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The welfare effects of extending unemployment benefits: Evidence from re-employment and unemployment transfers ☆

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ABSTRACT

This paper investigates the welfare effects of extending unemployment benefits by comparing the search effort responses to income transfers when employed (i.e. re-employment bonus) and unemployed (i.e. extended benefits). Specifically, we use administrative data on the universe of unemployment spells in Taiwan from 2001 to 2011 and evaluate effects of providing a re-employment bonus and extending unemployment benefits. Our results suggest that the provision of re-employment bonus increases job-finding hazards and results in the positive fiscal externality. The behavioral costs per New Taiwanese Dollar (NTD) of initial spending on bonuses is −0.61. In contrast, extending unemployment benefits reduces the rate of unemployment exit and generates the negative fiscal externality. We integrate the estimated policy effects with a search model with liquidity constraints to identify the value of extending unemployment benefits captured by the marginal rate of substitution (MRS) between consumption when unemployed and employed. We find that the estimated MRS of extending benefits is around 1.5 to 2.5 — the marginal value of transfers when unemployed is about two times larger than that when employed. Finally, the marginal value of public fund for extending UI benefits is between 1.3 and 2, suggesting the welfare gain from benefit extension is larger than its welfare cost by more than 30%.

1. Introduction

Unemployment insurance (UI) protects individuals from the risk of earnings loss during unemployment (i.e. insurance value), but it also distorts incentives to search for jobs (i.e. moral hazard). During a recession (e.g. the recent COVID-19 pandemic), extending UI benefits is one major policy tool that can be used to protect workers against adverse shocks. However, 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 (Card et al., 2007; Chetty, 2008; Landais, 2015). Empirical evidence regarding the value and welfare impacts of extending UI benefits (e.g. increase the potential duration of UI) is still scant.[1] On the other hand, to reduce the moral hazard effect of UI, some countries have offered re-employment bonuses as financial incentives to workers who find jobs quickly.[2] In fact, during the COVID-19

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[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.

[2] To our knowledge, Hungary (DellaVigna et al., 2017; Lindner and Reizer, 2019), Korea (Ahn, 2018), and Taiwan currently offer re-employment bonuses to UI recipients. Rotterdam, in the Netherlands, once offered a bonus program to long-term unemployed (Van der Klaauw and Van Ours, 2013), but it has now been removed.

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pandemic, the U.S. government has considered implementing a re-employment bonus (i.e. a back-to-work bonus), to reduce the financial burden on the UI system and get people to rejoin the workforce. While the disincentive effect of UI, measured by the elasticity of non-employment duration to UI benefits, has been estimated across a wide variety of UI contexts, the incentive effects of re-employment bonuses are less studied and still largely rely on early U.S. bonus experiments (Decker et al., 2001a; Woodbury and Spiegelman, 1987).

In this paper, we contribute to the current literature by examining the above under-studied issues. In the first part of the paper, we exploit two significant UI reforms in Taiwan, in order to investigate how workers' search efforts respond to income transfer when re-employed and unemployed. In 2003, the Taiwanese government introduced re-employment bonuses whereby people would be paid 50% of their remaining UI benefits after regaining employment. 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 offer for which workers were eligible as a function of the dates the UI spells started. Thus, we use a regression kink (RK) design herein to examine the effects of the re-employment bonus. On the other hand, after 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 effects of extended UI benefits, by comparing the outcomes of individuals just before and just after being 45 years old at the point of being laid off.

Our estimates using the RK design show that the provision of a re-employment bonus increases the monthly re-employment hazard by about 2 percentage points and significantly reduces benefits duration and non-employment duration by 6% to 9%. Moreover, we find that faster re-employment does not significantly affect the quality of a job match, such as post-unemployment wage and job tenure. Second, our RD estimates suggest that a three-month increase in potential benefit duration reduces middle-aged UI recipients' monthly re-employment hazard by 3 percentage points. The implied elasticity of non-employment duration (UI benefit duration) with respect to potential benefit duration is 0.27 (0.78). However, being eligible for longer potential benefit duration has little impact on job match quality.

In the second part of the paper, we integrate our reduced-form estimates with a search model provided by Chetty, 2008 and Landais, 2015, to conduct welfare analysis. First, we investigate the behavioral costs of one NTD's spending on the two policies. On the one hand, we find that the provision of re-employment bonus can induce the behavioral response (i.e. shortening spells of insured unemployment) that leads to a positive fiscal externality. The behavioral cost of one NTD of initial spending on the re-employment bonuses is -0.61—the behavioral response to one NTD of re-employment bonus enhances the government budget by 0.61 NTD so that only 0.39 NTD have to be raised to finance one NTD of re-employment bonus. On the other hand, our estimates suggest that extending UI benefits result in a negative fiscal externality. The behavioral cost of one NTD of spending on increasing potential benefit duration is 0.07, which is at the lower end of previous research (Schmieder and von Wachter, 2016).

Second, we exploit the two sources of income variation—re-employment bonuses and extended UI benefits—to estimate the value of UI extension. As Landais and Spinnewijn, 2021 point out, the value of UI is fully captured by the marginal rate of substitution between consumption when unemployed and employed (MRS), because the MRS describes how much consumption workers are willing to give up when employed, in order to increase one NTD of consumption when unemployed. Since workers' responses in re-employment hazard depend on their marginal utilities of consumption, the differential responses in the re-employment hazard to extended benefits and re-employment bonuses help us identify the MRS. An important issue we must address before estimating the MRS is that Taiwan's UI extension increases not only the potential benefit duration, but also the qualification period for re-employment bonuses—the UI extension increases workers' income not only when unemployed, but also when employed. To recover the effect of a pure unemployment transfer, we decompose the effect of the UI extension on the re-employment hazard into two effects: (1) the effect of an unemployment transfer and (2) the effect of a re-employment transfer. Our estimates suggest that the marginal value of an unemployment transfer during an extended benefit period is about two times larger than that of a re-employment transfer, i.e. an MRS around 1.5 to 2.5. Combining the estimated value of extended benefits with its behavioral cost, the marginal value of public fund (MVPF) (Hendren and Sprung-Keyser, 2020a) of extending UI benefits is about 1.3 to 2; thus, a welfare gain of one NTD's spending on extending potential benefit duration is about 1.3 to 2 NTD.

Our paper stands apart from the previous literature on unemployment insurance in the following ways. First, we provide one of the first pieces of evidence on the welfare gain (i.e. insurance value) 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 consumption-smoothing benefits (i.e. the welfare gains of UI) from UI is still scarce, and most existing studies focus on the welfare gain resulting from increasing the UI benefit level (Card et al., 2007; Chetty, 2008; Gruber, 1997; Landais, 2015). According to the survey by Schmieder and von Wachter, 2016, the welfare gain from increasing the benefit level is substantial—the consumption-smoothing benefits of UI range from 0.3 to 1.5, implying an MRS of between 1.3 and 2.5.[3] More recently, Landais and Spinnewijn, 2021 have shown that the marginal propensities to consume (MPC) when unemployed and employed can be used to provide a lower bound for the MRS.[4] Estimates from the MPC approach suggest a lower bound of 1.59 for the MRS. However, the welfare gain from extending the duration of UI benefits could be even greater than when increasing the UI benefit level, because extending UI benefits mainly affects 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

[3] Consumption-smoothing benefits and MRS are closely related concepts. The former represents the mark-up when workers are willing to transfer one NTD of consumption from employment to unemployment, while the MRS identifies the willingness to pay for one NTD of consumption when unemployed, one plus the mark-up—MRS is equal to one if the consumption- smoothing benefits of UI are zero. Previous literature mostly focuses on estimating consumption-smoothing benefits. Gruber, 1997 used panel data from the Panel Study of Income Dynamics and established variations in the 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 estimation of the risk aversion coefficient by using the sufficient statistic approach, whereby consumption-smoothing benefits equate 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. Overall, consumption-smoothing benefits using a consumption-based approach tend to be lower.

[4] Specifically, the ratio of the odds ratio for the MPC when unemployed to that when employed offers a lower bound for the MRS.

UI benefits affects social welfare (Schmieder and von Wachter, 2017; Kolsrud et al., 2018; Lindner and Reizer, 2019). The estimation of the consumption-smoothing benefit 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 the consumption-smoothing gain from extending the duration of UI benefits is greater than when 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. In contrast, the variation in eligibility for extended UI benefits we use in this paper is from an exogenous policy rule. Our estimates complement Ganong and Noel, 2019' estimates by comparing behavioral responses to income transfers when employed and unemployed that do not need to assume the value of the risk aversion coefficient. Moreover, our results are based on a quasi-experimental design so that selection bias is less concerned.

Second, this paper adds new evidence in terms of the effects of re-employment bonuses on unemployment duration and job match quality. Existing evidence on the effect of re-employment bonuses still relies on the U.S. field experiments conducted in the 1980s (Decker et al., 2001a; O'Leary et al., 1995; Woodbury and Spiegelman, 1987; Meyer, 1995). Although the designs of the experiments differ from each other, these experiments suggest that bonuses significantly reduce the insured duration of unemployment by about 1.5 weeks (i.e. around a 6% decrease from the baseline mean), with insignificant effects on re-employment earnings (O'Leary et al., 1995). Ahn, 2018 provides more recent evidence on this issue by using a quasi-experimental method and administrative data from South Korea.[5] He finds that an increase in the re-employment bonus for older workers can significantly reduce the duration of UI spells by 0.68 to 1.82 weeks (i.e. a 3.6% to 9.5% decrease), without affecting subsequent job match quality. Overall, our estimation results—indicating a modest effect on unemployment duration and an insignificant effect on job match quality—are consistent with the bonus literature.

Finally, to our knowledge, this paper provides the first estimates on the effects of UI using high-quality administrative data from Asian countries.[6] Most existing estimates on UI effects are in the context of U.S and European countries. According to Table 1 in Schmieder and von Wachter, 2016, the estimated elasticities of non-employment duration with respect to potential benefit duration in European countries and the United States range from 0.1 to 1, with a median around 0.37, while benefits duration elasticities range from 0.52 to 1.35, with a median of 0.58.[7] Our estimated non-employment duration elasticity of 0.27 and the benefits duration elasticity of 0.78 are therefore within the ranges of previous estimates. In addition, consistent with the previous literature, we also find that extending UI benefits has little impact on the quality of a new job.

The outline of this paper is as follows. In Section 2, we describe the Taiwanese UI system, and Section 3 describes our data and the estimation sample. In Section 4 and Section 5, we estimate the effects of the re-employment bonus and the effects of extended benefits. In Section 6, we estimate the behavioral costs of re-employment bonuses and extended UI benefits, and we evaluate the value of extending UI benefits by comparing the responses in re-employment hazard to these two income transfers. Section 7 summarizes the findings and discusses possible extensions to this paper.

2. Institutional Background

2.1. Unemployment Insurance in Taiwan

Unemployment benefits are part of the Taiwanese Employment Insurance (EI) program—a mandatory national system offering unemployment benefits, re-employment bonuses, vocational training living allowances, parental leave allowances, and national health insurance premium subsidies.[8] 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 was 43,900 NTD (1,460 US$) per month, so around 23% of the sample's monthly salaries are censored.

To be eligible for unemployment benefits, individuals who lose their jobs must have at least one year of employment history in the three years prior to the involuntary job loss.[9] 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.[10]

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 for UI benefits is 60% of the average insured salary during the six months prior to job loss for those without non-working dependants.[11] For UI recipients with one more non-

[5] The structure of Taiwan's bonus program is similar to South Korea's, albeit with three major differences in the design. First, while the length of the qualification period in Taiwan is as long as the potential benefit duration, the bonus qualification period is shorter in Korea, in that only those who find a job 30 days before exhausting their benefits are eligible for a bonus. Second, the re-employment period is longer in Korea, and workers have to find a job that lasts for at least 6 months (3 months in Taiwan) to be eligible. Third, the bonus offer for UI claimants aged 55 or older is two-thirds of their remaining entitlements, while the bonus offer in Taiwan is not a function of age.

