responses of the search effort to a UI extension and a re-employment bonus. We find that the estimated liquidity-to-moral hazard ratio of extending UI benefits is around 3.9, suggesting that the welfare gain of extending potential benefit duration is substantially larger than that of increasing benefit level.

*Po-Chun: National Chengchi University (huangpo5@msu.edu); Tzu-Ting: Academia Sinica (email: ttyang@econ.sinica.edu.tw). Po-Chun wishes to thank Steve Woodbury, Carl Davidson, Steven Haider, and Todd Elder for their guidance and support at Michigan State University. Tzu-Ting wants to thank excellent assistance from Yung-Yu Tsai and Yan-Yu Chiou. We thank Andrew Shephard, Yingying Dong, Betsey Stevenson, Taehyun Ahn, and participants at the North American Econometric Society Summer Meeting, the UM-MSU-UWO Labor Day Conference, the Canadian Labour Economics Forum, National Tax Association Annual Meeting, SOLE Annual Meeting, ASSA Annual Meeting, and the Learning from Taiwan Workshop at Seoul National University for their helpful comments. The unemployment insurance data were provided by the Bureau of Labor Insurance. This paper represents the views of the authors and does not reflect those of the Bureau of Labor Insurance. Po-Chun acknowledges financial support from the Chiang Chin-kuo Foundation and Ministry of Science and Technology (MOST-106-2410-H-004-187). Tzu-Ting acknowledges financial support from Ministry of Science and Technology (MOST-105-2410-H-001-014).

1 Introduction

Unemployment insurance (UI) protects individuals from the risk of earnings loss during unemployment, but it also distorts incentives to search for jobs. Chetty (2008) addresses this issue by showing that more generous UI decreases workers' search efforts through two channels with distinct welfare implications: a moral hazard effect and a liquidity effect. On the one hand, the moral hazard effect represents a welfare loss, since UI benefits decrease the opportunity cost of being unemployed and distorts workers' incentive to find a job. On the other hand, the liquidity effect is a welfare gain, because UI benefits increase workers' ability to smooth out and control consumption during unemployment and allows them more time to search for a job. Empirically, however, only a few papers (Chetty, 2008; Card et al., 2007; Landais, 2015) distinguish the liquidity effect from the moral hazard effect and provide estimates for welfare gains as a result of UI. Furthermore, as Schmieder et al. (2016) point out, although the duration of benefits as opposed to benefit levels is at the center of UI policy debates, most existing studies focus on the welfare effect of changing the UI benefit level (e.g. increase the replacement rate). The empirical evidence regarding the welfare effect of extending UI benefit (e.g. increase the potential duration of UI) is still scant.[1]

In this paper, we contribute to the current literature by estimating the welfare impacts of a UI extension. Specifically, we decompose the welfare effect of extending potential benefit duration into a moral hazard effect and a liquidity effect, using two natural experiments in Taiwan. The first natural experiment is the 2009 UI extension for middle-aged workers. Since 2009, workers who lost their jobs when aged 45 or over became eligible for 9 months of UI benefits instead of the 6 months offered to those under 45. We use a regression discontinuity (RD) design to examine the effect of extended UI benefits on re-employment hazard, by comparing the job finding rate of individuals just before and just after the 45 years old at the point of being laid off. Using a search model with borrowing constraints, we demonstrate that the effect of UI extension on search effort is a combination of a liquidity effect and a moral hazard effect.

We then use the second natural experiment—the introduction of a re-employment bonus pro-

[1] Schmieder and von Wachter (2016) provides an excellent survey of the literature on the effects of unemployment benefits on unemployment duration, re-employment wages, and welfare.

gram in January 2003—to estimate the moral hazard effect. In 2003, the Taiwanese government introduced re-employment bonuses whereby people would be paid 50% of their remaining UI benefits after they were re-employed. We can use the estimated effect of a re-employment bonus to identify the moral hazard effect of UI, because the bonus only increases the opportunity cost of being unemployed and does not change workers' income (cash-on-hand) during unemployment. In other words, we can recover the liquidity effect as long as we are able to provide credible estimates of labor supply responses to the UI extension and the re-employment bonus. The re-employment bonus program reached back to UI recipients receiving benefits when the program took effect in 2003. Therefore, it resulted in two kinks in the bonus offer for which workers were eligible as a function of the date UI spells started. Thus, we use a regression kink (RK) design to examine the effect of the re-employment bonus on the re-employment hazard of middle-aged UI recipients. Plugging the RD estimate (i.e. total effect) and the RK estimate (i.e. moral hazard effect) into the search model, we recover the liquidity effect of extending UI benefits and carry out welfare analysis.

We obtain three key findings. First, our RD estimates suggest that a three-month increase in potential benefit duration reduced middle-aged UI recipients' monthly re-employment hazard by 8 percentage points during the extended benefit period (i.e. the 7th to 9th months of an unemployment spell). In addition, we find that the extension of UI benefits also affected workers' search efforts before the UI extension was brought into effect—a three-month extension to UI reduces the monthly re-employment hazard, even in the regular benefit period (i.e. the first six months of an unemployment spell) by around 1.3 percentage points. Second, our estimates using the RK design show that being eligible for three months of UI benefits as a re-employment bonus increases the monthly re-employment hazard by about 1.8 percentage points for UI recipients aged between 40 and 50 at job loss.

Third, we decompose the effect of extending UI benefits and estimate the liquidity-to-moral hazard ratio by plugging the above estimates into a welfare decomposition formula. As discussed in a later section, the ratio of liquidity to moral hazard effect for the extended benefits period can

identify the consumption-smoothing gain of UI extension. Our results suggest that this gain is quite large for middle-aged UI recipients—the estimated liquidity-to-moral hazard ratio of extending potential benefit duration is around 3.9. Furthermore, following Landais (2015)'s argument, we provide a estimate of the consumption-smoothing gain from leveling UI benefits by decomposing the effect of the UI extension on search effort during the regular benefit period. We find the estimated liquidity-to-moral hazard ratio of increasing benefit levels is only 0.46, which is somewhat smaller than estimates found in the previous literature (Landais, 2015; Chetty, 2008). In sum, our results suggest that for middle-aged UI recipients the welfare gain of extending potential benefit duration is much greater than that of increasing benefit levels. The larger welfare gain from a UI extension is consistent with Ganong and Noel (2019) and Kolsrud et al. (2018)'s findings, in that consumption-smoothing benefits increase throughout the unemployment spell. Finally, we find this three-month UI extension is welfare-enhancing, even after incorporating welfare costs from fiscal externality (increased tax) due to job search distortions.

Our paper stands apart from previous literature on unemployment insurance in the following ways. First, we provide one of the first pieces of evidence on the welfare gain from extending UI benefits. There have been many convincing studies estimating the effects of UI benefits on unemployment duration (i.e. the welfare cost of UI), but empirical evidence on welfare gains (i.e. consumption-smoothing benefits) from UI is still scarce, and most existing studies focus on the welfare gain of increasing the UI benefit level (Gruber, 1997; Chetty, 2008; Card et al., 2007; Landais, 2015).[2] In general, these studies found that the welfare gain of increasing the benefit level is substantial, and its effect on search effort is driven almost as much by liquidity effects as by

[2] Gruber (1997) used panel data from the Panel Study of Income Dynamics and state variations in UI replacement rate. Gruber (1997)'s estimates suggest that a 10% increase in the replacement rate reduces the consumption drop during unemployment by 2.8%. However, these estimates are imprecise, and the estimated consumption-smoothing benefits are sensitive to the risk aversion coefficient. Chetty (2008) and Landais (2015) circumvent issues surrounding the estimation of risk aversion coefficient by using the sufficient statistic approach, whereby consumption-smoothing benefits equal 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, while Landais (2015) uses the difference in the behavioral response to extended benefits and an increase in benefit level to identify the moral hazard, which identifies the liquidity effect indirectly. These authors' estimates suggest that about half of the effect of UI on unemployment duration can be attributed to the liquidity effect, thereby suggesting significant consumption-smoothing benefits resulting from increasing the UI benefit level.

moral hazard effects, with estimated liquidity-to-moral hazard ratios of 0.88 (Landais, 2015) and 1.5 (Chetty, 2008), respectively. However, the welfare gain from extending duration of UI benefits could be even greater than that of increasing the UI benefit level, because extending UI benefits mainly affect consumption in the period when UI recipients run out of their benefits. A small but growing body of literature has developed a theory to demonstrate how extending the duration or changing the path of UI benefits affects social welfare (Schmieder and von Wachter, 2017; Kolsrud et al., 2018; Lindner and Reizer, 2019). The estimate of the consumption-smoothing gain as a result of extending the duration of UI benefits is a key parameter when implementing these theories, but almost no existing study provides such an estimate. One noticeable exception, however, is Ganong and Noel (2019), who used bank account data to investigate individuals' monthly spending during the course of unemployment and estimated the drop in consumption once the benefit was exhausted. They found that losing eligibility for UI benefits resulted in a 12 percent drop in monthly consumption. Based on this estimate, they concluded that the consumption-smoothing gain from extending duration of UI benefits is four times greater than from increasing the UI benefit level. However, their estimates were not based on a quasi-experimental design and might have involved selection bias, since the decision whether or not to exhaust UI benefits is an individual choice.[3] In contrast, the variation in eligibility for extended UI benefits we use in this paper is from an exogenous policy rule. Compared with Ganong and Noel (2019), we estimate the welfare effect of extending UI benefits by using a sufficient statistic approach with labor market data that does not need to assume the value of risk aversion coefficient. Moreover, our results are based on quasi-experimental design so that selection bias is less concerned.

Second, our paper provides the first evaluation of optimal UI in Asian countries. Most existing studies focus on UI policies in US or European countries. The welfare effect of UI in Asian countries could be different from existing evidence, since nations such as China, Japan, Korea, or

[3] Therefore, the characteristics of UI recipients could be quite different from those of UI exhaustees, which might affect estimating consumption drop at benefit exhaustion. For example, people who choose to run out of their UI benefits could have more household resources to support their consumption than those who find a new job before benefit exhaustion. Thus, the estimated drop in consumption during the UI-exhausted period could be underestimated.

Taiwan have low unemployment rates, high saving rates and close family ties.[4] Therefore, unemployed workers in these countries could have more available financial resources from their own savings or family transfers, so that might be less likely to be liquidity constrained. Consistent with these facts, our estimated liquidity-to-moral hazard ratio of increasing benefit level (i.e. 0.46) is lower than the existing estimates found in the US (e.g. Landais (2015): 0.88 and Chetty (2008): 1.5). Therefore, we expect that our estimates of the liquidity-to-moral hazard ratio for extending duration of UI benefits could serve a lower bound for estimates in US or European countries.

Finally, this paper adds new evidence in terms of the impact of re-employment bonuses on job searching. Existing evidence on the effect of re-employment bonuses still relies on the US field experiments conducted in the 1980s (Woodbury and Spiegelman, 1987; Decker et al., 2001; Meyer, 1995), and evidence in this regard from these studies is mixed and still inconclusive.[5] Ahn (2018) provides more recent evidence on this issue by using a quasi-experimental method and administrative data from South Korea. He finds that increases in the re-employment 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 for the effect of a re-employment bonus program in Taiwan. More importantly, we reinterpret the labor supply effect of bonuses as a countervailing force to the moral hazard effect of UI, following which we use it to recover the liquidity effect of UI extension.

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

[4] According to the World Economic Outlook Database, in 2012, Taiwan's saving rate was 32%, South Korea 34%, China 49%, and Japan 23%, all of which were higher than for the US (18%) and the UK (12%).

[5] Woodbury and Spiegelman (1987) find that a $500 bonus offered to Illinois UI claimants, who found new employment within 11 weeks, reduced the average duration of UI benefit receipt by more than a week. O'Leary et al. (1995)'s estimates in the bonus experiments for Pennsylvania and Washington suggested a smaller effect of bonuses on UI duration (about half a week).

2 Unemployment Insurance in Taiwan

Unemployment benefits are a part of the Employment Insurance (EI) in Taiwan, which is a mandatory national program that offers unemployment benefits, re-employment bonuses, vocational training living allowances, parental leave allowances, and national health insurance premium subsidies. EI covers all Taiwanese workers, excluding civil servants and the self-employed. Its premium is financed by 1% of the monthly insured salary: 20% is imposed on workers, 70% on employers, and the government pays the remaining 10%. Note that there is a cap on monthly insured salary. During our sample period, the salary cap is 43,900 NT$ per month, so around 23% of the sample's monthly salaries are censored.

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 involuntary job loss.[6] In order 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 benefit period begins. Since 2009, the maximum duration for benefits has been six months for workers aged below 45 at the time of job loss, and nine months for those aged 45 or older when they lost their job.[7]

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 re-employed before the end of a given 30-day interval, the amount of benefits paid in that month is prorated. The replacement rate of UI benefits is 60% of the average insured salary during the six months prior to job loss for those without non-working dependants.[8]

[6] Only workers losing their jobs involuntarily or due to the ending of a fixed-term contract are eligible. According to the Employment Insurance Act and the Labor Standard Act, involuntary separation from employment refers to separation from employment because the insured unit has closed down, relocated, suspended business, dissolved, filed for bankruptcy, or the business cycle has induced lay-offs and downsizing. Employment history is the number of days for which a worker has been enrolled in the employment insurance scheme. Since part-time workers must be insured according to the Employment Insurance Act, history as a part-time worker is included when determining eligibility.

[7] 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.

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

For UI recipients with one more non-working dependant, the replacement rate is increased by 10 percentage points, and it can reach as high as 80%.

