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The Effects of Extended Unemployment Benefits on the Reservation Wage: Evidence from a Benefit Extension in Taiwan

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This paper exploits an unemployment benefit extension for older workers in Taiwan to identify the effect of extended unemployment benefits on the reservation wage. Since 2009, the potential benefit duration for workers aged at least 45 at job loss has been 9 months rather than 6 months for those under 45. Using data from the Taiwanese Manpower Utilization Survey and a difference-in-differences design, I estimate that a three-month increase in potential duration of benefits increases the reservation wage for job losers by 4–5 percent. Moreover, the effect of extended benefits on the reservation wage is mainly driven by the reservation wage response during the extended benefit period and for workers who are most likely to exhaust their benefits. Combining the reduced-form estimates with a search model, my estimates suggest that it is welfare enhancing to increase the generosity of unemployment benefits for middle-aged workers.

Keywords: extended unemployment benefits, reservation wage, difference-in-differences
JEL classification: J64, J65, J68

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1 Introduction

The behavior of reservation wages plays an important role in standard job search models. Mortensen (1977) shows that increasing unemployment insurance (UI) benefits increase workers' value of being unemployed and raise their reservation wages. The increase in the reservation wage due to increased unemployment benefits was once viewed as a cost of increasing UI generosity because a higher reservation wage implies a lower probability of being employed and a longer unemployment duration (Chetty and Finkelstein, 2013; Feldstein and Poterba, 1984). Contrary to the traditional view, Shimer and Werning (2007) show that a higher sensitivity of reservation wages to UI benefits implies a larger welfare gain of UI. On one hand, an increase in UI generosity increases workers' pre-tax reservation wages, which is directly linked to workers' expected value when unemployed. On the other hand, workers have to pay more employment taxes to finance increased unemployment benefits. Therefore, the effect of unemployment benefits on the reservation wage encodes the marginal value of UI, and the after-tax reservation wage strikes the right balance between these two forces.

Empirically, quasi-experimental evidence from France and Germany suggests that increasing the potential duration of unemployment benefits has little impact on the reservation wage in the beginning of the unemployment spell (Le Barbanchon, Rathelot, and Roulet, 2019; Lichter, 2016). Based on the theory of Shimer and Werning (2007), the small reservation wage response to extended unemployment benefits appears inconsistent with the empirical evidence for the consumption-smoothing benefits of UI (Chetty, 2008; Ganong and Noel, 2017; Gruber, 1998; Huang and Yang, 2021; Kolsrud et al., 2018; Landais, 2015).[1] The welfare gain due to increasing this duration should be reliant on the consumption-smoothing benefits during the extended benefit period according to Schmieder and Wachter (2016). However, there are few estimates for the reservation wage response to a benefit extension during an extended benefit period.

In this paper, I exploit a natural experiment in Taiwan to study the effect of extended benefits on the reservation wage and investigate how the effect

[1] Gruber (1998), Chetty (2008), and Landais (2015) show evidence for the consumption-smoothing benefits of increasing the UI benefit level, while Ganong and Noel (2017) and Huang and Yang (2021) evaluate the consumption-smoothing benefits of increasing the potential benefit duration. Kolsrud et al. (2018) assess optimal UI design over the course of an unemployment spell.

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varies over the spell of unemployment. Since 2009, UI recipients aged at least 45 at the time of job loss have been eligible for 9 months of benefits rather than the 6 months available for those under 45. Because the potential duration for workers under 45 remained unchanged following the extension, I consider them as a control group for workers at least 45 years of age. I apply a difference-in-differences design that compares the reservation wage for these two groups of workers before and after 2009 to identify the effect of the benefit extension on the reservation wage.

Using data from the Taiwanese Manpower Utilization Survey (MUS) from 1999 to 2018, I obtain two key findings. First, I estimate that a three-month increase in unemployment benefits increases the reservation wage for job losers by 4–5 percent. I conduct several placebo tests to show that the eligibility for extended benefits is not correlated to age-specific trends or shocks. On one hand, I find that the benefit extension has no effect on workers not eligible for UI. On the other hand, I find no evidence that older and younger workers experienced different trends in terms of the reservation wage prior to the reform. Overall, these placebo tests suggest that the association between the eligibility for extended benefits and reservation wages can be attributed to the effect of extended benefits rather than to age-specific trends or shocks to reservation wages.

Second, the reservation wage response is estimated to be larger during an extended benefit period and for workers who are more likely to exhaust their benefits. In particular, I find that the effect of extended benefits on the reservation wage during the extended benefit period is almost three times larger than during the regular benefit period, consistent with Nekoei and Weber (2017) and Caliendo, Tatsiramos, and Uhlendorff (2013), who show findings that the effect of a benefit extension on accepted wages is stronger around UI exhaustion. In addition, the fact that the larger reservation wage response during the extended benefit period and for workers who are more likely to exhaust their benefits provides an explanation for why the effect of extended benefits on the reservation wage in Taiwan is stronger than in France and Germany. A shorter potential duration in Taiwan (6 or 9 months compared with more than 12 months) increases the likelihood that workers will exhaust their benefits and therefore makes them more responsive to the benefit extension.

The finding that extended benefits significantly increase workers' reservation wages has important theoretical implications. On one hand, according to job search theories, an increase in unemployment benefits reduces

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reemployment hazard through a reduction in search effort (search margin) and an increase in reservation wages (selectivity margin) (Burdett, 1979; Mortensen, 1977; Nekoei and Weber, 2017). While recent studies suggest that the selectivity margin plays a limited role in the effect of extended benefits on reemployment hazard (Le Barbanchon, 2016; Lichter, 2016; Schmieder, von Wachter, and Bender, 2016), my result suggests the selectivity margin can be substantial in a UI system with a short potential benefit duration. On the other hand, I incorporate the effect of extended benefits on the reservation wage, using the theoretical framework proposed by Shimer and Werning (2007) to conduct welfare analysis. On the assumption that the effect of extended benefits on the reservation wage is comparable to that of increasing benefit levels, I estimate that it is welfare-increasing to raise the UI benefit level.

This paper stands apart from the literature on UI in the following ways. First, it is one of the first papers using a quasi-experimental variation in UI benefits to estimate the effect of extended benefits on the reservation wage. Feldstein and Poterba (1984) use data from the Current Population Survey to estimate how unemployment benefits affect reservation wages, finding a large degree of elasticity in the reservation wage with respect to unemployment benefits. On the contrary, Krueger and Mueller (2016) use data from their survey of unemployed workers in New Jersey and find a small and insignificant relationship between the benefit level and the reservation wage.[2] However, the estimates from these studies might be subject to omitted variable bias because their empirical models do not isolate credibly exogenous variation of UI benefits. Le Barbanchon, Rathelot, and Roulet (2019) and Lichter (2016) are the only two research papers that I am aware of that use quasi-experimental empirical designs for the effects of unemployment benefits on the reservation wage. Le Barbanchon, Rathelot, and Roulet (2019) estimate the effects of an increase in potential benefit duration by exploiting a reform that simplifies the eligibility rule for potential benefit duration in France. While the estimates from Le Barbanchon, Rathelot, and Roulet (2019) suggest that the effect of extended benefits on the reservation wage is small for individuals overall, they find a significantly positive effect for low-tenure UI recipients. Lichter (2016) studies a reform in Germany that

[2]In addition, DellaVigna and Paserman (2005) estimates using data from the National Longitudinal Survey of Youth also show a small and insignificant correlation between a dummy of UI receipt and the reservation wage.

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increased the potential duration for older workers. He finds that this change reduced job search effort but had little effect on the reservation wage.

Second, to the best of my knowledge, this is the first study providing evidence regarding the effect of extended benefits on the reservation wage during an unemployment spell. Although previous studies have exploited discontinuous functions of eligibility rules for extended benefits and found that the decline in reemployment hazard due to extended benefits is larger during an extended benefit period than a regular benefit period (Caliendo, Tatsiramos, and Uhlendorff, 2013; Huang and Yang, 2021; Nekoei and Weber, 2017; Schmieder, von Wachter, and Bender, 2016), the effect on the reservation wage over the course of an unemployment spell has not been examined. Indeed, both Le Barbanchon, Rathelot, and Roulet (2019) and Lichter (2016) measure the reservation wage in the beginning of a UI spell. Since the effect of extended benefits on reservation during an extended benefit period should be more relevant for the welfare effect of the extended benefits, this paper evaluates the reservation wage response to the benefit extension not only during the regular benefit period but also during the extended benefit period.

Third, this paper is the first to estimate the effect of extended UI benefits on the reservation wage in Taiwan. Perhaps because of issues with data availability, evidence regarding UI benefits on labor market outcomes in Taiwan is still rare.[3] Hung and Lo (2008) provide a comprehensive survey of previous research in Taiwan and estimate the effect of UI benefits on unemployment duration. Using survey data on UI recipients and a hazard model, they find a higher reemployment hazard during the period of UI receipt. However, the higher reemployment rate during the period of UI receipt could be caused by duration dependence and heterogeneity, making the variation in UI benefits endogenous. I exploit the benefit extension for older workers in Taiwan to mitigate the endogeneity concern and find that extended benefits increase workers' reservation wages. Because a higher reservation wage implies a longer unemployment duration, my estimates are consistent with the consensus estimates that increasing UI generosity increases unemployment duration (Krueger and Meyer, 2002; Schmieder, von Wachter, and Bender,

[3]Without administrative data, an appropriate source of data to study the effects of UI is the MUS. However, using the MUS data to study the effects of UI on unemployment duration and reemployment outcomes is not feasible because the information about the month of interview is hidden, which makes it impossible to calculate monthly reemployment hazard.

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2016).