[6] Ahn, 2018 also uses UI administrative data from an Asian country (i.e. South Korea) but he focuses on the effect of re-employment bonus.

[7] See Schmieder and von Wachter, 2016 for an excellent survey.

[8] Note that individuals aged 15 to 65 can participate in EI. Before 2009, the oldest age of eligibility for EI was 60 years-old.

[9] 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.

[10] 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, very 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.

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

Table 1
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
Monthly re-employment hazard	0.064	0.070	0.063	0.064
	(0.25)	(0.26)	(0.24)	(0.24)
Benefit duration	156.84	150.40	158.26	160.71
	(45.65)	(49.86)	(46.57)	(43.37)
Non-employment duration	371.53	339.31	378.17	360.96
	(273.58)	(265.12)	(278.32)	(271.68)
Age	38.61	38.69	39.41	38.49
	(8.60)	(8.48)	(8.36)	(8.44)
Male	0.51	0.55	0.55	0.54
	(0.50)	(0.50)	(0.50)	(0.50)
Born in Taipei	0.11	0.12	0.09	0.10
	(0.32)	(0.33)	(0.28)	(0.30)
Monthly previous salary	28,813.86	28,214.81	27,178.74	27,880.62
	(9,636.09)	(9,454.81)	(9,211.21)	(9,270.46)
Work in manufacturing previously	0.44	0.42	0.55	0.55
	(0.50)	(0.49)	(0.50)	(0.50)
Temporary lay-off	0.05	0.04	0.04	0.04
	(0.21)	(0.20)	(0.19)	(0.20)
Number of previous UI spell	0.01	0.01	0.01	0.01
	(0.09)	(0.12)	(0.08)	(0.08)
Job tenure of the previous employment (days)	793.72	853.14	602.15	702.14
	(439.46)	(496.56)	(339.80)	(369.78)
Number of recipients	54,490	38,487	40,183	63,764
Number of observations	277,619	193,457	203,042	324,257

Notes: Data are from 2001–2003 unemployment benefits files and the employment insurance enrollee file. We focus on the UI recipients who are age 25 to 60 at job loss and lost their job 180 days before and after July 1, 2002 (or 2001). 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 unemployment spells. Except for re-employment hazard, which is at spell-month level, other variables are computed at spell level. Standard deviations are in parentheses.

working dependant, the replacement rate is increased by 10 percentage points, and it can reach as high as 80%.[12]

2.2. Re-employment Bonuses

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, whereby 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 exhausting their unemployment benefits, and who then accumulate at least three months of employment history after re-employment. In other words, Taiwan's re-employment bonus has a qualification period the same as the UI eligibility period, and a re-employment period of three months. For example, if a worker finds a job at the end of the second month of unemployment, she will receive two months of benefits as a one-off bonus after three months of 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.[13]

2.3. Benefits Extension for Older Workers

Taiwan's re-employment bonus program took effect in 2003, and the UI extension came into force on May 1, 2009.[14] As a result of these two changes, the potential benefit duration and the qualification period for bonuses increased simultaneously—Taiwan's UI extension not only extended unemployment benefits for older workers, but it also increased the bonuses for which those workers were potentially eligible. Consider two UI recipients: UI recipient 1 aged under 45 at job loss and UI recipient 2 aged 45 at job loss. Both recipients find a job at the end of the second month of their UI benefits period. UI recipient 1, aged under 45 at the time of job loss, however, is eligible for only six months of unemployment benefits and is thus eligible for an additional two months of benefits as a bonus, while UI recipient 2, aged 45 at job loss, eligible for nine months of unemployment benefits, qualifies for a bonus equivalent to an additional three and a half months of benefits. Therefore, the re-employment bonus creates a counteracting force that mitigates the moral hazard effect of the benefits extension.

3. Data and Sample

3.1. 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

[12] Unlike many European countries (e.g. Austria and Germany), Taiwan's UI program does not offer means-tested unemployment assistance after the benefits have been exhausted. However, job-losers in Taiwan are eligible for six months of vocational training subsidies, regardless of age, if they register with the employment service and participate in full-time vocational training. Like unemployment benefits, monthly training subsidies equate to 60% of the average insured wage during the six months prior to job loss. Workers are not eligible for unemployment benefits when they participate in vocational training; however, they are not prohibited from claiming unemployment benefits after completing a training program if still unemployed, or from participating in training after they have received benefits for a certain amount of time. During our study period, only 6.5% of UI benefits recipients participated in vocational training, and we find that the effects of extending UI benefits on the participation and duration of vocational training are small and insignificant.

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

[14] Additionally, the bonus program and the UI extension were approved in May 2002 and May 2008, respectively.

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) the cancellation of employment insurance (job separation) or wage changes, including the date of change, an enrollee's previous salary, and a scramble ID. We use the scramble ID to merge unemployment benefits files with employment insurance enrollee files, which means we can then use the date of new enrolments in employment insurance after job loss to represent the date of re-employment.

3.2. Sample

In order 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 eligible amount of re-employment bonus (i.e. length of bonus qualification period) and the starting date of a UI spell changes 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 1 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 the treatment group (i.e. 150 days) is shorter than that of the control group (i.e. 156 days). Similar results can be found in non-employment durations. For other characteristics, both groups are quite close. Columns (3) and (4) of Table 1 display the same comparison but using a pre-reform sample (i.e. 2001–2002 sample), namely, recipients who lost their jobs between six months (i.e. 180 days) before and after July 1, 2001.

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 aged 43 to 47 (i.e. two years before and after the age of 45).[15] The first two columns of Table 2 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 the first nine months of the unemployment spell (i.e. the regular benefit period plus the extended benefit period). Consistent with this finding, the benefit duration of the treatment group (i.e. 204 days) is much longer than that of the control group (i.e. 144 days). In addition, both groups have similar individual characteristics, except that the treatment group has a higher share of recipients working in the manufacturing sector than the control group. In Section B.4.3, we implement RD estimations for each pre-determined covariate, in order to examine formally the difference in sample composition around age 45.

[15] The original sample size is 193,263. In general, we only require sample's age at job loss has to be within two years before and after age 45. Therefore, the sample size becomes 20,483 individuals after applying this restriction.

The last two columns of Table 2 report the summary statistics of selected characteristics for the pre-reform sample (i.e. 2006–2008 sample), i.e. individuals who lost their job between May 2006 and July 2008. 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 this age (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.

4. Effects of Re-employment Bonuses

4.1. Regression Kink Design

In this section, we investigate the effect of the re-employment bonus on an individual's search effort. 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 (i.e. length of the bonus qualification period) increased as the starting date of the UI spell approached January 1, 2003.

Fig. 1 displays the relationship between an individual's potential amount of re-employment bonus, measured by the length of the bonus qualification period and the starting date of a UI claim. Three segments are distinguished by two cutoffs. The first cutoff is July 1, 2002, whereby 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 length of the bonus qualification period (i.e. six months/180 days). Note that UI recipients are paid 50% of their remaining UI benefits as a re-employment bonus. Therefore, being entitled to full re-employment bonus is equivalent to being eligible for three months' (90 days') UI benefits.[16] 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 one-day increase in the potential bonus qualification period. 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 1. The second kink is located on January 1, 2003, where the slope changes from 1 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 (e.g. re-employment hazard, benefit duration, or non-employment duration) to that in the treatment (i.e. re-employment bonus), if workers are similar around the kinks. To

[16] 180 days $\times$ 0.5 = 90 days.

Table 2
Descriptive Statistics for Extended Benefits Sample.

	2009–2011 sample		2006–2008 Sample	
	Age 43–44	Age 45–46	Age 43–44	Age 45–46
Monthly re-employment hazard	0.112	0.082	0.106	0.097
	(0.31)	(0.27)	(0.31)	(0.30)
Benefit duration	145.81	205.60	142.69	150.56
	(57.99)	(88.78)	(56.74)	(58.98)
Non-employment duration	270.77	309.11	289.09	311.66
	(252.49)	(262.33)	(264.60)	(274.07)
Age	44.00	45.97	43.99	45.99
	(0.57)	(0.58)	(0.58)	(0.58)
Male	0.50	0.49	0.54	0.54
	(0.50)	(0.50)	(0.50)	(0.50)
Born in Taipei	0.13	0.13	0.14	0.13
	(0.34)	(0.34)	(0.35)	(0.34)
Monthly previous salary	31,173.80	31,274.85	30,767.34	30,704.36
	(10,485.85)	(10,438.54)	(10,251.07)	(10,150.03)
Work in manufacturing previously	0.33	0.35	0.43	0.46
	(0.47)	(0.48)	(0.50)	(0.50)
Temporary lay-off	0.04	0.04	0.03	0.03
	(0.20)	(0.20)	(0.17)	(0.16)
Number of previous UI spell	0.20	0.20	0.14	0.14
	(0.47)	(0.47)	(0.38)	(0.37)
Job tenure of the previous employment (days)	1768.47	1845.38	1808.33	1824.01
	(1626.25)	(1653.40)	(1322.38)	(1318.57)
Number of recipients	10,043	10,440	9,687	8,890
Number of observations	59,016	65,763	57,678	54,756

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). Monthly re-employment hazard is the average monthly hazard rate during the 1st to 9th month of unemployment spells. Except for re-employment hazard, which is at spell-month level, other variables are computed at spell level. Standard deviations are in parentheses.

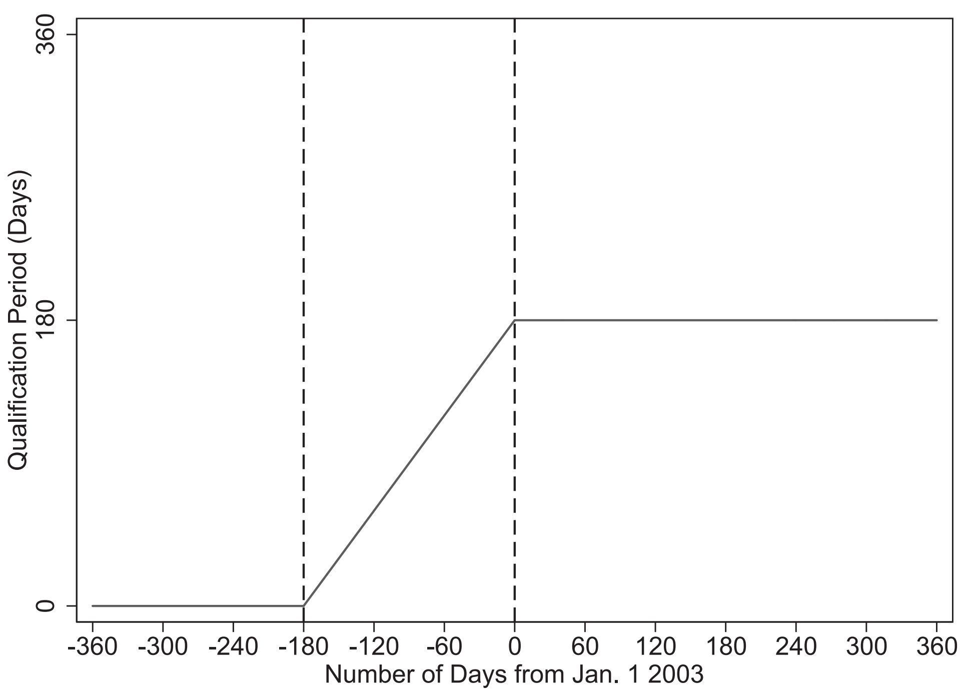


Fig. 1. Bonus Qualification Period and the UI Starting Date. *Notes:* This figure demonstrates the relationship between the bonus qualification period and the date UI spells started. UI recipients starting receiving benefits before July 1, 2002 are not eligible for any re-employment bonus. As the program phased in, UI recipients are potentially eligible for a more generous bonus offer, while the bonus qualification period is constant for UI recipients start receiving benefits after January 1, 2003.

formalize this idea, we implement a regression kink design (Nielsen et al., 2010; Card et al., 2015; Landais, 2015).

