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DO UNEMPLOYMENT INSURANCE BENEFITS IMPROVE MATCH QUALITY? EVIDENCE FROM RECENT U.S. RECESSIONS

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ABSTRACT

We present new evidence on the impact of more generous unemployment insurance (UI) on workers' ability to find jobs better suited to their skills. Using Longitudinal Employer-Household Dynamics data, we find the UI extensions introduced in the U.S. improved the quality of worker-job matches. Using Current Population Survey data, we also find that longer UI benefit durations decrease the mismatch between workers' educational attainments and the educational requirements of jobs. We find bigger effects of UI on match quality for those more likely to be liquidity—constrained women, non-whites and less-educated workers—,suggesting UI extensions improve the functioning of the labor market.

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1. Introduction

Unemployment insurance (UI) has long been seen as a government benefit that helps workers keep up their consumption through periods of unemployment until they can get a new job. A long-held view often highlighted by critics of UI is that when workers receive UI, they may put less effort into their search for a new job simply prolonging their periods of unemployment (i.e., the "moral hazard effect"). Recent work finds, however, that longer unemployment spells may also be due to credit-constrained workers searching longer for a better job match when they received more generous UI benefits (i.e., the "liquidity effect"). Card, Chetty and Weber (2007) and Chetty (2008) both find that the majority of the positive impact of UI on unemployment durations can be accounted by the "liquidity effect" and not the "moral hazard effect". In addition, contrary to most studies focusing on European data, Nekoei and Weber (2017) find a positive effect of UI on reemployment wages supporting the view that UI may improve subsequent match quality.

In this paper, we reexamine if UI increases wages by looking at the unemployment insurance extensions introduced during the past two recessions in the U.S. Importantly, we examine if the effect of UI on wages is due to better employee-employer matches from improved sorting; due to workers moving to higher paying firms; or due to stronger bargaining power of workers. Understanding these mechanisms can help explain why various studies find different average effects on wages. In addition, what mechanism is at work can also have different welfare implications of UI.

We exploit the fact that the UI extensions, during the past two recessions, had different durations in different states during each month, depending on whether each state qualified or not for the various extensions introduced by legislation. We first use the Longitudinal Employer-

Household Dynamics (LEHD) data to examine if, indeed, offering UI benefits for longer periods of time increases wages. Moreover, we examine if the effect comes from workers and firms matching better with each other or from workers moving to higher paying firms. An advantage of the LEHD is that it provides employer-employee data allowing us to control for many other factors that may affect earnings. We can focus on employers that are higher paying than others, even given their industry and location, and on individuals who are higher-paying, even though their education, age, gender, and race/ethnicity are the same. We use the Abowd-Kramars-Margolis (AKM) decomposition to estimate firm and worker fixed effects and then rank firms and workers by their percentile of pay and create a rank similarity index that captures how close the ranks of each firm and worker pair are.

We find that longer duration of UI benefits due to the introduction of the extensions increases the quality of employer-employee matches (i.e., reduces the difference in rank between the quality of the employer and employee). In particular, an increase in UI duration of 53 weeks, from 26 weeks (the average duration of UI in most states during normal times) to 79 weeks (close to the average UI benefit duration at the end of 2009) increases the similarity in the ranking by 1 percent. We also find this effect is greater for women than men, for minority than white workers, for less educated than more educated workers and for older than younger workers, all of whom are more likely to be credit-constrained. We also find that the increased duration of UI after the extensions increases the part of earnings that is unexplained by all other worker or firm characteristics, even after controlling for the bargaining power of workers as measured by unionization, pointing that this likely captures the match quality between the employer and employee. On the other hand, we do not find evidence that higher UI benefit duration during the Great Recession leads workers to move to higher paying firms.

We also use the Current Population Survey (CPS) to examine the impact of UI duration extensions during the Great Recession. Consistent with our LEHD results, we find that increasing UI benefit duration significantly decreases the mismatch between the educational attainment of workers and the educational requirements in their new jobs after their unemployment spells for women, non-whites, and less educated workers. Moreover, we find that increasing the weeks of UI benefit duration from 26 to 79 weeks increases the educational requirements in the occupation in the new job relative to the previous job by 14.4%. These results point towards workers finding better jobs (given firm quality) in which they are no longer over-educated.