Workers are required to search actively 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 re-employment bonus, offers 50% of any 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 re-employment. These three months do not have to be continuous, or with a single employer, and so a person who has worked for multiple employers for three months after re-employment will also qualify for the bonus.[9]

3 Theoretical Framework

In this section, we use a simple search model to show that the effect of extending potential benefit duration on the search effort can be decomposed into a liquidity effect and a moral hazard effect. Then, we argue that a moral hazard effect can be identified through estimating the effect of the re-employment bonus so that we can identify the liquidity effect indirectly. Finally, we show that the ratio of the liquidity effect to the moral hazard effect can identify the welfare gain from a UI extension. Here, we focus on intuition and leave the derivations to the Online Appendix A.

3.1 Model

Consider a discrete time search model modified from Chetty (2008) and Landais (2015). An individual becomes unemployed at period 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. Assume job searching results in a cost of $g(s_t)$, which is strictly increasing and convex. If she is unemployed

[9] The three-month re-employment period does not include recalls (work experience in the firm prior to lay-off).

at period t, she receives an unemployment benefit, b_t, with a potential benefit duration, P, that is

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

To match UI policy in Taiwan, we define each period as an interval of three months, such that the regular six months of potential benefit duration are equal to two periods.[10] Thus, in the later sections, the 1st to 6th months of an unemployment spell are defined as a "regular benefit period," since all workers are eligible for UI benefits during this period (i.e. $0 \leq t < P$).[11] With this timing definition, increasing the potential benefit duration from six months to nine months (i.e. extending UI benefits) is equivalent to a one-period increase in potential benefit duration. Note that the goal of this paper is to estimate the welfare effect of the UI extension so that we focus on job search behavior at the point of regular benefit exhaustion. Specifically, we define the 7th to 9th months (i.e. $t = P$) of an unemployment spell as an "extended benefit period," because workers under the age of 45 at lay-off are not eligible for UI benefits during this period, and those above 45 can receive extended benefits.[12] If an individual is re-employed before running out of UI benefits, she receives a re-employment bonus, r_t, equal to θ percent of the remaining benefits; otherwise, $r_t = 0$. Formally,

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

The worker's consumption at period t equals the difference in income and savings. Her income depends on her employment status, while savings are represented by changes in asset, $A_{t+1} - A_t$. Note that credit and insurance markets in this model are incomplete. Thus, an individual is liquidity-constrained and needs to keep her asset A_{t+1} above a lower bound on assets L (i.e. $A_{t+1} \geq L$). When employed, she earns a wage income, w_t, and pays a tax τ. Thus, the flow utility when employed at period t equals $u(c_t^e) = u(A_t - A_{t+1} + w_t + r_t - \tau)$, where c_t^e indicates the consumption

[10] This definition is similar to that of Card et al. (2007), who 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.

[11] In other words, all workers are eligible for benefits during period 0 (i.e. $t = 0$, 1st to 3rd months) and period 1 (i.e. $t = 1$, 4th to 6th months) so that the potential benefit duration P is equal to two periods.

[12] Under our time definition, workers under 45 at lay-off lost their regular benefits in period 2, and those above the age of 45 were still eligible for benefits due to one period (i.e. three-months) UI extension.

when employed at period t. Assuming the interest rate and the time discount rate are zero, the value of being employed in period t is

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

If a worker remains unemployed at period 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

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

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 at the beginning of period t maximizes

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

The intra-temporal first-order condition balances the marginal cost of searching and the difference between the value of being employed and unemployed at time t. That is,

$$g'(s_t) = V_t(A_t) - U_t(A_t). \tag{1}$$

We derive the effect of increasing potential benefit duration on search effort ($\frac{\partial s_t}{\partial P}$) by differentiating the intra-temporal first-order condition (i.e. equation (1)) with respect to P. Similarly, we differentiate the first-order condition with respect to A_t and w_t to obtain $\frac{\partial s_t}{\partial A_t}$ and $\frac{\partial s_t}{\partial w_t}$, respectively.

3.2 Decomposing the Effects of Extending UI Benefits

In the Online Appendix A, we show that if the credit constraint is not binding at time t, the effect of extended benefits on search intensity at period t can be written as

$$\frac{\partial s_t}{\partial P} = b \frac{\partial s_t}{\partial A_t} - (1 - \theta) S_{t+1}(P) b \frac{\partial s_t}{\partial w_t}; \forall t \leq P, \tag{2}$$

where $S_{t+1}(P) = (1 - s_{t+1})..(1 - s_P)$ is the probability of still being umeployed in period P conditional on being unemployed in period t. Similar to Chetty (2008) and Landais (2015), when an individual can only smooth consumption imperfectly, the effect of extending UI benefits on

search intensity ($\frac{\partial s_t}{\partial P}$) is a combination of the liquidity effect ($b\frac{\partial s_t}{\partial A_t}$) and the moral hazard effect ($-b\frac{\partial s_t}{\partial w_t}$). In equation 2, the moral hazard effect is multiplied by $S_{t+1}(P)$, because workers can only receive extended benefits if they exhaust their benefits. Moreover, the $(1-\theta)$ appears in the second term because the bonus counteracts the moral hazard effect of extended benefits, without changing the income stream during unemployment.[13]

Intuitively, the liquidity effect means that a UI extension can enhance an individual's ability to smooth consumption by raising cash in hand during unemployment, especially for an extended benefit period. Therefore, an individual feels less pressurized to find re-employment quickly, resulting in a lower search intensity. On the other hand, the moral hazard effect can also reduce search intensity, because UI benefits essentially decrease the net wage of a new job, which thus discourages the individual to seek employment opportunities. Both liquidity and moral hazard effects affect an individual's job search behavior in the same direction; however, two effects have distinct welfare implications: the liquidity effect is a behavioral response involving correcting credit and insurance market failures, which can increase social welfare, while, in contrast, the moral hazard effect can reduce social welfare, since it leads to excessive leisure by distorting the relative price of leisure and consumption.

Following Chetty (2008) and Landais (2015), in the Online Appendix A, we show that the ratio of liquidity to the moral hazard effect R_t (i.e. $-\frac{b\frac{\partial s_t}{\partial A_t}}{b\frac{\partial s_t}{\partial w_t}}$) identifies consumption-smoothing gain of UI at period t, namely, the gap in marginal utility between unemployed and employed states induced by a consumption reduction during the unemployment period. That is,

$$R_t = -\frac{b\frac{\partial s_t}{\partial A_t}}{b\frac{\partial s_t}{\partial w_t}} = \frac{u'(c_t^u) - u'(c_t^e)}{u'(c_t^e)}.$$

Therefore, the consumption-smoothing benefit of UI is larger when this gap becomes bigger. A UI extension mainly targets the extended benefit period (i.e. $t = P$), when consumption is much lower and marginal utility is much higher, due to long-term unemployment. We expect the estimated

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

liquidity-to-moral hazard ratio for the extended benefit period is higher than for the regular benefit period (i.e. $t < P$).

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), $-b\frac{\partial s_t}{\partial w_t}$ (i.e. moral hazard effect), and $b\frac{\partial s_t}{\partial A_t}$ (i.e. liquidity effect). In Section 5, we estimate the total effect of extended benefits ($\frac{\partial s_t}{\partial P}$) by exploiting age discontinuity in the eligibility rule, but we have information on neither the asset amount nor exogenous variations in wage offers. To address this issue, we recognize that variations in wage income and in the bonus offer affect the search intensity in the same way, so that $\frac{\partial s_t}{\partial w_t} = \frac{\partial s_t}{\partial r_t}$. A similar argument can be found in Ahn (2018). In Section 6, we estimate the effect of the re-employment bonus on search effort $b\frac{\partial s_t}{\partial r_t}$ (i.e. the moral hazard effect), using the introduction of the bonus program. Then, we can recover $b\frac{\partial s_t}{\partial A_t}$ (i.e. liquidity effect) indirectly.

4 Data and Sample

4.1 Data

We use two sources of administrative data from the Bureau of Labor Insurance (BLI) in Taiwan: unemployment benefits files and employment insurance enrollee files, 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 salary in the six months prior to lay-off (hereafter, previous salary), a scramble individual identifier (ID), 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 measure precisely our key variable, namely, the recipient's age at the time of job loss.

For the employment insurance enrollee files, each entry represents a change in the employment record: (1) New enrolments 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 salary, and a scramble ID. We use the scramble ID to merge unemployment benefits files

with employment insurance enrollee files, and we use the date of new enrolments in employment insurance after job loss to represent the date of re-employment.

4.2 Sample

To examine the effect of the UI extension, since the extended benefits reform took effect on May 1, 2009, we use those recipients who lost their job during May 2009 and July 2011 as our post-reform sample (i.e. 2009-2011 sample) and focus on individuals losing their jobs when they were age 43 to 47 (i.e. two years before and after the age of 45). The first two columns of Table 1 report the summary statistics of selected characteristics for the control group (column (1), age 43-44) and the treatment group (column (2), age 45-46), respectively.

We find that the treatment group has a lower re-employment hazard than the control group in both the regular benefit period (i.e. 1st to 6th month) and the extended benefit period (i.e. 7th to 9th month). Consistent with this finding, the benefit duration of treatment group (i.e. 204 days) is much longer than that of control group (i.e. 144 days). In addition, both groups have similar individual characteristics, except that treatment group has a higher share of recipients working in the manufacturing sector than the 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): individuals who lost their job between May 2006 and July 2008. The last two columns of Table 1 report the summary statistics of selected characteristics for the pre-reform sample. We find that the difference in the re-employment hazard between individuals losing their job before age 45 (column (3), age 43-44) and those losing their job after the age of 45 (column (4), age 45-46) is quite small before the benefits extension reform. Similar to the post-reform sample, the only difference in selected characteristics is that the older group also has a higher share of recipients working in the manufacturing sector than the younger group.

To investigate the impact of the re-employment bonus, we utilize recipients who lost their job around six months (i.e. 180 days) before and after July 1, 2002, as the reform sample (i.e. 2002-2003 sample). Since the relationship between the potential re-employment bonus and the starting

date of a UI spell change for recipients starting their UI spell after July 1, 2002, we define this cohort as the treatment group, and those starting their UI spell before July 1, 2002, are seen as the control group. The first two columns of Table 2 suggest that the treatment group (column (2)) has a higher re-employment hazard than the control group (column (1)). Consistent with this result, the benefit duration of treatment group (i.e. 156 days) is shorter than that of control group (i.e. 164 days). For other characteristics, both groups are quite similar. Moreover, in order to control for any seasonality in relation to employment opportunities, we also include a pre-reform sample for a placebo test (2001-2002 sample), namely, recipients who lost their jobs between six months (i.e. 180 days) before and after July 1, 2001, in our main analysis (see columns (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 "age discontinuity" to identify the UI effect on labor market outcomes (Lalive, 2008; Schmieder et al., 2016; Nekoei and Weber, 2019). To quantify the effects of extending UI benefits, we estimate the following regression:

$$y_{im} = \alpha_m + \beta^{Age45} Age45_i + f(a_i) + v_{im} \tag{3}$$

where y_{im} represents our outcome variable—the re-employment 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 lay-off and is measured in days. The variable $Age45_i$ is a treatment dummy indicating an individual is eligible for a three-month (90-days) extended UI benefits (i.e. from six to nine months), namely, 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 the re-employment hazard. In our main analysis, we specify $f(a_i)$ as a linear function that allows for different slopes below and above the age cutoff.[14] v_{im} is an error term that reflects all of the other factors affecting the outcome of interest.

[14] Specifically, we use the following linear function: $f(a_i) = \pi_1(a - c_{45}) + \pi_2 Age45_i \times (a - c_{45})$, where c_{45} is the age cutoff of interest (i.e. age 45 at lay-off). Note that c_{45} is also measured in days.

Our primary interest is in β^{Age45}, which measures any deviation from the continuous relation between age at lay-off and the re-employment hazard if an individual has an involuntary job loss after the age of 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 45th birthday at lay-off, so that β^{Age45} can be interpreted as the causal effect of a three-month extended UI benefits period on the re-employment hazard.

In order to control for potential discontinuity around the age cutoff prior to extended benefits reform, we then combine the RD design with difference-in-differences (DID) by exploiting policy changes (i.e. RD-DID model). Specifically, we include those 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 age 45. The RD-DID specification can substantially increase the credibility of our research design and is estimated as follows:

$$y_{im} = \alpha_m + \kappa_1 Age45_i + \kappa_2 PostEB_i + \beta^{PostEB} PostEB_i \times Age45_i + g(a_i) + \varepsilon_{im} \quad (4)$$

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 laid-off after the age of 45, so κ_1 can represent any change in re-employment hazard for those who lose their jobs around 45, which is not induced by the extended UI benefits during the pre-reform period. Again, we use a smooth function $g(a_i)$ to control the age profile of the re-employment hazard, and we specify it as a linear function of age at job loss that fully interacts with $Age45_i$ and $PostEB_i$.[15] The key variable in this specification is $PostEB_i \times Age45_i$, indicating individuals losing their job after the age of 45 in the post-reform period. Its coefficient β^{PostEB} 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 (3) and (4) locally within a bandwidth of two years (i.e. 730 days), before and after the age 45 at lay-off. In a later section, we examine whether our main results are sensitive to different bandwidth choices and specifications.