$$E\left(\frac{\partial y}{\partial RB(t)}|t = c_k\right) = \frac{\lim_{t\to c_k^+}\frac{dE(y|t)}{dt} - \lim_{t\to c_k^-}\frac{dE(y|t)}{dt}}{\lim_{t\to c_k^+}\frac{dRB(t)}{dt} - \lim_{t\to c_k^-}\frac{dRB(t)}{dt}} \quad (1)$$

where t represents the starting date of the UI spell. $E\left(\frac{\partial y}{\partial RB(t)}|t=c_k\right)$ is the causal effect of interest: The effect of re-employment bonus $RB(t)$ on the conditional expectation y (i.e. outcomes of interest) around cutoff date c_k (i.e. c_1 is July 1, 2002, and c_2 is January 1, 2003). We can express this 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 1 for the first kink and -1 for the second one. On the other hand, the numerator in Eq. (1) is estimated. We first use the month-spell level data and the following model to estimate the effect of the re-employment bonus on job search efforts.

$$\mathrm{E}[y_{im}|t] = \mu_m + \sum_{s=1}^{S}\gamma_s^{Kink}Kink_i \times (t-c_k)^s + \sum_{s=1}^{S}\delta_s(t-c_k)^s + X_i\phi \qquad (2)$$

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. Since we want to examine the effects of the re-employment bonus, we focus on the re-employment hazard within the bonus qualification period (i.e. the first six months of the unemployment spell). 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 Eq. (1)) can be measured by γ_1^{Kink}. Combining the slope change in potential re-employment bonus $RB(t)$ (i.e. the denominator of Eq. (1)), the effect of being eligible for one-day bonus qualification period on the re-employment hazard can be represented by $\frac{\gamma_1^{Kink}}{1}$. In the following analysis, we multiply $\frac{\gamma_1^{Kink}}{1}$ by 180 (i.e. $\gamma_1^{Kink}\times 180$) to match the effect of being eligible for the full length (six months/180 days) of the bonus qualification period, which is equivalent to three months' (90 days') UI benefits. In our main specification, we estimate Eq. (2) locally within a bandwidth of 180 days, before and after the cutoff dates (i.e. c_k). X_i denotes a rich set of observed characteristics, including gender, birth place, previous work history, and number of previous UI spells.

In order to compare our results with classical estimates in other studies on the topic (Woodbury and Spiegelman, 1987; Decker et al., 2001a; Meyer, 1995; Ahn, 2018), we also examine bonus effects on the duration of unemployment and subsequent labor market outcomes by using spell-level data and estimating the following regression:

$$\mathrm{E}[y_i|t] = \mu + \sum_{s=1}^{S}\gamma_s^{Kink}Kink_i \times (t-c_k)^s + \sum_{s=1}^{S}\delta_s(t-c_k)^s + X_i\phi \qquad (3)$$

All notations are the same as those in Eq. (2). The only difference is that we estimate it at the spell level. Finally, in order to account for any potential correlation in errors within the individuals who claim UI on the same date, we cluster standard errors by UI starting date when estimating Eqs. (2) and (3).

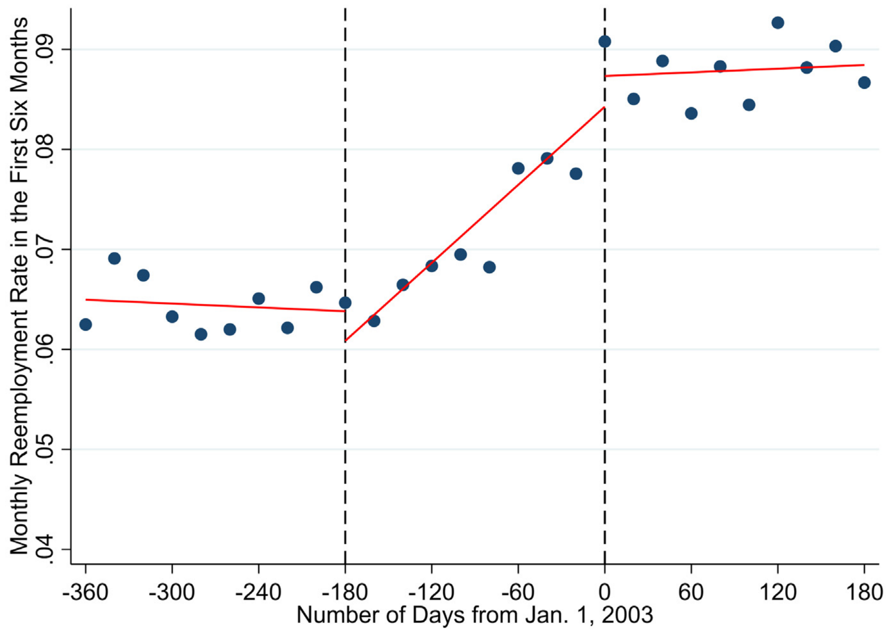
4.2. Estimation Results

Fig. 2 presents the relationship between the outcomes of interest and the starting date of UI benefits, using the main sample and a placebo sample. The main sample includes UI recipients aged 25 to 60 and starting the UI spell at some point between January 2002 and June 2003. The placebo has the same ages, but the starting dates for the UI spell are during January 2001 to June 2002. Each bin represents the total number of UI spells, starting within a 20-day interval. Fig. 2a displays the monthly re-employment hazard for the first six months of the spell. We find that for individuals who started their UI between July 1, 2002 and January 1, 2003 (i.e. partially eligible for the re-employment bonus), the monthly re-employment hazard increased as their UI starting date approached January 1, 2003. On average, the monthly re-employment hazard increased substantially from 0.06 (those who started UI around July 1, 2002) to 0.08 (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. eligible for a full 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 Fig. 1.

Fig. 2b displays the relationship between the monthly re-employment hazard and the starting date of UI benefits for the placebo sample (i.e. individuals are unaffected by the re-employment bonus reform, since they started their UI between January 2001 and June 2002). In sharp contrast to Fig. 2a, Fig. 2b 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 0.06. Consistent with the above finding, the benefits duration (see Fig. 3a) and non-employment duration (see Fig. 3c) also exhibit slope changes on July 1, 2002. That is, for those who started their UI between July 1, 2002 and January 1, 2003, the benefits duration and non-employment duration decreased as their UI starting date approached January 1, 2003. We do not find similar patterns when using the placebo sample (see Fig. 3b and 3d).

So far, we have summarized the effect of the re-employment bonus on search behavior in a single statistic—either mean durations or the average re-employment hazard over the first six months of the spell. In order to understand the mechanisms and distinguish between theoretical explanations for the observed effects, Fig. 4 plots monthly and weekly re-employment hazards for individuals fully eligible (circle symbol) or ineligible (square symbol) for a bonus, in order to show how the effect of the re-employment bonus varies in line with the existing non-employment duration. Fig. 4a and c indicate that being eligible for full length of bonus qualification period increases the job-finding rate before UI benefits are exhausted (i.e. month 6 and week 24). In contrast, Fig. 4b and d show that there are no such phenomena when using the placebo sample. Note that there is a big spike in the re-employment hazard at month 6 (i.e., week 24) when the benefits are exhausted. The magnitude of spike is much larger and longer lasting than other estimates found in developed economies like Europe or the U.S. For example, Fig. 4a indicates that the monthly job-finding rates for bonus-ineligible people are 4.5% and 11.7% in the 4th and 6th month after layoff, respectively (around 160% increase). We conjecture that this pattern is related to the fact that Taiwan has a significant share of self-employed workers (i.e., around 14% of labor), who are not enrolled in EI program. The UI recipients could get re-employed by becoming self-employed workers but still continue claiming UI since government

(a) Monthly Re-employment Hazard: 2002-2003



(b) Monthly Re-employment Hazard: 2001-2002

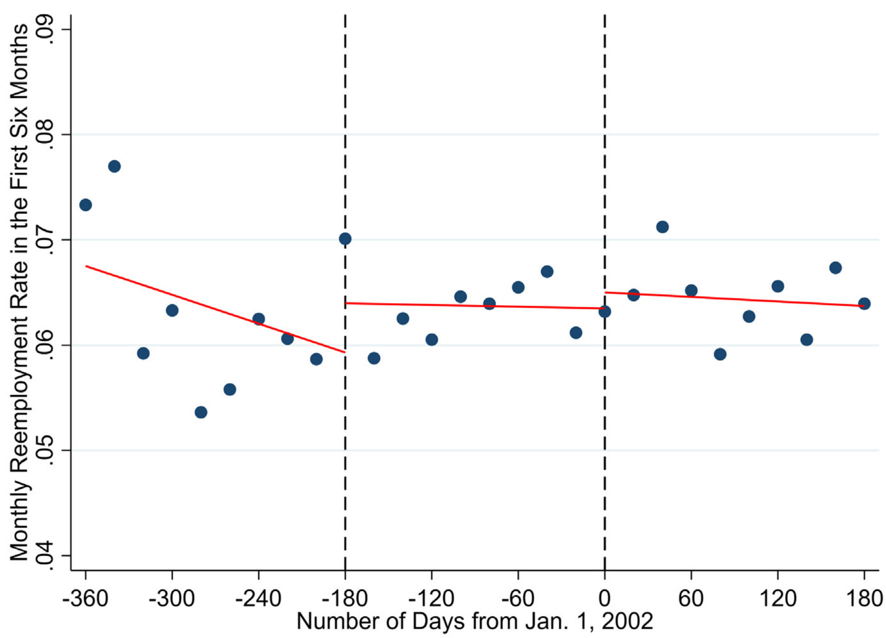


Fig. 2. Effects of Re-employment Bonus on Monthly Re-employment Hazard. *Notes:* Fig. 2a plots the average monthly re-employment hazard during the 1st to 6th of an unemployment spell over the number of days between January 1, 2003 and the date UI spells started. Each bin represents the average monthly 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. Fig. 2b is for a placebo test. It plots the average monthly re-employment hazard during the 1st to 6th of an unemployment spell over the number of days between January 1, 2002 and the date UI spells started. The first dash line indicates July 1, 2001, 6 months before the "placebo" bonus program began. The second line indicates January 1, 2002.