Finally, this paper is related to the literature on the effect of UI benefits on reemployment wages. Most studies exploiting discontinuous functions of eligibility for extended benefits to identify the effects of extended benefits on reemployment wages have obtained insignificant estimates for the effects of extended benefits on the reemployment wage for overall UI recipients (Card, Chetty, and Weber, 2007; Huang and Yang, 2018; Lalive, 2007; Le Barbanchon, 2016; Van Ours and Vodopivec, 2008). Schmieder, von Wachter, and Bender (2016) and Nekoei and Weber (2017) are notable exceptions to this trend. Schmieder, von Wachter, and Bender (2016) find that a six-month increase in potential duration lowered the reemployment wage by roughly 0.7%, whereas Nekoei and Weber (2017) estimates showed that a nine-week increase in potential duration increased the reemployment wage by about 0.5%. My findings of a substantial reservation wage response to extended benefits do not contradict the small UI wage effects if an unemployed worker's reservation wage declines over her unemployment spell.[4] Based on Huang and Yang (2018), the elasticity of nonemployment duration to potential benefit duration in Taiwan is about 0.3–0.4, larger than the estimates of Schmieder, von Wachter, and Bender (2016) and Nekoei and Weber (2017). Therefore, it is not surprising that Huang and Yang (2018) did not find a wage gain for UI recipients.

The next section introduces the UI extension in Taiwan and the data set. I introduce the econometric framework in Section 3. Section 4 shows the intention-to-treat estimates for the effect of extended benefits on the reservation wage, and Section 5 conducts robustness checks. In Section 6, I estimate the probability of UI receipt and discuss the implications of the empirical results on optimal UI design. Section 7 concludes the paper.

2 Institutional Background and Data

2.1 Institutional Background

UI in Taiwan was inaugurated in January 1999 as part of the Labor Insura-

[4] In the classic job search model of Burdett (1979) and Mortensen (1977), increasing unemployment benefits shift the reservation wage curve outward; the reservation wage declines before the exhaustion point and stays constant afterwards. Learning about the unknown wage distribution (Burdett and Vishwanath, 1988) and skill depreciation (Pissarides, 1992) also can generate a declining reservation wage.

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nce Law. The original unemployment benefits replaced 50% of workers' average insured earnings for at most 3 months. In August 1999, the replacement rate increased to 60%, and the potential duration also increased to 6 months. In January 2003, UI became part of the overall employment insurance program, which is a mandatory national program that offers unemployment benefits, reemployment bonuses, vocational training living allowances, parental leave allowances, and national health insurance premium subsidies. It covers all Taiwanese workers, excluding civil servants and the self-employed.[5] Since May 2009, the maximum duration of benefits has been six months for workers under age 45 at the time of job loss and nine months for those aged 45 or older when they lost their job.[6] The empirical strategy in this paper is to exploit this reform in potential benefit duration to analyze the effect of benefit extension on the reservation wage.

Workers aged 15 to 65 who satisfy three UI eligibility criteria are eligible to receive UI benefits. First, only workers losing their jobs involuntarily or due to the ending of a fixed term contract are eligible. Involuntary separation from employment refers to separation from employment because the insured unit has closed down, relocated, suspended business, dissolved, or filed for bankruptcy. Second, workers have to have at least one year of employment history in the three years prior to the job loss. Third, workers are required to actively search for a job while receiving benefits. Specifically, they have to list at least two job contacts for each continued claim. Additionally, 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 insured period begins (up to the maximum duration of benefits the claimant is entitled to).

2.2 Data and Sample

The data of this study are from the Taiwanese MUS from 1999 to 2018, conducted by the Taiwanese Directorate General of Budget, Accounting, and Statistics, Executive Yuan. The MUS is a supplementary survey of the

[5] Employment insurance is financed by 1% of the monthly insured wage: 20% is imposed on workers, 70% on employers, and the government pays the remaining 10%.

[6] There were two other reforms in 2009: (1) an age limit extension from 60 to 65 and (2) an increase in the replacement rate up to 80% for UI recipients with non-working dependents.

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Manpower Survey that is conducted every May. The Manpower Survey is a national monthly survey of roughly 20,000 households that collects the employment status of each member of the household aged at least 15. Each individual completes the survey for two consecutive months and another two consecutive months the following year. The MUS provides supplementary information on interviewees' current employment, such as their monthly earnings and hours of work, or the potential job opportunities and expected monthly pay of a jobless person. The reason I chose to use data from the MUS is that the reservation wage is not available in the UI administrative records. The MUS asked unemployed workers: "Expected monthly pay should be at least?" I use the answer to this question as a measure of the reservation wage. I transform reservation wages into constant 2016 NTD using the annual CPI from the DGBAS.

To estimate the effect of extended benefits on the reservation wage, an appropriate sample would be workers who are eligible for UI. Since there is no direct information on UI eligibility in the MUS, I follow Rothstein (2011), Farber, Rothstein, and Valletta (2015), and Farber and Valletta (2015), who classify job losers as workers eligible for UI. I apply three restrictions to the dataset to select the necessary sample of job losers potentially eligible to receive UI benefits.

In the MUS data, job losers can be sorted into two types based on their answer to the question "What was the main reason you left the last job?": those whose answer is "Business shrunk or establishment closed" and those who answer "Seasonal or temporary work finished." I consider workers whose answer is something other than these two types as job leavers.[7] I restrict the sample to job losers who answer "Business shrunk or establishment closed" since temporarily laid off workers may not be searching for jobs (Card, Chetty, and Weber, 2007). Second, I restrict the sample to workers who worked in private industry prior to job loss because public or self-employed workers are not eligible for unemployment benefits in Taiwan. Finally, I restrict the sample to workers ages 25–60 since workers over 60 were not eligible for UI before 2009 and workers younger than 25 might not have accumulated sufficient working experience to be eligible for UI. Under these three sample restrictions, the main estimation sample contains 6,221 job losers between 1999 and 2018.

[7] Specifically, those answers include "Unsatisfied to that job," "Ill health," "Got married or gave birth," "Retired," "Busy in housekeeping," or "Others".

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Importantly, losing a job at a private firm is a necessary condition for the eligibility for unemployment in Taiwan. A job loser who did not have at least one year of work experience in the three years prior to layoff would not be eligible for benefits. Moreover, workers eligible for UI benefits might not choose to receive benefits. Therefore, due to data limitations and the incomplete take-up of UI benefits, the difference-in-differences estimates below will be estimates for the effect of extended benefits on the reservation wage for job losers (an intention-to-treat estimate) rather than the effect of extended benefits for UI recipients (a treatment effect on the treated). To recover the treatment effect on the treated, I estimate the ratio of the number of UI claimants to the number of involuntary unemployed workers in Section 6.1. I then divide the intention-to-treat estimate by the percentage of job losers receiving UI benefits to estimate the treatment effect on the treated.

3 Econometric Framework

In this section, I introduce the difference-in-differences design in a regression framework. The difference-in-differences design compares the evolution of the reservation for workers at least 45 years old and those below 45 years of age around the time of the benefit extension. The difference-in-differences estimates will identify the effect of extended benefits on the reservation wage if there were no other reasons for a change in the relative reservation wage of older and younger workers at the time of the extension.

Note that this paper does not use a regression discontinuity (RD) design, which compares workers around the age 45 cutoff. Although the RD helps eliminate the concern that business cycles affect younger and older workers differently, I use the difference-in-differences instead for two reasons. First, the RD design requires many observations around the age 45 cutoff, but the MUS only records individual age by year, limiting the number of observations for an RD design. Second, although UI administrative data contain age information by day, there is no reservation wage information in the UI administrative data. Therefore, I decided to use the difference-in-differences design in this paper.

To implement the difference-in-differences design, I consider the following two-way fixed effects model:

$$\log R_{iat} = \alpha_a + \lambda_t + \beta \mathrm{EB}_{at} + X'_{iat}\delta + \epsilon_{iat}, \tag{1}$$

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where R_{iat} is the reservation wage for worker i who was a years old when interviewed in year t. The age effects, α_a, control for fixed differences between age groups; survey-year fixed effects, λ_t, capture trends in reservation wages common to all age groups. EB_{at} indicates whether worker i is eligible to receive extended benefits, equal to 1 if worker i is at least 45 years of age when interviewed and was interviewed after the reform (2009–2018). In other words, EB_{at} is a variable that interacts whether worker i is at least 45 years of age when interviewed and whether she was interviewed after the reform. X_{iat} are control variables, including gender, marital status, educational attainment, unemployment duration in progress (months), whether an individual lives in Taipei, and previous job characteristics, such as industry, occupation, and firm size. β is the coefficient of interest, which measures the deviation of the trend in the reservation wages of workers at least 45 years of age from the trend of the observed outcomes for those below the age of 45.

I attribute the difference in the evolution of reservation wages between the two groups to the intention-to-treat effect of the extended benefits on the reservation wage if the following identification assumptions hold. First, the defining assumption of the difference-in-differences design is that the reservation wage for job losers aged at least 45 and those under 45 would have exhibited a similar trend in the absence of the benefits extension (after controlling for X_{iat}). This assumption will be violated if the benefits extension is correlated with older-worker-specific trends or shocks to reservation wages. In particular, the Great Recession leads to rarely seen high unemployment rates in 2009 and 2010 (5.85% and 5.21%). If younger workers lower their reservation wages more than older workers during an economic downturn, the difference-in-differences estimates will be upwardly biased. I investigate whether the estimated effect of extended benefits on the reservation wage is confounded by the business cycle (or other age- and time-varying factors) using placebo tests in Section 4.4 and narrower age groups in Section 5.

Second, I assume the benefits extension for older workers has no effect on the reservation wage for younger workers. If extended benefits increase older workers' reservation wages, firms might substitute younger workers for older ones. This will increase the labor demand for younger workers and might increase their reservation wages, attenuating the difference-in-differences estimates toward zero. Since workers just below the age of 45 are more likely to be affected by such spillovers, in next section I use workers younger than 40 as control groups and pretend that there may be spillovers

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from workers at least 45 to those between 40 and 44 years of age. In Section 5, I also report results excluding individuals close to 45 years of age. I find no evidence for the spillover effects of extended benefits.[8]

4 Estimation Results

I next turn to assessing whether and to what extent the benefit extension affects the reservation wage. Before showing the estimated effects of the benefit extension, I present summary statistics and examine the common trend assumption. I focus on the intention-to-treat estimates in this section and estimate the probability of receiving UI benefits to recover the treatment effect on the treated in Section 6.1.

4.1 Descriptive Statistics and Examinations of Pre-Trends

Table 1 displays the summary statistics of the reservation wage (in 2016 NTD) and all individual characteristic variables for job losers age 25–60 between 1999 and 2018. I break down the sample into four groups based on the age 45 cutoff and the year of the reform. The last column displays the simple difference-in-differences estimates for the reservation wage and individual characteristics (column 6 minus column 3). The difference-in-differences estimate suggests that the extended benefits increase the reservation wage for older workers by 1,736 NTD, about 5.2% of the pre-reform average. While some observed characteristics between older and younger workers exhibit differential trends, I provide the following evidence that it does not invalidate the difference-in-differences design.