Our results from the LEHD and CPS both indicate that increasing the generosity of unemployment insurance improves the quality of employee-employer matches and the functioning of the labor market. We do not find that higher UI iss associated with higher firm-effects during the Great Recession. Instead, we find that UI generosity allows workers to search longer and eventually find jobs better suited to their skills. This means that while some workers may move to better firms others may move to worse firms, depending on their individual rankings. Thus, while the "liquidity effect" is welfare enhancing, the fact that it works through improved employee-employer matches may mean that UI is less welfare enhancing than if UI uniformly increased employer quality for all workers.

In addition, the increased reallocation from more generous UI is likely to improve efficiency and to free up other jobs a worker would have taken, but which would not have been a good fit for this worker. In turn, this UI recipient who finds a job well-suited for herself opens these jobs up for other workers. This likely generates a chain reaction that allows all other workers to also match up with better employers given their qualifications and to improve labor market efficiency. If such an externality is present, then UI would be even more welfare enhancing.

The rest of the paper is structured as follows. In the next section, we present a brief survey of the related literature. In Section 3, we provide the institutional background on federal unemployment insurance programs in the U.S. and their extensions during the last few recessions. We describe the LEHD and CPS data in Section 4. We explain our empirical methodology in Section 5 and present our results from the LEHD in Section 6 and from the CPS in Section 7. Finally, we conclude in Section 8.

2. Literature Review

A robust finding in the empirical literature studying UI benefits is that the length of unemployment spells is positively related to the duration of UI benefits.[1] Despite the consensus that higher UI benefits lead to longer durations of unemployment, however, the magnitude of the effect varies across studies (see Card et al., 2007; Lalive, 2007; Van Ours and Vodopivec, 2008; Card et al., 2015; Farber and Valletta, 2015; Farber et al., 2015; Schmieder et al., 2016; and Nekoei and Weber, 2017).

There are two interpretations of the impact of UI benefit duration on longer unemployment spells. The first interpretation is that UI benefits decrease the job search efforts of workers because it raises the reservation wage of job seekers and/or because it generates moral hazard. Earlier research such as Moffitt (1985), Katz and Meyer (1990), Meyer (1990), and Card and Levine (2000) found support for this interpretation. Recent studies such as Rothstein (2011), Farber and Valletta (2015), and Farber, Rothstein and Valletta (2015) have examined the disincentive effects of UI extensions during and after the Great Recession. They found small but

[1] There are also a large literature examining the impact of UI benefit amounts on unemployment duration (Card et al., 2015; Hunt, 1995; Johnston and Mas, 2018; Landais, 2015), which we do not review in detail here since we do not focus on UI benefit amounts in our paper.

significant decreases in the probability of leaving unemployment, mostly driven by the decrease in the likelihood of moving out of the labor force and with little impact on the decline in the likelihood of exiting to employment. Mulligan (2012) argues that the decline in job search effort due to the extensions in the UI benefits may partially explain the slow recovery and the persistence of high unemployment rates. However, Kroft and Notowidigdo (2016) find that the moral hazard cost of UI is pro-cyclical, while the consumption smoothing benefit is acyclical. Thus, the adverse effects of UI benefits might be quantitatively small in a deep recession like the Great Recession when labor demand was weak, and the return to job search was lower.[2]

The second interpretation of the relationship between UI benefits and unemployment spells is that extensions in UI benefits allow workers to search for better jobs. Under this perspective, UI extensions correct distortions in the labor market and lead to better employment outcomes and higher productivity. Acemoglu and Shimer (1999) develop a theoretical model in which more generous UI encourages risk-averse workers to seek higher productivity jobs. Marimon and Zilibotti (1999) investigate the case in which agents are risk-neutral. Their search-matching model with risk-neutral agents and two-sided ex-ante heterogeneity predicts that UI benefits reduce employment, but also improve job matches. Chetty (2008) also shows that prolonged unemployment spells due to more generous UI can partly be attributed to liquidity-constrained individuals being able to search for longer. He finds that 60% of the increased duration of unemployment spells is due to the liquidity effect and the remaining 40% to the moral hazard effect. Card, Chetty and Weber (2007) also find that that the representative job searcher is much closer to credit-constrained behavior than to the permanent income hypothesis.