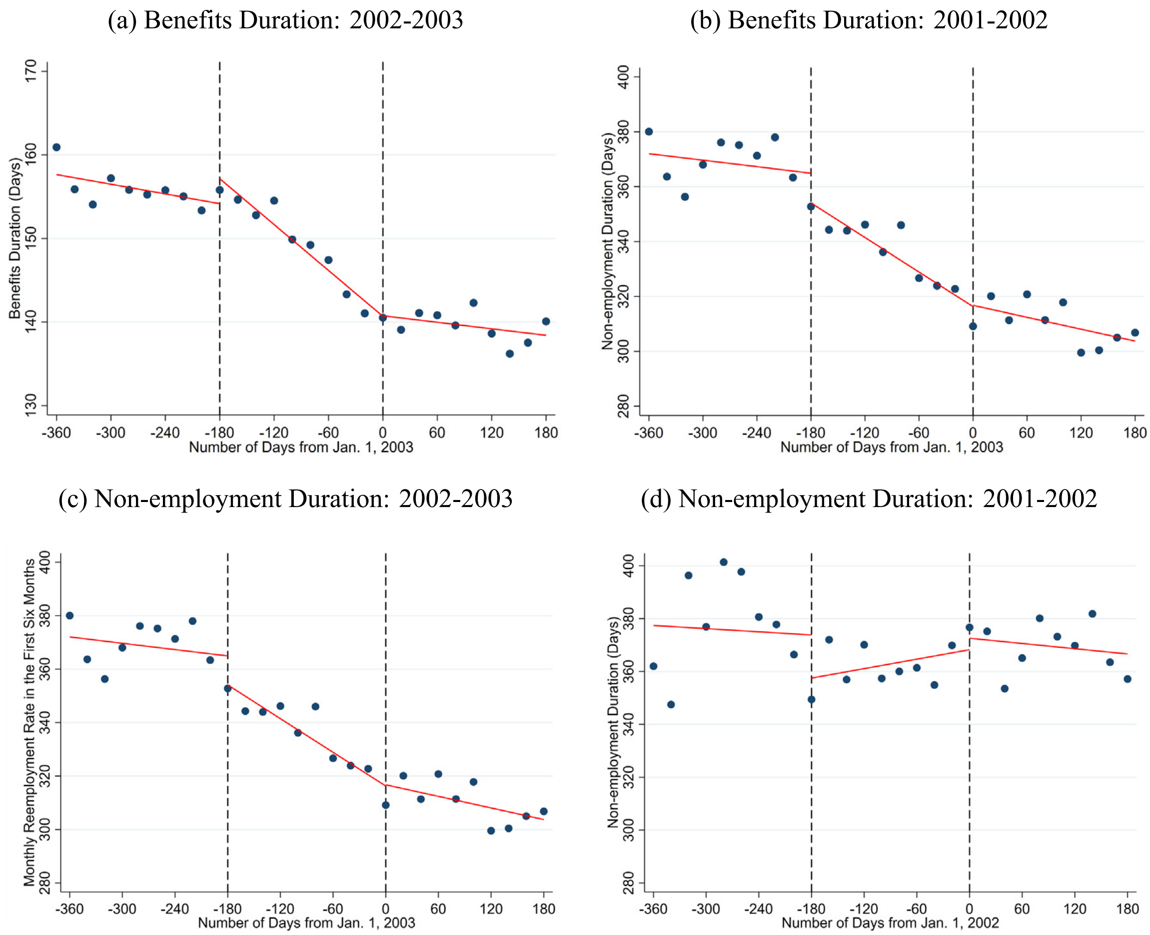


Fig. 3. Effects of Re-employment Bonus on Benefits Duration and Non-employment Duration. *Notes:* Fig. 3a and 3c plot the average benefits duration and non-employment duration over the number of days between January 1, 2003 and the date UI spells started. Each bin represents the average monthly re-employment 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. Fig. 3b and 3d are placebo tests and plot the the average benefits duration and non-employment duration over the number of days between January 1, 2002 and the date UI spells started. The first dash line indicates July 1, 2001, 6 months before the "placebo" bonus program began. The second line indicates January 1, 2002.

has a limited ability to monitor their job status. Similar patterns can be found in the countries that have large informal labor market, such as Brazil (Gerard and Gonzaga, 2021).[17]

Panel A of Table 3 reports our main estimates for the effect of the re-employment bonus on the re-employment hazard during the first six months of the spell. 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 full re-employment bonus equivalent to three months' (90 days') UI benefits can increase the monthly re-employment hazard by 2 percentage points. Column (2) includes basic covariates (e.g. dummy for male, dummy for being born in Taipei city, dummies for five main industries of previous job, and number of previous UI spells) in a linear specification.[18] We find the estimate is quite similar. In Column (3), we use a quadratic function on either side of the cutoff to control the time trend of the re-employment hazard. The estimate is not statistically significant, due to a large standard error, but the magnitude (i.e. 1.6 percentage points increase) is similar to the one in Column (2). Moreover, both AIC and BIC suggest that the quadratic specification is dominated by a linear specification. 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 this specification gives a similar result (i.e. 2 percentage points increase).

Finally, we further control the variables related to previous employment history, such as monthly previous salary and the duration of the previous job, using a set of flexible polynomial terms (i.e. fifth-degree polynomials). Since unobserved characteristics (e.g. ability, attachment to the labor market) that affect job search efforts should be highly correlated with these observed variables, flexibly including them could help us indirectly reduce differences in unobserved attributes between eligible and ineligible workers. The estimate in Column (5) suggests that flexibly controlling the variables that are proxies for unobserved ability or labor market attachment does not overturn our results. To sum up, our preferred estimate in Column (5) suggests that being eligible for the full length of the bonus qualification period (i.e. equiv-

[17] We use the 1999–2011 Manpower Utilization Survey, which is similar to the Current Population Survey in the U.S., to provide some basic facts about labor market in Taiwan. Among the workers whose ages are between 25 and 60, 63.7% of them are private-sector employees, 10.6% are public-sector employees, 14.4% are self-employed, 4.8% are employers, and 6.6% are non-paid family workers. The definition of a non-paid family worker is an individual who works in his/her family business but does not receive any payment and works for more than 15 h per week. If we restrict our sample to unemployed workers who get re-employed, around 83% of them become private-sector employees again, which are covered by EI program. Around 11% of them become self-employed or non-paid family workers.

[18] The five industries are mining, construction, manufacturing, retailing, logging, and transportation. We use other services as the reference group.

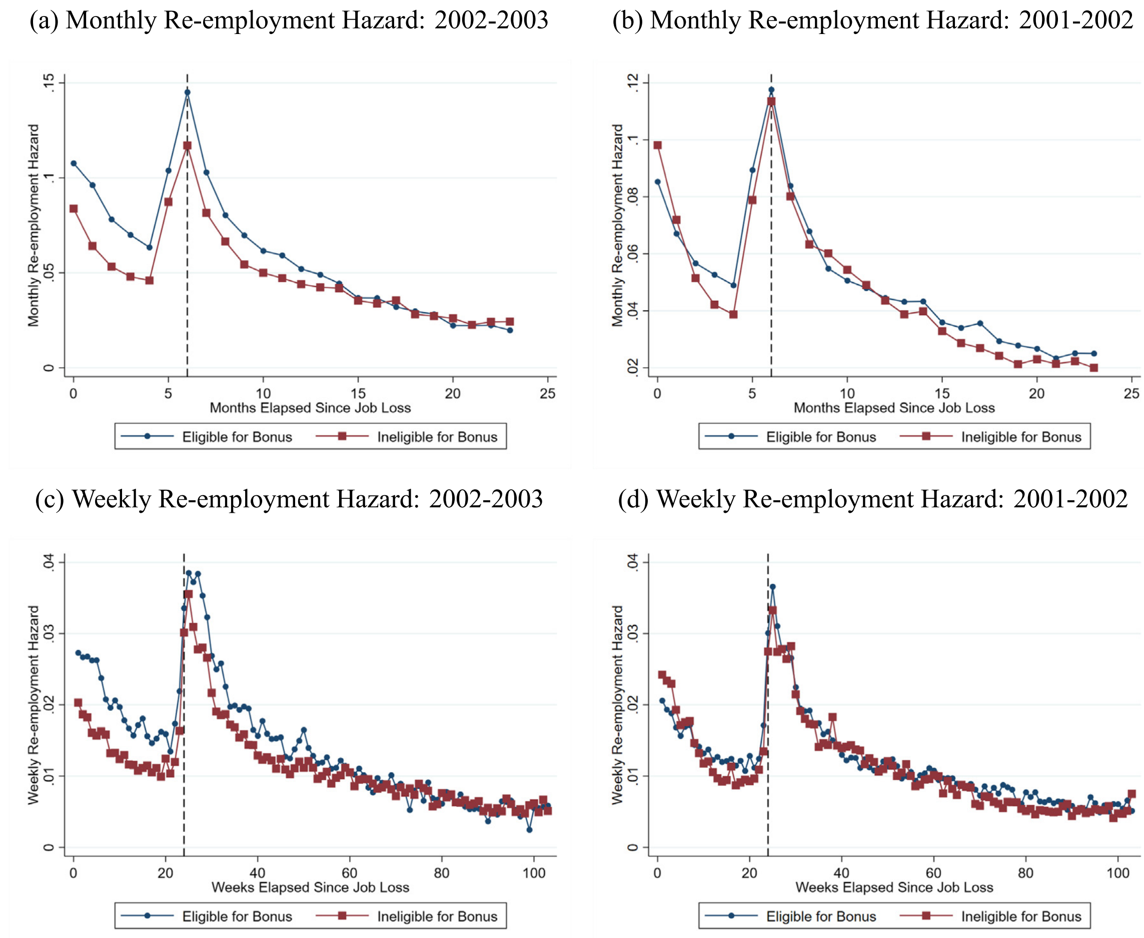


Fig. 4. Re-employment Hazard over Time: Re-employment Bonus Sample. *Notes:* Fig. 4a and 4c plot monthly and weekly re-employment hazard for individuals fully eligible (circle symbol) or ineligible (square symbol) for bonus, respectively. Fig. 4b and 4d display the results using placebo sample. Each bin represents the average monthly/weekly re-employment hazard. The x-axis is months/weeks elapsed since job loss. The dash vertical line represents the timing of benefits exhaustion (i.e. week 24 and month 6).

alent to three months' (90 days') UI benefits) can increase the monthly re-employment hazard by 1.9 percentage points. Compared to the baseline mean (around 6.4%), this represents a 30% increase in the average monthly re-employment hazard. In our later analysis, we use this estimate as our main result.

One may be concerned that our analysis of the re-employment hazard could be biased, due to dynamic selection. Since as people find jobs they drop from the dataset, the different selection induced by treatment (i.e. re-employment bonus) on the two sides of the cutoff might change the initial composition of the treatment and control groups. In the Online Appendix A.1, we examine this issue by comparing treated and untreated individuals' differences in pre-determined characteristics at the beginning of the spell with those in the last month of the UI spell (i.e. the sixth month of the spell). Our results suggest that differences in the characteristics of the treatment and control groups are largely the same over the unemployment spell (see Column (3) of Table A1).[19] In addition, following Chetty, 2008 and Card et al., 2007, we implement the RK design, using a Cox proportional-hazards model, which includes all of the estimation samples (see Online Appendix A.2). Table A3 suggests that the estimate based on a Cox regression provides consistent results.

Panels B and C of Table 3 report the estimated effect of the re-employment bonus on benefit duration and non-employment duration, respectively. We find that being eligible for a full re-employment bonus equivalent to three months' (90 days') UI benefits can significantly decrease benefit duration by 9.6 days, which accounts for a 6% decline from the baseline mean (i.e. 156.8 days). Furthermore, our results suggest that being eligible for a full length of bonus qualification period can reduce non-employment duration by 33.7 days (i.e. an 9% decline from the baseline mean). Finally, in the Online Appendix A.3, we examine the effects of the re-employment bonus on job quality, such as post-unemployment wage and post-unemployment job tenure. Our results indicate that a shortened benefit (non-employment) duration—induced by the provision of a re-employment bonus—has little impact on the quality of the new job (see Table A4 and Fig. A1).

In the Online Appendix A.5, we examine the robustness of our main estimates using different specifications, sample criteria, choices of bandwidth. In addition, we investigate the validity of RK design. Basically, we find that our results are quite robust and the identification assumption is valid.

4.3. Comparison to the Previous Literature

Although several countries (e.g. Korea, Taiwan, Netherlands, and Hungary) around the world have implemented re-

[19] Table A2 compares treated and untreated individuals' *differences in slope* of pre-determined characteristics at the beginning of the spell with those in the last month of the UI spell (i.e. the sixth month of the spell). We find that *differences in slope* of the characteristics for the treatment and control groups are largely the same over the unemployment spell.

Table 3
Effects of Re-employment Bonus on Job Search Efforts.

