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Abstract

Many countries have a two-tiered unemployment compensation system which provides earnings-related unemployment insurance for a limited period of time and less generous unemployment assistance thereafter. This study evaluates the effects of a reform in Finland that increased the level of unemployment assistance by 22%. The reform led to a drop of 9% in the unemployment exit hazard, which can be attributed to fewer exits to both employment and inactivity. The implied elasticities suggest that a 10% increase in unemployment assistance reduces the unemployment exit hazard by 4% and the job finding hazard by 6%. These effects are relatively small compared to the existing evidence on the effects of unemployment insurance benefits.

Keywords: Unemployment assistance, labor market subsidy, hazard rate, unemployment duration.

JEL codes: J64, J68.

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1 Introduction

Many countries have adopted a two-tiered unemployment compensation system which provide earnings-related unemployment insurance (UI) for a limited period of time and less generous unemployment assistance thereafter. In the European Union, 19 member states have such a two-tier system in place (Esser et al. 2013). In Finland, job losers who meet certain eligibility conditions are entitled to UI benefits up to 300, 400 or 500 days, depending on their age and employment history. Unemployed individuals who exhaust their UI benefits and those without sufficient contribution history can claim labor market subsidy, the Finnish version of unemployment assistance. This benefit is much lower than the average UI benefit, but it is available for an indefinite time.[1] Labor market subsidy recipients are a large and growing group: their share of all unemployment benefit recipients grew from 45% in 2000 to 56% in 2018 (Statistical Yearbook on Unemployment Protection in Finland 2018).

While a large body of literature has studied the effects of unemployment insurance,[2] much less is known about the effects of unemployment assistance. Since unemployment assistance recipients are either long-term unemployed or labor market entrants, their employment prospects can be relatively weak compared to the UI recipients with stronger labor market attachment. They are also more likely to receive means-tested income transfers such as housing allowance or social assistance. For these reasons, this group may be less responsive to changes in their benefits than UI recipients.

This paper studies the effects of labor market subsidy on unemployment exits in Finland. On January 1, 2012, the labor market subsidy was raised by 22%, 121 Euros a month. The reform affected all unemployment benefit recipients, indirectly also the UI recipients. By implication, there is no unaffected group of the unemployed that could be used as a comparison group. However, since the reform also affected ongoing spells, I estimate the effect of the benefit level by comparing the unemployment exit hazards of labor market subsidy recipients before and after the day when the reform came into effect. To remove the seasonal component from the hazard rate, I use the change in the hazard rate at the beginning of 2011 when the benefit level did not change as a counterfactual. The analysis is based on comprehensive register data on the entire population of labor market subsidy recipients.

The results point to a decline of 9% in the unemployment exit hazard following the benefit hike. This decline is mainly driven by lower hazard rates to employment and inactivity. There is also a negative effect on the hazard rate to subsidized employment

[1] In 2018, the average UI benefit was 61.83 Euros a day, while the average labor market subsidy was 34.67 Euros a day, 44% less (Statistical Yearbook on Unemployment Protection in Finland 2018).

[2] For surveys, see Holmlund (1998), Tatsiramos and van Ours (2014) and Schmieder and von Wachter (2016). Kyyrä et al. (2017) summarize the existing evidence on the effects of UI in the context of the Finnish labor market.

but the effect is statistically significant only at the 10% risk level. The results imply that a 10% increase in labor market subsidy reduces the unemployment exit hazard by 4% and the job finding hazard by 6%. The analysis of different groups shows that the overall effect on unemployment exits increases with age but is roughly the same for women and men. For workers under the age of 35, the overall effect is driven by a reduction in the job finding hazard. In the older groups, the drop in the job finding hazard plays a smaller role, while reductions in the hazard rates to inactivity and subsidized employment account for a significant part of the overall decline in unemployment exits. While one-half of the overall effect for men is due to the reduction in the job finding hazard, women's job finding rate is hardly affected by higher labor market subsidy. For women, the overall decline in unemployment exits results from lower hazard rates to subsidized employment and inactivity.

This paper contributes to the literature on the effects of unemployment benefit generosity. Unlike most other studies, I analyze the effects of unemployment assistance for long-term unemployed and labor market entrants. The results show that also labor market subsidy recipients do respond to benefit changes but their response is somewhat weaker than the response of UI recipients to the similar benefit changes found in other studies (e.g. Carling et al. 2001, and Uusitalo and Verho 2010). Also, it is worth emphasizing that only one quarter of job seekers exiting from the labor market subsidy find a job in the open labor market, while one-half leave the labor force and one quarter end up in subsidized employment. The low share of exits to unsubsidized employment suggests that benefit generosity may not be the most severe barrier to employment for this group.

I am aware of only a few studies on the role of unemployment assistance, all of which examine the effects of benefit cuts. Doris et al. (2020) analyze the effect of a 51% cut in unemployment assistance for 18- and 19-year-olds in Ireland during the Great Recession. They find a substantial reduction in the average duration of unemployment assistance receipt, which corresponds to an elasticity of around 1. Exits to both training and work turned out to be important components of the overall reduction in the benefit duration. Given the age groups and exceptional time period analyzed, the general applicability of these results is limited. Wolfgang and Weber (2016) and Prize (2019) study the Hartz IV reform in Germany that cut unemployment assistance for long-term unemployed and therefore increased the drop in benefits at the time when UI benefits expire. Both studies find that the reform increased the job finding hazard of UI recipients as the time of the benefit exhaustion approaches. Prize (2019) also finds an increase in the job finding rate after the UI benefit exhaustion, but Wolfgang and Weber (2016) find such an effect only for women and for those under the age of 30.[3]

[3]The reform took place in 2005. Wolfgang and Weber (2016) compare unemployment spells that started in 2000–2003 (pre-reform period) to those that started in 2007–2010 (post-reform period). Although they control for macroeconomic indicators and balance the pre- and post-reform samples on observed individual characteristics, the results may be biased due to changes in economic conditions and

A large number of studies have examined the effects of UI benefits on unemployment duration. Using data for Finland, Uusitalo and Verho (2010) estimate that a 10% increase in UI benefit reduces the unemployment exit hazard by 6% and the job finding hazard by 11%.[4] These effects are 50% and 83% larger than the corresponding point estimates of labor market subsidy effects found in this study. The estimated effects of UI benefits for Finland in the same ballpark as the UI estimates for other countries. Carling et al. (2001) and Rebollo-Sanz and Rodríguez-Planas (2020) report larger effects on the job finding rate for Sweden and Spain, respectively, while Røed and Zhang (2005) find a smaller effect for Norway.

The differences in the effects of UI benefits and labor market subsidy are related to the discussion on the optimal time profile of the benefits. The two-tier benefit system produces a declining profile as the benefit level drops at the time when UI benefits expire. Most of the existing literature on the optimal time profile has been theoretical, and the results from these studies have been mixed (see e.g. Shavell and Weiss 1979, Cahuc and Lehmann 2000, and Shimer and Werning 2008). The empirical evidence is scarce and also mixed. Using data for Sweden, Kolsrud et al. (2018) find that unemployed individuals respond more strongly to changes in their UI benefits paid earlier in the unemployment spell than to the benefits paid later in the spell. They also provide evidence that the consumption-smoothing value of UI benefits is higher for the long-term unemployed. These results suggest that the declining benefit profile is not optimal but a flat or even increasing benefit profile might be more desirable. Linder and Balázs (2020) came to the opposite conclusion by studying a UI reform in Hungary. Their findings give support for the declining benefit profile. The results of this study are perhaps more in line with Kolsrud et al. (2018). Since the consumption-smoothing value of the benefits is likely to be higher for labor subsidy recipients than for UI recipients, the relatively weak effect of labor market subsidy on the job finding hazard does not give support at least for a deeply declining benefit profile. However, this interpretation is only suggestive and should be treated with caution, as we do not know how the UI recipients respond to changes in labor market subsidy in the Finnish labor market.