[2] Other studies have, instead, examined the impact of UI benefit generosity on labor demand and have found mixed results. Hagedorn et al. (2015) exploit the discontinuity of UI benefits at state borders and find that higher reservation wages due to extended benefits reduce firm vacancy creation rates. These findings are, however, in sharp contrast with Marinescu's (2015) work using online job postings, which finds no effects of UI extensions on labor demand.

The key implication from the theory that higher UI should increase wages due to liquidity-constrained individuals' ability to search for longer and find better jobs has also been explored empirically by a handful of papers using data from European countries. Among them, Lalive (2007), Schmieder et al. (2016), and Nekoei and Weber (2017) use different variations of age-based regression discontinuity designs by comparing individuals around an age cutoff that makes them eligible for more generous UI benefits durations.[3] On the other hand, Card, Chetty and Weber (2007) use a tenure-based regression discontinuity design comparing Austrian workers just before and after the 36 month cutoff for UI benefit eligibility. Finally, Van Ours and Vodopivec (2008) exploit the 1998 reform of the Slovenian UI system to identify the effect of the reduction in UI benefits. In particular, they compare the outcomes pre- and post-reform for workers who were employed for more than a year and a half prior to their unemployment spell, and who were exempted from the reform, and for others who were not exempt.

These previous studies find mixed results on re-employment wages. Card et al. (2007), Lalive (2007), and Van Ours and Vodopivec (2008) find no effect of UI on re-employment wages, while Schmieder et al. (2016) find a negative impact of UI on wages. The paper by Nekoei and Weber (2017) is the only one of these studies that finds positive and statistically significant estimates of the impact of UI on re-employment wages, which they interpret as being due to increased job searches by the unemployed. Nekoei and Weber(2017) reconcile their results with previous results by explaining that depending on the heterogeneity of the population the "liquidity effect" or the "moral hazard effect" may dominate.

[3] The age cutoffs and the number of UI weeks granted are country-specific. Nekoei and Weber (2017) use Austrian data and compare individuals around 40 years old that represent the cutoff after which individuals are eligible for 39 weeks of UI benefits instead of 30 weeks. Schmieder et al. (2016) use German administrative data and compare individuals younger and older than 42 years old, the age cutoff for eligibility for 18 months of UI benefits rather than 12 months. Finally, Lalive (2007) exploits the discontinuous changes in UI benefit duration at age 50 which are embedded in the Austrian UI system implemented in the regional extended benefit program.

Our paper contributes to this last branch of the literature by investigating the impacts of extensions in the duration of UI benefits on wages, the quality of employer-employee matches (as predicted by Marimon and Zilibotti (1999)), and the quality of jobs (as predicted by Acemoglu and Shimer (1999)) in the U.S. labor market, with a particular focus on the Great Recession. Our paper differs from the previous literature we use direct measures of similarity between worker and firm fixed effects as well as a direct measure of whether the firm uniformly pays high wages to investigate how UI benefits affect sorting and firm quality.[4] Nekoei and Weber (2017) provide some evidence suggesting that the positive effects of UI on wages are due to unemployed workers finding new jobs in larger firms, but their estimates are imprecise. They also examine the impacts of UI on the likelihood of moving firms, industries, occupations and geographic location and find no effects, though their occupation measure is very coarse. While they interpret this as lack of evidence on the impact of UI on reallocation, this analysis does not capture employer-employee matches. To our knowledge, our paper is the first paper that studies the effect of UI benefits on direct measures of employer-employee match quality and firm quality. Similar to Nekoei and Weber (2019), we may also expect to find bigger effects on wages and our match quality and firm quality measures than other studies, given the small effects of UI on the likelihood of exiting into employment found by Farber, Rothstein, and Valletta (2015).