	(1)	(2)	(3)	(4)	(5)
Panel A: *Re-employment Hazard*					
$\gamma_1^{Kink} \times 180$	0.020***	0.020***	0.016	0.020**	0.019***
	(0.004)	(0.004)	(0.015)	(0.008)	(0.004)
Baseline mean			0.064		
Sample size	471,052	471,052	471,052	471,052	471,052
Panel B: *Benefits Duration*					
$\gamma_1^{Kink} \times 180$	−9.31***	−9.95***	−6.55	−9.42**	−9.64***
	(1.58)	(1.51)	(5.91)	(3.72)	(1.43)
Baseline mean			156.84		
Sample size	92,977	92,977	92,977	73,088	92,977
Panel C: *Non-employment Duration*					
$\gamma_1^{Kink} \times 180$	−34.69***	−35.13***	−66.39**	−38.30*	−33.69***
	(8.04)	(8.08)	(35.81)	(20.23)	(8.01)
Baseline mean			371.44		
Sample size	92,977	92,977	92,977	88,091	92,977
RKD	Yes	Yes	Yes	Yes	Yes
Basic Control	–	Yes	Yes	–	Yes
Working History	–	–	–	–	Yes
Poly. model	linear	linear	quadratic	linear	linear
Bandwidth (days)	180	180	180	CCT	180

Notes: Each column displays the estimated coefficient γ_1^{Kink} on $Kink \times (t - c_k)$ in Eq. (2) or (3). We multiply them by 180 to give effects of being eligible for a full length of bonus qualification period (i.e. equivalent to three-month UI benefits). The outcome variables are monthly re-employment hazard during the 1st to 6th of an unemployment spell (Panel A), benefit duration (Panel B), and non-employment duration (Panel C). Column (1) displays a basic RK estimate using a linear function to control the effect of UI starting date on outcome variables. Column (2) includes basic covariates (e.g. dummy for male, dummy for being born in Taipei city, dummies for five main industries of previous job, and number of previous UI spells) in a linear specification. Column (3) uses a quadratic function on either side of the cutoff to control the effect of UI starting date on outcome variables. 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) further controls the variables related to previous employment history, such as monthly previous salary and the duration of the previous job using a set of flexible polynomial terms (i.e. fifth degree polynomials). Except column (4), the bandwidth choice is 180 days. All standard errors are clustered by UI starting date. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

employment bonus programs, empirical evidence regarding the bonus effect on duration and job match quality is still limited.[20] Existing evidence suggests that the offer of a re-employment bonus can modestly speed up the transition to employment (Woodbury and Spiegelman, 1987; Decker et al., 2001a; Meyer, 1995; Ahn, 2018) and does not affect the quality of the new job. For example, Woodbury and Spiegelman, 1987 found that providing a 500 US$ re-employment bonus (i.e. about four weeks of unemployment benefits) to UI recipients in Illinois significantly reduced the average benefit duration by 1.15 weeks (i.e. around a 6% decline from the baseline mean) and did not affect re-employment wages.[21] More recently, using a regression discontinuity design and administrative data from South Korea, Ahn, 2018 determined that increasing the re-employment bonus can significantly shorten benefit duration by 0.68 to 1.82 weeks (i.e. around a 3.6% to a 9.5% decline from the baseline mean) and does not affect employment stability. In general, our results indicate that the duration response to the provision of a re-employment bonus is modest (i.e. a 6% to an 9% decrease), and the quality of the job match is not significantly affected. This implies that the elasticity of benefit duration (non-employment duration) with respect to potential re-employment bonus is about −0.03 (−0.06).[22]

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 on re-employment hazard, we estimate the following regression at the month-spell level:

$$y_{im} = \alpha_m + \beta^{Age45} Age45_i + f(a_i) + X_i\phi + v_{im} \tag{4}$$

where y_{im} and α_m are defined in the same way as in Eq. (2). In order to evaluate the overall impact of the UI extension on unemploy-

[20] Several European countries also implement re-employment programs. For example, Van der Klaauw and Van Ours, 2013 estimate the effect of the re-employment bonus program on welfare exit in Rotterdam. Their hazard model estimates suggest the re-employment bonus is not effective in increasing the numbers of people leaving the welfare system. However, the bonus in Rotterdam is offered to welfare recipients who have been unemployed for at least one year, so their estimates may not be comparable to the effects of Taiwan's bonus program. Lindner and Reizer, 2019 discuss the bonus program in Hungary and show that it has a limited impact on job searches, because the participation rate in the program is low.

[21] Four random experiments were conducted in New Jersey, Illinois, Washington, and Pennsylvania. Although the bonus designs did differ from each other, these experiments suggest bonuses significantly reduce the insured duration of unemployment by about one-half a week (O'Leary et al., 1995). New Jersey's design was the most similar one to Taiwan's. It provided 50% of the remaining entitlement, and the amount declined by 10% per week. Anderson, 1992 established that the New Jersey bonus increased the re-employment hazard early in the offer period, and the effect diminished over time. However, the bonus offers were made after seven weeks of insured unemployment, and participants did not know an offer would be made before that time. Hence, the New Jersey experiment is not externally valid, because in a real program, individuals would know that a bonus offer would be made in week seven. O'Leary et al., 1995's estimates in the bonus experiments for Pennsylvania and Washington suggest a smaller effect of bonuses on UI duration (about half a week).

[22] Since workers are ineligible for bonus before the reform (i.e. receive zero bonus), we compute the price elasticity using an arc-elasticity calculated as $((D_2 - D_1)/((D_1 + D_2)/2))/((B_2 - B_1)/((B_1 + B_2)/2))$, where D_1 and B_1 denote, respectively, the baseline duration and re-employment bonus (i.e., control group's average duration and bonus), and D_2 and B_2 are the duration and re-employment bonus affected by the re-employment bonus reform.

ment exit, before and after the point at which the extended benefits become available, we focus on the re-employment hazard before extended benefits are exhausted (i.e. the first nine months). 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 three-month (90-day) extended 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.[23] X_i denotes a rich set of observed characteristics, including gender, birth place, previous work history, and number of previous UI spells. v_{im} is an error term that reflects all of the other factors affecting the outcome of interest.

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 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 benefits period on the hazard of transition to employment. Following Schmieder et al., 2016, we estimate Eq. (4) locally within a bandwidth of two years (i.e. 730 days), before and after the age of 45 at lay-off. In a later section, we examine whether our main results are sensitive to different bandwidth choices and specifications.

Similar to the evaluation of the re-employment bonus, we also examine several classical outcomes in the UI literature, such as benefits duration, non-employment duration, and job quality. Understanding the effects of a UI extension on these outcomes is useful for comparing our results with previous studies. In addition, these estimates are key parameters used in the welfare analysis of Section 6. Specifically, we estimate the following regression at the spell level:

$$y_i = \alpha + \beta^{Age45} Age45_i + f(a_i) + X_i\phi + v_i \tag{5}$$

All notations are the same as ones in Eq. (4); the only difference is that we estimate it at the spell level. Finally, in order to account for any potential correlation in errors within age group, we cluster standard errors by age at job loss when estimating Eqs. (4) and (5).

5.2. Estimation Results

In this section, we first examine the effect of a three-month extended benefit on re-employment hazard, following which we discuss benefit effects on other outcomes. Fig. 5 displays how the monthly re-employment hazard during the first nine months of an unemployment spell varies in accordance with an individual's age at job loss.[24] We show the graphical results based on the main sample (see Fig. 5a) and the placebo sample (see Fig. 5b).

Fig. 5a suggests the monthly re-employment hazard shows a discernible drop at the age 45 by about 3 percentage points. To examine any confounding factors affecting our results, 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 Fig. 5b was mainly driven by an extended benefit. In sharp contrast to Fig. 5a, we find no visible change in re-employment hazard at the age of 45 when using pre-reform data.

Fig. 6 shows the graphical evidence for duration outcomes. Consistent with the re-employment hazard results, we find that both benefits duration (See Fig. 6a) and non-employment duration (See Fig. 6c) increase substantially at age 45. Figs. 6b and 6d suggest that such discontinuity did not exist before the 2009 UI extension, which thus provides clear evidence that the change in outcomes at age 45 is driven exclusively by the extension of UI benefits.

Again, the graphical evidence so far summarizes the average effects of extended benefits on search behavior. In order to understand how extending UI benefits affects search behaviors as the spell evolves, Fig. 7 plots monthly and weekly re-employment hazards for individuals with job losses before and after age 45. Consistent with previous studies (Meyer, 2002; Schmieder et al., 2012), Figs. 7a and 7c illustrate that there are clear spikes in the job-finding rate at the month 6 (i.e. week 24) and month 9 (week 36)—the benefit exhaustion points for individuals below (square symbol) and above (circle symbol) the age of 45, respectively. Furthermore, people eligible for extended benefits also have substantially lower re-employment hazards than those ineligible for it during the period covered by the UI extension (i.e. months 7 to 9/weeks 25 to 36). In contrast, Figs. 7b and 7d show that there is no such phenomena when using the placebo sample (i.e. 2006–2008 data). Finally, Figs. 7a and 7c also indicate that the re-employment hazard of individuals above 45 years of age drop slightly prior to month 6 (i.e. week 24), i.e. before they actually receive any additional income. This result provides clear evidence that at least some individuals are forward-looking and take into account their future expected income stream.

Table 4 reports our main estimates for the effects of extending UI benefits. Panel A shows the results for monthly re-employment hazard during the first nine months of an unemployment spell. Similar to the structure used to present the results in Table 3, we first display a basic RD estimate, using a linear function to control the age profile of the re-employment hazard (see Column (1)). The result suggests that a three-month increase in potential benefit duration significantly reduces the monthly re-employment hazard by 2.9 percentage points. Then, Columns (2) to (5) present the results based on different settings. In general, the RD estimates are quite stable across specifications. Our preferred estimate is Column (5), which controls variables related to previous employment history, such as monthly previous salary and the duration of the previous job, using a set of flexible polynomial terms (i.e. fifth-degree polynomials). The result suggests that a three-month extension in UI benefits reduces the monthly re-employment hazard by 3 percentage points. Compared to the baseline mean (around 11%), this represents a 27% reduction in the average monthly re-employment hazard.

Similar to RK design, one caveat of re-employment hazard analysis is that when people find jobs, they drop from the dataset, which might change sample composition at different times during an unemployment spell. Again, Online Appendix B.1 examines whether the differences in the observed characteristics of the treated and untreated workers vary for different durations in the unemployment spell. Our results suggest that differences in the sample composition of the treatment and control group do not change over time. Furthermore, we include all estimation samples and implement the RD design within a framework in the Cox proportional-hazards model (Card et al., 2007; Chetty, 2008). Online Appendix B.2 suggests that the estimate based on Cox regression provides consistent results (see Table B2).

[23] 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.

[24] 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.