The rest of the paper proceeds as follows. The next section describes relevant institutions and the 2012 reform. Section 3 discusses the research design. Section 4 describes the data and sample selection. Section 5 discusses macroeconomic conditions around the time of the reform. Section 6 reports empirical hazard rates. Section 7 presents the

other labor market reforms that were implemented during the 10-year period. Prize (2019) applies a more elaborate approach and exploits individual-level variation in the potential duration of UI benefits for identification, which is likely to produce more reliable results. Both studies rely on the same register data which had a flaw that information on unemployment assistance payments was mainly missing for the first two post-reform years due to the administrative transition process related to the benefit reform.

[4] Uusitalo and Verho (2010) study a reform in 2003 that abolished the severance pay and increased UI benefits for unemployed workers with at least 20 years of employment history. The average increase in the UI benefits was 15% and the increased benefit was paid over the first 150 days of unemployment.

results of the difference-in-differences analysis. The final section concludes.

2 Institutional setting

The empirical analysis covers the years 2009–2014. At that period the entitlement period of earnings-related UI benefits (*ansiopäiväraha*) was 500 payment days, about two calendar year as the benefits are paid for five days a week.[5] The UI benefit can be paid to the unemployed individual who belongs to an unemployment fund and who satisfies certain eligibility conditions regarding employment and contribution history.[6] Unemployed who are not entitled to UI benefits may qualify for labor market subsidy (*työmarkkinatuki*), which is paid by the Social Insurance Institution. This benefit is means-tested against certain other benefits and spouse's income (up until 2012) but, as long as the eligibility conditions are met, it can be received for an indefinite period. At the beginning of each year, the level of the labor market subsidy is adjusted according to an index measuring the inflation.

In 2011, the labor market subsidy was 25.74 Euros a day, without child supplements. On January 1, 2012, it was raised by 21.8% to 31.36 Euros a day. The increase in the monthly benefit was 121 Euros, of which 100 Euros was due to the reform and 21 Euros due to the annual index adjustment of 3.8%. At the same time also the income limits for housing allowance were adjusted in such a way that the benefit hike did not affected the amount of the housing allowance the labor market subsidy recipient was possibly entitled to. In most cases, the 2012 reform therefore raised the income of labor market subsidy recipients by the full amount of the benefit increase. According to the government's law proposal to the parliament, the aim of the reform was to reduce poverty and marginalization among labor market subsidy recipients, and to reduce income inequality.

In the econometric analysis, I model unemployment exits of labor market subsidy recipients who had no children, whose benefits were not reduced due to means-testing or other benefits and who were unemployed close to the turn of the years 2011 ("comparison period") and 2012 ("reform period"). For the labor market subsidy recipients during the considered periods, there were no other notable changes than the 21.8% increase in the benefit level at the beginning of 2012. As a robustness check, I also provide evidence on changes in unemployment exits at the turn of other years, although these results should be interpreted with caution as they may be confounded to some extent by other changes in the benefits, especially at the beginning of 2013.

In 2010, 2011 and 2014, the labor market subsidy was raised by less than 1% due to

[5]In 2014, the maximum duration of UI benefits was reduced to 400 days for new UI recipients with less than three years of employment history .

[6]Membership in the unemployment funds is voluntary but as many as 90% of employed workers were members in 2015.

the low inflation (see table 1). In 2013, the index adjustment was somewhat higher, 3.5%. On December 15, 2010, means-testing was slightly relaxed by increasing the threshold above which the spouse's income reduces the labor market subsidy. As of January 1, 2013, the spouse' income has not affected the amount of the labor market subsidy. Also other changes took place in January 2012: new labor market programs were introduced and an employment bonus experiment begun in 60 municipalities. In the experiment municipalities, the labor market subsidy recipients were allowed to keep their subsidy for one extra month if they took up a new job with the duration of at least three months.

3 Research design

Since the UI benefit is defined as the sum of a basic component equal to labor market subsidy plus an earnings-related supplement, the 2012 reform did not only increase the subsidy but also the UI benefits and, therefore, all unemployment benefit recipients were affected by the reform. Due to the lack of the comparison group not affected by the 2012 reform, I have to rely on time series variation in the unemployment exit hazard around the day of the benefit hike. This approach is motivated by the fact that the reform affected all ongoing benefit spells from January 1, 2012 onward.

One could compare the hazard rate in the last week of 2011 to the hazard rate in the first week of 2012 but there are confounding factors that would bias such comparison. First, the flows into and out of labor market subsidy are subject to seasonal variation. New jobs (and training and education programs) often start at the beginning of the month, whereas old jobs typically terminate at the end of the month. Therefore, the unemployment exit hazard is elevated in the end of each month while the composition of the unemployed may change at the beginning of the month due to the discrete change in the inflow. Moreover, at the turn of the year, this kind of seasonal variation can be pronounced due to the Christmas and New Year. Second, in anticipation of the forthcoming benefit increase, some unemployed may have reduced their search effort and/or increased reservation wages at the end of 2011, in which case the unemployment exit hazard may have been abnormally low just prior to the reform. Third, since the unemployment benefits are paid in four-week or one-month periods, and the benefits are claimed afterwards, the higher benefits did not show up on individuals' bank account on January 1, 2012, but on some later day in January or at the beginning of February, depending on an individual-specific payment schedule. Thus, unemployed workers who were not aware of the reform may have noticed the benefit increase with some delay, in which case the unemployment exit hazard may have been "too high" just after the reform.

Because of these reasons, the effect of the reform cannot be identified from a discrete change in the hazard rate on January 1, 2012. Instead, I will compare the average hazard rates before and after the benefit increase within the 13-week window around the turn of 2012. Due to the possible anticipation effects and delayed responses, it might be advisable to ignore the observations in near vicinity of the reform day, although it is not obvious which time interval should be excluded from the analysis. On the one hand, a longer time interval for excluded observations around the reform day eliminates more surely the confounding effects due to anticipatory behavior and delays in the benefit payments. On the other hand, when the exit hazards far away from each other are compared, the hazard rates are more likely to differ due to the seasonal variation and business cycle conditions. I can estimate the seasonal component of the hazard rates by comparing the hazard rates around the first week of 2011 when the benefit level remained nearly constant. While the resulting difference-in-differences (DID) approach eliminates the common seasonal component, it does not eliminate possible asymmetric changes in the business cycle conditions within the time periods used in the analysis. To mitigate this concern I mainly use the relatively short window of 13 weeks in the analysis. As robustness checks, I also use a shorter window of nine weeks as well as a "donut-hole" sample that excludes observations within four weeks of the reform day.