3. Institutional Background

In the U.S., the UI system is a joint federal-state program. The Federal government sets minimum taxes, benefits, and standards, but each state is free to go beyond these minima. To qualify for UI, workers must have paid into the UI system through their employer, usually for the last four quarters before the start of the claim. They must also have received a minimum level of

[4] Our measures of sorting between employers and employees at the individual match level are computed using the estimation of employer and employee fixed-effects in the spirit of Abowd et al. (1999) as explained in Section 5.

earnings in the base period, which varies across states.[5] In addition, to qualify for benefits, workers must be currently totally or partially unemployed and must have lost their jobs other than for cause. Finally, the unemployed must be looking for work, and, in many states, they must show they have applied for jobs during the week for which they are claiming UI benefits.

In the vast majority of states, unemployed workers can claim UI benefits for up to 26 weeks. Eleven states offer 26 weeks of benefits uniformly to all workers. In another 31 states, the duration of UI benefits is capped at 26 weeks, but unemployed workers can only claim UI benefits for fewer weeks, depending on their contributions. Only Montana and Massachusetts provide more than 26 weeks of benefits. Montana pays benefits for up to 28 weeks, and Massachusetts pays benefits for up to 30 weeks, but only during periods of high unemployment. There are also 10 states which only provide benefit payments for less than 26 weeks. Florida and North Carolina provide 12 weeks of benefits; Alabama and Georgia provide up to 14 weeks of benefits; Arkansas and Kansas up to 16 weeks; and Michigan, Missouri, and South Carolina provide benefit payments for up to 20 weeks.[6]

Additional weeks of benefits are granted during recessions through Federal programs to ensure that workers who lose their jobs do not suffer massive drops in their income and consumption. There are two major Federal programs used to extend the duration of UI benefits during recessions. The first is the Extended Benefits (EB) program, which is a permanently authorized program established by Congress in 1970. The Federal-State Extended Unemployment Compensation Act establishes the provision of financial support to extend benefits for individuals who have exhausted their state UI benefits when unemployment rates are

[5] State-by-state laws with regards of minimum contributions are included in this report: https://oui.doleta.gov/unemploy/comparison/2010-2019/comparison2019.asp.

[6] These states are included in the analysis that use the CPS data. However, of these states, only Arkansas and Kansas are included in the analysis based on LEHD data, as we do not have access to data from the other states through the Census Research Data Center.

high. The EB Program grants an extension of UI benefits by 13 and 20 additional weeks if the 3-month state average unemployment rate exceeds 6.5% and 8.5%, respectively. The additional weeks of UI benefits can be claimed once the regular UI benefits are exhausted or once the extended weeks granted in the temporary programs are exhausted.

The second type of programs that typically extend the duration of UI benefits during recessions are federally-funded temporary UI benefit extension programs, which have been introduced during each recession since the late 1950s. Although these programs have changed over time in name, duration, and conditions to qualify, they all provide additional weeks for unemployment insurance benefits on top of the ordinary state UI benefits. Temporary Unemployment Compensation (TUC) programs were introduced between June of 1958 and June of 1959, between April 1961 and June of 1962, and between January of 1972 and March of 1973 for an additional period of 13 weeks. Federal Supplemental Benefits or Compensation (FSB or FSC) programs were introduced between January 1975 and January of 1978, and between September of 1982 and June of 1985 for various lengths of time. In the recession of the early 2000s, between March of 2002 and March of 2004, the TUC program was re-introduced, granting up to 13 or 26 additional weeks of benefits. Finally, the temporary benefit programs were renamed Emergency Unemployment Compensation (EUC) Programs between November 1991 and April 1994, and during and in the aftermath of the Great Recession, between July of 2008 and January of 2014.

The TUC 2002 Program consisted of two separate levels or tiers. Individuals who exhausted their regular state UI benefits were automatically eligible for 13 weeks of benefits. By contrast, the second tier contains a trigger mechanism and established a threshold requirement related to the state's unemployment rate. States with a 3-month average unemployment rate above 6.5% were eligible for an additional 13 weeks of UI benefits. The TUC 2002 Program jointly with the EB Program guaranteed up to a maximum of 72 weeks of UI benefits.