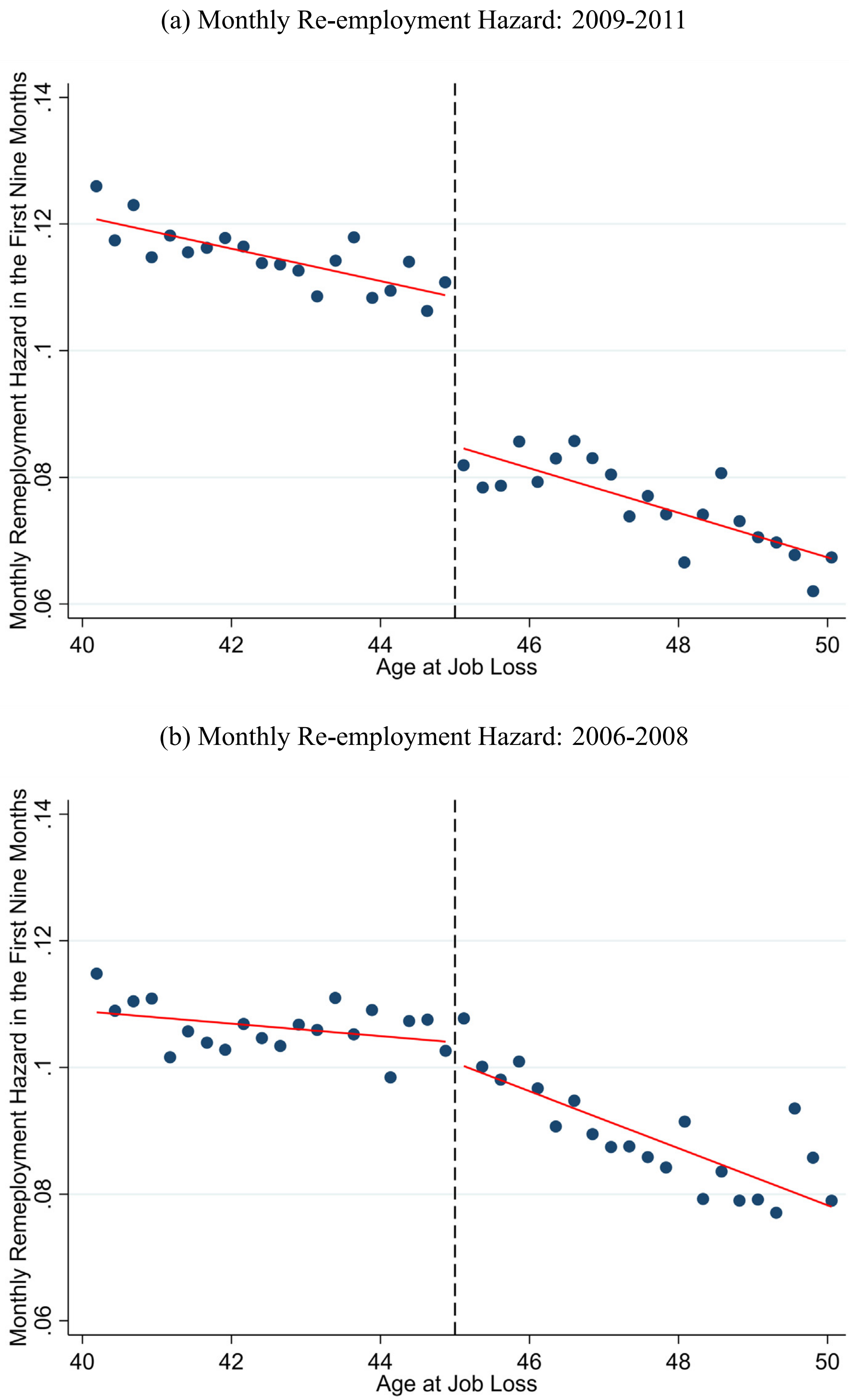


Fig. 5. Effects of Extended UI Benefits on Monthly Re-employment Hazard. *Notes:* Fig. 5a plots the average monthly re-employment hazard during the 1st to 9th of an unemployment spell over age 40 to 50 using 2009–2011 sample. Fig. 5b is a placebo test and plots the average monthly re-employment hazard during the 1st to 9th of a unemployment spell over age 40 to 50 using 2006–2008 sample. Each bin represents the average monthly re-employment hazard within 90 days (3 months) interval. The dash line represents age 45. The solid lines are fitted values from a linear regression on either sid.e of the cutoff.

Panels B and C of Table 4, respectively, report the estimated effect of the UI extension on benefit duration and non-employment duration. We find that being eligible for a three-month (90-day) UI extension can significantly increase benefit duration by 56.9 days (i.e. a 39% change from the baseline) and non-employment duration by 36.9 days (i.e. a 14% change from the baseline). Finally, in the Online Appendix B.3, we also examine the effect of UI extension on job match quality. Our results suggest that a longer job search induced by a UI extension has little impact on job match quality (see Table B3 and Fig. B1).

Online Appendix B.4 investigates the robustness of our RD estimates using different specifications, sample criteria, choices of

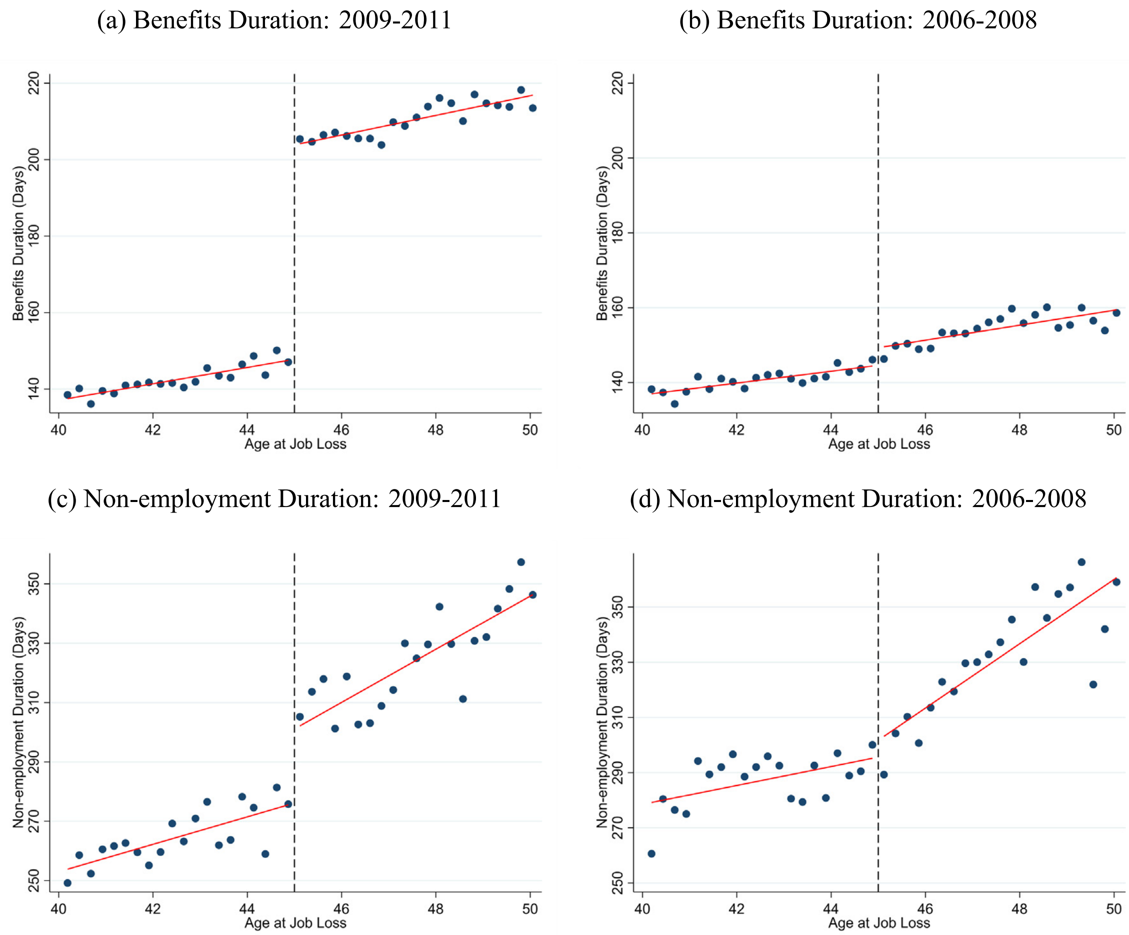


Fig. 6. Effects of Extended UI Benefits on Benefits Duration and Non-employment Duration. *Notes:* Fig. 6a and 6c plot the average benefits duration and non-employment duration over age 40 to 50 using 2009–2011 sample. Fig. 6b and 6d are placebo tests and plot the the average benefits duration and non-employment duration over age 40 to 50 using 2006–2008 sample. Each bin represents the average duration within 90 days (3 months) interval. The dash line represents age 45. The solid lines are fitted values from a linear regression on either side of the cutoff.

bandwidth. We also examine the validity of RD design. In sum, our results suggest that the estimates are quite robust and the identification assumption is valid.

5.3. Comparison to the Previous Literature

Most existing estimates on UI effects are based on U.S. and European countries. To the best of our knowledge, we provide one of the first pieces of evidence on UI effects, using administrative data from Asian countries that have different labor market characteristics (e.g. low unemployment rates). Therefore, it is interesting to compare our results with estimates from U.S and European countries. In this section, we use the estimates in Section 5.2 to calculate two commonly reported parameters: The marginal effect and duration elasticity. Since most of the previous literature has studied non-employment duration (Schmieder and von Wachter, 2016), our comparison is based on this outcome.

First, our results indicate that the marginal effect of UI extension on non-employment duration is 0.41, suggesting that for a one-month increase in potential benefit duration, non-employment duration increases by about 12 days. Schmieder and von Wachter, 2016 provide a comprehensive review of the estimates from Europe and U.S. According to their Table 1, which excludes two outliers at the top and bottom, the mean marginal effect on non-employment duration is 0.23, ranging from 0.05 to 0.65. Our estimates are at the higher end of these estimates. Since both non-employment durations and potential benefit durations are quite different across countries, it is better to make a comparison using duration elasticity. Our findings suggest that the implied elasticity of non-employment duration with respect to potential benefit duration is 0.27, which is lower than the median value of previous estimates, 0.37, provided in Table 1 of Schmieder and von Wachter, 2016. In other words, after considering the large baseline mean of non-employment duration and the relatively short potential benefit duration in Taiwan, our estimated duration elasticity is not particularly high. Furthermore, for benefits duration, we obtain relatively large marginal effects 0.63 and duration elasticity 0.78, compared with the estimates offered by Schmieder and von Wachter, 2016. Since our sample consist of the individuals whose age at job loss is around 45 years-old, we compare our estimates further with those of Schmieder et al., 2012, who also study benefits extensions for middle-age workers (i.e. age 42 to 49) in Germany. Our estimated non-employment duration (benefit duration) elasticity of 0.27 (0.78) is a bit larger than their estimates of 0.12 to 0.14 (0.54 to 0.67).

Finally, our results indicate that the UI effect on the quality of a new job is small, which is consistent with estimates in the prior literature that found UI wage effects are not significantly different from zero. According to the meta-analysis in Nekoei and Weber, 2019, there is a negative relation between the marginal effects of UI on post-unemployment wage and non-employment duration. Since we find that the duration response to UI extension (i.e. mar-

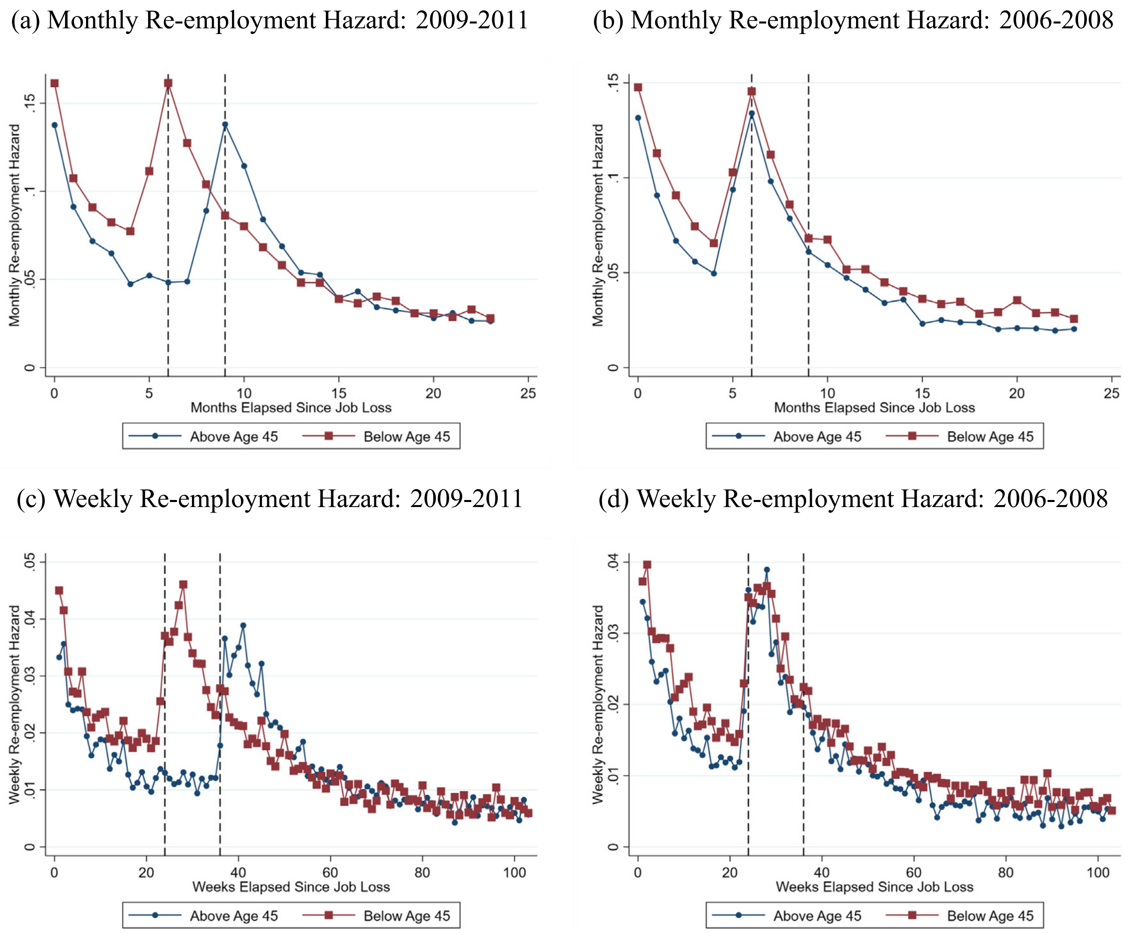


Fig. 7. Re-employment Hazard over Time: Extended Benefits Sample. *Notes:* Fig. 7a and 7c plot monthly and weekly re-employment hazard for individuals eligible or ineligible for extended benefits (i.e. individuals above or below age 45), respectively. We use circle symbol (square symbol) to represent individuals who are above (below) age 45. Fig. 7b and 7d display the results using placebo sample. Each bin represents the average monthly/weekly re-employment hazard. The x-axis is months/weeks elapsed since job loss. Two dash lines represents the timings of regular benefits exhaustion (i.e. week 24 and month 6) and extended benefits exhaustion (i.e. week 36 and month 9).

ginal effects of UI extension) is relatively large, it is reasonable that we get a small UI wage effect.