First, to assess the potential bias from the small amount of selection on observed characteristics, I predict the log reservation wage using the observed characteristics in Table 1. If the small amount of selection on observed characteristics leads to much bias in the estimated effect of extended benefits on the log reservation wage, we should observe that the predicted log reservation wages for older and younger workers are different in their trends. However, this is not the case — I find that older and younger work-

[8] Lalive, Landais, and Zweimüller (2015) study a "three-year" benefit extension for workers at least 50 years old in Austria. They find that this large benefit extension increased the job finding rate for workers ineligible for the extended benefits. This market externality seems less likely in the case of Taiwan's extension because the length of extension is only one-tenth of Austria's.

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Table 1: Descriptive Statistics

	Age 45–60			Age 25–44			
	1999–2008 (1)	2009–2018 (2)	Diff. (3)	1999–2008 (4)	2009–2018 (5)	Diff. (6)	Diff.-in-Diff. (7)
Reservation Wage (2016 NTD)	33,586	31,486	−2,101***	34,275	30,437	−3,837***	1,736**
	(15,667)	(11,671)	(648)	(12,461)	(8,032)	(408)	(754)
Log Reservation Wage	10.37	10.31	−0.060***	10.40	10.29	−0.111***	0.052***
	(0.27)	(0.31)	(0.014)	(0.32)	(0.31)	(0.009)	(0.017)
Age (years)	50.35	50.78	0.437**	34.52	34.56	0.040	0.397
	(4.06)	(4.11)	(0.183)	(5.90)	(5.71)	(0.240)	(0.298)
Female	0.22	0.24	0.017	0.25	0.33	0.080***	−0.064***
	(0.41)	(0.42)	(0.014)	(0.43)	(0.47)	(0.014)	(0.020)
Married	0.75	0.65	−0.098***	0.44	0.32	−0.116***	0.018
	(0.44)	(0.48)	(0.020)	(0.50)	(0.47)	(0.022)	(0.029)
College Degree	0.10	0.17	0.069***	0.19	0.37	0.179***	−0.110***
	(0.30)	(0.38)	(0.013)	(0.39)	(0.48)	(0.017)	(0.021)
Duration in Progress (months)	9.05	7.50	−1.550***	7.68	6.98	−0.708***	−0.842**
	(8.11)	(6.81)	(0.350)	(7.31)	(6.56)	(0.209)	(0.403)
Live in Taipei	0.08	0.10	0.018	0.07	0.09	0.016	0.003
	(0.27)	(0.30)	(0.011)	(0.26)	(0.29)	(0.009)	(0.014)
Live in Kaohsiung	0.07	0.09	0.024	0.08	0.09	0.014	0.010
	(0.25)	(0.29)	(0.015)	(0.27)	(0.29)	(0.011)	(0.018)
Worked in Manufacturing Ind.	0.44	0.33	−0.110***	0.38	0.38	−0.003	−0.107***
	(0.50)	(0.46)	(0.022)	(0.49)	(0.48)	(0.017)	(0.027)

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Table 1: Descriptive Statistics (Continued)

	Age 45–60			Age 25–44			
	1999–2008 (1)	2009–2018 (2)	Diff. (3)	1999–2008 (4)	2009–2018 (5)	Diff. (6)	Diff.-in-Diff. (7)
White Collar	0.18	0.25	0.067***	0.27	0.35	0.075***	−0.007
	(0.38)	(0.47)	(0.018)	(0.44)	(0.48)	(0.011)	(0.020)
Previous Firm Size > 10	0.51	0.56	0.017	0.52	0.50	0.058***	−0.041*
	(0.50)	(0.50)	(0.018)	(0.50)	(0.50)	(0.013)	(0.022)
Predicted Log Reservation Wage	10.37	10.51	0.006	10.40	10.53	−0.003	0.009
	(0.16)	(0.15)	(0.005)	(0.14)	(0.14)	(0.004)	(0.007)
Number of observations	1,115	980	2,095	1,692	2,434	4,126	6,221

Note: This table displays the descriptive statistics for the main variables for job losers age 25–60 from the MUS between 1999 and 2018. I cut the sample by the age 45 cutoff and the year of the reform (2009). Columns 3 and 6 show the differences between average values in post-reform and pre-reform periods for workers age 45–60 and those age 25–44, respectively. Column 7 reports the difference between the values in columns 3 and 6. Job losers are workers who report their reason for unemployment is "Business shrunk or establishment closed". Job leavers are workers who report their reason for unemployment other than "Business shrunk or establishment closed" and "Seasonal or temporary work finished". Standard deviations are in parentheses of columns 1, 2, 4 and 5. Standard errors are in parentheses of columns 3, 6 and 7. All standard errors are clustered at age levels. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.



Figure 1: Log Reservation Wages (Residualized) over Time

Note: Figure 1 shows the residualized log reservation wage for workers age 45–60 and 25–44 using MUS data from 1999 to 2018. I construct the residualized log reservation wage using the residuals in a regression of the log reservation wage on all observed characteristics in Table 1.

ers experience a similar trend in their predicted log reservation wages, suggesting that the small degree of selection on observables has limited impact on the estimated effect of extended benefits on the log reservation wage. Second, the sample using job losers age 35–55 show similar observed characteristics (Table A1), and the difference-in-differences estimates using tighter age groups (discussed in Section 5) are comparable to the results using job losers age 25–60.

Third, the common trend assumption is satisfied visually. Figure 1 shows the residualized log reservation wage from 1999 to 2018 for workers age 45–60 and 25–44, respectively. I construct the residualized log reservation wage using the residuals in a regression of the log reservation wage on all observed characteristics in Table 1. There is no clear difference in the

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residualized log reservation wage between the two age groups prior to 2009, and we can observe that the residualized log reservation wage has been consistently higher for older workers since 2009. As an alternative way to check the common trend assumption, I estimate the following regression using data prior to the benefit extension:

$$\log R_{iat} = \alpha_a + \lambda t + \delta Age45_i \times t + \epsilon_{iat}, \tag{2}$$

where α_a is the age fixed effect. t is a variable controlling for linear time trends, equal to 1 for 1999, 2 for 2000, $\cdots$, and 10 for 2008. I allow the possibility that older and younger workers follow different trends in reservation wages by including $Age45_i \times t$. Therefore, λt represents the baseline trend for workers below age 45 prior to the reform, while $(\lambda + \delta)t$ is the pretreatment trend for workers at least 45. If δ is significantly different from zero, the common trend assumption would be violated. Table 2 presents the estimation results for equation 2. Columns 1 and 2 show the estimates using data from 1999 to 2008 with and without covariates. The reservation wage exhibits a downward trend for younger workers, and the time trend for older workers is not significantly different. Similarly, columns 3 and 4 display the results using shorter sample periods from 2003 to 2008. The reservation wage for younger workers appears to increase over time during this period, and I cannot reject that the trend in the reservation wage for older workers is the same as that for younger ones.

4.2 Main Results

Table 3 presents the estimates of the effect of extended benefits on reservation wages for job losers. Column 1 shows the difference-in-differences estimate without additional controls. This estimate suggests that a three-month increase in potential benefit duration increases the reservation wage by 4.9 percent for job losers. However, it is possible that the benefit extension in 2009 may have induced compositional change in UI claimants. For

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Table 2: Examininations of Pre-Treatment Trends

	(1)	(2)	(3)	(4)
t	−0.021	−0.018	−0.013	−0.009
	(0.002)	(0.002)	(0.003)	(0.003)
$Age45_i \times t$	−0.001	0.001	0.002	−0.001
	(0.004)	(0.004)	(0.006)	(0.006)
Female		−0.250		−0.234
		(0.015)		(0.017)
Married		0.048		0.033
		(0.009)		(0.012)
College Degree		0.110		0.100
		(0.014)		(0.014)
Duration in Progress (months)		−0.0003		−0.0002
		(0.0006)		(0.0007)
Live in Taipei		0.116		0.121
		(0.016)		(0.023)
Live in Kaohsiung		−0.028		−0.031
		(0.017)		(0.022)
Worked in Manufacturing		−0.066		−0.043
		(0.008)		(0.011)
White Collar		0.087		0.094
		(0.014)		(0.017)
Previous Firm Size > 10		0.018		0.019
		(0.009)		(0.013)
R-squared	0.066	0.281	0.046	0.273
Sample size	3,549	3,549	1,992	1,992
Time Period	1999–2008	1999–2008	2003–2008	2003–2008

Note: This table examines whether older and younger workers share a common trend in reservation wage prior to the benefits extension in 2009. I estimate equation 2 using data from the MUS between 1999 and 2018. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

example, if higher-wage workers became more likely to claim UI benefits after the benefit extension, it could have increased the average reservation wage for UI recipients and biased the estimate in column 1.