[15] Specifically, we use the following linear function: $g(a_i) = \rho_1(a - c_{45}) + \rho_2 Age45_a \times (a - c_{45}) + \rho_3 PostEB_i \times (a - c_{45}) + \rho_4 PostEB_i \times Age45_a \times (a - c_{45})$

5.2 Estimation Results

In this section, we examine the effect of a UI extension on re-employment hazard. We first discuss the estimate for the extended benefit period followed by the regular benefit period. Figure 2 displays how the monthly re-employment hazard during the extended benefit period (i.e. 7th to 9th months of a UI spell) varies in accordance with an individual's age at job loss.[16] As shown in Figure 2a, for those who lose their job after the age of 45, the average re-employment hazard in the extended benefit period shows a discernible drop at the cutoff by about 7 percentage points. To examine any confounding factors affecting our estimates, we repeat the above analysis by using pre-reform data (i.e. 2006-2008 sample) as a placebo test. Since workers above the age of 45 at job loss were not eligible for the extended UI benefit during this period, we should not observe any discernible drop in our outcomes if discontinuity at the cutoff in Figure 2a was mainly driven by an extended UI benefit. In sharp contrast to Figure 2a, we find no visible change in re-employment hazard at the age of 45 when using pre-reform data (see Figure 2b). This provides clear evidence that the change in the monthly re-employment hazard at the of age 45 is driven exclusively by the extension of UI benefits.

Table 3 reports our main estimates for the effect of a UI extension on the re-employment hazard during the extended benefit period. The first four columns display the results based on RD models, and the last column shows the estimate for the RD-DID model. Column (1) displays a basic RD estimate, using a linear function to control the age profile of the re-employment hazard. The result suggests that a three-month increase in potential benefit duration significantly reduces the monthly re-employment hazard by 7.7 percentage points during an extended benefit period. In Column (2), we use a quadratic function on either side of the cutoff to control the age profile of the re-employment hazard and find the estimate decreases slightly to 6.8 percentage points. However, all quadratic terms are not significantly different from zero, and model selection criteria such as Akaike information criterion (AIC) and Bayesian information criterion (BIC) suggest that we

[16] We plot the monthly re-employment hazards within the 10 years before and after the age of 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.

should choose a linear function as the main specification. Therefore, column (3) reports the RD estimate based on a linear specification with covariates such as gender, birth place, monthly insured salary of previous job, industry of previous job, and number of previous UI spell. We find the estimate is quite similar, suggesting the estimated effect is not driven by the observed difference between eligible and ineligible workers. Column (4) reports a bias-corrected estimate using a local linear regression with a triangular kernel and the optimal bandwidth according to an algorithm proposed by Calonico et al. (2014). In addition, its standard error is adjusted for bias correction. We find the estimate is robust in relation to this specification.

The estimates in columns (1) to (4) are based on the 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 workers who were unemployed before the reform (i.e. those who lost their job during 2006-2008), since workers above the age of 45 were not eligible for the extended UI benefit during this period. In practice, we implement the RD-DID model (i.e. equation (4)) by subtracting from our RD estimates any changes in the re-employment hazard at 45 when the age-based extended benefits policy had not been introduced. The estimated coefficient on $PostEB_i \times Age45_i$ suggests that a three-month extension in UI benefits reduces the monthly re-employment hazard by 8 percentage points during the period of extended benefits. Compared to the baseline mean (around 13%), it represents a 62% reduction in the average monthly re-employment 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 re-employment hazard during the regular benefit period (i.e. 1st to 6th months of a UI spell), if UI recipients are forward-looking. Compared with estimates for the extended benefit period, the results in Figure 3 and Panel B of Table 3 suggest the extension of UI benefits has a much smaller effect on UI recipients' search efforts during the regular benefit period (i.e. when they have not run out of their UI 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 re-employment hazard during the regular benefit period by only

1.3 percentage points (i.e. around a 10% reduction from the baseline mean). Consistent with the findings in the previous literature (Card et al., 2007; Nekoei and Weber, 2019; Caliendo et al., 2013), our result shows that at least some UI recipients can anticipate a longer duration for UI benefits and reduce their effort in 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 displays the results using different specifications and sample criteria. We first discuss the results for the extended benefit period (see Panel A) and then conduct the same robustness checks for the regular benefit period (see Panel B).

In columns (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.[17] The sample size decreases substantially by 72% due to this restriction. We find the estimated effect of the extended UI benefits increases to 8.7 percentage points, which is not significantly different from our main estimate in Table 3. Following Card et al. (2007) and Nekoei and Weber (2019), column (2) displays the estimate based on the sample that excludes temporary lay-offs. That is, we eliminate individuals recalled to their prior firms. This restriction reduces sample size by 17% and increases the estimated effect to 9.5 percentage points. However, it 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 any potential correlation in errors within age group by clustering the standard errors by age at job loss. We find the standard error of the estimated coefficient on $PostEB_i \times Age45_i$ changes slightly as a result of this setting.

Next, we examine the robustness of our main estimate to bandwidth choices. Figure B1 of the Online Appendix B displays RD-DID estimates over various bandwidth choices (from 100 to 1,900 days before and after the age of 45 cutoff). The result in Figure B1 suggests our estimates are quite robust to bandwidth choices, as most are within the 6 to 8 percentage point range.

Finally, we investigate the validity of our research design, which depends on whether UI recip-

[17] We define such a sample by using those recipients whose maximum gap between two UI claims is within 3 days.

ients 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 the job loss rather than on when claiming benefits.[18] It seems possible, however, that some firms might be willing to delay laying off workers for a certain period of time, so that they will qualify for extended benefits. If many firms were doing this, we would likely see a larger-than-expected number of workers just above the age of 45 claiming UI benefits. Furthermore, if these workers or employers fell into certain types of industries, then this sorting would not be random and would thus 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 age ranges.

Figure B3 of the Online Appendix B shows the number of UI recipients for each age at lay-off (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 1,000 to 1,200 new claimants within each age interval, albeit this number decreases in line with age at job loss. Similar to Schmieder et al. (2012), we find that there are about 167 more workers losing their jobs within the first three months (i.e. 90 days) after the age of 45 than just before that cutoff, and the number of new claimants within a few months past 45 is still slightly higher than for just before 45. This increase in the number of UI recipients at and just above the cutoff is significant at the 5% level when 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 the job loss, and it is thus unlikely to invalidate our main estimate.

In order to understand any potential bias resulting from this small discontinuity in the numbers of UI recipients around the cutoff, we conduct the following robustness checks. First, we check for the possibility of non-random sorting. In Table B1 of the Online Appendix B, we use the pre-determined observables as dependent variables in the RD-DID model (i.e. equation (4)) to examine any change in sample composition after extended benefit reform. Most estimates are either

[18] The eligibility rules for extended benefits in Germany 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 they addressed this concern by using a variety of methods, including adding covariates, a donut RD, and bounding.

insignificant or small. 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 a worker's re-employment hazard based on his/her observable characteristics. Then, we examine the effect of extended UI benefits on this predicted re-employment hazard, using equation (4). 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 the re-employment hazard, and 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 those UI recipients with the longest (shortest) non-employment 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 that a three-month benefits extension reduces the monthly re-employment hazard by 6 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 originating from employers experiencing multiple lay-offs and 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 lay-offs.

Finally, inspired by Barreca et al. (2016), we implement the donut RD-DID model in Table B2 of the Online Appendix B. We exclude observations within 15 to 105 days around the cutoff to examine how selective lay-offs around the cutoff affect the results. The estimates in Table B2 suggest that the removal of observations around the cutoff does not overturn our results. To sum up, the above results suggest that the impact of manipulating age at lay-offs on our estimates could be limited.

We also replicate the same robustness checks for the regular benefit period in Panel B of Table 4, Panel B of B2, and Figure B2. Although some estimates are not statistically significant, most of

the estimated magnitudes are quite similar to our main results.

6 Effects of Re-employment Bonuses

6.1 Regression Kink Design

In this section, we investigate the effect of the re-employment bonus on an individual's search effort, in order to identify the moral hazard effect of the UI. The re-employment bonus program in Taiwan offers 50% of remaining benefits to UI recipients re-employed before the exhaustion point and holding a new job 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 applies not only to workers starting their UI spells after January 1, 2003, but also to those with UI spells spanning across January 1, 2003. Therefore, depending on when a recipient started to receive UI benefits, his/her potential re-employment bonus increased as the starting date of UI spell approached January 1, 2003.

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

Based on the above observation, to estimate the effects of re-employment bonuses, we look for induced kinks in the relationship between the starting dates of UI spells and re-employment outcomes around the cutoffs, and we then compare the magnitude of the kinks at the cutoffs in the outcome to that of the potential bonus amount. The idea is that we can attribute the slope change in the outcome (i.e. re-employment hazard) to that in the treatment (i.e. re-employment 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).

$$E(\frac{\partial y}{\partial RB(t)}|t = c_k) = \frac{\lim_{t \to c_k^+} \frac{dE(y|t)}{dt} - \lim_{t \to c_k^-} \frac{dE(y|t)}{dt}}{\lim_{t \to c_k^+} \frac{dRB(t)}{dt} - \lim_{t \to c_k^-} \frac{dRB(t)}{dt}} \tag{5}$$

where t represents the starting date of the UI spell. $E(\frac{\partial y}{\partial RB(t)}|t = c_k)$ is the causal effect of interest: the effect of re-employment bonus $RB(t)$ on the conditional expectation of y (i.e. re-employment 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 re-employment bonus $RB(t)$. In this case, the denominator is straightforward to calculate, since the slope change in $RB(t)$ at these two kinks is deterministic. Specifically, the slope change is 0.5 for the first kink and -0.5 for the second one. To find the estimated numerator for equation (5), we can estimate the following model:

$$\mathrm{E}[\mathrm{y}_{im}|t] = \mu_m + \sum_{s=1}^{S} \gamma_s^{Kink} Kink_i \times (t - c_k)^s + \sum_{s=1}^{S} \delta_s (t - c_k)^s \tag{6}$$

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 focus on the first kink, since UI recipients should have limited scope for manipulating their UI benefit starting date right after the policy announcement. We also use the results based on the second kink as a robustness check. $t - c_k$ measures the difference between the UI starting date and the cutoff date. γ_s^{Kink} and δ_s represent the coefficients on the polynomial terms. We use a linear model (i.e. $S = 1$) as our main specification

and conduct a robustness check, using a quadratic model (i.e. $S = 2$). The slope change in $E(y|t)$ with respect to UI starting date t (i.e. the estimated numerator for equation (5)) can be measured by γ_1^{Kink}. Combining the slope change in potential re-employment bonus $RB(t)$ (i.e. the denominator of equation (5)), the effect of being eligible for a re-employment bonus equivalent to one day of UI benefits on the re-employment hazard can be represented by $\frac{\gamma_1^{Kink}}{0.5}$. To match the estimated effect of a three-month (90 days) UI benefit extension, in the following analysis we multiply $\frac{\gamma_1^{Kink}}{0.5}$ by 90 (i.e. $\gamma_1^{Kink} \times 180$) to give the effects of eligibility for a re-employment bonus that is equivalent to three-months (90-days) UI benefits. In our main specification, we estimate equation (6) locally within a bandwidth of 180 days, before and after the cutoff dates (i.e. c_k).

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

$$\begin{aligned} \mathrm{E}[\mathrm{y}_{im}|t] &= \mu_m + \kappa PostRB_i + \sum_{s=1}^{S} \lambda_s Kink_i \times (t - c_k)^s + \sum_{s=1}^{S} \gamma_s^{PostRB} PostRB_i \times Kink_i \times (t - c_k)^s \\ &+ \sum_{s=1}^{S} \delta_s (t - c_k)^s + \sum_{s=1}^{S} \theta_s PostRB_i \times (t - c_k)^s \end{aligned} \tag{7}$$

where $PostRB_i$ indicates an individual i belonging to the reform sample: those who started their UI spell between January 2002 and June 2003 (i.e. $PostRB_i = 1$). The coefficient λ_1 on $Kink_i \times (t - c_k)$ can measure any slope change in the relationship, not due to the re-employment bonus, between the UI starting date and re-employment hazards for those who started their UI benefits around July 1 (January 1). The parameter of interest in the RK-DID model is γ_1^{PostRB}, which can capture the causal effect of the re-employment bonus on the hazard of transitioning to employment. Again, in order to represent the effect of a re-employment bonus equivalent to three-months (90-

days) UI benefits on the monthly re-employment hazard, we first divide γ_1^{PostRB} by 0.5 and then multiply it by 90 (i.e. $\gamma_1^{PostRB} \times 180$).

6.2 Estimation Results

Figure 5a presents the relationship between the average monthly re-employment hazard during a regular benefit period (i.e. the 1st to 6th months of a UI spell) and the starting date of UI benefits for UI recipients aged 40 to 50 for January 2002 to June 2003. Each bin represents the total number of UI spells starting within a 20-day interval. We find that the monthly re-employment hazard for those UI recipients who started their spells between July 1, 2002, and January 1, 2003 (i.e. partially eligible for re-employment bonus) increased as their UI starting date approached January 1, 2003. On average, the monthly re-employment hazard increased substantially from 0.05 (those who started UI around July 1, 2002) to 0.07 (those who started UI around January 1, 2003). For those who started their UI spells before July 1, 2002 (i.e. ineligible for a re-employment bonus), or after January 1, 2003 (i.e. fully eligible for a re-employment bonus), we find that their monthly re-employment hazard changed relatively little in relation to their UI starting date. In other words, there are two changes in the slope of the average re-employment hazard against workers' UI starting date, which is consistent with the relationship between the potential bonus offer and the UI starting date, as depicted in Figure 4.

Figure 5b displays the relationship between the average monthly re-employment hazard during a regular benefit period and the starting date for UI benefits for UI recipients aged 40 to 50 between January 2001 and June 2002 (i.e. placebo sample, individuals unaffected by the re-employment bonus reform). In sharp contrast to Figure 5a, Figure 5b suggests that the monthly re-employment hazard for the placebo sample did not change according to the UI spell's starting date and was constantly around 0.05.