6. Welfare Implications

We have estimated the effects of re-employment bonuses and extended benefits. In this section, we use these estimates to evaluate the behavior costs of the two policies, following which we compare the response in re-employment hazard to re-employment bonuses and extended benefits to estimate the value of extended UI benefits. We focus on the intuition and leave the derivations to the Online Appendix C.

6.1. Model Settings

Consider a discrete time search model modified from Chetty, 2008 and Landais, 2015. An individual becomes unemployed in 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 in period t, she receives an unemployment benefit, b_t, with a potential benefit duration, P_b, that is

$$b_t = \begin{cases} b, & \text{if } 0 \leqslant t < P_b \\ 0, & \text{if } t \geqslant P_b \end{cases}$$

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.[25] If an individual finds a job before the end of the bonus qualification period, P_θ, and accumulates three months of re-employment, she receives a re-employment bonus, r_t, equal to $100 \cdot \theta$ percent of her remaining UI benefits. Formally,

$$r_t = \begin{cases} \theta b \sum_{t=0}^{P_\theta - 1} [1 - S_0(t)], & \text{if } \sum_{t=0}^{T} S_0(t) < P_\theta \text{ and } t = \sum_{t=0}^{T} S_0(t) + 1 \\ 0, & \text{otherwise} \end{cases}$$

where $S_0(t)$ is the probability that workers remain unemployed at time t. A lump-sum transfer $(\theta b \sum_{t=0}^{P_\theta - 1} [1 - S_0(t)])$ is paid to workers who find a job before the end of the qualification period and accumulate three months of re-employment, where $\theta = 0.5$ in Taiwan. Note that the bonus qualification period is the same as the UI eligibility period in Taiwan, i.e. $P_\theta = P_b = P$. Thus, the qualification period for the bonus (P_θ) mechanically increases as the potential benefit duration increases (P_b).

[25] This definition is similar to the one offered by 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.

Table 4
Effects of Extended UI Benefits on Job Search Efforts.

	(1)	(2)	(3)	(4)	(5)
Panel A: *Re-employment Hazard*					
β^{Age45}	−0.029***	−0.030***	−0.031***	−0.029***	−0.030***
	(0.003)	(0.003)	(0.005)	(0.003)	(0.003)
Baseline mean			0.112		
Sample size	124,774	124,693	124,693	185,690	124,693
Panel B: *Benefits Duration*					
β^{Age45}	56.97***	57.15***	56.64***	55.10***	56.91**
	(2.02)	(1.99)	(2.92)	(1.75)	(1.96)
Baseline mean			145.81		
Sample size	20,483	20,473	20,473	46,753	20,473
Panel C: *Non-employment Duration*					
β^{Age45}	35.46***	36.83***	35.93***	30.68***	36.90***
	(6.99)	(6.94)	(9.64)	(5.40)	(6.90)
Baseline mean			270.83		
Sample size	20,483	20,473	20,473	52,595	20,473
RDD	Yes	Yes	Yes	Yes	Yes
Basic Controls	–	Yes	Yes	–	Yes
Working History	–	–	–	–	Yes
Poly. model	linear	linear	quadratic	linear	linear
Bandwidth (days)	730	730	730	CCT	730

Notes: Each column displays the estimated coefficients β^{Age45} on *Age*45 using Eq. (4) or (5). The outcome variables are monthly re-employment hazard during the 1st to 9th of an unemployment spell (Panel A), benefit duration (Panel B), and non-employment duration (Panel C). Column (1) displays a basic RD estimate using a linear function to control the effect of age on outcome variables. Column (2) includes basic covariates (e.g. dummy for male, dummy for being born in Taipei city, dummies for five main industries of previous job, and number of previous UI spells) in a linear specification. Column (3) uses a quadratic function on either side of the cutoff to control the effect of age on outcome variables. 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) further controls the variables related to previous employment history, such as monthly previous salary and the duration of the previous job using a set of flexible polynomial terms (i.e. fifth degree polynomials). Except column (4), the bandwidth choice is 730 days. All standard errors are clustered by age at job loss. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

When employed, the worker earns a wage income, w, and pays a tax τ. Thus, the flow utility when employed in period t equals $v(c_t^e) = v(A_t - A_{t+1} + w + r_t - \tau)$, where c_t^e indicates consumption when employed in period t, and A_t is her asset level at time t. Note that the credit and insurance markets in this model are incomplete. Thus, an individual needs to keep her asset A_{t+1} above a lower bound on assets L (i.e. $A_{t+1} \geqslant L$), in order to borrow and save. As in Chetty, 2008, we maintain the assumption that workers are liquidity-constrained, so that they do not change their saving in response to income transfers. Therefore, $u(c_t^e) = u(w + r_t - \tau)$ and $u(c_t^u) = u(b_t)$. This is a crucial assumption for our approach to identifying the value of extended UI benefits. We discuss the implications of the liquidity-constraint assumption in Section 6.4.

6.2. Behavioral Costs of the Re-employment Bonus

Our estimates in Section 4 suggest that the provision of a re-employment bonus substantially reduces benefit duration and non-employment duration. Do the re-employment bonuses pay for themselves? In other words, is the reduction in UI expenditure (and increased tax revenues) caused by the shortened unemployment duration being larger than the increased bonus payment? In this section, we examine the behavioral "gain" of re-employment bonuses and their cost-effectiveness by combining our reduced-form estimates and model implications.

Before discussing the behavioral cost of re-employment bonuses, it is important to clarify from where the empirical variation in the re-employment bonus comes from. According to the re-employment bonus formula, r_t, we can see that there are two types of policy instruments that can increase bonus generosity—an increase in θ and an increase in the bonus qualification period, P_θ. As we show in Fig. 1, workers who start their UI spell closer to January 1, 2003, are eligible for longer re-employment bonuses, while θ remains fixed. Therefore, the following analysis focuses on the effects of extending the bonus qualification period (P_θ) rather than increasing the bonus level (θ).

The government budget balance (G) in the model is given by $bB + \theta b \sum_{t=0}^{P_\theta - 1}[1 - S_0(t)]$, which is financed by $(T - D)\tau$, where D is the non-employment duration. In the Online Appendix C.3, we demonstrate that the effect of extending re-employment bonuses on the government budget balance is

$$\frac{dG}{dP_\theta} = \underbrace{\theta b[1 - S_0(P_\theta)]}_{\text{Mechanical Cost}} \qquad \underbrace{-\theta b \sum_{t=0}^{P_\theta - 1} \frac{dS_0(t)}{dP_\theta} + b\frac{dB}{dP_\theta} + \tau\frac{dD}{dP_\theta}}_{\text{Behavioral Cost}},$$

The first term $\theta b[1 - S_0(P_\theta)]$ is bonus spending in the absence of a behavioral response, where $1 - S_0(P_\theta)$ is the probability that an unemployed worker will find a job before exhausting their UI benefits (i.e. workers who are eligible for bonuses). The remaining three terms are driven by behavioral responses: $-\theta b \sum_{t=0}^{P_\theta - 1} \frac{dS_0(t)}{dP_\theta}$ and $b\frac{dB}{dP_\theta}$ are the increases in bonus spending and the decreases in UI expenditures due to more rapid re-employment, respectively, and $\tau\frac{dD}{dP_\theta}$ is increases in tax revenues due to a decrease in the non-employment duration. Based on formula (6), the behavioral cost per NTD of the re-employment bonus is the ratio of the behavioral cost of extending bonuses $(-\theta b \sum_{t=0}^{P_\theta - 1} \frac{dS_0(t)}{dP_\theta} + b\frac{dB}{dP_\theta} + \tau\frac{dD}{dP_\theta})$ to the mechanical cost of extending them $(\theta b[1 - S_0(P_\theta)])$.

Our estimates from Table 3 suggest that being eligible for the full bonus qualification period (i.e. six months (180 days)) reduces benefit duration B and non-employment duration D by 9.6 and 34 days, respectively. We plug in these estimates into formula (6) and find that the behavioral cost per NTD of the re-employment bonus is -0.61. That is, the behavioral response to the re-employment bonus enhances the government budget by 0.61 NTD so that only 0.39 NTD have to be raised to finance one NTD of re-employment bonus, thus suggesting the bonus program is not cost-effective when the bonus take-up is complete. We provide a detailed calculation of the behavioral cost per NTD of bonus spending in the Online Appendix C.1 and offer additional welfare analysis in the Online Appendix C.3.

6.3. Behavioral Costs of Extending UI Benefits

We have estimated the behavioral "gain" of the re-employment bonus. A natural question to ask is, how does a re-employment bonus interact with UI? Intuitively, the re-employment bonus reduces the moral hazard cost of UI (without affecting the value of UI) (Ahn, 2018) and makes a UI extension a more appealing policy option. In this section, we estimate the behavioral cost of the UI extension.

In the Online Appendix C.4, we show - balance is:

$$\frac{dG}{dP} = \underbrace{bS_0(P) + \theta b[1 - S_0(P)]}_{\text{Mechanical Cost}} + \underbrace{b\sum_{t=0}^{P-1} \frac{dS_0(t)}{dP} - \theta b \sum_{t=0}^{P-1} \frac{dS_0(t)}{dP} + \tau \frac{dD}{dP}}_{\text{Behavioral Cost}},$$

where $bS_0(P) + \theta b[1 - S_0(P)$ is the mechanical cost deriving from UI and bonus spending. Furthermore, extending benefits reduces workers' search effort, thereby increasing UI payment ($b \sum_{t=0}^{P-1} \frac{dS_0(t)}{dP}$), decreasing bonus payment ($-\theta b \sum_{t=0}^{P-1} \frac{dS_0(t)}{dP}$), and increasing the non-employment duration ($\frac{dD}{dP}$). As a result, the government needs to raise taxes in a shorter employment period, to finance increased expenditure caused by job search distortions. Note that the above formula takes into account the re-employment bonus: A marginal increase in benefit duration reduces the bonus payment by θb, so the re-employment bonus reduces the behavioral cost of the UI extension.[26]Combining estimates in Table 4 with formula (7), we can calculate the behavioral cost per NTD of spending on extending the potential benefit duration. Table 4 suggests that extending UI benefits by three months (90 days) increases benefit duration B and non-employment duration D by 57 and 37 days, respectively. Plugging in these estimates, we find that the behavioral cost per NTD of government spending on the UI extension is about 0.07, near the lower end of estimates in the literature (Schmieder and von Wachter, 2016).[27]

6.4. Value of Extended UI Benefits

In this section, we estimate the value of extended UI benefits by comparing the responses in re-employment hazard to extending UI benefits and re-employment bonuses. As shown in Online Appendix C.4, the value of UI is captured by the MRS between consumption when unemployed and employed (Landais and Spinnewijn, 2021). It measures a worker's willingness to pay to move one NTD of consumption when employed to her consumption when unemployed (Hendren, 2017).