The EUC Program introduced in the last recession was extended several times. It became increasingly more complicated by adding more tiers over time. To simplify the exposition, Figure 1 provides a schematic summary of the available benefits during the Great Recession. Initially, EUC was launched from July 2008 to March 2009 and allowed the unemployed to claim an additional 13 weeks of benefits. In November 2008, the program expanded to allow 20 instead of 13 weeks of benefits. Between March and December 2009, the program changed from a one tier to a two-tier program, with those in the first tier continuing to receive 20 weeks and those in the second tier receiving an additional 13 weeks of benefits. Between December of 2009 and February 2010, the program consisted of four tiers. The first two tiers were activated unconditionally for all states and offered 20 and 14 additional weeks, respectively. The last two tiers were conditioned on state-specific unemployment rates. The third tier offered 13 extra weeks in states with a 3-month average unemployment rate above 6%, while the last tier offered 6 additional weeks of unemployment benefits to individuals residing in states with a 3-month average unemployment rate above 8.5%. This same four-tier program was then extended six more times through new legislation until May 2012. From May to September 2012, the program went back to a four-tier system with benefits of up to 20, 14, 13, and 6 weeks in Tiers 1 through 4, respectively. Starting in June 2012, the second tier required a 3-month total unemployment rate above 6% in the state to qualify, and the third and fourth-tiers now required unemployment rates above 7% and 9%, respectively. Finally, between September 2012 and January 2014, the program changed the four Tiers maximum benefit weeks to 14, 14, 9, and 10, respectively.

Since states could trigger "on" and "off" tiers 3 and 4 due to changes in their unemployment rates, UI benefit durations varied in a given year within each state and also varied across states. Since June 2012, states could also trigger "on" or "off" from Tier 2. Together with ordinary UI benefits and the additional weeks granted under the EB program, the EUC allowed individuals to use up to 99 weeks of UI benefits.

Figure 2 shows heatmaps constructed by using the monthly maximum amount of UI benefit weeks in three different periods.[7] The first row shows the ordinary UI benefits between January 2000 and December 2001, and between January 2005 and July 2008. The middle row refers to the early 2000s recession extension between January 2002 and December 2004. Finally, the bottom row reports the heatmaps for the Great Recession extension between August 2008 and December 2013. The maps on the left side include all U.S. states, while the maps on the right focus on the 20 states we use in the LEHD analysis.[8] Comparing the top to the two bottom panels in Figure 2 shows that the duration of UI benefits is always extended during downturns. In our analysis, we not only take advantage of variations in UI durations between recessions and expansions, but we also exploit variation on a monthly or quarterly during a recession. Figure 2 also shows a significant increase in the dispersion of UI benefits duration during recessions. While most states in our sample grant 26 weeks of UI benefits in normal times, the maps during the early 2000s recession and the Great Recession show much more variation in terms of duration of UI benefits.

Figures 3 and 4 examine each source of variation −across states and over time− further. Figure 3 reports the variation in weeks of UI benefits across states by quarter for the full sample

[7] The UI benefit weeks are grouped in Figure 2 based on the availability of weeks from the EB Program and the Four-Tiered EUC Program between December 2009 – May 2012 as showed in Figure 1.
[8] For graphical purposes, we omit Alaska and Hawaii in the maps.

of states (Panel A) and the 20 states in the LEHD data (Panel B).[9] There are two main takeaways from this figure. First, there is little variability in UI durations during expansionary periods, but the variation grows widely during quarters in recessionary periods when the UI extensions were introduced. Second, there is little variation in UI durations during the early 2000s, while the interquartile range in the duration of UI benefits was much wider during the Great Recession. While the maximum weeks of UI benefits the unemployed could obtain were 79 weeks by the second half of 2009, the maximum duration of UI benefits increased to 99 weeks for most states during the aftermath of the Great Recession. Yet, there are a few states with significantly lower benefit durations. The minimum amount of UI benefit weeks during the financial crisis is about 15 weeks below the number of weeks at the 25% percentile.

Figure 4 shows the variation in the duration of UI benefits over the period 2000-2013 for each individual state (red states are in both the LEHD and CPS data, while blue states are only included in the CPS data). We observe a large variation in UI benefits within states. The range between the minimum and the maximum amount of UI weeks is between 26 and 99 weeks for 65% of the states – 33 out of the 50 states and the District of Columbia. Figure 4 also highlights that the availability of large extensions of UI benefits is a rare event. Indeed, the maximum amount of UI benefit weeks is a little over 20 weeks for most states at the 75% percentile. Importantly, this figure shows that there is substantial variation across states.