Since there is no information on previous earnings records in the MUS data, I include a variety of covariates that are correlated to workers' earning ability. Column 2 includes gender, education, martial status, unemployment

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Table 3: Effect of Extended Benefits on Log Reservation Wage for Job Losers (UI-Eligibles)

	(1)	(2)	(3)	(4)	(5)	(6)
EB	0.049	0.056	0.045	0.048	0.046	0.049
	(0.016)	(0.014)	(0.014)	(0.014)	(0.014)	(0.015)
1[Age $\geq$ 45] $\times$ 1[2006 $\leq$ Year $\leq$ 2008]				0.013		0.013
				(0.020)		(0.020)
1[40 $\leq$ Age $<$ 45] $\times$ 1[Year $\geq$ 2009]					0.005	0.005
					(0.015)	(0.015)
Female		−0.225	−0.239	−0.239	−0.239	−0.239
		(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
Married		0.047	0.046	0.046	0.046	0.046
		(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
College Degree		0.184	0.126	0.126	0.126	0.126
		(0.013)	(0.011)	(0.011)	(0.011)	(0.011)
Duration in Progress (months)		0.00016	−0.00005	−0.00004	−0.00005	−0.00004
		(0.00053)	(0.00056)	(0.00056)	(0.00056)	(0.00056)
Live in Taipei		0.135	0.105	0.105	0.105	0.105
		(0.012)	(0.011)	(0.012)	(0.011)	(0.011)
Live in Kaohsiung		−0.031	−0.037	−0.037	−0.037	−0.037
		(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Worked in Manufacturing			−0.048	−0.049	−0.048	−0.048
			(0.006)	(0.006)	(0.006)	(0.006)

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Table 3: Effect of Extended Benefits on Log Reservation Wage for Job Losers (UI-Eligibles) (Continued)

	(1)	(2)	(3)	(4)	(5)	(6)
White Collar			0.098	0.098	0.098	0.098
			(0.010)	(0.010)	(0.010)	(0.010)
Previous Firm Size > 10			0.025	0.025	0.025	0.025
			(0.007)	(0.007)	(0.007)	(0.007)
R-squared	0.070	0.277	0.302	0.302	0.302	0.302

Note: This table shows the estimated effect of extended benefits on the reservation wage for job losers. I estimate equation 1 using 6,221 job losers from the MUS between 1999 and 2018. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

duration in progress (months), and whether workers live in Taipei as covariates.[9] The estimated effect of extended benefits on the reservation wage for job losers increases to 5.6 percent with increased precision. Column 3 further controls for previous industry, occupation, and firm size; this addition to the estimation reduces the estimated benefits extension's effect to 4.5 percent, but the estimate is comparable to the estimates in columns 1 and 2.

The key assumption of the difference-in-differences design is that workers between 25–44 and workers between 45–60 share a parallel trend in reservation wages in the absence of the reform. Column 4 adds an interaction between whether an individual was at least 45 when interviewed (1[Age $\geq$ 45]) and whether the interview year is between 2006 and 2008 (1[2006 $\leq$ Year $\leq$ 2008]). If the two groups of workers experienced differential trends in reservation wages prior to the extension, this coefficient of this interaction term should be different from zero. This is not the case, as the coefficient of this interaction variable is insignificantly different from zero, and the estimated effect of extended benefits on the reservation wage

[9]Unemployment duration in progress is measure in weeks in the MUS. I assume one week is equal to 0.25 month when calculating duration in months.

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changes little.

In addition, if job losers below age 45 are a valid control group, the benefit extension should have no effect on their reservation wages. Column 5 examines the quality of the control group by including an interaction variable: $1[40 \leq \text{Age} < 45] \times 1[\text{Year} \geq 2009]$. The estimated effects of the benefit extension on workers between ages 40–45 is small and statistically indistinguishable from zero, offering support for the quality of the control group. Column 6 includes both interaction variables, and the results are similar to those when entering each variable separately. Overall, the results in columns 4–6 offer preliminary support for the internal validity of the difference-in-differences estimates.

4.3 Reservation Wage Responses over Unemployment Spell

The significant and positive effect of the benefit extension on the the reservation wage in Taiwan is at odds with the small and insignificant effects found in Le Barbanchon, Rathelot, and Roulet (2019) and Lichter (2016). I argue that this is related to to the fact that the potential benefit duration in Taiwan is shorter than in France and Germany, and the reservation wages in their studies are recorded at the beginning of the UI spell.[10] I offer empirical support for this explanation by showing that (1) the effect of extended benefits on the reservation wage is stronger during the extended benefit period (7th–9th month) than during the regular benefit period (1st–6th month), and (2) the effect of extended benefits on the reservation wage is stronger for those who are more likely to exhaust their benefits.

I estimate variants of equation 1, which distinguishes the effects in the regular benefit and extended benefit periods:

[10] The mean potential duration of Le Barbanchon, Rathelot, and Roulet (2019) sample is more than 12 months (413 days), while Lichter (2016) examines the effects of increasing potential duration from 12 to 15 months. Both are significantly longer than the 6–9 month (180–270 days) potential duration of Taiwan's UI.

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$$\log R_{iat} = \alpha_a + \lambda_t + \beta \mathrm{EB}_{at} + \psi \mathrm{EB}_{at} \times 1[30 < \mathrm{Dur.} \leq 43] + \theta \mathrm{EB}_{at} \times 1[\mathrm{Dur.} > 43] + X'_{iat}\delta + \epsilon_{iat}, \quad (3)$$

According to UI administrative data from the Bureau of Labor Insurance, the median number of days between workers losing their jobs and the first date of UI receipt is about 24 days (about 3–4 weeks) for UI recipients age 25–60 between 1999 and 2012. Therefore, I categorize 31–43 weeks as the extended benefit period. As seen in column 2 of Table 4, the benefit extension is estimated to increase the reservation wage during the regular benefit period (1st–6th month) by 3.6 percent. Moreover, the effect of the benefit extension on the reservation wage is 5.7 percent larger during the extended benefit period than during the regular benefit period, implying an effect of 9.3 percent during the extended benefit period.

One important concern about the estimated response to extended benefits over the unemployment spell is a dynamic selection problem. Since the benefit extension increases unemployment duration (Huang and Yang, 2021), it may change the composition of unemployed workers over the unemployment spell. The dynamic selection problem arises if the composition of the treatment group and the control group changes differently, leading to biased estimates for the reservation wage responses to extended benefits over the spell. In Table A2 of the Online Appendix, I investigate whether the benefit extension changes the difference between the treatment group and the control group during the extended benefit period. Column 7 shows that the dynamic selection problem is limited—the majority of the observed differences between the treatment and the control group during the extended benefit period do not change significantly after the extension. Furthermore, the predicted log reservation wages for the treatment and the control group also change in a similar way after the extension, offering some evidence that the results in column 2 of Table 4 are not driven by dynamic selection.

While the above analysis suggests that there is no evidence for the selection for observables, I cannot rule out the possibility of the selection of un-

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Table 4: Effect of Extended Benefits on Log Reservation Wage for Job Losers over Unemployment Spells

	(1)	(2)	(3)
EB	0.045	0.036	0.041
	(0.013)	(0.015)	(0.013)
EB × 1[30 < Duration ≤ 43]		0.057	
		(0.028)	
EB × 1[Duration > 43]		0.021	
		(0.023)	
EB × Predicted Exhaustion Rate			0.221
			(0.093)
1[30 < Duration ≤ 43]		−0.014	
		(0.010)	
1[Duration > 43]		−0.006	
		(0.010)	
Female	−0.239	−0.239	−0.236
	(0.013)	(0.013)	(0.013)
Married	0.046	0.045	0.049
	(0.006)	(0.006)	(0.006)
College Degree	0.126	0.126	0.127
	(0.011)	(0.011)	(0.011)
Live in Taipei	0.105	0.106	0.106
	(0.011)	(0.011)	(0.011)
Live in Kaohsiung	−0.037	−0.037	−0.039
	(0.011)	(0.011)	(0.011)
Worked in Manufacturing	−0.048	−0.049	−0.049
	(0.006)	(0.006)	(0.006)
White Collar	0.098	0.098	0.096
	(0.010)	(0.010)	(0.010)
Previous Firm Size > 10	0.025	0.025	0.024
	(0.007)	(0.007)	(0.007)
R-squared	0.302	0.302	0.303

Note: This table shows the estimated effect of extended benefits on the reservation wage for job losers over unemployment spells. The estimated samples are 6,221 job losers from the MUS between 1999 and 2018. Column 1 replicates the results from column 3 of Table 3 for comparison. Columns 2 estimates equation 3 to investigate how the effect of extended benefits on reservation wage varies over the uenmployment spell. Column 3 estimates equation 4 to examine how the effect of extended benefits on reservation wage varies with predicted exhaustion rates. I demean the predicted exhaustion rate such that the interpretation of the coefficient on EB is in reference to the average predicted exhaustion rate. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

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observables. To avoid the concern about dynamic selection issues, I examine whether the effect of extended benefits on the reservation wage is stronger for those who are more likely to exhaust their benefits. Specifically, I predict the exhaustion rate for each worker using data prior to the extension and estimate a linear probability model of whether a worker's unemployment duration in progress is longer than 30 weeks on worker characteristics and previous job characteristics. Column 3 of Table 4 examines how the effect of extended benefits on the reservation wage varies with predicted exhaustion rates by estimating the following equation:

$$\log R_{iat} = \alpha_a + \lambda_t + \beta \mathrm{EB}_{at} + \theta \mathrm{EB}_{at} \times \text{Predicted Exhaustion Rate} + X'_{iat}\delta + \epsilon_{iat}, \tag{4}$$

where I demean the predicted exhaustion rate so the interpretation of the β coefficient is in reference to the average predicted exhaustion rate. The estimate of β suggests an effect of 4.1 percent for a worker at average predicted exhaustion rate. Meanwhile, the estimate of θ implies that a 10 percentage point rise in the likelihood of exhausting benefits increases the effect of extended benefits on the reservation wage by 2.21 percent.

The findings shown in Table 4 are consistent with the theoretical predictions from the job search models in Card, Chetty, and Weber (2007) and Nekoei and Weber (2017), which predict that workers' behavioral responses to extended benefits should be larger closer to the exhaustion point and for workers who are more likely to exhaust their benefits. Furthermore, these findings offer an explanation for why the effect of extended benefits on the the reservation wage in Taiwan is stronger than in France and Germany — a shorter potential duration in Taiwan (6 or 9 months compared with more than 12 months) increases the likelihood that workers will exhaust their benefits and therefore makes them more responsive to the benefit extension.

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4.4 Placebo Tests

One important concern for the difference-in-differences estimates is that the estimates would be biased up if the 2009 financial crisis lowers the reservation wage more for younger workers than for older workers. In this subsection, I conduct several placebo tests to alleviate this concern. First, I estimate the effect of extended benefits on the reservation wage for workers who voluntarily leave their jobs. If the positive estimated effect of extended benefits on the reservation wage for job losers is driven by unobserved age-specific shocks, those shocks would also likely induce a positive correlation between the EB dummy and the reservation wage for job leavers. As shown in Table 5, in all three specifications, the estimated association between the treatment and the reservation wage (after controlling for age and year effects) for job leavers is economically small and statistically indistinguishable from zero at conventional significance levels. This placebo test supports the identification assumption that the eligibility for extended benefits is orthogonal to unobserved determinants, such as macroeconomic conditions, of the reservation wage.