4 Data

The registers of the Social Insurance Institution provide detailed information on labor market subsidies and some background characteristics for benefit recipients. These records are complemented by merging information on active labor market programs and employment spells from the registers of Ministry of Economic Affairs and Employment and of the Finnish Centre for Pensions. This supplementary information is used to detect exits to job placement programs (i.e. subsidized employment) and employment in the open labor market. The benefit spell is defined as the time the worker receives labor market subsidy, allowing for short breaks of four or fewer weeks in benefit receipt within the spell. Individuals who take up a new job that starts within four weeks from the end of the benefit period and lasts for at least four weeks are classified to be exited to employment.

I restrict the analysis to labor market subsidy recipients aged 25 to 60 with no children, and consider their benefit spells that were ongoing in the period 2009–2014. I drop spells during which the claimant's benefit were reduced due to means-testing or receipt of other benefits. The resulting sample contains 194,651 individuals, of whom 53,635 were unemployed at the time of the 2012 reform. Given the sample restrictions, the benefit level of these individuals increased 21.8% on January 1, 2012.

I model the behavior of benefit recipients by means of weekly exit hazards. The number of unemployed during a calendar week varies between 45,805 and 77,113, with an average of 57,032 over the period 2009–2014. Sample members are on average 41 years old, and 44% of them are female. The average duration of labor market subsidy spells is over one year.

5 Economic environment

Unfortunately, the observation period was rather turbulent. In 2009, the gross domestic product plunged 8.1% due to the global financial crisis. In 2010 and 2011, the economy grew 3.2% and 2.5%, respectively, but then turned on a declining path again and in 2012 the economy shrunk 1.4%. As seen in figure 1, the volume of production remained rather stable in 2011 but started to shrink at the beginning of 2012, that is, at the time when the reform came into effect. This poses a potential problem for the evaluation task due to the lack of the comparison group not affected by the reform.

From figure 2 we see that the labor force share of UI recipients follows the aggregate growth rates: it declined in 2010 and 2011 but turned on an increasing path at the beginning of 2012. For the labor market subsidy recipients, the pattern is quite different. Except for a short decline in the first half of 2010, their labor force share increases smoothly over the observation period, and the turn in the business cycle conditions in 2012 has no visible impact on the trend.

Figure 3 shows the flow into labor market subsidies, i.e. the number of new benefit spells in our sample. Apart from seasonal variation, the inflow increased smoothly up to the end of 2013, after which it stabilized. There are no large seasonal variation in the inflow. Importantly, we do not see evidence of excess inflow at the end of 2012, which would have raised concerns about anticipatory behavior. This is not very surprising given that many new labor market subsidy recipients are those whose UI benefits expired and, thus, they have been unemployed for the two preceding years. Overall, it appears that the stock of labor market subsidy recipients is less responsive to changes in the business cycle conditions than that of UI recipients.

6 Hazard rates

I consider weekly hazard rates for unemployment exits, as well as for exits to employment, subsidized employment and inactivity. The sum of the last three hazards equals the overall exit hazard, so that they provide a useful decomposition for the overall exit hazard. The weekly hazards exhibit a large degree of variation across months within years as well as across weeks within months. As a result, long series of weekly hazards are very noisy and therefore difficult to interpret. To ease the detection of breaks in the trends of the hazard rates, figure 4 shows seasonally-adjusted average weekly hazards.[7]

[7]The seasonal variation was removed as follows: first, weekly exits were regressed against calendar month dummies, using individual-level data and linear probability model; second, the mean of the residuals from this regression for each month was computed; and, finally the average exit rate over the

In the years 2009–2011, the seasonally-adjusted unemployment exit hazard varies around 0.015, suggesting that on average 1.5% of individuals receiving labor market subsidy on Monday exited from unemployment by the end of the week. From the cause-specific hazard rates we see that about one-half of them left the labor force, roughly one quarter found a new job in the open labor market, and one quarter took up a subsidized job. That is, unlike for the UI recipients, entering employment is not the main reason of unemployment exit. While the job finding hazard is increasing and other hazards are relative stable until the end of 2011, all the hazards decline smoothly over the later part of the observation period, reflecting the change in the business cycle conditions. However, we are interested in changes at the time of the reform. The unemployment exit hazard evolves smoothly at the beginning of 2010, 2011, 2013 and 2014, but shows a clear, discrete drop in January 2012, which is consistent with the hypothesis that the 2012 reform reduced exits from unemployment. A similar yet less clear drop can also be seen in the hazards to inactivity and subsidized employment, while the picture for the job finding hazard is more blurry.

In the subsequent analysis, I focus on changes in the weekly hazard rates around the turn of the years. Figures 5–7 show the unadjusted weekly hazards within the 13-week window from the first week of a given year (= the first complete calendar week in January), which corresponds to the value of 0 on the horizontal axis.[8] Note that all the hazard rates are multiplied by 100. Figure 5 compares changes in the hazard rates around the turn of 2010 and 2011, that is, over the two periods when the benefit level remained roughly constant. The hazard rates are almost overlapping and the marginally higher average levels of unemployment exit and job finding hazards in the time interval around the turn of 2011 are due to better business cycle conditions. More importantly, changes in the hazard rates from the last 13 weeks of the year to the first 13 weeks of the next year are similar in the two periods: the increase in the average unemployment exit hazard is 0.092 (7.5%) and 0.125 (9.6%) at the turn of 2010 and 2011, respectively. This roughly 0.10 increase in the hazard rate at the beginning of the year describes the seasonal variation in unemployment exits, which we need to take into account when estimating the effect of the benefit hike at the beginning of 2012.

In figure 6, the unemployment exit hazard remains stable from the last 13 weeks of 2011 to the first 13 weeks of 2012 (a decrease of 0.003 or 0.2%), whereas the unemployment exit hazard increases by 0.125 (9.6%) from the last 13 weeks of 2010 to the first 13 weeks of 2011. Thus, unlike at the beginning of the two previous years, the unemployment exit hazard did not increase at the beginning of 2012. This finding is consistent with the hypothesis that the benefit hike in January 2012 reduced unemployment exits.

years 2009–2014 was added to the mean residuals. Thus, the curves show the average weekly hazard rates during a given month, net of systematic seasonal variation across calendar months.

[8]The first week begins on the 4th, 3rd, 2nd, 7th and 6th day in January 2010, 2011, 2012, 2013 and 2014 respectively.

Figure 7 compares the hazards in the two periods after the 2012 reform. The hazard rates in these periods are highly similar. The increase in the unemployment exit hazard from the last 13 weeks of 2013 to the 13 first weeks of 2014 is 0.068 (7.0%), which is very close to the changes at the beginning of 2010 and 2011. However, at the beginning of 2013 the unemployment exit hazard increases somewhat less, by 0.038 (3.6%). The smaller increase over the first weeks of 2013 may not come as a surprise because the benefit level was raised by 3.5% in January 2013 (see table 1). Also the employment bonus experiment and the introduction of new labor market programs at the beginning of 2013 may have affected the difference in the hazard rates between the last weeks of 2012 and the first weeks of 2013. Because of these changes, the time interval around the firsts week of 2013 is a less valid comparison period than other considered periods.