Since we do not have data on consumption, we cannot use the MPC approach to provide a lower bound on MRS (Landais and Spinnewijn, 2021) or the consumption-based approach that identify MRS directly (Gruber, 1997). Instead, we compare labor supply responses to an income transfer when unemployed (i.e. extended benefits) and an income transfer when employed (i.e. re-employment bonus), to identify the value of extended benefits. Since workers' responses in the re-employment hazard depend on their marginal utilities of consumption, we can infer the MRS between consumption when unemployed and employed by comparing the differential responses in the re-employment hazard to extended benefits and re-employment bonuses. This argument requires the assumption that consumption responses to income transfers are comparable between unemployed states and employed states. Following Chetty, 2008 and Landais, 2015, we assume workers are liquidity-constrained so that their consumption responses to income transfers when employed and unemployed are the same. However, if the consumption response is larger when unemployed than when employed, our result may provide an upper bound estimate for MRS (Landais and Spinnewijn, 2021).

Decomposition To illustrate our approach to identifying the MRS, we first differentiate the intra-temporal first-order condition (Eq. C.1 in the Online Appendix C.2) with respect to P_θ to obtain $\frac{\partial s_t}{\partial P_\theta}$.

$$\frac{\partial s_t}{\partial P_\theta} = \theta b S_{t+1}(P_\theta) \frac{v'(w-\tau)}{g''(s_t)}; \forall t \leqslant P_\theta. \tag{8}$$

A worker who finds a job at time t expects to receive an additional θb of bonuses. On the other hand, an unemployed worker at time t expects that she will receive an additional $[1 - S_{t+1}(P_\theta)]\theta b$ of bonuses, because she has to find a job before exhausting her benefits to be eligible for bonuses. As increasing the bonus qualification period has a larger effect on the value of employment, i.e. more than the value of unemployment, extending re-employment bonuses increases the marginal benefits of a search; therefore, we predict that extending re-employment bonuses increases job search effort. Importantly, the effect of extending bonus qualification period on a job search is proportional to workers' marginal utility of consumption when employed. We provide a detailed derivation for the effect of extending bonus qualification period on job search in the Online Appendix C.2.

Second, we derive the effect of extending potential benefit duration on a job search. Note that increasing P_b mechanically increases P_θ. Let $P_\theta = P_b = P$, in Appendix C.2, in which case we demonstrate that

$$\begin{aligned} \frac{\partial s_t}{\partial P} &= -bS_{t+1}(P)\frac{u'(c_P^u)}{g''(s_t)} + \theta b S_{t+1}(P)\frac{v'(w-\tau)}{g''(s_t)} \\ &= \frac{\partial s_t}{\partial P_b} + \frac{\partial s_t}{\partial P_\theta}; \forall t \leqslant P. \end{aligned} \tag{9}$$

Eq. (9) shows that the effect of the UI extension on job search in a UI system with bonuses ($\frac{\partial s_t}{\partial P}$) is a combination of two effects: (1) Income transfer when unemployed ($\frac{\partial s_t}{\partial P_b}$), and (2) Income transfer when employed ($\frac{\partial s_t}{\partial P_\theta}$).[28] From Eq. (9), we can see that subtracting

[26] Our formula for the welfare effect of the UI extension is similar to Eq. (1) in Schmieder et al., 2012, but we extend it by incorporating re-employment bonuses into the model.

[27] Specifically, the behavioral cost per NTD of transfer is

$$\begin{aligned} &\frac{1}{bS_0(P)+\theta b[1-S_0(P)]}\left\{b\sum_{t=0}^{P-1}\frac{dS_0(t)}{dP} - \theta b\sum_{t=0}^{P-1}\frac{dS_0(t)}{dP} + \tau\frac{dD}{dP}\right\} \\ = \ &\frac{1}{90 \cdot 0.62 + 45 \cdot 0.38} \cdot [57 - (0.62 \cdot 90) - 0.5 \cdot (57 - (0.62 \cdot 90)) + 0.12 \cdot 37] \\ = \ &0.07. \end{aligned}$$

[28] The formula also shows that the re-employment bonus counteracts the moral hazard effect by offering θ remaining benefits for workers re-employed before the exhaustion point, thereby suggesting that 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.

$\frac{\partial s_t}{\partial P_\theta}$ from $\frac{\partial s_t}{\partial P}$ yields the effect of an unemployment transfer ($\frac{\partial s_t}{\partial P_b}$). Therefore, we can write the MRS during the extended benefit period as follows:

$$\frac{u'(c_P^u)}{v'(c_P^e)} = -\frac{\frac{\partial s_t}{\partial P} - \frac{\partial s_t}{\partial P_\theta}}{2 \cdot \frac{\partial s_t}{\partial P_\theta}}.$$

Eq. (10) shows that we can estimate the MRS by estimating the ratio of the effect of an unemployment transfer (the numerator) to the effect of a re-employment transfer (the denominator). Since the bonus amount is 50% of any remaining entitlement, the effect of a three-month increase in the bonus qualification period has to be multiplied by two, such that it is comparable to the effect of three months' worth of UI benefits. From Eq. (10), we can see that our approach does not require information on the value of the risk aversion coefficient. However, it does require the assumption that other consumption-smoothing means hold fixed when there are changes in income transfers.

In order to implement Eq. (10), we need estimates for $\frac{\partial s_t}{\partial P_\theta}$ and $\frac{\partial s_t}{\partial P}$. We estimate these two unknown statistics using the hazard rate responses to a three-month increase in bonus qualification period (i.e. $180 \times \gamma_1^{Kink}$ of Eq. (2 and a three-month increase in potential benefit duration (i.e. β^{Age45} of Eq. (4)), respectively. Importantly, we recognize that variations from extending the bonus qualification period and the potential benefit duration apply to different individuals at different times. For instance, compared to workers who are affected by the UI extension, those who are affected by the re-employment bonuses tend to have shorter unemployment spells. To address this concern, we re-weight the bonus sample such that it is more comparable to extended benefits sample (column (5) of Table C1), the RK estimate from the re-weighted sample is similar to the estimates from unweighted sample.[29] Plugging in the RD estimate of -0.03 and the RK estimate of 0.019 into Eq. (10), we estimate that the value of one NTD UI transfer at exhaustion is $-\frac{-0.03-\frac{0.019}{2}}{0.019} = 2.08$. On the other hand, although the re-weighted bonus sample is similar to extended benefits sample in observed characteristics, macroeconomic conditions around 2003 are also stronger than in 2009 and 2010. Therefore, we also use the RD estimate from the 2011 sample (column (5) of Table C1) for implementing Eq. (10). The estimated value of extended benefits increase to 2.5, implying workers are willing to pay an additional 1.5 NTD mark-ups to move one NTD of consumption when employed to consumption when unemployed. In the Online Appendix C.5, we take time preferences and incomplete take-up of bonuses into account. We find that the estimated MRS is about 1.96 when we take time preferences into account and around 1.53 when incomplete take-up of bonuses is taken into consideration.

6.5. Welfare Effects of Extending UI Benefits

Having estimated the behavioral costs of extending UI benefits and the MRS, we can evaluate the welfare effects of extending UI benefits, using the MVPF (Hendren and Sprung-Keyser, 2020a), which measures a policy's bang-for-buck through the ratio of the beneficiaries' willingness to pay and the net cost to government. For the extension of UI benefits, the MVPF can be written as

$$MVPF_{EB} = \frac{WTP_{EB}}{1 + FE_{EB}}.$$

The numerator is the beneficiaries' willingness to pay for one NTD spending on the UI extension. The denominator is the marginal cost per NTD spending on the UI extension, which is the sum of one NTD mechanical cost and its fiscal externalities (FE_{EB}) on the government budget. Note that only mechanical transfers are valued for utility-maximizing workers, due to the envelope theorem. As seen in Eq. (7), there are two mechanical transfer components: A mechanical transfer when unemployed ($bS_0(P)$) and a mechanical transfer when re-employed ($\theta b[1 - S_0(P)]$). Suppose the income transfer when re-employed does not offer an additional consumption-smoothing value, we only need estimates for the MRS. Using our estimated behavioral cost (0.07) and the value of extended UI benefits (2.5), the estimated $MVPF_{EB}$ is about 2.01[30]—the welfare gain of one NTD spending on the UI extension is about 2.01 NTD, larger than the available MVPF estimates for UI in the literature (Hendren and Sprung-Keyser, 2020b).[31] In the Online Appendix C.5, we estimate a $MVPF_{EB}$ of 1.33 if the incomplete take-up of bonuses is taken into account.

7. Conclusion

In the 2000s, Taiwan's UI experiences two significant reforms—the re-employment bonus program and the UI extension to older workers. We investigate the effects of these two reforms using the variation in the bonus offer around the time when bonus was introduced, and age discontinuity in eligibility for extended benefits. Since the bonuses and extended UI benefits represent income transfers during re-employment and unemployment, workers' responses in re-employment hazard to the two income transfers allow us to indirectly identify the MRS that captures the value of extended UI benefits. Our estimates suggest the behavioral costs of per NTD of initial spending on bonuses and extending UI benefits are -0.61 and 0.07, respectively. Comparing the effects of bonuses and extended benefits on re-employment hazard, we find that the MRS for extended UI benefits is about 1.5 to 2.5, and the MVPF for extending UI benefits is around 1.3 to 2.

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

[29] 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. We restrict sample to those who are age 40 to 50 and match the following observable characteristics: dummy for male, dummy for being born in Taipei city, dummies for five main industries of previous job, and number of previous UI spells, and job finding rate of the first nine months of a spell.

[30] In the Online Appendix C.3, we also estimate that the MVPF of extending bonus qualification period ($MVPF_{RB}$) is 2.56, which is larger than the estimated $MVPF_{EB}$. Does it suggest that it is welfare enhancing to reduce spending on the UI extension and increase spending on re-employment bonuses? Note that the beneficiary group of extended UI benefits is long-term unemployed, while the bonuses are given to the re-employed workers after a short- term unemployment spell. Therefore, the answer not only depends on the relative MVPFs but also the social welfare weights assigned to beneficiary groups of the two policies (i.e. how citizens and the government feel about the trade-offs), thereby involving making judgment calls.

[31] Specifically, the willingness to pay for the UI extension has two components—the mechanical transfer from extended UI benefits ($bS_0(P)$) and the mechanical transfer from bonuses ($\theta b[1 - S_0(P)]$). Therefore, we estimate that

$$\begin{aligned} WTP_{EB} &= \frac{bS_0(P)}{bS_0(P)+\theta b[1-S_0(P)]} \cdot MRS + \frac{\theta b[1-S_0(P)]}{bS_0(P)+\theta b[1-S_0(P)]} \\ &= \frac{90 \cdot 0.62}{90 \cdot 0.62 + 45 \cdot 0.38} \cdot 2.5 + \frac{45 \cdot 0.38}{90 \cdot 0.62 + 45 \cdot 0.38} \\ &= 2.15, \end{aligned}$$

and the $MVPF_{EB} = 2.15/1.07 = 2.01$.

be longer if the macroeconomic externalities of extended benefits are taken into account.

Finally, recent literature has documented that workers eligible for more UI benefits are more likely to be laid off using administrative data from Brazil (Carvalho et al., 2018; Cravo et al., 2020; Van Doornik et al., 2018). This is an unintended side effect of UI which may add to the cost of extending UI benefits in Taiwan. Like Brazil, Taiwan's UI is not experience-rated, increasing the possibility of workers-firms collusion. While this unintended side effect is an important topic to study, our data only includes individuals who receive UI benefits so that we cannot compare the layoff risk between older and younger workers (around age 45 years). Future studies that incorporates the effect of extending UI benefits on layoff risks into Baily-Chetty framework will be valuable.

Declaration of Competing Interest

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Appendix A. Supplementary material

Supplementary data associated with this article can be found, in the online version, at https://doi.org/10.1016/j.jpubeco.2021.104500.

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