Although Table 5 provides some reassurance that the estimated effect of extended benefits on the reservation wage for job losers is not driven by unobserved determinants of the reservation wage, it is possible that these unobserved determinants affect job losers and job leavers differently. To alleviate this concern, in Table 6, I estimate model 1 using job losers younger than 45 and that pretending the potential benefit duration increased for workers age at least 35 years of age (at least 40) in column 1 (column 2). Neither of the estimates is statistically different from zero, suggesting that the financial crisis did not lower the reservation wage more for younger workers than for older workers.

In addition, to examine whether older and younger workers follow similar trends in the reservation wage prior to reform, I conduct a falsification exercise using data before the reform and pretend that the benefit exten-

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Table 5: Effect of Extended Benefits on Log Reservation Wage for Job Leavers (UI-Ineligibles)

	(1)	(2)	(3)
EB	−0.016	0.015	0.008
	(0.022)	(0.017)	(0.017)
Female		−0.163	−0.178
		(0.012)	(0.013)
Married		0.038	0.032
		(0.009)	(0.009)
College Degree		0.195	0.138
		(0.013)	(0.009)
Duration in Progress (months)		−0.0012	−0.0013
		(0.0005)	(0.0005)
Live in Taipei		0.103	0.087
		(0.014)	(0.013)
Live in Kaohsiung		−0.023	−0.023
		(0.013)	(0.012)
Worked in Manufacturing			−0.002
			(0.008)
White Collar			0.101
			(0.011)
Previous Firm Size > 10			0.030
			(0.007)
R-squared	0.084	0.275	0.302

Note: This table shows the estimated effect of extended benefits on the reservation wage for job leavers. Job leavers are workers who report their reason for unemployment as something other than "Business shrunk or establishment closed" and "Seasonal or temporary work finished". I estimate equation 1 using 5,419 job leavers from the MUS between 1999 and 2018. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

sion was enacted in 2003 (Column 3) or 2006 (Column 4). Since the benefits had not been extended for older workers before 2009, the placebo estimates should not be significantly different from zero. However, if the placebo estimates are significantly positive, it would suggest that the estimation results are partly driven by differential reservation wage trends for older

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Table 6: Effect of Extended Benefits on Log Reservation Wage: Placebo Tests

	(1)	(2)	(3)	(4)
1[Age ≥ 35] × 1[Year ≥ 2009]	−0.002			
	(0.013)			
1[Age ≥ 40] × 1[Year ≥ 2009]		0.004		
		(0.016)		
1[Age ≥ 45] × 1[Year ≥ 2003]			0.002	
			(0.019)	
1[Age ≥ 45] × 1[Year ≥ 2006]				0.008
				(0.020)
Female	−0.203	−0.246	−0.251	−0.235
	(0.016)	(0.017)	(0.015)	(0.015)
Married	0.046	0.047	0.047	0.032
	(0.008)	(0.010)	(0.009)	(0.012)
College Degree	0.105	0.120	0.109	0.100
	(0.011)	(0.015)	(0.014)	(0.015)
Duration in Progress (months)	−0.0004	−0.0010	−0.0002	−0.0003
	(0.0006)	(0.0007)	(0.0006)	(0.0008)
Live in Taipei	0.098	0.110	0.117	0.122
	(0.011)	(0.012)	(0.016)	(0.023)
Live in Kaohsiung	−0.025	−0.031	−0.029	−0.029
	(0.009)	(0.010)	(0.017)	(0.022)
Worked in Manufacturing	−0.049	−0.059	−0.066	−0.045
	(0.008)	(0.010)	(0.008)	(0.010)
White Collar	0.075	0.093	0.088	0.094
	(0.009)	(0.010)	(0.014)	(0.017)
Previous Firm Size > 10	0.034	0.039	0.019	0.019
	(0.009)	(0.012)	(0.009)	(0.013)
R-squared	0.295	0.333	0.285	0.275
Age Groups	20–44	30–44	25–60	25–60
Time Periods	1999–2018	1999–2018	1999–2008	2003–2008
Treatment Group	35–44	40–44	45–60	45–60
Sample Size	4,677	3,098	3,549	1,992

Note: This table conducts four falsification exercises to examine the validity of difference-in-differences design. Column 1 uses job losers between ages 20–44 and pretends the potential benefit duration increased since 2009 for workers age 35 or above. Column 2 uses job losers between ages 30–44 and pretends the potential benefit duration increased since 2009 for workers age 40 or above. Column 3 uses data from 1999 to 2008 and pretends the potential benefit duration increased for workers age 45 or older since 2003. Column 4 uses data from 2003 to 2008 and pretends the potential benefit duration increased for workers age 45 or above since 2006. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

and younger workers. In particular, there was a recession in 2003, and the placebo estimates would be positive if younger workers were more affected

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by the recession. Columns 3 and 4 show that the reservation wage for older workers exhibits a comparable trend to that for younger workers, increasing the credibility of the common trend assumption.

In Table 7, I present the difference-in-difference-in-differences (DDD) estimates that use job leavers as an additional control group. The coefficient on EB × Job Loser captures the reservation wage effect of extended benefits on job losers relative to that on job leavers, while the coefficient on EB now represents the reservation wage effect of extended benefits on job leavers. The DDD estimates suggest that the benefit extension increases the reservation wage by 4–6 percent, and they are significantly different from zero at a 5% significance level.[11]

In sum, my main results suggest that a three-month increase in the potential duration of benefits is associated with a 4–5 percent increase in the reservation wage for job losers. The placebo exercises provide evidence showing that this association is driven by the eligibility for extended benefits rather than by differential trends or shocks for older and younger workers.

5 Robustness Checks

In this section, I conduct extensive robustness checks to solidify the empirical results.

5.1 Pseudo Panel of MUS

First, I estimate equation 1 using the MUS Pseudo Panel Data, which contains data on individuals for two consecutive years.[12] While the MUS Pseudo Panel Data includes fewer observations, it is useful to control for

[11] Note that the DDD estimates are less precise than the difference-in-differences estimates because the triple difference introduces an additional source of noise.

[12] The MUS Pseudo Panel Data are available from the Survey Research Data Archive, Academia Sinica. It is a pseudo panel data set because the observations in two consecutive years are matched using individual characteristics rather than individual IDs, which are not available in the MUS.

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Table 7: Effect of Extended Benefits on Log Reservation Wage: Difference-in-Difference-in-Differences

	(1)	(2)	(3)
EB× Job Loser	0.062	0.041	0.043
	(0.028)	(0.020)	(0.021)
EB	−0.013	0.016	0.005
	(0.022)	(0.017)	(0.017)
Job Loser	−0.027	−0.020	−0.008
	(0.009)	(0.007)	(0.007)
1[Age ≥ 45]× Job Loser	−0.010	−0.002	−0.002
	(0.016)	(0.016)	(0.016)
1[Year ≥ 2009]× Job Loser	0.007	0.010	0.002
	(0.012)	(0.009)	(0.009)
Female		−0.193	−0.209
		(0.013)	(0.013)
Married		0.044	0.041
		(0.006)	(0.005)
College Degree		0.191	0.134
		(0.011)	(0.008)
Duration in Progress (months)		−0.0004	−0.0006
		(0.0004)	(0.0004)
Live in Taipei		0.117	0.0926
		(0.009)	(0.009)
Live in Kaohsiung		−0.027	−0.030
		(0.008)	(0.009)
Worked in Manufacturing			−0.030
			(0.004)
White Collar			0.100
			(0.008)
Previous Firm Size > 10			0.025
			(0.005)
Sample Size	11,640	11,640	11,640
R-squared	0.072	0.270	0.294

Note: This table shows the DDD estimates for the effect of extended benefits on the reservation wage for job losers using data from the MUS between 1999 and 2018. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

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Table 8: Effect of Extended Benefits on Log Reservation Wage: Using Pseudo Panel of MUS

	(1)	(2)	(3)	(4)
EB	0.046	0.052	0.066	0.058
	(0.021)	(0.018)	(0.020)	(0.020)
Female		−0.241	−0.219	−0.227
		(0.015)	(0.017)	(0.018)
College		0.185	0.149	0.120
		(0.015)	(0.018)	(0.016)
Married		0.050	0.046	0.048
		(0.010)	(0.010)	(0.009)
Duration in Progress (months)		−0.00004	0.003	0.003
		(0.00059)	(0.002)	(0.002)
Live in Taipei		0.111	0.086	0.072
		(0.022)	(0.024)	(0.022)
Live in Kaohsiung		−0.023	−0.032	−0.037
		(0.015)	(0.021)	(0.021)
Log Previous Wage			0.136	0.129
			(0.048)	(0.047)
Worked in Manufacturing				−0.044
				(0.013)
White Collar				0.052
				(0.018)
Previous Firm Size > 10				0.011
				(0.012)
Sample Size	2,686	2,686	1,705	1,705
R-squared	0.045	0.264	0.351	0.361

Note: This table shows the estimated effect of extended benefits on the reservation wage for job losers. I estimate equation 1 using data from the Pseudo MUS Panel Data between 1999-2018. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

previous earnings, which is an effective way to control individual heterogeneity (Heckman and Hotz, 1989). Specifically, I include workers' monthly earnings in the first year as a control variable. Table 8 reports the difference-in-differences estimates using workers aged 25–60 between 1999 and 2018

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from the MUS Pseudo Panel Data. The estimated effect of extended benefits on the reservation wage is similar to the estimate in Table 3.

5.2 Narrower Age Groups and Shorter Sample Periods

Second, I examine the sensitivity of my estimates to narrower age groups or shorter sample periods. Table 9 shows the results using workers aged 25–60, 35–55, and 40–50, respectively. These results eliminate the influence of younger workers, who might not have accumulated sufficient work experience to be eligible for UI, and older workers, who are close to retirement. Both estimated effects of extended benefits on the reservation wage using the MUS and the Pseudo MUS are robust to limiting the sample to tighter age groups — the magnitudes of the estimates are similar and not statistically distinguishable from each other. The fact that the difference-in-differences estimates are robust to using narrower age groups offers additional evidence that the estimated impact of extended benefits on the reservation wage is not driven by the possibility that different age groups were differentially affected by the recession.

Table 10 investigates the robustness of the difference-in-differences estimates to shorter time periods. Columns 1–3 report the estimates using MUS, and columns 4–6 present the estimates using the Pseudo MUS. Although the standard errors become larger when fewer years are used to estimate the effect of extended benefits on the reservation wage, these estimates are in the same ballpark and insignificantly different from each other.