7 Statistical analysis

7.1 Model specification and identification

Next, I test statistical significance of the hazard changes between the year pairs in figures 5–7, using a simple difference-in-differences model. To assess whether the benefit hike in 2012 affected the hazard rates I compare changes in the hazards at the beginning of 2012 to changes at the beginning of 2011 within the 13-week windows. The difference between these changes eliminates the effect of the seasonal variation around the turn of the years (the systematic difference between the last 13 weeks of the year and the first 13 weeks of the next year) but not the possible effect of differently changing economic conditions *within* the 26-week time intervals between the two periods (+/–13 weeks around the turn of 2011 and 2012). To mitigate the possible effect of the business cycle conditions I only use observations that are at most 13 weeks away from the turn of the year.

I estimate the following equation:

$$Y_{it} = \alpha + \beta Post_{it} + \gamma Reform_i + \delta_{DID} \left(Reform_i \times Post_{it} \right) + \varepsilon_{it},$$

where i indexes individual and t, $-13 \leq t \leq 12$, denotes the calendar week since the first week of the relevant year ($t = 0$). $Y_{it} = 1$ if individual i who was unemployed at the beginning of week t left unemployment by the end of the week, and $Y_{it} = 0$ otherwise. $Post_{it} = 1\{t \geq 0\}$ is a dummy variable for the first weeks of the year. $Reform_i$ is a dummy for the period during which the reform took place (i.e. the time interval around the turn of 2012).

It follows that α is the average hazard rate at the end of 2010, and β captures its change at the beginning of 2011, i.e. the difference in the average hazards between the last 13 weeks of 2010 and the first 13 weeks of 2011. That is, β measures the seasonal component of the hazard rate, which is assumed to be the same in the comparison and reform periods. γ is the difference in the average hazard rates between the comparison and reform periods, for example, due to different business cycle conditions. The parameter of interest is δ_{DID}, which is equal to the change in the hazard rate at the beginning of 2012, net of the seasonal component. Under the assumptions that the seasonal variation is constant across the years and that the economic environment evolved similarly *within* the 13-week windows around the turn of 2011 and 2012, this parameter identifies the causal effect of the benefit increase. It is worth emphasizing that the economic environment, and thus the level of the hazard rates, can be different between the two periods, as captured by γ, but divergent changes in the business cycle conditions *within* the considered time intervals between the comparison and reform periods are ruled out by assumption. This is a rather strong assumption in the light of the evidence presented in section 5. However, the possible bias due to divergent pattern of the business cycle conditions is likely to be small when the short windows are used.

As a robustness check, I also estimate the model using data around the turns of 2010 and 2011 as well as data around the turns of 2013 and 2014, treating the latter period as a placebo reform period. In these cases, the estimate of γ_{DID} provides a simple test for the underlying assumptions of the DID setting; it should be close to zero if the seasonal component is roughly constant across the years and the business cycle conditions do not change differently within the 13-week windows around the turn of different years.

7.2 Results

The point estimates of δ_{DID} along with their 95% confidence intervals are shown in figure 8. The effect of the benefit increase in 2012 on the unemployment exit rate is -0.1275, which corresponds to a decline of 9% from the counterfactual hazard of 1.4217 over the first 13 weeks of 2012. Since the benefit level was raised by 22%, the implied elasticity of the unemployment exit hazard with respect to labor market subsidy is –0.4.

The decline in unemployment exits after the benefit increase in 2012 is mainly due to lower hazard rates to employment and inactivity. There is also a negative effect on the hazard rate to subsidized employment, but the effect is statistically significant only at the 10% level (p-value 0.0765). The estimated effects are generally rather imprecise, even though the number of worker-week observations is as large as 2.7 million.Due to the wide and overlapping confidence intervals, we cannot really say whether the effects of the benefit hike on the hazard rates to employment, subsidized employment and inactivity differ from each others.

It should be stressed the estimated effects measure *absolute* changes in the underlying hazard rates. This is convenient as the absolute changes in the cause-specific hazards add up to the change in the unemployment exit hazard, providing a useful decomposition for the overall effect. While the estimated absolute changes in the cause-specific hazards are on the same ballpark, they do imply different *relative* changes due to differences in the levels of the underlying counterfactual hazards (see table 2). Namely, the estimates correspond to a decline of 14% in the job finding hazard and declines of 7% in the hazards to subsidized employment and inactivity. It follows that the elasticity of the job finding rate with respect to labor market subsidy is –0.6, while the elasticity of other two cause-specific hazards rates is only –0.3.

In figure 8, the placebo effects on the unemployment exit hazard obtained from the periods 2010–2011 and 2013–2014 do not differ from zero at the 5% risk level. Except for the effect on the inactivity hazard from the year pair of 2013 and 2014, also the placebo effects on the cause-specific hazards are close to zero and statistically insignificant. Recall that the time period around the turn of 2013 is a problematic period due to the benefit increase of 3.5% and certain other reforms that became into effect in January 2013. That said, the results of the placebo reforms are reassuring as they give support for the validity of the research setting.

Table 2 reports the results of additional robustness checks. Our baseline DID estimates (the red points in figure 8) are shown on the top of the table. These are followed by DID estimates obtained from two different model specification and from two different estimation samples. First, I add a set of control variables to the analysis. These include the elapsed duration of the ongoing unemployment spell (9 categories), gender, age (7 categories), region (19 provinces), the log of the municipal unemployment rate, and a dummy for the last week of the month. Second, I allow distinct linear trends in the hazard rates over the comparison and reform periods. This specification relaxes the assumption that the business cycle conditions did not change differently within the 13-week windows around the turn of 2011 and 2012. Third, I exclude observations that are closer than four weeks to the first week of the year. Estimates from this "donout-hole" sample should be less affected by possible anticipatory behavior or delayed responses. Finally, I exclude the furthermost observations from the analysis using a shorter window of nine weeks around the turn of the years. These estimates should be less sensitive with respect to possible asymmetric changes in the business cycle conditions between the reform and comparison periods (yet more sensitive with respect to possible anticipatory behavior and delayed responses). As seen in table 2, all of these estimates are rather close to the corresponding baseline estimates reported on the top of the table. This observation together with the small and mainly insignificant placebo effects in figure 8 show that the results are reasonably robust. In particular, since controlling for the local unemployment rate, allowing for period-specific linear trends and using the shorter time window do not change the results, it seems that the estimates are not significantly biased due to differently changing business cycle conditions within the two time intervals.

The DID estimates by gender, age and the elapsed duration of labor market subsidy receipt are reported in table 3. The overall effect of the benefit hike on unemployment exits is roughly the same for women and men. However, whereas one-half of the overall effect for men stems from the reduction in the job finding hazard (with an elasticity of –0.9), women's job finding rate is hardly affected as the decline in unemployment exits comes from lower hazards to subsidized employment and inactivity (with the elasticities of –0.5 and –0.3).

For workers between the ages of 20 and 34, the only effect is a reduction of the job finding rate. In the older groups, the overall effect on unemployment exits is larger, while the drop of the job finding rate plays smaller role. Among the oldest workers, the change in the job finding rate explains less than one-fifth of the overall effect on the unemployment exit hazard. The smaller role of the changes in the job finding hazard for the oldest group is due to a lower counterfactual level of the job finding hazard (0.21 compared to 0.34 and 0.63 for the age group 35–49 and 25–34, respectively). In relative terms, the employment responses do not vary by age: the elasticity of the job finding hazard with respect to labor market subsidy is between –0.6 and –0.7 for all age groups. The counterfactual level of the unemployment exit hazard declines with age. Taken together with the larger absolute overall effect on unemployment exits for the older group this suggest that older workers respond more strongly to the level of labor market subsidy. The elasticities of the unemployment exit hazard are –0.2, –0.5 and –0.6 for the age groups 25–34, 35–49 and 50–60, respectively.