4. Data

In this section, we first describe the data sources used in the empirical analysis and, then, present the descriptive statistics for the population of interest.

[9] Although we have collected weeks of UI benefits at the monthly frequency, this figure shows the quarterly average of the duration of UI benefits coming from the monthly data.

4.1. Longitudinal Employer-Household Dynamics (LEHD) Data

Our primary source of data is the Longitudinal Employer-Household Dynamics (LEHD). The LEHD infrastructure File system consists of restricted microdata information on earnings disbursements paid by employers to their employees as part of unemployment insurance. Those records cover nearly all private sector employment as well as state and local government employment. The coverage does not extend to self-employment or workers hired by the federal government.

The Census Bureau collaborates with its state partners through the Local Employment Dynamics (LED) cooperative federal-state program to compile the LEHD data infrastructure. Each LED partner state collects earnings data from Unemployment Insurance (UI) administrative files. The LED partner states also extract information from the Bureau of Labor Statistic's Quarterly Census of Employment and Wages (QCEW) administrative files. The Census Bureau receives these files and complements them with information on individual characteristics that are internally maintained.[10]

The records in the LEHD infrastructure correspond to any individual's UI-covered earnings paid by an employing tax-payer entity identified by a State Employer Identification Number (SEIN).[11] An individual is included in a quarter in the LEHD infrastructure if at least one employer reports earnings of at least one dollar for that individual during that quarter. We denote individuals to be employed in a quarter if they receive a non-zero earning from at least one employer in that quarter. We classify individuals to be non-employed in quarters in which they do not receive earnings from an employer during those quarters and assign them zero

[10] See Abowd et al. (2009) for a detailed description of the LEHD infrastructure File system.

[11] SEINs identify tax-payer entities. There is no one-to-one correspondence between tax-payer entities and firms. Although most firms have a unique tax identifier, there are few firms, usually large firms, that may have multiple SEINs due to tax advantages.

earnings.[12] We then construct our earnings measure by deflating the current earnings using the CPI Index adjusting for the Regional Price Parities Index to account for the difference in purchasing power between states.[13]

Each record is also completed with information on employers and individuals. The employers' data comes from each state's Department of Employment Security administrative files that are collected as part of the Covered Employment and Wages (CEW) program, jointly administered between the BLS and the Employment Security Agencies. Individuals' demographic characteristics integrated into the LEHD infrastructure come from two administrative data sources, the Person Characteristics File (PCF) and the Composite Person Record (CPR), both maintained by the Census Bureau. Individuals are uniquely identified by a Protected Identification Key (PIK) that tracks them across states and time.

The LEHD infrastructure contains matches between each employer-employee pair. This feature enables us to explore the effects of UI benefits extensions on labor market sorting during the last two recessions. We assign to each individual the maximum number of statutory UI weeks available based on the state of either their last or current employer. As the LEHD records are reported on a quarterly basis, and we have collected UI benefits weeks at a monthly frequency, we aggregate the weeks of UI benefits by averaging monthly benefit weeks to the quarterly level.

The Census Bureau has granted us access to 20 U.S. states, which account for about 50% of the U.S. labor market. The states included in our analysis are Arizona, Arkansas, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Iowa, Kansas, Maine,

[12] Since we only have access to 20 U.S. states, we cannot distinguish between non-employed individuals and individuals who have moved to one of the states which we do not have access to. We abstract from this possibility and we classify all these individuals as non-employed.

[13] The BEA calculates the RPP starting only in 2008. As the Regional Price Parities Index has very little volatility over time, we use an average over the available years to adjust the years before 2008.