5.3 Absolute Levels of Reservation Wage as the Dependent Variable

Third, in Table 11, I replace the natural logarithm of the reservation wage with the absolute value of the reservation wage as the dependent variable in equation 1.[13] The estimates are consistent with the results using the natural

[13] Note that the reservation wage is inflation adjusted using the annual CPI (2016 constant NTD).

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Table 9: Effect of Extended Benefits on Log Reservation Wage: Using Narrower Age Groups

	MUS			Pseudo MUS		
	(1)	(2)	(3)	(4)	(5)	(6)
EB	0.045	0.038	0.036	0.058	0.058	0.060
	(0.014)	(0.014)	(0.019)	(0.020)	(0.026)	(0.042)
Female	−0.239	−0.279	−0.283	−0.227	−0.242	−0.258
	(0.013)	(0.008)	(0.010)	(0.018)	(0.022)	(0.022)
Married	0.046	0.041	0.043	0.048	0.046	0.049
	(0.006)	(0.009)	(0.012)	(0.009)	(0.012)	(0.015)
College	0.126	0.156	0.133	0.120	0.173	0.183
	(0.010)	(0.014)	(0.018)	(0.016)	(0.022)	(0.033)
Duration in Progress (months)	−0.00005	−0.0009	−0.0016	0.003	0.003	0.0008
	(0.00056)	(0.0008)	(0.0009)	(0.002)	(0.002)	(0.0025)
Live in Taipei	0.105	0.097	0.100	0.072	0.047	0.041
	(0.011)	(0.015)	(0.020)	(0.022)	(0.030)	(0.043)
Live in Kaohsiung	−0.037	−0.046	−0.037	−0.037	−0.050	−0.051
	(0.011)	(0.016)	(0.011)	(0.021)	(0.029)	(0.032)
Worked in Manufacturing	−0.048	−0.059	−0.056	−0.044	−0.049	−0.039
	(0.006)	(0.006)	(0.009)	(0.013)	(0.014)	(0.019)
White Collar	0.098	0.115	0.128	0.052	0.046	0.054
	(0.010)	(0.014)	(0.017)	(0.018)	(0.030)	(0.039)

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Table 9: Effect of Extended Benefits on Log Reservation Wage: Using Narrower Age Groups(Continued)

	MUS			Pseudo MUS		
	(1)	(2)	(3)	(4)	(5)	(6)
Previous Firm Size > 10	0.025	0.033	0.033	0.011	0.022	0.016
	(0.007)	(0.009)	(0.011)	(0.012)	(0.015)	(0.018)
Log Previous Wage				0.129	0.147	0.130
				(0.047)	(0.060)	(0.062)
Observations	6,221	3,846	2,444	1,705	1,028	672
R-squared	0.302	0.346	0.342	0.361	0.406	0.379
Age 25–60	Yes	–	–	Yes	–	–
Age 35–55	–	Yes	–	–	Yes	–
Age 40–50	–	–	Yes	–	–	Yes

Note: This table shows the estimated effect of extended benefits on the reservation wage for job losers using narrower age groups. I estimate equation 1 using data from the MUS (columns 1–3) and the MUS Pseudo Panel (columns 4–6) between 1999 and 2018. Column 1 and column 4 replicate the results from column 3 of Table 3 and column 1 of Table 8 for comparison. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

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Table 10: Effect of Extended Benefits on Log Reservation Wage: Using Shorter Time Periods

	MUS			Pseudo MUS		
	(1)	(2)	(3)	(4)	(5)	(6)
EB	0.045	0.031	0.048	0.058	0.055	0.106
	(0.014)	(0.020)	(0.027)	(0.020)	(0.024)	(0.038)
Female	−0.239	−0.232	−0.224	−0.227	−0.180	−0.188
	(0.013)	(0.015)	(0.015)	(0.018)	(0.016)	(0.015)
Married	0.046	0.041	0.050	0.048	0.006	0.026
	(0.006)	(0.009)	(0.011)	(0.009)	(0.013)	(0.013)
College	0.126	0.124	0.124	0.120	0.077	0.072
	(0.011)	(0.012)	(0.015)	(0.016)	(0.024)	(0.029)
Duration in Progress (months)	−0.00005	−0.0010	−0.0017	0.003	0.004	0.0003
	(0.00056)	(0.0007)	(0.0009)	(0.002)	(0.002)	(0.0025)
Live in Taipei	0.105	0.010	0.102	0.072	0.084	0.084
	(0.011)	(0.019)	(0.021)	(0.022)	(0.026)	(0.028)
Live in Kaohsiung	−0.037	−0.027	0.008	−0.037	−0.025	0.003
	(0.011)	(0.018)	(0.017)	(0.021)	(0.029)	(0.027)
Worked in Manufacturing	−0.048	−0.038	−0.048	−0.044	−0.045	−0.056
	(0.006)	(0.008)	(0.010)	(0.013)	(0.018)	(0.022)
White Collar	0.098	0.111	0.095	0.052	0.029	0.025
	(0.010)	(0.013)	(0.013)	(0.018)	(0.018)	(0.020)

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Table 10: Effect of Extended Benefits on Log Reservation Wage: Using Shorter Time Periods (Continued)

	MUS			Pseudo MUS		
	(1)	(2)	(3)	(4)	(5)	(6)
Previous Firm Size > 10	0.025	0.039	0.054	0.011	0.025	0.027
	(0.007)	(0.009)	(0.012)	(0.012)	(0.014)	(0.017)
Log Previous Wage				0.129	0.325	0.316
				(0.047)	(0.034)	(0.040)
Year 1999–2018	Yes	–	–	Yes	–	–
Year 2003–2014	–	Yes	–	–	Yes	–
Year 2007–2010	–	–	Yes	–	–	Yes

Note: This table shows the estimated effect of extended benefits on the reservation wage for job losers using shorter time periods. I estimate equation 1 using data from the MUS (columns 1–3) and the MUS Pseudo Panel (columns 4–6) between 1999 and 2018. For comparison, column 1 and column 4 replicate the results from column 3 of Table 3 and column 1 of Table 8, respectively. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

Table 11: Effect of Extended Benefits on Log Reservation Wage: Using Absolute Level of Reservation Wage

	(1)	(2)	(3)
EB	1,665	1,963	1,571
	(712)	(648)	(639)
Female		−6,934	−7,429
		(431)	(425)
Married		2,112	2,088
		(256)	(256)
College Degree		6,506	4,558
		(573)	(471)
Duration in Progress (months)		−4.62	−11.22
		(21.23)	(21.94)
Live in Taipei		5,279	4,227
		(567)	(527)
Live in Kaohsiung		−1,127	−1,322
		(357)	(375)
Worked in Manufacturing			−1,859
			(267)
White Collar			3,328
			(395)
Previous Firm Size > 10			863
			(251)
R-squared	0.052	0.181	0.198

Note: This table shows the estimated effect of extended benefits on the reservation wage for job losers. I replace the log reservation wage with the absolute level of reservation wages in equation 1 and estimate the model using data from the MUS between 1999-2018. Column 1 replicates the results from column 3 of Table 3 for comparison. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

logarithm of the reservation wage as the dependent variable — the benefit extension increases the reservation wage by 1,571–1,963 NTD, about 4.7–5.8% of the baseline mean (33,586 NTD).

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5.4 Additional Sample Restrictions

Finally, in Table 12, I employ a variety of sample restrictions to assess the robustness of the empirical results. For these additional results, I present estimates using the same specification as in column 3 of Table 3.

In column 1, I replicate the results from column 3 of Table 3 for com parison. Column 2 checks the sensitivity of my estimate to the exclusion of job losers who were 45 when interviewed. The eligibility for extended benefits depends on workers' age at job loss rather than on their age when interviewed. Since some workers were 45 when interviewed but lost their job before they turned 45, these workers were not eligible for extended benefits and may contaminate the treatment group. Fortunately, the estimate changes little when job losers who were 45 when interviewed are excluded. Furthermore, the estimate increases to 4.8 percent when excluding workers between 42 and 47, but it is not meaningfully different from the baseline estimate, mitigating the concern that workers just below 45 are affected by the benefit extension. In column 4, I exclude the sample year of 2009 to mitigate the influence of the U.S. financial crisis. The estimated coefficient is insensitive to the removal of the observations in 2009. In column 5, I exclude job losers who have been unemployed for more than 40 weeks (about 9 months). These workers were likely to have exhausted their benefits and unlikely to be affected by extended benefits. Again, the estimate changes little. The estimated coefficient decreases to 4 percent but remains similar to the main result.

6 Discussion

6.1 The Probability of UI Receipt

So far, I have estimated that the extended benefits increase the reservation wage for job losers from the private sectors by 4–5 percent. Note that this estimate is an intention-to-treat effect. To arrive at the average treatment

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Table 12: Effect of Extended Benefits on Log Reservation Wage: Additional Robustness Checks

	(1)	(2)	(3)	(4)	(5)
EB	0.045	0.048	0.061	0.045	0.040
	(0.014)	(0.014)	(0.015)	(0.014)	(0.017)
Female	−0.239	−0.237	−0.226	−0.241	−0.252
	(0.013)	(0.013)	(0.015)	(0.013)	(0.016)
Married	0.046	0.047	0.049	0.045	0.044
	(0.006)	(0.006)	(0.007)	(0.007)	(0.008)
College Degree	0.126	0.127	0.123	0.124	0.105
	(0.011)	(0.011)	(0.012)	(0.012)	(0.010)
Duration in Progress (months)	−0.00005	0.00009	0.0004	0.0002	0.0033
	(0.00056)	(0.00056)	(0.0006)	(0.0006)	(0.0016)
Live in Taipei	0.105	0.103	0.104	0.104	0.109
	(0.011)	(0.011)	(0.014)	(0.014)	(0.014)
Live in Kaohsiung	−0.037	−0.039	−0.036	−0.046	−0.027
	(0.011)	(0.011)	(0.014)	(0.013)	(0.015)
Worked in Manufacturing	−0.048	−0.048	−0.039	−0.050	−0.047
	(0.006)	(0.006)	(0.007)	(0.008)	(0.008)
White Collar	0.098	0.095	0.090	0.100	0.104
	(0.010)	(0.010)	(0.011)	(0.012)	(0.011)
Previous Firm Size > 10	0.025	0.025	0.023	0.022	0.034
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)