The results by the elapsed duration of benefit receipt are surprising. For individuals who have been on labor market subsidy for 6–24 months, the overall effect on unemployment exits is very small and does not differ from zero, while much stronger effects are found for those who have received the benefits for shorter or longer time (the elasticities of –0.5 and –0.7). It is also surprising that one-half of the overall effect for the long-term unemployed with the elapsed benefit duration of over 24 months can be attributed to the decline in the job finding rate. For this group the counterfactual job finding hazard is very low (0.10 compared to 0.36 for the entire population) and exits to employment account only for 15% of all exits without the reform. The benefit hike due to the reform decreases the job finding finding rate by 50%, dropping the share of exits to employment to 8%. The implied elasticity of the job finding rate with respect to the benefit level is very large in absolute terms, –2.3. Thus, although finding a job in the open labor market is a very unlikely outcome for those individuals who have already been on labor market subsidy for over two years, their low employment probability appears to be rather sensitive to the benefit level.

8 Concluding remarks

This paper studies the effects of labor market subsidy, the Finnish version of unemployment assistance. According to the results, the unemployment exit hazard declined by 9% at the beginning of 2012 due to the benefit increase of 22%. The implied elasticity suggests that a 10% increase in the labor market subsidy reduces the unemployment exit hazard by 4%. The overall effect on unemployment exits can be decomposed into the effects on the job finding hazard, subsidized employment hazard, and inactivity hazard. In absolute terms, the effects of labor market subsidy on the job finding and inactivity hazards are roughly of the same magnitude, while the effect on the subsidized employment hazard is smaller and statistically significant only at the 10% level. However, since exits to employment are relatively rare among labor market subsidy recipients, the results imply that a 10% increase in the labor market subsidy reduces the job finding hazard by 6%, twice as much as the subsidized employment and inactivity hazards. All in all, the labor market subsidy recipients respond to the benefit increase by exiting from unemployment to all considered destinations at lower rates.

These results are based on a simple difference-in-differences setting where the changes in the hazard rates at the beginning of 2012 are compared to the changes at the beginning of 2011. The approach eliminates the seasonal variation in the hazard rates and the average effect of different business cycle conditions between the two time intervals but not possible asymmetric changes in the business cycle conditions within the time intervals. This is a matter of concern as aggregate unemployment turned on a growing path at the time of the reform, which may bias the behavioral effects upward (i.e. implying too strong behavioral responses). Although the results proved to be robust with respect to various deviations from the basic setup, the conservative interpretation is that the estimates provide upper bounds for the true behavioral responses. Compared to the existing evidence on the effects of UI benefits on unemployment exit and re-employment, the estimated behavioral responses to the level of labor market subsidy, especially if they are slightly biased upward, are relatively small. This implies that benefit increases lead to relatively small undesired behavioral changes among labor market subsidy recipients. Analogously, benefit cuts are unlikely to lead to large employment effects for this group.

This study only provides evidence on the changes in behavior of labor market subsidy recipients right following the benefit increase. An important question for future work is how the UI recipients possibly respond to changes in the labor market subsidy.

References

Cahuc, P. and Lehmann, E. (2000). Should Unemployment Benefits Decrease with the Unemployment Spell? *Journal of Public Economics* 77:135–153.

Carling, K., Holmlund, B. and Vejsiu, A. (2001). Do Benefit Cuts Boost Job Finding? Swedish Evidence from the 1990s. *The Economic Journal*, 111: 766–790.

Doris, A., O´Neill, D. and Sweetman, O. (2020). Does Reducing Unemployment Benefits during a Recession Reduce Youth Unemployment? Evidence from a 50 Percent Cut in Unemployment Assistance. *Journal of Human Resources*, 55: 902–925.

Esser, I., Ferrani, T., Nelson, K., Palme, J. and Sjöberg, O. (2013). Unemployment Benefits in EU Member States. The European Commission, Directorate-General for Employment, Social Affairs and Inclusion.

Holmlund, B. (1998). Unemployment Insurance in Theory and Practice. *Scandinavian Journal of Economics*, 100:113–141.

Kolsrud, J., Landais, C., Nilsson, P. and Spinnewijn, J.(2018). The Optimal Timing of Unemployment Benefits: Theory and Evidence from Sweden. *American Economic Review*, 108: 985–1033.

Kyyrä, T., Pesola, H. and Rissanen, A. (2017). Unemployment Insurance in Finland: A Review of Recent Changes and Empirical Evidence on Behavioral Responses. VATT Research Reports 184.

Lindner, A. and Balázs, R. (2020). Frontloading the Unemployment Benefit: An Empirical Assessment. *American Economic Journal: Applied Economics*, 12:140–74.

Price, B. M. (2019). The Duration and Wage Effects of Long-Term Unemployment Benefits: Evidence from Germany's Hartz IV Reform, unpublished manuscript.

Duration of Unemployment, and Job-Match Quality. *The Journal of Human Resources*, 55:199–163.

Røed, K. and Zhang, T. (2005). Unemployment Duration and Economic Incentives — A Quasi Random-Assignment Approach. *European Economic Review*, 49:1799–1825.

Schmieder, J. F. and von Wachter, T. (2016). The Effects of Unemployment Insurance Benefits: New Evidence and Interpretation. *Annual Review of Economics*, 8:547–581.

Shavell, S. and Weiss, L. (1979). The Optimal Payment of Unemployment Insurance Benefits over Time. *Journal of Political Economy,* 87:1347–1362.

Shimer, R. and Werning, I. (2008). Liquidity and Insurance for the Unemployed. *American Economic Review*, 98: 1922–1942.

Statistical Yearbook on Unemployment Protection in Finland 2018. www.kela.fi/statistics.

Tatsiramos, K. and van Ours, J. C. (2014). Labor Market Effects of Unemployment Insurance Design. J*ournal of Economic Surveys*, 28:284–311.

Uusitalo, R. and Verho, J. (2010). The Effect of Unemployment Benefits on Re-Employment Rates: Evidence from the Finnish Unemployment Insurance Reform. *Labour Economics*, 17, 643-654.

Wolfgang, N. and Weber, M. (2016). Stuck in a Trap? Long-Term Unemployment under Two-Tier Unemployment Compensation Schemes. Ifo Working Paper No. 231.