Table 5 reports our main estimates for the effect of the re-employment bonus on the re-employment hazard during a regular benefit period. The first four columns display the results based on RK models, and the last column shows the estimate of the RK-DID model. Column (1) displays a basic

RK model using a linear function to control the relationship between the UI starting date and the re-employment hazard. The result suggests that being eligible for a re-employment bonus equivalent to three-months (90-days) UI benefits can increase the monthly re-employment hazard by 1.4 percentage points. In Column (2), we use a quadratic function on either side of the cutoff to control the age profile of the re-employment hazard, and we find that the estimate increases substantially to 5.2 percentage points. However, all quadratic terms are not significantly different from zero. In addition, both AIC and BIC suggest that the quadratic specification is dominated by a linear specification. Thus, we use the linear function as our main specification. Column (3) includes covariates (e.g. gender, birth place, monthly insured salary of previous job, industry of previous job, and number of previous UI spells) in a linear specification. We find the estimate is quite similar, suggesting the estimated effect is not driven by the observed difference between eligible and ineligible workers. Column (4) reports a bias-corrected estimate, using a local linear regression with a triangular kernel, and its standard error is adjusted for bias correction. The selection of an optimal bandwidth is based on an algorithm proposed by Calonico et al. (2014). We find the local linear estimate is still statistically significant but a bit higher (4.1 percentage points).

The estimates in columns (1) to (4) are based on the RK model, which assumes that the there is no kinked relationship between the UI starting date and the monthly re-employment hazard before introducing a re-employment bonus. To examine this assumption, column (5) includes workers who were unemployed before the reform (i.e. those who started their UI spell during January 2001 to June 2002), since no worker was eligible for the re-employment bonus during this period. In practice, we implement the RK-DID model (i.e. equation (7)) by subtracting from the RK estimates that captured the kinked relationship in the UI starting date and the re-employment hazard before the reform. The estimated coefficient on $PostRB_i \times Kink_i \times (t - c_k)$ suggests that a re-employment bonus equivalent to three-months (90-days) UI benefits can increase the monthly re-employment hazard by 1.8 percentage points. Compared to the baseline mean (around 4.3%), this represents a 42% increase in the average monthly re-employment hazard. In our 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 the re-employment bonus. Table 6 displays the results, using different specifications and sample criteria. In columns (1) and (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.[19] The estimated effect of the re-employment bonus based on this sample increases slightly to 3 percentage points. Again, following Card et al. (2007) and Nekoei and Weber (2019), column (2) displays the estimate using the sample excluding temporary lay-offs. 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 the UI starting date. We find the standard error of the estimated coefficient on $PostRB_i \times Kink_i \times (t - c_k)$ changes slightly due to this setting. In the later analysis, we combine the main estimate for the effect of re-employment bonus with our estimate of the extended benefits effect, in order to conduct a welfare analysis of the UI extension. An important caveat to our analysis is that we use an early-2000s sample to estimate the re-employment bonus effect. Following Schmieder et al. (2012), we use the typical procedure to re-weight the re-employment bonus sample to match the distribution of observable characteristics in the extended benefits sample.[20] The re-weighting estimate is shown in Column (4) and makes sure our main result is robust to changes in observable characteristics of UI recipients over time. Column (5) displays the estimates for the second kink. Similarly, we divide γ_1^{PostRB} by -0.5 and then multiply it by 90 (i.e. $\gamma_1^{PostRB} \times -180$) to represent the effect of a re-employment bonus equivalent to three-months (90-days) UI benefits on the monthly re-employment hazard. We find that the estimate of 1.6 percentage points is similar to the first kink.

Next, we examine the robustness of our main estimate to bandwidth choices. Figure C1 of the

[19] Similar to the analysis for extended UI benefits, we define such a sample by using those recipients whose maximum gap between two UI claims was within three days.

[20] We match the following observable characteristics: gender, age, birth place, monthly insured salary of previous job, industry of previous job, and job finding rate.

Online Appendix C displays the RK-DID estimates over various bandwidth choices (from 100 to 240 days before and after July 1). The result in Figure C1 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 manipulate their UI starting date, to make it eligible for a re-employment bonus. Figure C2 in the Online Appendix C 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 C1 of the Online Appendix C, we use observable characteristics as dependent variables in the RK-DID model (i.e. equation (7)) to examine any change in sample composition after re-employment bonus reform. We find majority of RK-DID estimates are either insignificant or small but some are statistically significant and large, such as share of UI recipients working in manufacturing sector previously. We do not think it is a big concern for our main result due to the following reasons: First, our main specification has already controlled these covariates; Second, similar to the last section, we also predict a worker's re-employment hazard based on his/her observable characteristics and examine the effect of re-employment bonus reform on this predicted re-employment hazard, using equation (7). In contrast to our main result, estimated coefficient in Column (7) is very small and insignificantly different from zero.

7 Welfare Implications

7.1 Estimated Liquidity-to-Moral Hazard Ratio: Increasing Benefit Level

Since most existing studies focus on the liquidity-to-moral hazard ratio (i.e. the consumption-smoothing gain) of increasing benefit levels, to our knowledge there is no corresponding estimate in the literature for extending potential benefit duration. In order to make a comparison with the previous literature, we provide an estimate of the liquidity-to-moral hazard ratio of increasing the benefit level, before introducing our estimate for extending UI benefits.

Although we have no variations in the benefit level, following the arguments in Landais (2015), the liquidity effect of a 50% increase in potential duration (i.e. an increase from six months to nine

months) on search effort in the regular benefit period is equivalent to that of a 50% increase in the benefit level if workers are not totally liquidity-constrained. In the dynamic labor supply model (MaCurdy, 1981), the wealth effect of a future wage increase is the same as that of a current wage increase. Similarly, the liquidity effect of a future benefit increase is the same as the liquidity effect of a current benefit increase. Therefore, we can back out the liquidity effect of increasing the benefit level, using the decomposition of search effort responses to extending UI benefits during a regular benefit period (i.e. evaluating equation (2) at period $t < P$) and the reduced-form estimates in Panel B of Table 3 and Table 5.

According to Table 3, a three-month increase in potential benefit duration $\frac{\partial s_t}{\partial P}$ is estimated to decrease the monthly re-employment hazard during the regular benefit period (i.e. the 1st and 6th months of the unemployment spell) by 1.3 percentage points, which is a combination of a liquidity effect and a moral hazard effect. Based on Table 5, eligibility for the three-month benefit as a re-employment bonus is estimated to increase the monthly re-employment hazard by 1.8 percentage points (i.e. moral hazard effect, $b\frac{\partial s_t}{\partial r_t}$). Plugging the reduced-form estimates into equation (2), we get

$$\begin{aligned} b\frac{\partial s_t}{\partial A_t} &= \frac{\partial s_t}{\partial P} + 0.5 S_{t+1}(P) b \frac{\partial s_t}{\partial r_t} \\ &= -0.013 + 0.5 \cdot 0.52 \cdot 0.018 \\ &= -0.0083. \end{aligned}$$

Since the UI exhaustion rate $S_{t+1}(P)$, the percentage of UI recipients exhausting their regular (six-month) benefits, is 0.52 for UI-recipients aged 43-44, and the re-employment bonus counteracts the moral hazard effect of extended benefits by 50%, the moral hazard effect of extended benefits becomes $0.5 \cdot 0.52 \cdot 0.018 \approx 0.0047$. Subtracting the moral hazard effect from the total effect, the

estimated liquidity effect in the regular benefit period ($b\frac{\partial s_t}{\partial A_t}$) is -0.0083.

$$\begin{aligned} R_{t,t<P} &= -b\frac{\partial s_t}{\partial A_t}/b\frac{\partial s_t}{\partial r_t} \\ &= 0.0083/0.018 \\ &= 0.46 \end{aligned}$$

Therefore, the estimated liquidity-to-moral hazard ratio of increasing benefit level (i.e. $t < P$) is about 0.46, which is somehow smaller than existing estimates from Chetty (2008) and Landais (2015) on the liquidity-to-moral hazard ratio of increasing benefit level (i.e. 0.88-1.5). The relatively low liquidity-to-moral hazard ratio of increasing UI benefit level in Taiwan is consistent with the fact that the saving rate in Taiwan (32% in 2012) is higher than in the US (18% in 2012), thus increasing individuals' ability to smooth consumption during unemployment.

7.2 Estimated Liquidity-to-Moral Hazard Ratio: Extending Potential Benefit Duration

Next, we decompose the effect of extended UI benefits to a liquidity effect and a moral hazard effect by evaluating equation (2) at period P.[21]

$$\frac{\partial s_P}{\partial P} = b\frac{\partial s_P}{\partial A_P} - (1-\theta)b\frac{\partial s_P}{\partial r_P} \tag{8}$$

Note that we do not have the estimate of the moral hazard effect for the extended benefit period ($\frac{\partial s_P}{\partial r_P}$), since individuals were not eligible for extended UI benefits before 2009. To circumvent this issue, we assume that the re-employment response to the re-employment bonus during the extended benefit period is comparable to that during the regular benefit period (i.e. $\frac{\partial s_P}{\partial r_P} = \frac{\partial s_t}{\partial r_t}$). Two recent studies (Kolsrud et al., 2018; Ahn, 2018) suggest that the moral hazard effect could decline over the unemployment spell.[22] Based on their results, we are likely to underestimate the consumption

[21] The survival rate in period P is conditional on being unemployed in period t and is equal to one (i.e. $S_{t+1}(P) = 1$).

[22] Kolsrud et al. (2018) estimate that the duration response to unemployment benefits decreases over the spell, while consumption smoothing benefits increase, thus implying that the moral hazard effect declines over the spell. On the other hand, Ahn (2018) estimates that the effect of an increase in bonuses for older workers on the job-finding hazard declines over the unemployment spell, thereby providing some evidence that the moral hazard is stronger earlier in the spell.

smoothing gain from extending UI benefits. Therefore, our estimated welfare gain from extending UI benefits could serve a lower bound of the true welfare gain.

The estimates in Table 3 suggest that a three-month increase in potential benefit duration decreases the monthly re-employment hazard during the extended benefit period ($\frac{\partial s_P}{\partial P}$) by 8 percentage points. Assume the moral hazard effect is constant over the unemployment period so that we can use the estimate in Table 5. Plugging these estimates into equation (8) implies the estimated liquidity effect in the extended benefit period is -0.071.[23] Therefore, the estimated liquidity to the moral hazard ratio of extending UI benefits is

$$\begin{aligned} R_P &= -b\frac{\partial s_P}{\partial A_P}/b\frac{\partial s_P}{\partial r_P} \\ &= 0.071/0.018 \\ &= 3.9. \end{aligned}$$

Compared to the ratio of increasing benefit levels (0.46) found in section 7.1, the estimated ratio of 3.9 for extending potential benefits duration is quite large. This finding is consistent with Ganong and Noel (2019)'s results suggesting the consumption-smoothing gain of extending potential benefits duration is substantially greater than that of increasing benefit level.[24]

7.3 Welfare Effect of Extending Potential Benefit Duration

Finally, we analyze the welfare effect of the UI extension.[25] Suppose a social planner chooses the potential benefit duration (P) to maximize the expected utility of a job loser W_0 at the beginning

[23] Plugging the reduced-form estimates into equation (8), we get

$$\begin{aligned} b\frac{\partial s_P}{\partial A_P} &= \frac{\partial s_P}{\partial P} + 0.5b\frac{\partial s_P}{\partial r_P} \\ &= -0.080 + 0.5 \cdot 0.018 \\ &= -0.071. \end{aligned}$$

[24] Ganong and Noel (2019) found that the welfare gain from extending potential benefits duration is about four times greater than from increasing benefits level. Note that the welfare gain they estimated is for all UI recipients. However, our estimated liquidity-to-moral hazard ratio represent the welfare gain for the UI recipients who do not exhaust their benefits at the period t. Therefore, we need to consider exhausting rate of regular UI benefits when calculating welfare gain of extending benefit duration for all UI recipients. After taking this issue into account, we find that the welfare gain of extending UI benefits is about five times larger than that of increasing benefit levels.

[25] A detailed derivation can be found in the Online Appendix D.

of period 0, subject to the worker's optimal search behavior and government budget constraints, $Bb + (P - B)\theta b = (T - D)\tau$, where B and D are benefit duration and non-employment duration, respectively. In the Online Appendix D, we show that the welfare effect of a balanced-budget increase in unemployment benefits at period P (normalized by $u'(c_P^e)$, the marginal utility of consumption at period P if individuals were employed) is:

$$\frac{dW_0}{dP}/u'(c_P^e) = bS_0(P)R_P - b\{(1-\theta)[\frac{dB}{dP} - S_0(P)] + \frac{(1-\theta)B + \theta P}{T - D} \cdot \frac{dD}{dP}\}, \tag{9}$$

where $S_0(P)$ is the probability of exhausting benefits for a worker without a job at the beginning of period 0. $\frac{dB}{dP}$ and $\frac{dD}{dP}$ are the effect of increasing potential benefit duration on benefit duration and non-employment duration, respectively. $b\frac{(1-\theta)B+\theta P}{T-D}$ is the tax implied by the balanced budget constraint.

The welfare impacts of the UI extension balances two forces: (1) the welfare gain from the increased ability to maintain consumption during extended benefit period (i.e. $bS_0(P)R_P$); (2) the welfare loss due to higher taxes levied on employment (i.e. $b\{(1-\theta)[\frac{dB}{dP} - S_0(P)] + \frac{(1-\theta)B+\theta P}{T-D} \cdot \frac{dD}{dP}\}$).