Table 12: Effect of Extended Benefits on Log Reservation Wage: Additional Robustness Checks (Continued)

	(1)	(2)	(3)	(4)	(5)
R-squared	0.302	0.300	0.291	0.298	0.311
Sample Size	6,221	6,028	4,776	5,268	4,341
Exclude Workers Age 45	–	Yes	–	–	–
Exclude Workers Age 42–47	–	–	Yes	–	–
Exclude Year 2009	–	–	–	Yes	–
Exclude Duration > 9 Months	–	–	–	–	Yes

Note: This table conducts additional robustness checks for the estimated effect of extended benefits on the reservation wage for job losers. I estimate equation 1 using data from the MUS between 1999–2018. Column 1 replicates the results from column 3 of Table 3 for comparison. Each specification controls for age and year fixed effects. Standard errors in parentheses are clustered at age level.

effect on the treated, it must be divided by the probability of UI receipt for job losers from private industries. I estimate the probability of UI receipt using the formula from Blank and Card (1991):

$$\text{UI Receipt Rate} = \frac{\text{Fraction of Insured Unemployment}}{\text{Fraction of Eligible Unemployment}}.$$

The fraction of insured unemployment (FIU) is the fraction of unemployed workers who receive unemployment benefits. The fraction of eligible unemployment (FEU) is the fraction of unemployed workers potentially eligible for unemployment benefits. I collect the number of UI recipients from the Bureau of Labor Insurance, while the total unemployment and involuntary unemployment numbers are taken from the Directorate General of Budget, Accounting and Statistics, Executive Yuan.

Table 13 and Figure 2 show the estimates for FIU, FEU, and the probability of UI receipt from 1999 to 2018. It is clear that the estimates for all three are higher during the recession in the early 2000s and the Great Recession. Overall, since the FIU increased slightly over the past 20 years

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Table 13: Estimated Take-Up Rates, Fraction of Insured Unemployment, and Fraction of Eligible Unemployment over Time

Year	Fraction of Insured Unemployment	Fraction of Eligible Unemployment	Probability of UI Receipt
1999	1.5	24.6	6.1
2000	3.7	25.7	14.3
2001	11.1	39.3	28.3
2002	11.6	41.6	27.8
2003	6.5	37.7	17.1
2004	4.7	28.1	16.8
2005	6.0	25.7	23.4
2006	6.9	23.6	29.2
2007	7.3	25.8	28.4
2008	8.8	30.0	29.5
2009	17.3	49.1	35.3
2010	8.4	36.9	22.6
2011	5.9	26.5	22.5
2012	6.6	25.9	25.5
2013	7.1	26.2	27.0
2014	6.4	25.5	25.1
2015	7.0	22.5	31.0
2016	7.2	23.4	31.0
2017	8.0	23.9	33.6
2018	8.3	20.3	40.9

Note: Data are publicly available from the Bureau of Labor Insurance and Directorate-General of Budget, Accounting, and Statistics of Executive Yuan.

and the FEU has gradually declined, the probability of UI receipt follows an upward trend. In 1999, when unemployment benefits were first offered to unemployed workers, the probability of UI receipt was lower than 0.1, rising to 0.4 in 2018. On average, the estimated probability of UI receipt is 0.26 between 1999 and 2018 and 0.29 between 2009 and 2018.[14]

[14]The estimated probability of UI receipt among job losers in Taiwan is on average smaller

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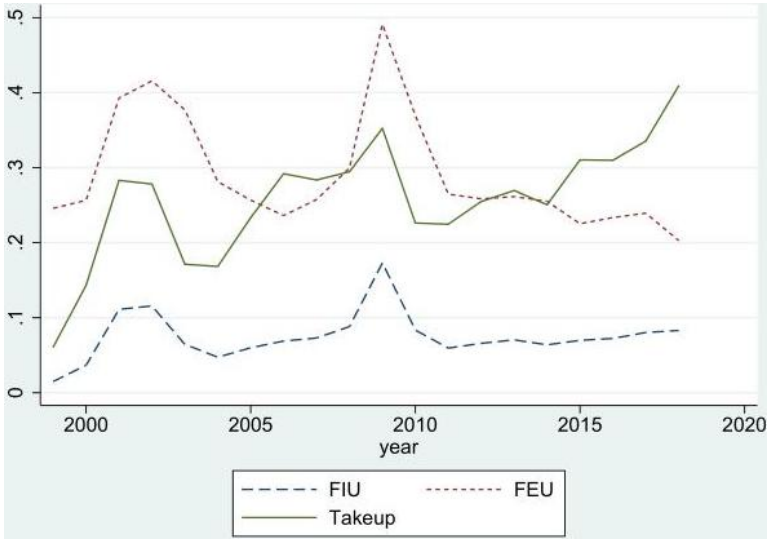


Figure 2: Estimated Probability of UI Receipt (Takeup), Fraction of Insured Unemployment (FIU), and Fraction of Eligible Unemployment (FEU) over Time

Note: Data are publicly available from the Bureau of Labor Insurance and Directorate-General of Budget, Accounting, and Statistics of Executive Yuan.

According to column 3 of Table 3, the estimate of the effect of extended benefits on the reservation wage for job losers is 0.045, indicating that the effect of extended benefits on the reservation wage for UI recipients is 0.173 (0.045/0.26). That is, a three-month increase in unemployment benefits is estimated to increase the reservation wage for UI-recipients by 17.3 percent. In the next section, I incorporate the estimated effect of extended benefits on the reservation wage for UI recipients with a simple search model to study the welfare implications of the estimate.

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than the estimates in the U.S. The estimates from Farber and Valletta (2015) and Rothstein (2011) using the Current Population Survey suggest that the probability of UI receipt among job losers was about 50 percent in 2009.

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6.2 Welfare Implications

Finally, I relate my reduced-form estimates to the analysis of Shimer and Werning (2007), who show that an increase in the generosity of UI is welfare-enhancing if increasing UI benefits increases after-tax reservation wages. They derive their formula from a model where the social planner chooses a unemployment benefit b to maximize the unemployed worker's welfare subject to a balanced budget constraint. They demonstrate that the optimal unemployment benefit b satisfies the following condition:

$$\epsilon_{R,b} \cdot \frac{R}{b} = ur \cdot \left(1 + \frac{\epsilon_{ur,b}}{1 - ur}\right), \tag{5}$$

where $\epsilon_{R,b}$ is the elasticity of the reservation wage to benefit level. $\epsilon_{ur,b}$ is the elasticity of the unemployment rate with respect to benefit level. On one hand, the larger the effect of unemployment benefits on the pre-tax reservation wages, the larger the marginal value of unemployment benefits. On the other hand, an increase in pre-tax reservation wages raises the unemployment rate, ur, and the elasticity of the unemployment rate with respect to UI benefits, $\epsilon_{ur,b}$, resulting in higher taxes that the worker must pay when she is employed.[15] Therefore, if the left-hand side of equation 5 is higher than the right-hand side, a marginal increase in benefit level will increase welfare. Since $\epsilon_{ur,b}$ is unknown, I use the approximation from Shimer and Werning (2007): $\epsilon_{ur,b} = (1 - ur) \cdot \epsilon_{D,b}$, where $\epsilon_{D,b}$ is the elasticity of unemployment duration to unemployment benefits. The above formula is robust to whether or not a worker is liquidity-constrained and a variety of extensions.[16]

Next, I incorporate the reduced-form estimates into the welfare formula to estimate the welfare effect of UI benefits. The effect of extended benefits

[15] Note that the above welfare formula is slightly different from that of Chetty and Finkelstein (2013). The elasticity of the unemployment rate to the unemployment benefits in Shimer and Werning (2007) holds the tax fixed, while that in Chetty and Finkelstein (2013). does not. I adopt Shimer and Werning (2007) definition in this paper.

[16] See Section VI of Shimer and Werning (2007) for additional discussion.

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on the reservation wage for UI recipients is estimated to be 0.173, implying an elasticity of the reservation wage to the potential duration of 0.346. According to Huang and Yang (2018), the average monthly benefit is 18,568 NTD for workers age 45 and 46 at job loss, so the ratio of the reservation wage to the level of UI benefits is $(30,450/18,568) = 1.640$. Taken together, this implies that $(dR/db) = 0.567$. That is, a one NTD increase in UI benefits is estimated to increase workers' reservation wages by 0.567 NTD. Regarding the welfare cost, the average unemployment rate between 2009 and 2018 for workers aged between 45 and 49 is 0.028. Based on Huang and Yang (2018), the elasticity of the insured duration of unemployment to the potential duration of benefits is about 0.8. Therefore, the right-hand side of equation 5 gives the value of 0.052.[17] This indicates that the welfare benefits of a marginal increase in UI benefits are substantially larger than its welfare cost. However, since the welfare analysis relies on estimates from other studies, it is more of an illustrative exercise rather than an analysis to reach a firm conclusion.

It is also important to note that the welfare analysis above relies on several assumptions, and the violations of these assumptions might alter the sign or magnitude of the welfare effect. First, the effects of an increase in the level of benefits are not necessarily the same as the effects of an increase in the potential duration of benefits. According to Krueger and Meyer (2002) and Schmieder and Wachter (2016), the elasticity of unemployment duration to potential duration is somewhat smaller than the elasticity of unemployment duration to benefit level. However, it is empirically unclear whether the effect of extended benefits on the reservation wage is comparable to the effect of an increase in benefit level. Second, if the benefit extension increases the reservation wage, this may discourage firms to create vacancy because they may find it harder to meet the wage demand of job applicants. This labor

[17] $ur \cdot (1 + (\epsilon_{ur,b}/1 - ur)) = ur \cdot (1 + ((1 - ur) \cdot \epsilon_{D,b}/1 - ur)) = 0.028 \cdot (1 + (1/0.972) \cdot 0.972 \cdot 0.8) = 0.052$.

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demand externality will further increase unemployment and generate an additional cost for increasing UI generosity. Further investigation is needed to fully understand how these factors change the welfare calculation.