Table 1: Benefits in 2009–2014

			Child supplements		
Year	Daily benefit	YoY change	1 child	2 children	3+ children
2009	25.63		4.86	7.13	9.19
2010	25.63	0%	4.86	7.13	9.19
2011	25.74	0.43%	4.88	7.16	9.23
2012	31.36	21.83%	5.06	7.43	9.58
2013	32.46	3.51%	5.24	7.69	9.92
2014	32.66	0.62%	5.27	7.74	9.98

Table 2: Robustness of estimated treatment effect ($\times 100$)

			Hazard rates to		
	Number of observations	Unemployment exit hazard	Employment	Subsidized employment	Inactivity
	(1)	(2)	(3)	(4)	(5)
Counterfactual hazard ($\times 100$)		1.4217	0.3619	0.3175	0.7423
Baseline estimate	2,704,254	-0.1275***	-0.0496***	-0.0235*	-0.0545***
		(0.0272)	(0.0127)	(0.0133)	(0.0205)
Elasticity		-0.4115	-0.6284	-0.3392	-0.3367
Robustness checks					
Control variables	2,704,254	-0.1095***	-0.0444***	-0.0212	-0.0439**
		(0.0276)	(0.0127)	(0.0133)	(0.0207)
Linear trend	2,704,254	-0.1414**	-0.0304	-0.0178	-0.0932**
		(0.0556)	(0.0250)	(0.0270)	(0.0423)
Donut-hole	1,765,964	-0.1401***	-0.0630***	-0.0326*	-0.0446*
		(0.0339)	(0.0165)	(0.0167)	(0.0251)
Shorter window	1,873,712	-0.1274***	-0.0409***	-0.0248	-0.0618**
		(0.0323)	(0.0147)	(0.0155)	(0.0247)

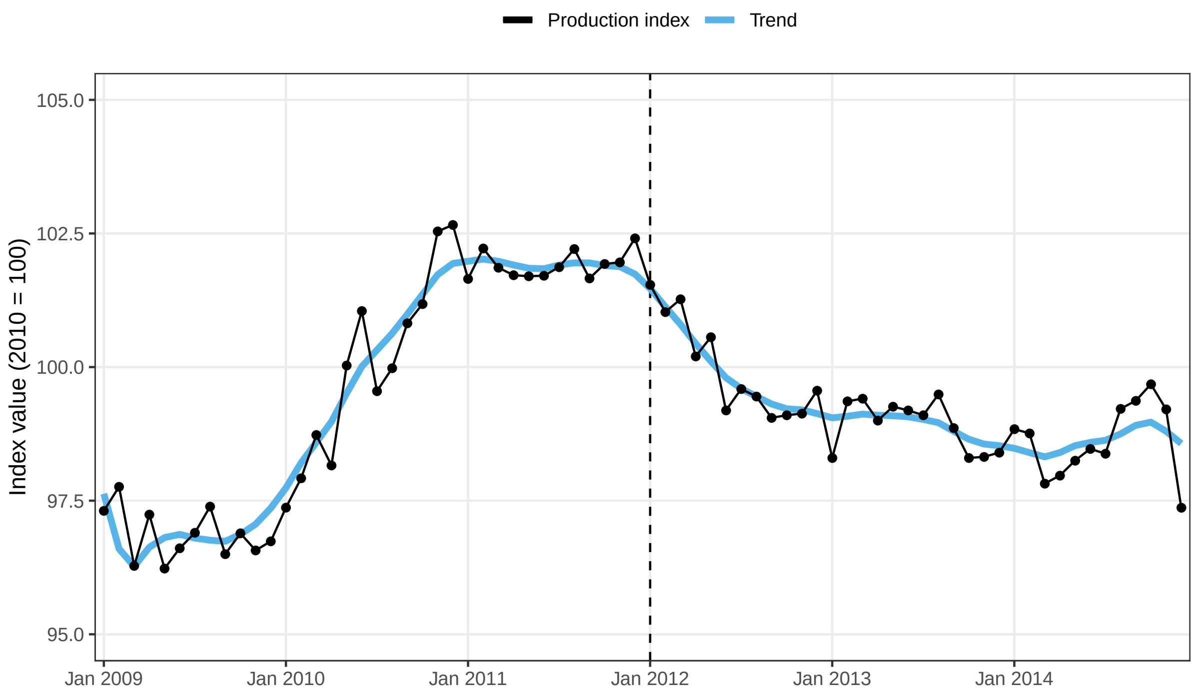
Notes: The estimating sample contains observations within the 13-week window from the turn of 2011 and 2012 (the last specification uses a narrower window). The number of observations refers to the number of worker-week observations, i.e. the total number of workers at risk of exiting from unemployment over the period used in the estimations. The counterfactual hazard is the predicted average hazard over the first 13 weeks of 2012 in the absence of the reform ($= \alpha + \beta + \gamma$) based on our baseline specification. The first set of the estimates are from our baseline model specified in equation (1). The same estimates are also illustrated in figure 8. The second set of the estimates are from an otherwise similar model but with controls for the elapsed duration of unemployment, gender, age, region, local unemployment rate and a dummy for the last week of the month. The subsequent estimates are from a model (without controls) that allows for distinct linear trends over the reform and comparison periods. These are followed by the estimates from a donut-hole sample excluding the closest 4 weeks of the data from the turn of the year. The last set of the estimates were obtained using a shorter window of 9 weeks around the turn of the year. The standard errors clustered at the individual level in parentheses. Significance levels: *** 1%, ** 5% and * 10%.

Table 3: Estimated treatment effects ($\times 100$) by group

	Number of observations	Unemployment exit hazard	Hazard rates to		
			Employment	Subsidized employment	Inactivity
	(1)	(2)	(3)	(4)	(5)
Baseline estimate	2,704,254	-0.1275***	-0.0496***	-0.0235*	-0.0545***
		(0.0272)	(0.0127)	(0.0133)	(0.0205)
Women	1,037,664	-0.1145**	-0.0127	-0.0427*	-0.0591*
		(0.0469)	(0.0216)	(0.0241)	(0.0347)
Men	1,666,590	-0.1366***	-0.0726***	-0.0120	-0.0520**
		(0.0330)	(0.0155)	(0.0154)	(0.0253)
Age 25–34	699,351	-0.0761	-0.0825**	0.0037	0.0028
		(0.0624)	(0.0340)	(0.0244)	(0.0474)
Age 35–49	833,608	-0.1336***	-0.0533**	-0.0016	-0.0787**
		(0.0478)	(0.0216)	(0.0238)	(0.0360)
Age 50–60	1,171,295	-0.1545***	-0.0289**	-0.0550***	-0.0705**
		(0.0374)	(0.0144)	(0.0209)	(0.0281)
Unemployed $<$ 6 months	834,750	-0.2564***	-0.0843**	-0.0774**	-0.0948**
		(0.0634)	(0.0329)	(0.0308)	(0.0458)
Unemployed 6–24 months	950,538	-0.0184	-0.0129	0.0097	-0.0152
		(0.0445)	(0.0198)	(0.0215)	(0.0339)
Unemployed $\geq$ 24 months	918,966	-0.1071***	-0.0532***	-0.0075	-0.0464*
		(0.0332)	(0.0102)	(0.0159)	(0.0275)