On the one hand, extending benefits increases utility by smoothing out income (consumption) stream between employed and unemployed statuses in the extended benefit period (i.e. R_P). Note that only people who use up regular benefit can receive extended benefit so that we need to multiply R_P by the exhaustion rate of regular benefit (i.e. $S_0(P)$) to get expected welfare gain of extending potential benefit duration. 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 UI extension leads to the fiscal externality from behavioral responses to extended benefits. Specifically, extending benefits reduces workers' search effort, thereby increasing the UI payment ($b[\frac{dB}{dP} - S_0(P)]$), decreasing bonus payment ($-\theta b[\frac{dB}{dP} - S_0(P)]$), and increasing the non-employment duration ($\frac{dD}{dP}$). As a result, the government needs to raise the taxes in a shorter employment period to finance increased expenditure caused by job search distortions ($(1-\theta)b[\frac{dB}{dP} - S_0(P)]$). Note that our welfare formula takes the re-employment bonus into account: a marginal increase in benefit duration reduces bonus payment by θb, so the reemployment bonus

reduces the behavioral cost of the UI extension.[26]

According to equation (9), in order to estimate the welfare cost of extending UI benefits, we have to calculate the following parameters: (1) $\frac{dB}{dP}$; (2) $\frac{dD}{dP}$; (3) $\frac{B}{T-D}$; (4) $\frac{P}{T-D}$. Firstly, based on the estimates from our companion paper (Huang and Yang, 2018) using the same extended benefits sample, we know a three-month (90-days) extension in UI benefits can increase benefit duration and non-employment duration by 54.55 days and 41.73 days, respectively. Therefore, we estimate that $\frac{dB}{dP}$ and $\frac{dD}{dP}$ equal $\frac{54.55}{90}$ and $\frac{41.73}{90}$, respectively. Secondly, $\frac{B}{T-D}$ is the insured unemployment rate, which we approximate by using the total number of UI recipients divided by the total number of employees paying payroll taxes in the wage records (Landais, 2015). This yields $\frac{B}{T-D} \approx 0.1$ from 2009-2011. Lastly, we can compute $\frac{P}{T-D}$ using $\frac{P}{B} \times \frac{B}{T-D}$. According to Table 1, the average benefit duration B for workers aged 43-44 is 144.56 so that $\frac{P}{B} = \frac{180}{144.56}$. Plugging these estimates into the equation (9), we get

$$\begin{aligned}
\frac{dW_0}{dP}/u'(c_P^e) &= bS_0(P)R_P - b\{(1-\theta)[\frac{dB}{dP} - S_0(P)] + \frac{(1-\theta)B + \theta P}{T-D} \cdot \frac{dD}{dP}\} \\
&= b0.52 \cdot 3.9 - b\{0.5 \cdot (\frac{54.55}{90} - 0.52) + (0.5 \cdot 0.1 + \frac{180}{144.56} \cdot 0.1 \cdot 0.5) \cdot \frac{41.73}{90}\} \\
&= 1.93b.
\end{aligned}$$

This means that increasing potential duration is estimated to enhance middle-aged UI recipients' welfare because the expected consumption-smoothing benefits are greater than the increased welfare cost arising from the lengthened benefit and non-employment durations. Specifically, we estimate that the welfare increase caused by increasing potential duration is 1.93 times larger than an equivalent increase in the consumption when employed.

8 Conclusion

This paper evaluates the welfare effects of extending UI benefits, using two policy changes in Taiwan: a benefit extension for workers aged at least 45, and the introduction of a re-employment bonus program. Extending duration of UI benefits affects workers' search efforts through two

[26] Our formula for the welfare effect of the UI extension is similar to equation (1) in Schmieder et al. (2012). We extend their formula by incorporating reemployment bonuses into the search model.

channels and with distinct welfare implications: a moral hazard effect and a liquidity effect. Our strategy exploits the fact that the re-employment bonus affects an individual's search effort only through the moral hazard effect. Using the variation in the bonus offer around the time when bonus was introduced, and age discontinuity in eligibility for extended benefits, we separately identify the moral hazard effect and the liquidity effect of extending UI benefits. We find that the liquidity-to-moral hazard ratio of extending UI benefits is about 3.9. Furthermore, our result suggests it is welfare-enhancing to increase the potential benefit duration, even after incorporating the welfare cost from fiscal externality due to job search distortions.

Note that our welfare calculation assumes a flat labor demand, and every unemployed worker is eligible for UI benefits. However, in the search model including the reservation wage, when the government increases the generosity of its UI benefits, workers will increase their selectivity, so the firms might be less willing to open up vacancies (i.e. wage externality). On the other hand, since workers eligible for more generous benefits will decrease their search efforts, those who are ineligible will have a better chance of being employed and will be more willing to put effort into searching for a job (i.e. congestion externality). Recent evidence from Lalive et al. (2015) suggests the congestion externality dominates the wage externality, implying the optimal potential benefit duration should be longer if the macroeconomic externalities of extended benefits are taken into account.

Finally, this paper can be extended to study the design of an optimal age-dependent UI. Intuitively, an optimal UI should consider heterogeneity in consumption-smoothing benefits and in moral hazard costs, thereby 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 and such they will want a job that builds 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 ex-

tended benefits and re-employment bonuses vary over ages will be useful for optimal UI designs over a life cycle.

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Tables

Table 1: Descriptive Statistics for Extended Benefits Sample

	2009-2011 sample		2006-2008 Sample	
	Age 43-44	Age 45-46	Age 43-44	Age 45-46
Re-employment hazard (1st to 6th month)	0.11	0.09	0.09	0.08
	(0.31)	(0.28)	(0.29)	(0.28)
Re-employment hazard (7th to 9th month)	0.13	0.07	0.11	0.10
	(0.34)	(0.25)	(0.31)	(0.30)
Benefit duration	144.56	204.68	143.01	152.18
	(58.54)	(88.84)	(56.73)	(59.07)
Age	44.00	45.97	44.01	45.99
	(0.57)	(0.59)	(0.57)	(0.58)
Male	0.51	0.49	0.54	0.54
	(0.50)	(0.50)	(0.50)	(0.50)
Born in Taipei	0.13	0.13	0.14	0.14
	(0.34)	(0.34)	(0.35)	(0.34)
Work in manufacturing previously	0.31	0.33	0.39	0.44
	(0.46)	(0.47)	(0.49)	(0.50)
Monthly insured salary of previous job	30,724.22	30,825.05	30,437.92	30,638.09
	(10,577.20)	(10,544.97)	(10,358.22)	(10,270.16)
Temporary lay-off	0.11	0.12	0.09	0.10
	(0.32)	(0.33)	(0.29)	(0.30)
Number of previous UI spell	1.24	1.20	1.21	1.19
	(0.51)	(0.50)	(0.44)	(0.43)
Number of recipients	8,258	8,575	6,372	6,128
Number of observations	48,402	54,534	39,001	39,203

Notes: Data are from 2006-2011 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 the age of 45. This table displays the means and standard deviations of our outcome variable and related individual characteristics for the extended benefits sample (2009-2011 sample) and corresponding placebo sample (2006-2008). re-employment hazard is the average monthly hazard rate during the 1st to 6th or 7th to 9th month of UI spells. Standard deviations are in parentheses.

Table 2: Descriptive Statistics for Re-employment Bonus Sample

	2002-2003 Sample		2001-2002 Sample	
	Before July 1	After July 1	Before July 1	After July 1
Re-employment hazard (1st to 6th month)	0.045	0.059	0.045	0.048
	(0.21)	(0.24)	(0.21)	(0.21)
Benefit duration	164.07	156.05	164.27	166.55
	(39.41)	(45.89)	(40.77)	(37.20)
Age	44.85	44.83	45.00	44.77
	(2.86)	(2.86)	(2.88)	(2.86)
Male	0.50	0.53	0.53	0.51
	(0.50)	(0.50)	(0.50)	(0.50)
Born in Taipei	0.11	0.12	0.09	0.09
	(0.32)	(0.32)	(0.28)	(0.28)
Work in manufacturing previously	0.49	0.48	0.62	0.62
	(0.50)	(0.50)	(0.49)	(0.49)
Monthly insured salary of previous job	28,691.33	27,938.97	27,245.51	27,624.68
	(10,242.91)	(10,003.98)	(9,820.85)	(9,868.52)
Temporary lay-off	0.15	0.13	0.14	0.13
	(0.36)	(0.34)	(0.35)	(0.34)
Number of previous UI spell	1.01	1.03	1.01	1.01
	(0.12)	(0.17)	(0.11)	(0.12)
Number of recipients	9,871	7,605	7,410	11,184
Number of observations	52,664	39,249	39,296	59,290

Notes: Data are from 2001-2003 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 re-employment bonus sample (2002-2003 sample) and corresponding placebo sample (2001-2002 sample). Re-employment hazard is the average monthly hazard rate during the 1st to 6th month of UI spells. Standard deviations are in parentheses.

Table 3: The Effect of Extended Benefits on Monthly Re-employment Hazard

	(1)	(2)	(3)	(4)	(5)
Panel A: *Extended benefit period*					
β^{Age45}	-0.077***	-0.068***	-0.078***	-0.071***	
	(0.008)	(0.011)	(0.008)	(0.006)	
β^{Post45}					-0.080***
					(0.012)
Baseline mean			0.132		
Sample size	25,683	25,683	25,683	58,625	45,042
Panel B: *Regular benefit period*					
β^{Age45}	-0.014***	-0.019***	-0.014***	-0.015***	
	(0.004)	(0.006)	(0.004)	(0.004)	
β^{Post45}					-0.013**
					(0.006)
Baseline mean			0.106		
Sample size	77,253	77,253	77,253	149,234	136,044
RD	Yes	Yes	Yes	Yes	–
RD+DID	–	–	–	–	Yes
Covariates	–	–	Yes	–	Yes
Poly. model	linear	quadratic	linear	linear	linear
Bandwidth (days)	730	730	730	CCT	730

Notes: The first four columns display the estimated coefficients β^{Age45} on $Age45$ using equation (3). The outcome variable is monthly re-employment hazard in the 7th to 9th of a UI spell (Panel A) or that in the 1st to 6th of a UI spell (Panel B). Column (1) displays a basic RD estimate using a linear function to control age profile of re-employment hazard. Column (2) uses a quadratic function on either side of the cutoff to control age profile of re-employment hazard. Column (3) includes covariates, such as gender, birth place, monthly insured salary of previous job, industry of previous job, and number of previous UI spell in a linear specification. 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 β^{Post45} on $PostEB \times Age45$ using equation (4). 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 Re-employment Hazard: Robustness Checks

	(1) Continuous UI	(2) Non-temporary lay-off	(3) Clustered at Age	(4) Lower bound	(5) Upper bound	(6) Mass lay-off
Panel A: *Extended benefit period*						
β^{Post45}	-0.087***	-0.095***	-0.080***	-0.060***	-0.080***	-0.076***
	(0.022)	(0.014)	(0.011)	(0.011)	(0.012)	(0.024)
Sample size	12,851	37,448	45,042	45,042	45,042	10,620
Panel B:*Regular benefit period*						
β^{Post45}	-0.018	-0.013*	-0.013*	-0.003	-0.031***	-0.015
	(0.013)	(0.007)	(0.006)	(0.006)	(0.006)	(0.013)
Sample size	43,851	118,909	136,044	136,044	136,045	32,605

Notes: This table displays the estimated coefficients β^{Post45} on $PostEB \times Age45$ in equation (4) using different sample criteria and specifications. The outcome variable is monthly re-employment hazard in the 7th to 9th of a UI spell (Panel A) or that in the 1st to 6th of a UI spell (Panel B). All columns use a linear function to control age profile of re-employment 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 lay-offs. 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 lay-offs. 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 Re-employment Bonus on Monthly Re-employment Hazard

	(1)	(2)	(3)	(4)	(5)
$\gamma_1^{Kink} \times 180$	0.014***	0.052**	0.015***	0.041**	
	(0.005)	(0.021)	(0.005)	(0.017)	
$\gamma_1^{PostRB} \times 180$					0.018***
					(0.007)
Baseline mean			0.043		
Sample size	91,913	91,913	91,913	61,354	190,499
RK	Yes	Yes	Yes	Yes	–
RK+DID	–	–	–	–	Yes
Covariates	–	–	Yes	–	Yes
Poly. model	linear	quadratic	linear	linear	linear
Bandwidth (days)	180	180	180	CCT	180

Notes: The first four columns display the estimated coefficients γ_1^{Kink} on $Kink \times (t - c_k)$ in equation (6) and multiply them by 180 to give effects of eligibility for re-employment bonus that is equivalent to three-month UI benefits. The outcome variable is monthly re-employment hazard in the in the 1st to 6th of a UI spell. Column (1) displays a basic RK estimate using a linear function to control the effect of UI starting date on re-employment hazard. Column (2) uses a quadratic function on either side of the cutoff to control the effect of UI starting date on re-employment hazard. Column (3) further includes covariates, such as gender, birth place, monthly insured salary of previous job, industry of previous job, and number of previous UI spell in a linear specification. 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 γ_1^{PostRB} on $PostRB_i \times Kink_i \times (t - c_k)$ in equation (7) and multiply it by 180 to give effects of eligibility for re-employment 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.