7 Conclusion

This paper presents quasi-experimental evidence on the effect of extended benefits on the reservation wage, using the 2009 UI policy reform in Taiwan, which increased the potential benefit duration for workers at least 45. My estimates using a difference-in-differences design and data from the MUS suggest that a three-month increase in potential benefit duration increases the reservation wage for job losers by 4–5 percent. The response of reservation wages to the benefit extension is mainly driven by the effect on the reservation wage during extended benefit period — the effect of extended benefits on the reservation wage during the extended benefit period is more than two times larger than during the regular benefit period.

Compared with Le Barbanchon, Rathelot, and Roulet (2019) and Lichter (2016), who exploit natural experiments in France and Germany, respectively, the estimated effect of extended benefits on the reservation wage in this paper are larger. One possible explanation for this is that the potential duration in Taiwan is only about half of those in France and Germany, thereby increasing the likelihood of workers receiving extended benefits. The findings that the effect of extended benefits on the reservation wage is particularly large during the extended benefit period and for those who are most likely to exhaust their benefits provide empirical support for this explanation.

The main limitation of this paper is the dynamic selection issues related to the estimated reservation wage responses to the benefit extension over the unemployment spell. While my estimates suggest that the reservation wage response to extended benefits during the extended benefit period is significantly larger than during the regular benefit period, I cannot completely

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rule out the possibility that the results are partly driven by the differential changes in unobserved characteristics of the treatment and control groups. Future work that explores the dynamic selection process during the course of an unemployment spell and provides more evidence regarding the effect of extended benefits on the reservation wage over the unemployment spell would be valuable.

Based on the welfare formula of Shimer and Werning (2007) and the assumption that the effects of increasing potential duration are comparable to the effects of increasing benefit level, the estimated reservation wage response to unemployment benefits in Taiwan suggests a large potential welfare gain from increasing the level of unemployment benefits for middle-aged workers. Since the welfare effect of an increased potential duration is not necessarily the same as that of an increased benefit level, another potential goal of future research is to develop a formal welfare formula for the case of an increase in the potential duration of benefits using Shimer and Werning (2007) reservation wage approach. Combining the formula with the estimates in this paper would shed light on the welfare effects of extended unemployment benefits.

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Table A1: Descriptive Statistics: Workers Age 35–55

	Age 45–55			Age 35–44			
	1999–2008 (1)	2009–2018 (2)	Diff. (3)	1999–2008 (4)	2009–2018 (5)	Diff. (6)	Diff.-in-Diff. (7)
Reservation Wage (2016 NTD)	34,070	31,638	−2,431***	34,879	31,260	−3,619***	1,188
	(16,357)	(11,469)	(672)	(12,461)	(9,508)	(573)	(862)
Log Reservation Wage	10.38	10.31	−0.068***	10.42	10.31	−0.110***	0.042**
	(0.32)	(0.31)	(0.013)	(0.28)	(0.29)	(0.014)	(0.019)
Age (years)	49.17	49.56	0.386**	39.59	39.66	0.074	0.312
	(2.94)	(3.07)	(0.178)	(2.91)	(2.91)	(0.079)	(0.190)
Female	0.22	0.24	0.028**	0.24	0.30	0.065***	−0.037
	(0.41)	(0.43)	(0.013)	(0.43)	(0.46)	(0.019)	(0.023)
Married	0.75	0.64	−0.105***	0.61	0.44	−0.170***	0.066*
	(0.44)	(0.50)	(0.022)	(0.49)	(0.50)	(0.027)	(0.034)
College Degree	0.10	0.18	0.073***	0.14	0.29	0.150***	−0.077***
	(0.31)	(0.38)	(0.015)	(0.35)	(0.46)	(0.016)	(0.022)
Duration in Progress (months)	9.08	7.40	−1.68***	8.21	7.38	−0.828**	−0.850
	(8.11)	(6.81)	(0.39)	(7.72)	(6.94)	(0.342)	(0.506)
Live in Taipei	0.08	0.10	0.022	0.08	0.11	0.028**	−0.006
	(0.27)	(0.30)	(0.013)	(0.27)	(0.31)	(0.011)	(0.016)
Live in Kaohsiung	0.07	0.09	0.027	0.08	0.08	0.004	0.023
	(0.25)	(0.29)	(0.017)	(0.27)	(0.27)	(0.012)	(0.020)
Worked in Manufacturing Ind.	0.44	0.35	−0.095	0.41***	0.37	−0.044*	−0.052
	(0.50)	(0.47)	(0.023)	(0.49)	(0.48)	(0.022)	(0.031)

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Table A1: Descriptive Statistics: Workers Age 35–55 (Continued)

	Age 45–55			Age 35–44			
	1999–2008 (1)	2009–2018 (2)	Diff. (3)	1999–2008 (4)	2009–2018 (5)	Diff. (6)	Diff.-in-Diff. (7)
White Collar	0.19	0.25	0.069***	0.25	0.32	0.062***	0.008
	(0.39)	(0.47)	(0.021)	(0.44)	(0.46)	(0.016)	(0.025)
Previous Firm Size > 10	0.52	0.54	0.023	0.50	0.53	0.032	−0.010
	(0.50)	(0.50)	(0.020)	(0.50)	(0.50)	(0.020)	(0.028)
Predicted Log Reservation Wage	10.39	10.39	0.003	10.43	10.43	−0.002	0.005
	(0.14)	(0.15)	(0.005)	(0.14)	(0.14)	(0.006)	(0.008)
Number of observations	963	833	1,796	1,231	819	2,050	3,846

Note: This table displays the descriptive statistics for the main variables for job losers age 35-55 from the MUS between 1999 and 2018. I cut the sample by the age 45 cutoff and the year of the reform (2009). Columns 3 and 6 show the differences between average values in post-reform and pre-reform periods for workers age 45–55 and those age 35–44, respectively. Column 7 reports the difference between the values in columns 3 and 6. Job losers are workers who report their reason for unemployment is "Business shrunk or establishment closed". Job leavers are workers who report their reason for unemployment other than "Business shrunk or establishment closed" and "Seasonal or temporary work finished". Standard deviations are in parentheses of columns 1, 2 ,4, and 5. Standard errors are in parentheses of columns 3, 6 and 7. All standard errors are clustered at age levels. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

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Table A2: Examination of Dynamic Selection

	Age 45–60			Age 25–44			
	1999–2008 (1)	2009–2018 (2)	Diff. (3)	1999–2008 (4)	2009–2018 (5)	Diff. (6)	Diff.-in-Diff. (7)
Female	0.17	0.16	−0.010	0.20	0.38	0.183**	−0.193*
	(0.38)	(0.37)	(0.051)	(0.40)	(0.49)	(0.082)	(0.096)
Married	0.67	0.73	0.060***	0.38	0.32	−0.058	0.119
	(0.47)	(0.45)	(0.087)	(0.48)	(0.47)	(0.048)	(0.097)
College Degree	0.13	0.20	0.067***	0.19	0.30	0.104*	−0.037
	(0.43)	(0.40)	(0.074)	(0.40)	(0.46)	(0.052)	(0.089)
Duration in Progress (months)	9.13	8.87	−0.261*	9.14	9.09	−0.050	−0.211
	(0.88)	(0.91)	(0.131)	(0.90)	(0.89)	(0.097)	(0.161)
Live in Taipei	0.08	0.09	0.011	0.06	0.16	0.102**	−0.092
	(0.27)	(0.29)	(0.039)	(0.23)	(0.37)	(0.043)	(0.057)
Live in Kaohsiung	0.04	0.15	0.105*	0.07	0.07	0.006	0.100
	(0.20)	(0.36)	(0.057)	(0.23)	(0.26)	(0.028)	(0.063)
Worked in Manufacturing Ind.	0.49	0.40	−0.093	0.41	0.33	−0.076	−0.017
	(0.50)	(0.49)	(0.109)	(0.49)	(0.47)	(0.064)	(0.124)
White Collar	0.19	0.31	0.122	0.26	0.27	0.003***	0.119
	(0.39)	(0.47)	(0.090)	(0.44)	(0.44)	(0.076)	(0.117)
Previous Firm Size > 10	0.55	0.55	−0.001	0.51	0.56	0.050***	−0.051
	(0.50)	(0.50)	(0.070)	(0.50)	(0.50)	(0.070)	(0.098)

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Table A2: Examination of Dynamic Selection (Continued)

	Age 45–60			Age 25–44			
	1999–2008 (1)	2009–2018 (2)	Diff. (3)	1999–2008 (4)	2009–2018 (5)	Diff. (6)	Diff.-in-Diff. (7)
Predicted Log Reservation Wage	10.42	10.41	−0.012	10.38	10.35	−0.028*	0.016
	(0.18)	(0.27)	(0.053)	(0.17)	(0.18)	(0.014)	(0.054)
Number of observations	75	55	130	175	94	269	399

Note: This table investigate the dynamic selection issue by comparing the observed characteristics for job losers during the extended benefit period between 45–60 years of age (treatment group) and those between 25–44 years of age (control group). I cut the sample by the age 45 cutoff and the year of the reform (2009). Columns 3 and 6 show the differences between average values in post-reform and pre-reform period for the treatment and control group, respectively. Column 7 reports the difference between the values in columns 3 and 6. Job losers are workers who report their reason for unemployment is "Business shrunk or establishment closed". Job leavers are workers who report their reason for unemployment other than "Business shrunk or establishment closed" and "Seasonal or temporary work finished". Standard deviations are in parentheses of columns 1, 2 ,4, and 5. Standard errors are in parentheses of columns 3, 6 and 7. All standard errors are clustered at age levels. *** significant at the 1 percent level, ** significant at the 5 percent level, and * significant at the 10 percent level.

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延長失業給付對保留工資的影響: 來自台灣失業給付延長的證據

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本文利用台灣延長中高齡失業勞工失業給付的政策變革來估計延長失業給付對保留工資的影響。從2009年開始, 45歲以上失業勞工失業給付期限由6個月延長爲9個月, 而45歲以下仍維持6個月。利用人力運用調查資料和差異中的差異法, 我估計增加3個月的失業給付使得非自願失業勞工的保留工資上升約4–5%, 而傾向於用完6個月失業給付的勞工的保留工資上升尤其明顯。最後, 將估計結果結合搜尋模型, 本文估計增加失業給付的寬厚程度會提高中高齡勞工的福利水準。

關鍵詞: 延長失業給付, 保留工資, 差異中的差異
JEL 分類代號: J64, J65, J68

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