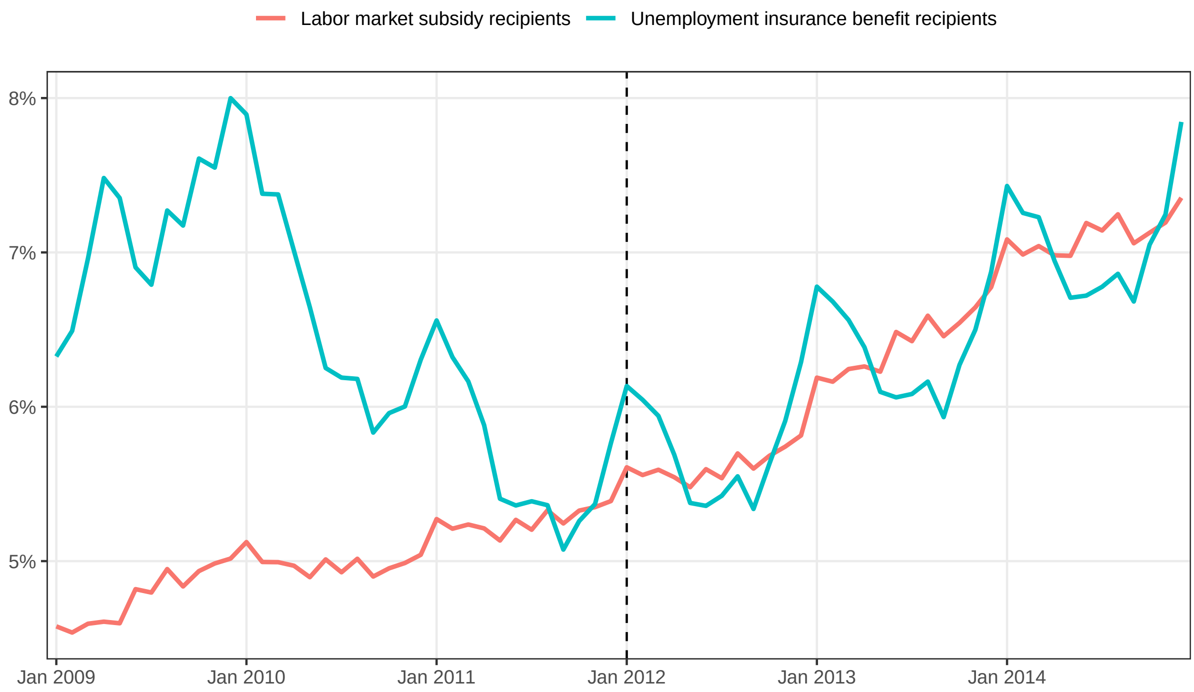
Notes: The estimating sample contains observations within the 13-week window from the turn of 2011 and 2012. The number of observations refers to the number of worker-week observations, i.e. the total number of workers at risk of exiting from unemployment over the period used in the estimations. All estimates are obtained from our baseline model specified in equation (1). The standard errors clustered at the individual level in parentheses. Significance levels: *** 1%, ** 5% and * 10%.

Figure 1: Seasonally and working day adjusted volume of production



Source: Statistics Finland.

Figure 2: Labor force share of unemployment insurance and labor market subsidy recipients by month



Notes: The numbers of labor market subsidy and unemployment insurance benefit recipients are based on data obtained from the benefit registers of Social Insurance Institution and Financial Supervisory Authority, respectively. These numbers are then divided by the size of the labor force obtained from the Labor Force Survey of Statistics Finland.

Figure 3: The number of new labor market subsidy recipients in the estimation sample by calendar month

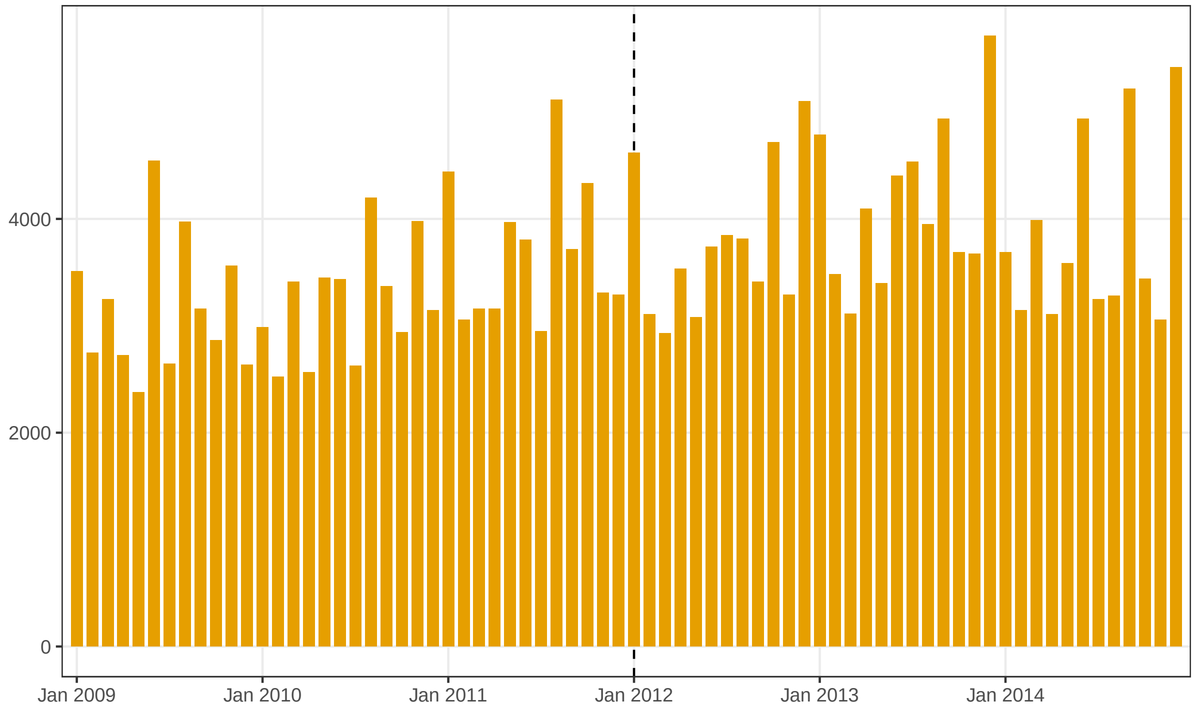


Figure 4: Seasonally adjusted average weekly hazards by month

Notes: The adjusted hazard equals the mean of residuals for a given month from the regression of weekly exits on calendar month dummies, plus the average weekly hazard rate over the whole observation period