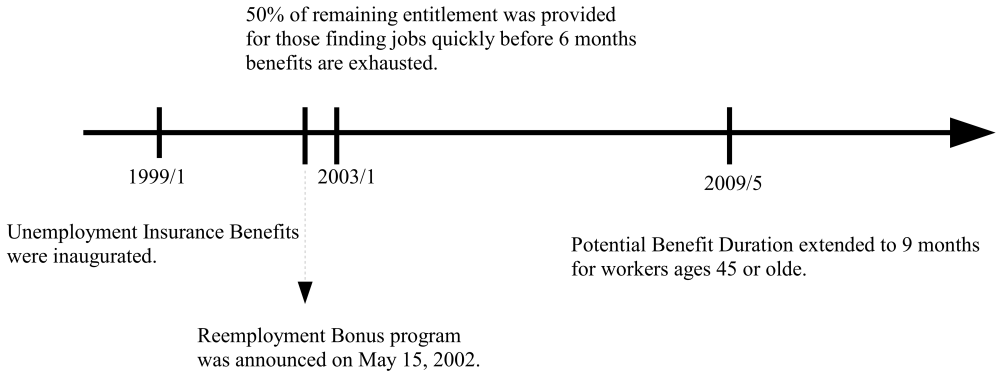
Table 6: The Effect of Re-employment Bonus on Monthly Re-employment Hazard: Robustness Checks

	(1) Continuous UI	(2) Non-temporary lay-off	(3) Clustered at UI starting date	(4) Reweighting	(5) Kink 2
$\gamma_1^{PostRB} \times 180$	0.030***	0.017**	0.018**	0.021*	
	(0.010)	(0.008)	(0.009)	(0.012)	
$\gamma_1^{PostRB} \times (-180)$					0.016**
					(0.008)
Baseline mean			0.043		
Sample size	106,418	162,675	190,499	186,100	190,495
RK+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 γ_1^{PostRB} on $PostRB \times Kink \times (t - c_k)$ in equation (7) and multiply them by 180 to give effects of eligibility for re-employment bonus that is equivalent to three-month UI benefits. The outcome variable is monthly re-employment hazard in the in the 1st to 6th of a UI spell. All columns use a linear function to control age profile of re-employment 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 lay-offs. Column (3) adjusts for potential correlation in errors within age group by clustering the standard errors by UI starting date. Column (4) reports reweighting estimate based on the reweighting sample that match the observable characteristics of extended benefits sample. Column (5) displays the estimate for the second kink. 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

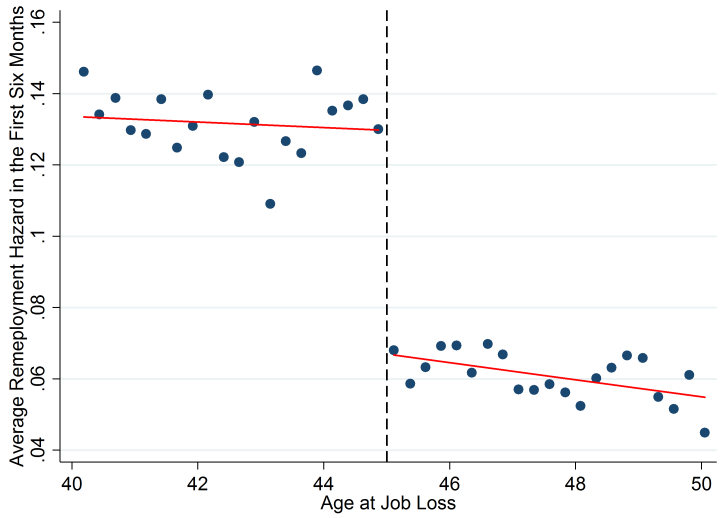
Figure 1: UI Policy in Taiwan



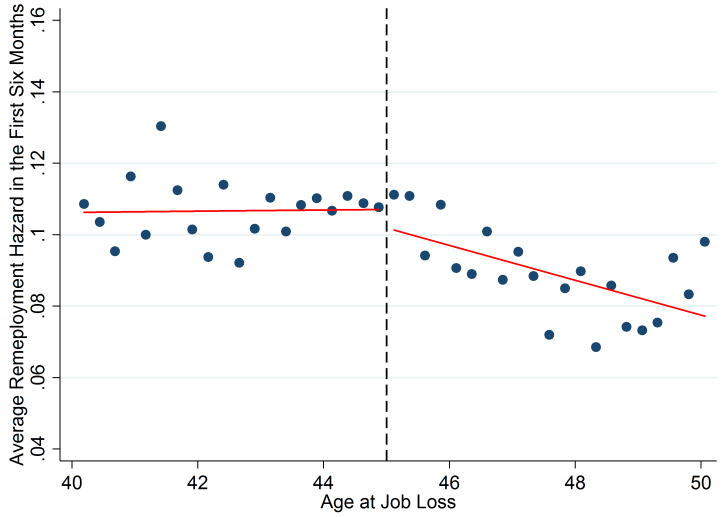
Notes: This figure summarizes the evolution of Taiwan's UI policy. UI in Taiwan was inaugurated in Jan 1999. On May 15, 2002, the re-employment bonus program was announced. On January 1, 2003, a bonus, equal to 50% of remaining benefits, began to offer for UI recipients who find jobs before exhausting 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: Extended Benefit Period

(a) Monthly Re-employment Hazard in the 7th to 9th of a UI spell: 2009-2011



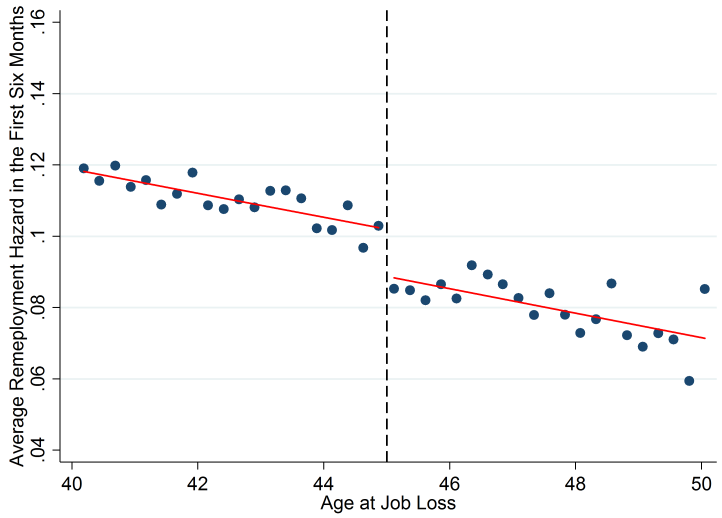
(b) Monthly Re-employment Hazard in the 7th to 9th of a UI spell: 2006-2008



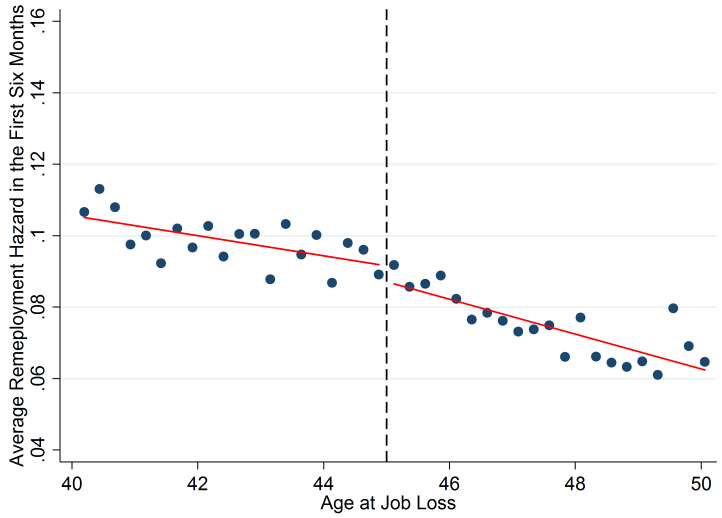
Notes: Data are from 2006-2011 unemployment benefits files and the employment insurance enrollee file. Figure 2a displays result for 2009-2011 sample and figure 2b is for 2006-2008 sample. Both figures plot the average monthly re-employment hazard in the 7th to 9th of a UI spell for UI recipients aged 40 to 50 at job loss. Each bin represents the average monthly re-employment 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: Regular Benefit Period

(a) Monthly Re-employment Hazard in the 1st to 6th of a UI spell: 2009-2011

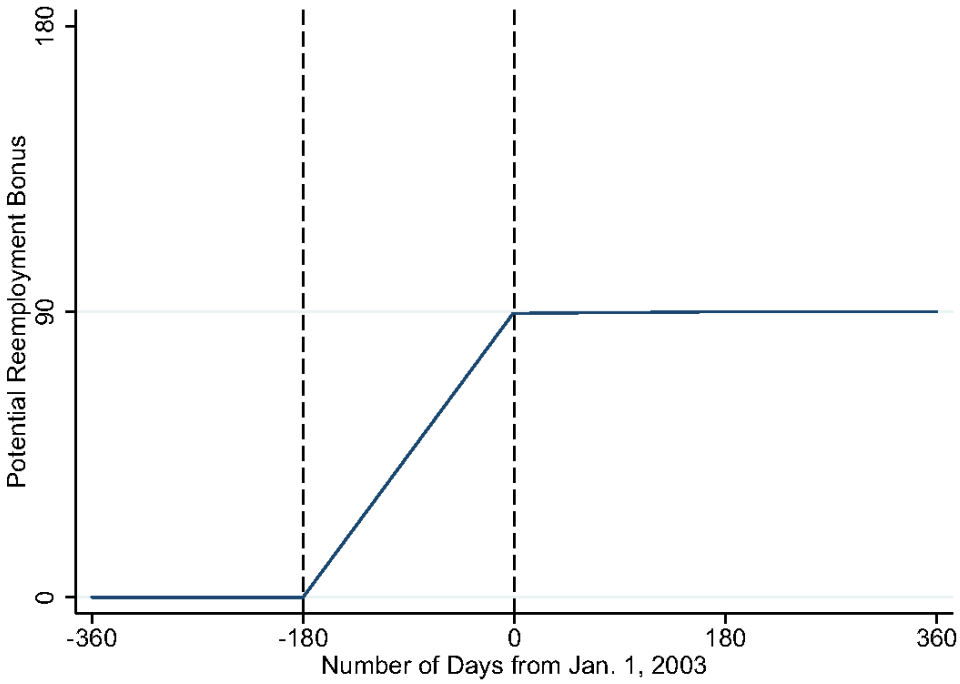


(b) Monthly Re-employment Hazard in the 1st to 6th of a UI spell: 2006-2008



Notes: Data are from 2006-2011 unemployment benefits files and the employment insurance enrollee file. Figure 3a displays result for 2009-2011 sample and figure 3b is for 2006-2008 sample. Both figures plot the average monthly re-employment hazard in the 1st to 6th of a UI spell for UI recipients aged 40 to 50 at job loss. Each bin represents the average monthly re-employment hazard within 90 days (3 months) interval. The solid lines are fitted values from a linear regression on either side of the cutoff.

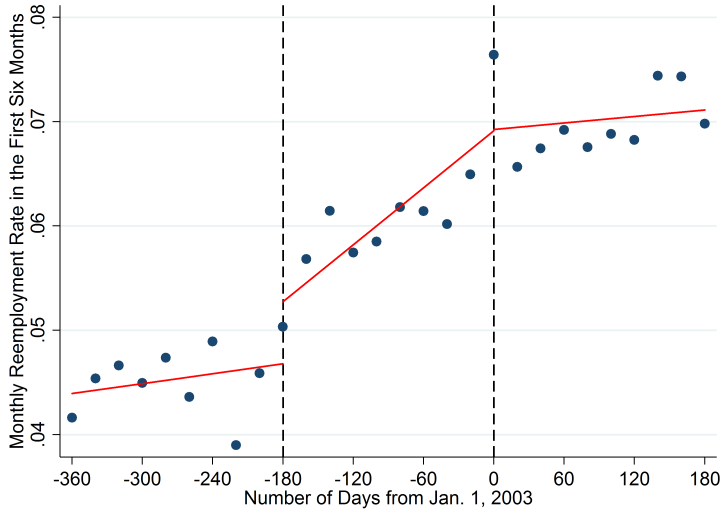
Figure 4: Potential Re-employment Bonus and the UI Starting Date



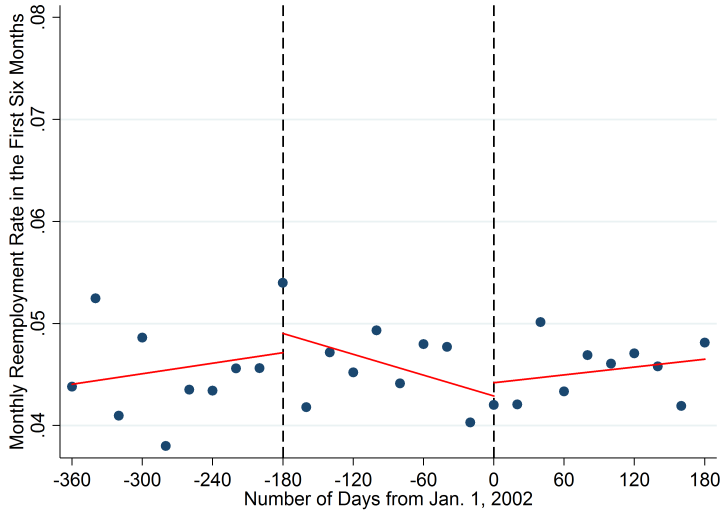
Notes: This figure demonstrates the relationship between the potential re-employment bonus (measured in number of days of UI benefits) and the date UI spells started. UI recipients starting receiving benefits before July 1, 2002 are not eligible for any re-employment bonus. As the program phased in, UI recipients are potentially eligible for a more generous bonus offer, while the potential re-employment bonus is constant for UI recipients start receiving benefits after January 1, 2003.