Figure 5: Hazard rates around the time of the turn of years 2010 and 2011

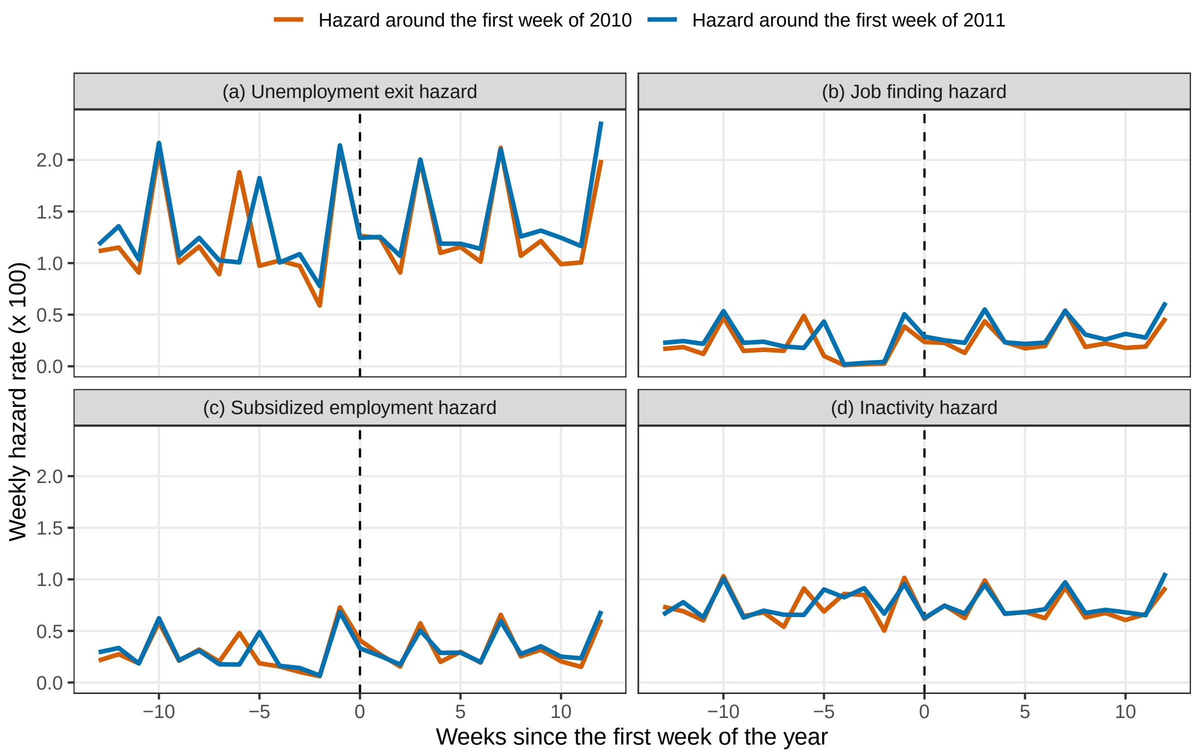


Figure 6: Hazard rates around the time of the turn of years 2011 and 2012

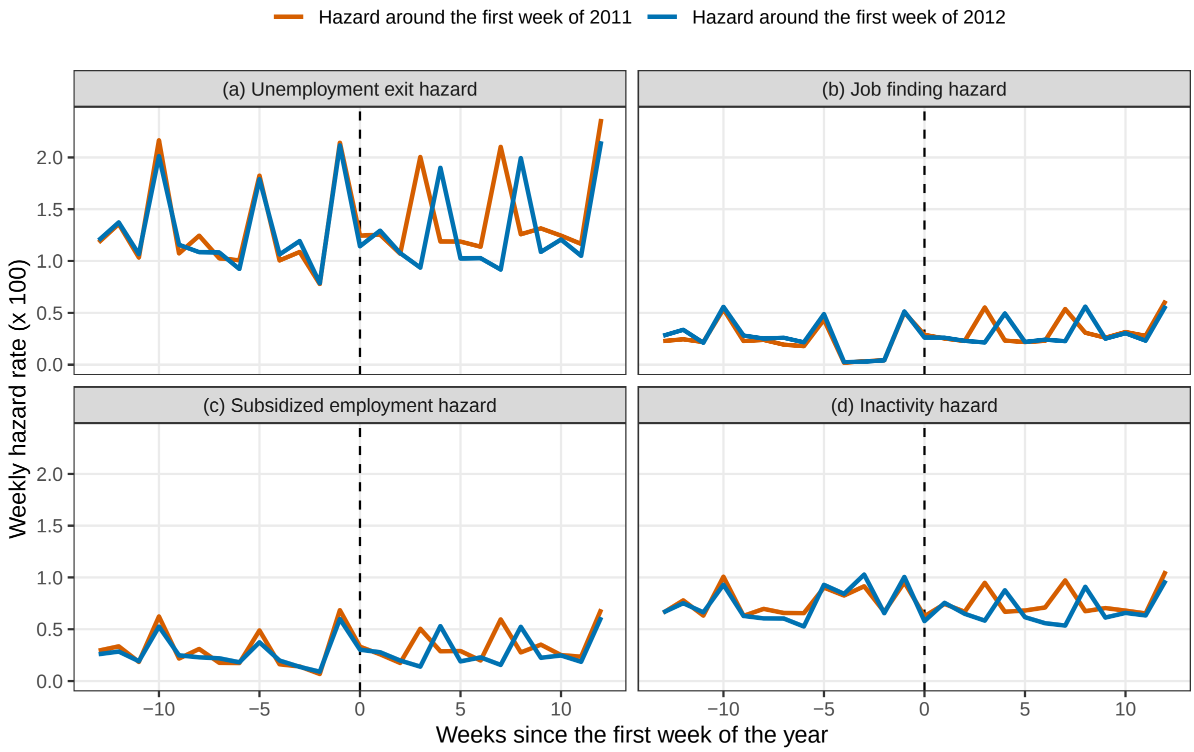


Figure 7: Hazard rates around the time of the turn of years 2013 and 2014

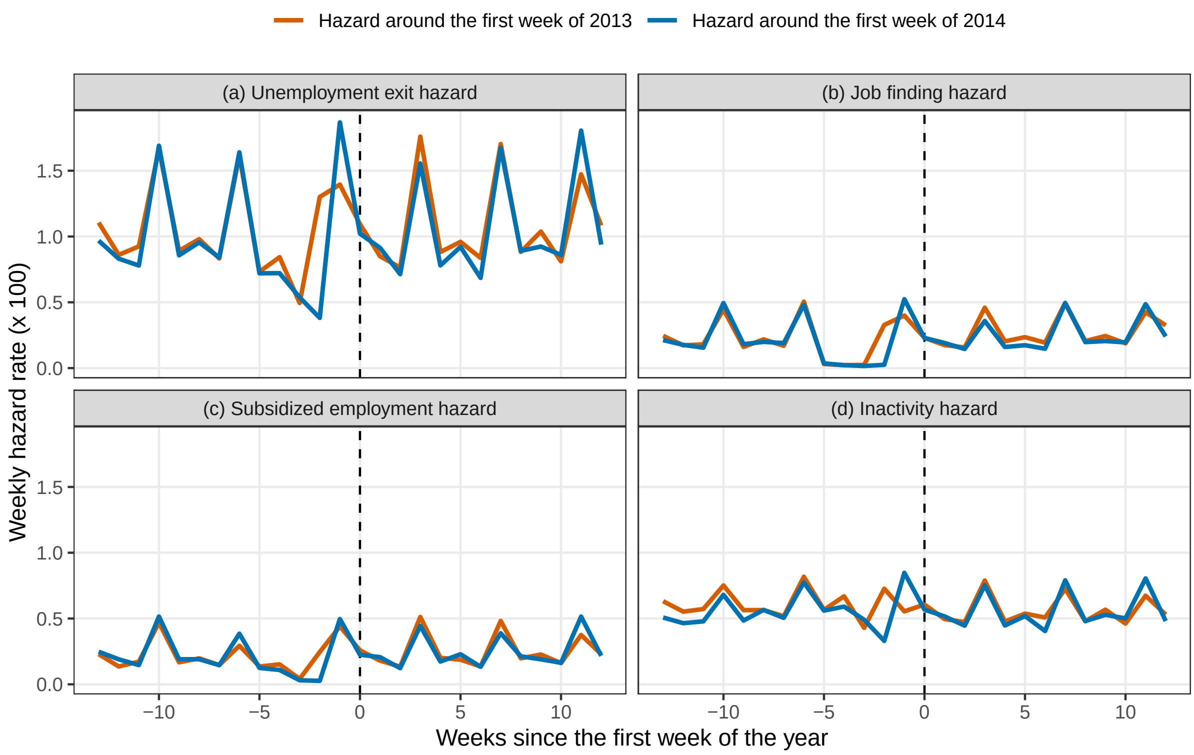


Figure 8: The effects of the 2012 reform and placebo reforms ($\times 100$) with the 95% confidence intervals using the 13-week window