Figure 5: Effects of Re-employment Bonus

(a) Average Monthly Re-employment Hazard in the 1st to 6th of a UI spell: 2002-2003



(b) Average Monthly Re-employment Hazard in the 1st to 6th of a UI spell: 2001-2002



Notes: Figure 5a plots the average monthly re-employment hazard over the number of days between January 1, 2003 and the date UI spells started. The sample includes every UI spell started within 360 days from January 1, 2003. Each bin represents the average monthly re-employment 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 re-employment 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 "placebo" bonus program began. The second line indicates January 1, 2002.

Online Appendix: For Online Publication

A Decomposition of the Effect of Extended Benefits

In this section, we provide detailed derivations for equation (2). We first derive the intertemporal first order conditions as Landais (2015). Then, we combine the intertemporal and intratemporal first order conditions to decompose the effect of extending potential benefit duration to a liquidity effect and a moral hazard effect.

In our model, since individuals 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, individuals set consumption at time t equal to after tax wage rate. Formally, we can write the intertemporal first order condition when employed as follows:

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

Similarly, if individuals 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

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

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

$$u'(c_t^e) = u'(c_P^e); \tag{A.1}$$

$$u'(c_t^u) = [1 - S_{t+1}(P)]u'(c_P^e) + S_{t+1}(P)u'(c_P^u). \tag{A.2}$$

Recall 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.

$$g'(s_t) = V_t(A_t) - U_t(A_t). \tag{A.3}$$

The effect of extended 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.

$$g''(s_t)\frac{\partial s_t}{\partial P} = \frac{\partial V_t(A_t)}{\partial P} - \frac{\partial U_t(A_t)}{\partial P}. \tag{A.4}$$

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

$$\frac{\partial V_t(A_t)}{\partial P} = b\theta u'(c_t^e). \tag{A.5}$$

An increase in potential benefit duration P increases the value of unemployment in period t through two channels. On the one hand, the value of being unemployed in period t increases since individuals are eligible for extended benefits in period P. On the other hand, the value of being employed in period t also increases because individuals might get higher reemployment bonuses.

$$\frac{\partial U_t(A_t)}{\partial P} = (1-s_{t+1})..(1-s_P)bu'(c_P^u) + s_{t+1}b\theta u'(c_{t+1}^e) + .. + (1-s_{t+1})..s_P b\theta u'(c_P).$$

Assuming the liquidity constraint is not yet binding at time $P-1$, the above equation can be simplified using equation (A.1) and (A.2),

$$\frac{\partial U_t(A_t)}{\partial P} = S_{t+1}(P)bu'(c_P^u) + [1-S_{t+1}(P)]b\theta u'(c_t^e). \tag{A.6}$$

Plugging equations (A.5) and (A.6) into equation(A.4), we can write the effect of extending potential benefit duration on search effort at time t as below

$$\begin{aligned}
\frac{\partial s_t}{\partial P} &= b\frac{\theta u'(c_t^e) - \{S_{t+1}(P)u'(c_P^u) + \theta u'(c_t^e)[1-S_{t+1}(P)]\}}{g''(s_t)} \\
&= b\frac{-S_{t+1}(P)u'(c_P^u) - \theta u'(c_t^e)S_{t+1}(P)}{g''(s_t)} \\
&= b\frac{[1-S_{t+1}(P)]u'(c_t^e) - u'(c_t^u) + \theta S_{t+1}(P)u'(c_t^e)}{g''(s_t)} \\
&= b\frac{u'(c_t^e) - u'(c_t^u)}{g''(s_t)} - b(1-\theta)S_{t+1}(P)\frac{u'(c_t^e)}{g''(s_t)}; \forall t \leq P
\end{aligned} \tag{A.7}$$

The liquidity effect and the moral hazard effect of extending potential benefit duration on search effort at time t are captured by $b\frac{\partial s_t}{\partial A_t}$ and $b\frac{\partial s_t}{\partial w_t}$, respectively. To derive $b\frac{\partial s_t}{\partial A_t}$ and $b\frac{\partial s_t}{\partial w_t}$, we differentiate

equation A.3 on A_t and w_t on both sides of the equation:

$$\begin{aligned} g''(s_t)\frac{\partial s_t}{\partial A_t} &= \frac{\partial V_t(A_t)}{\partial A_t} - \frac{\partial U_t(A_t)}{\partial A_t} \\ &= u'(c_t^e) - u'(c_t^u); \end{aligned}$$

$$\begin{aligned} g''(s_t)\frac{\partial s_t}{\partial w_t} &= \frac{\partial V_t(A_t)}{\partial w_t} - \frac{\partial U_t(A_t)}{\partial w_t} \\ &= u'(c_t^e). \end{aligned}$$

Therefore, we get $b\frac{\partial s_t}{\partial A_t}$ and $b\frac{\partial s_t}{\partial w_t}$ as below

$$b\frac{\partial s_t}{\partial A_t} = b\frac{u'(c_t^e) - u'(c_t^u)}{g''(s_t)}; \tag{A.8}$$

$$b\frac{\partial s_t}{\partial w_t} = b\frac{u'(c_t^e)}{g''(s_t)}. \tag{A.9}$$

Combining equations (A.7), (A.8), and (A.9), we get the decomposition formula in Section 3.2:

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

Further, using equations (A.8) and (A.9), we can see that the ratio of a liquidity effect and a moral hazard effect in period t identifies the consumption-smoothing benefits in period t:

$$R_t = -\frac{b\frac{\partial s_t}{\partial A_t}}{b\frac{\partial s_t}{\partial w_t}} = \frac{u'(c_t^u) - u'(c_t^e)}{u'(c_t^e)}.$$

B Robustness Checks for Regression Discontinuity Design

Table B1: Examining Smoothness of Predetermined Covariates for Extended Benefits Sample

	(1) Female	(2) Born in Taipei	(3) Manu. Sector	(4) Previous Earnings	(5) Recall Job	(6) Number of Unemployment	(7) Predicted Reemp. hazard
β^{Post45}	0.007	-0.009	-0.017	595.7	-0.007	0.034	0.002
	(0.023)	(0.016)	(0.023)	(487.3)	(0.014)	(0.022)	(0.001)
Sample size	29,350	29,357	29,357	29,357	29,357	29,357	45,042

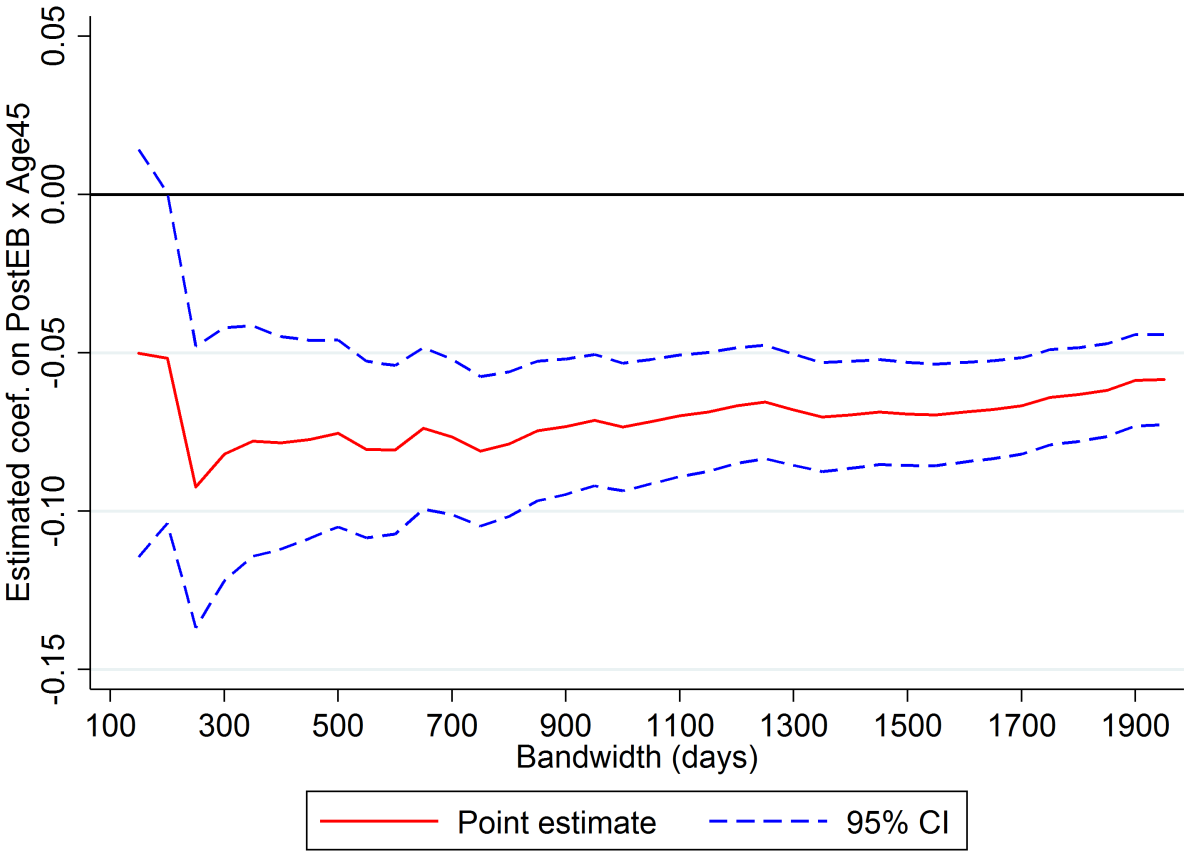
Notes: This table displays the estimated coefficients β^{Post45} on $PostEB \times Age45$ in equation (4) using different covariates as outcomes. All columns use a linear function to control age profile of re-employment hazard and the bandwidth choice is 730 days. Standard errors are in parentheses. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

Table B2: The Effect of Extended Benefits on Monthly Re-employment Hazard: Donut RD Analysis

	Monthly Re-employment Hazard							
Size of Donut around age 45 (days)	0	15	30	45	60	75	90	105
Panel A: *Extended benefit period*								
β^{Post45}	-0.080***	-0.082***	-0.091***	-0.089***	-0.069***	-0.064***	-0.080***	-0.093***
	(0.012)	(0.013)	(0.014)	(0.015)	(0.017)	(0.019)	(0.021)	(0.023)
Sample size	45,042	43,108	41,156	39,308	37,371	35,452	33,502	31,563
Panel B: *Regular benefit period*								
β^{Post45}	-0.013**	-0.011	-0.008	-0.007	-0.014	-0.012	-0.015	-0.015
	(0.006)	(0.007)	(0.008)	(0.008)	(0.009)	(0.010)	(0.012)	(0.013)
Sample size	136,044	130,238	124,459	118,765	113,074	107,291	101,531	95,684

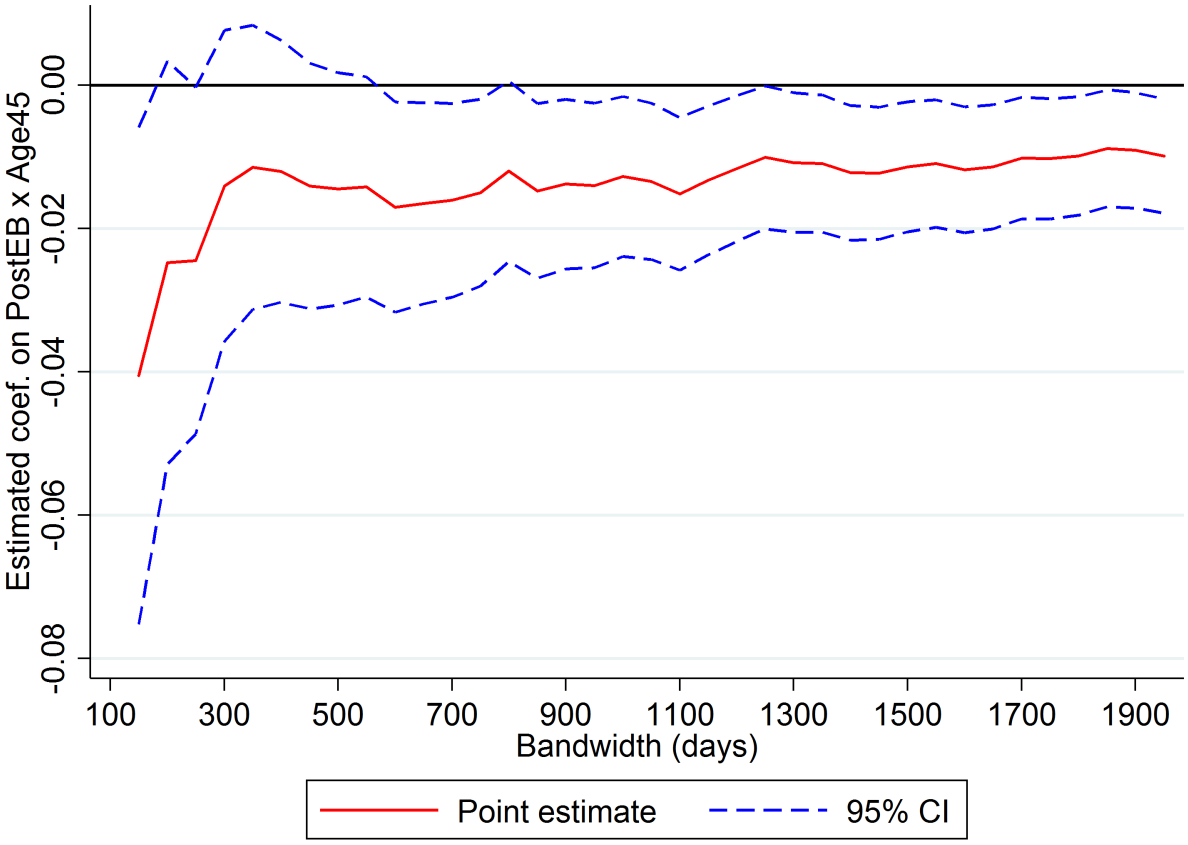
Notes: This table displays the estimated coefficients β^{Post45} on $PostEB \times Age45$ in equation (4) by excluding different number of days around age 45. The outcome variable is monthly re-employment hazard in the 7th to 9th of a UI spell (Panel A) or that in the 1st to 6th of a UI spell (Panel B). All columns use a linear function to control age profile of re-employment 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.

Figure B1: RD-DID Estimates with Varying Bandwidths: Extended Benefit Period



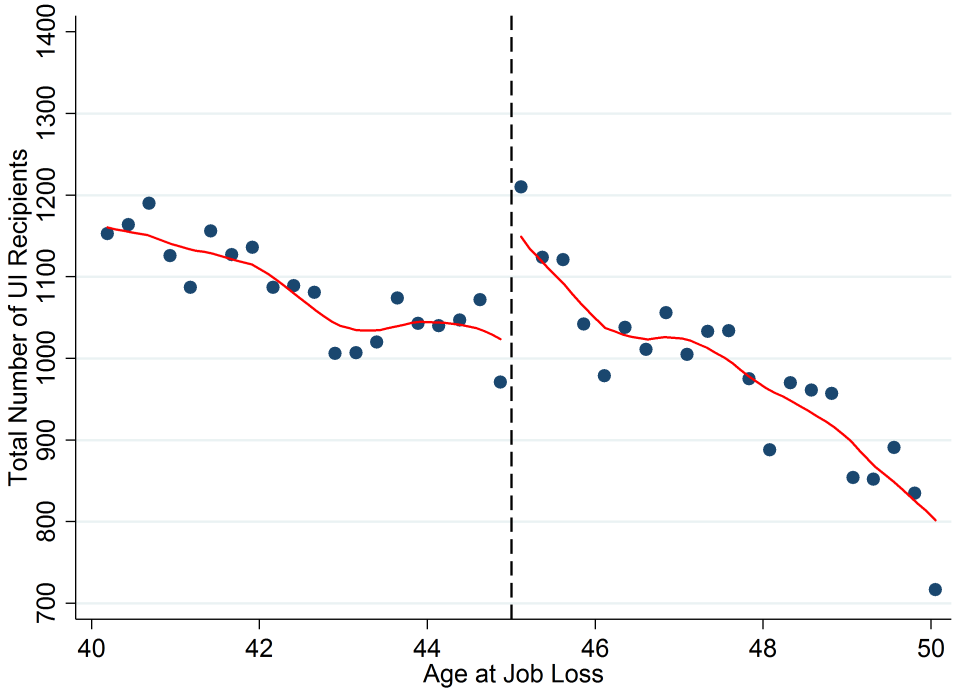
Notes: This figure plots the estimated coefficients β^{Post45} on $PostEB \times Age45$ in equation (4) using different bandwidth choices ranging from 100 to 1900 days. The solid line indicates the point estimates, and the dash lines are corresponding 95% confidence intervals.

Figure B2: RD-DID Estimates with Varying Bandwidths: Regular Benefit Period



Notes: This figure plots the estimated coefficients β^{Post45} on $PostEB \times Age45$ in equation (4) using different bandwidth choices ranging from 100 to 1900 days. The solid line indicates the point estimates, and the dash lines are corresponding 95% confidence intervals.

Figure B3: Validity of RD Design: Density Test



Notes: This figure plots the number of workers starting UI spells between May 1, 2009 and July. 31, 2011, conditional on age at job loss. Each bin corresponds to the total number of workers starting UI spells within a 90 days interval.

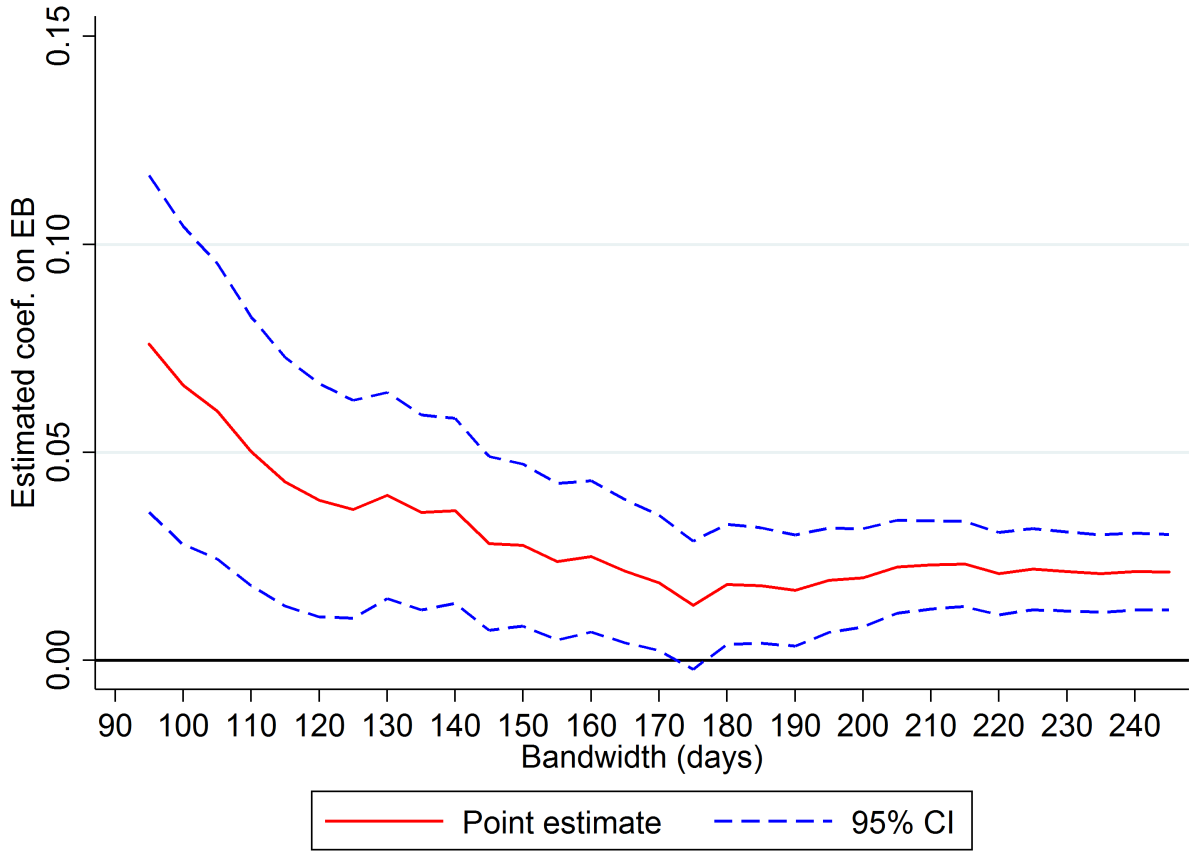
C Robustness Checks for Regression Kink Design

Table C1: Examining Smoothness of Predetermined Covariates for Re-employment Bonus Sample

	(1) Female	(2) Born in Taipei	(3) Manu. Sector	(4) Previous Earnings	(5) Recall Job	(6) Number of Unemployment	(7) Predicted Reemp. hazard
$\gamma_1^{PostRB} \times 180$	0.0783**	-0.0359	0.178***	-117.5	0.0339	-0.0186*	-0.0194
	(0.0368)	(0.0224)	(0.0363)	(735.5)	(0.0254)	(0.0105)	(0.107)
Sample size	36,076	36,076	36,076	36,076	36,076	36,076	88,954

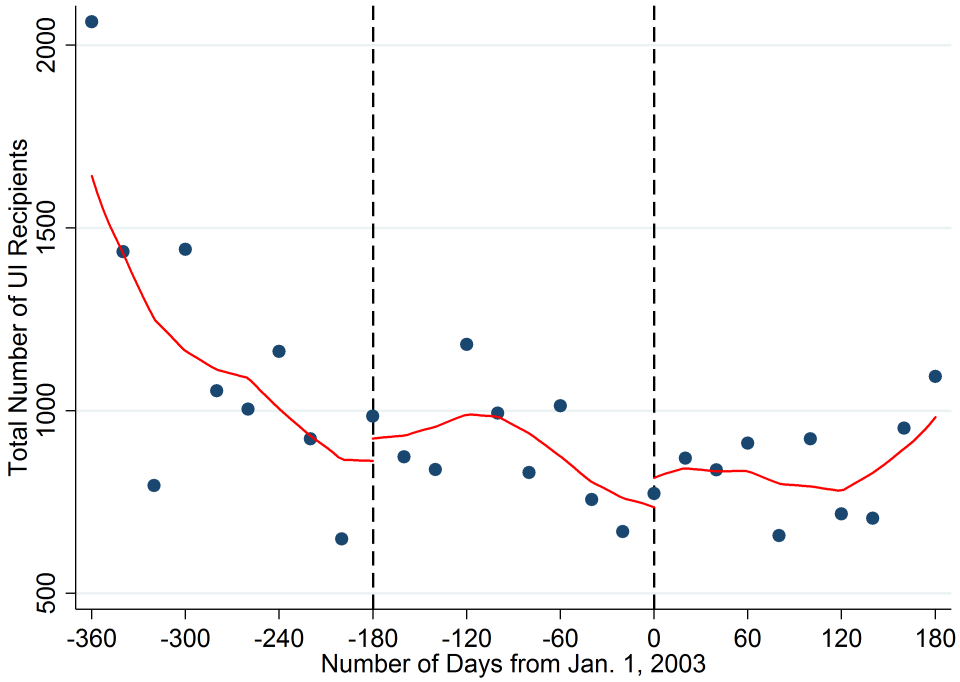
Notes: This table displays the estimated coefficients γ_1^{PostRB} on $PostRB \times Kink \times (t - c_k)$ in equation (7) and multiply it by 180 using different covariates as outcomes. All columns use a linear function to control age profile of re-employment hazard and the bandwidth choice is 180 days. 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 C1: RK-DID Estimates with Varying Bandwidths Using Kink 1



Notes: This figure plots the estimated coefficients γ_1^{PostRB} on $PostRB \times Kink$ in equation (7) and multiply it by 180 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 C2: Validity of RKD: Density Test



Notes: This figure plots the number of workers starting UI spells between January 1, 2002 and June 30, 2003, conditional on age at job loss. Each bin corresponds to the total number of workers starting UI spells within a 20 days interval.

D Welfare Effects of Extending Benefits Duration

In this section, we offer detailed derivations for equation (9). 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

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

Using envelope theorem and Euler equations, we obtain $\frac{\partial U_0}{\partial P}$, $\frac{\partial V_0}{\partial P}$, $\frac{\partial U_0}{\partial w}$, and $\frac{\partial V_0}{\partial w}$ as follows:

$$\begin{aligned}\frac{\partial U_0}{\partial P} &= bS_1(P)u'(c_p^u) + bs_1\theta u'(c_1^e) + .. + b(1 - s_1)..(1 - s_{P-1})s_P\theta u'(c_P^e) \\ &= bS_1(P)u'(c_p^u) + b[1 - S_1(P)]\theta u'(c_P^e);\end{aligned}$$

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

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

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

Note that

$$\begin{aligned}(1 - s_0)\frac{\partial U_0}{\partial w} + s_0\frac{\partial V_0}{\partial w} &= (1 - s_0)\sum_{t=1}^{T-1}[\prod_{i=1}^{t-1}(1 - s_i)]s_t(T - t)u'(c_t^e) + s_0 Tu'(c_0^e) \\ &= (T - D)u'(c_p^e),\end{aligned}$$

where the second equality uses equation (A.1). Using the results above, we obtain

$$\begin{aligned}\frac{dW_0}{dP} &= bS_0(P)u'(c_p^u) + b(1 - s_0)[1 - S_1(P)]\theta u'(c_p^e) + bs_0\theta u'(c_p^e) - (T - D)u'(c_p^e)\frac{d\tau}{dP} \\ &= bS_0(P)u'(c_p^u) + b[1 - S_0(P)]\theta u'(c_p^e) - (T - D)u'(c_p^e)\frac{d\tau}{dP} \qquad \text{(D.1)}\end{aligned}$$

An increase in P increases the tax rate by

$$\frac{d\tau}{dP} = \frac{b}{T-D}[(1-\theta)\frac{dB}{dP} + \theta + \frac{\tau}{b}\frac{dD}{dP}], \tag{D.2}$$

where

$$\frac{\tau}{b} = \frac{(1-\theta)B + \theta P}{T-D}. \tag{D.3}$$

In other words, the tax rate, τ, increases because a longer potential benefit duration raises the benefits and bonus payment in a shorter period of employment. Plugging equations (D.2) and (D.3) into equation (D.1), we obtain

$$\begin{aligned}\frac{dW_0}{dP} &= bS_0(P)u'(c_p^u) + b[1-S_0(P)]\theta u'(c_p^e) - bu'(c_p^e)[(1-\theta)\frac{dB}{dP} + \theta + \frac{(1-\theta)B+\theta P}{T-D}\frac{dD}{dP}] \\ &= bS_0(P)[u'(c_p^u) - u'(c_P^e)] - bu'(c_p^e)\{(1-\theta)[\frac{dB}{dP} - S_0(P)] + \frac{(1-\theta)B+\theta P}{T-D}\frac{dD}{dP}\}.\end{aligned}$$

Finally, normalizing the welfare effect of extending potential benefit duration by the marginal utility of consumption in period p if individuals were employed, we get the formula in Section 7.3 as below:

$$\frac{dW_0}{dP}/u'(c_P^e) = bS_0(P)R_P - b(1-\theta)[\frac{dB}{dP} - S_0(P)] + \frac{(1-\theta)B+\theta P}{T-D}\frac{dD}{dP}\},$$

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