Maryland, Nevada, Oklahoma, Pennsylvania, Tennessee, Texas, Washington, and Wisconsin. These 20 states have joined the LEHD microdata at different times.[14]

The core of the empirical analysis focuses on the effect of UI on the quality of employer-employee matches in the aftermath of the Great Recession. This was the period with the largest extension of UI benefits and the greatest variation across states. In this sub-period, 2008-2013, all 20 states participate in the LEHD. We also present the empirical results for the period from the first quarter of 2000 to the fourth quarter of 2013. DC and Arkansas entered the LEHD infrastructure after the beginning of this period – DC in the second quarter of 2002 and Arkansas in the third quarter of 2002.[15]

We restrict our analysis to the main job, which we define as the job with the employer at which that worker earns most of her earnings in a quarter. We restrict the sample only to workers who have had at least two different employers over the period 2000-2013. The sample includes both employed and non-employed workers. We restrict our sample to workers of working-age between 20 and 65 years old. The final sample for the empirical analysis consists of approximately 100,000,000 individuals for the period 2000-2013, and approximately 82,000,000 individuals for the period 2008-2013.[16]

Table 1 reports the summary statistics of the U.S. working population for the full period from 2000-2013 (Columns (1)-(3)) as well as for our main period of analysis −the period 2008-2013 (Columns (4)-(6)). Column (1) reports the summary statistics for the 20 U.S. states we have access to from 2000-2013, while Columns (2) and (3) provide summary statistics for the states with UI duration above and below the average UI duration across all 20 states for the full

[14] See Vilhuber (2018) for the exact dates in which each state has joined the LEHD program.
[15] The empirical results we present below are robust to the exclusion of these two states.
[16] Due to Census disclosure rules, we cannot report the exact number of observations in our dataset.

period.[17] The shares of male and female workers are about the same (52% vs. 48%). The working population in our sample is predominantly composed of white workers (80%). The share of workers with some college or more is higher than the percentage of workers with high school or less (57% vs. 43%).[18] There are more educated workers in states with UI benefit duration above the average in all states compared to those in states offering shorter UI benefit periods. Otherwise, workers in more and less generous UI states are similar in terms of gender, race, and age. The average quarterly earnings are around $4,600 in the full sample and in the sub-samples of workers living in states that offer UI benefits for longer and shorter periods.

The population characteristics are very similar for the full period (2000-2013) and the sub-period 2008-2013. Columns (4)-(6) shows summary statistics for the period 2008-2013 for all workers and workers in states with UI benefits duration above and below the average. The share of men (51% vs. 52%), whites (79% vs. 80%), and those with college (56% vs. 57%) are somewhat lower during the 2008-13 sub-period compared to the full period, but the average age is slightly higher (40 vs. 39 years of age). As for the full time period, the states with more and less generous benefits look very similar in terms of worker characteristics, except that workers in the more generous states are more educated. Quarterly earnings at around $4,700 during the period of the Great Recession and similar for workers living in more than less generous states.

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[17] The average number of weeks of UI is computed over the period 2000-2013 among the 20 states available to us through the LEHD program. The average UI is equal to 46 weeks. For comparability, we classify the states in the both periods, 2000-2013 and 2008-2013, using the same cutoff. Although the average UI for the sub-period 2008-2013 is higher (around 63 weeks), the classification between states above and below UI average is substantially unchanged by using the average over the period 2000-2013 or the average over the period 2008-2013. That is the case because most of the variation in the UI benefits that determines the classification between states above and below UI average comes from the sub-period 2008-2013. We classify states as above the mean as the states whose average over the 2000-2013 is above the mean across all states. States below the mean are classified following a similar procedure. The states with UI benefits above the average are: Arizona, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Nevada, Pennsylvania, Tennessee, and Washington. The states with UI benefits below the average are: Arkansas, Iowa, Kansas, Maine, Maryland, Oklahoma, Texas, and Wisconsin.

[18] Our estimates are similar to the ones reported from the Current Population Survey (CPS). The main difference is in level of education attainment. As the education achievements in the LEHD are imputed and based on the 2000 education distribution, it is thus not surprising that the share of more educated workers is lower in the LEHD than recent estimates from the CPS.

4.2. CPS Data

We also use the Monthly Public Data files from the Current Population Survey to examine the impact of the UI benefits extensions on wages and the job skill requirements for workers after unemployment spells. Households in the CPS are interviewed four months, then let go for eight months, then interviewed again for another four months. Every month one-eighth of the households enter the sample, and a similar number leaves the sample. Households are asked questions about wage income and hours worked in their fourth and eighth interviews, and we only use information from these months when analyzing outcomes. We leverage the longitudinal and rotating structure of the CPS sample to construct longitudinal histories of workers. In addition to the unique structure, the monthly CPS also collects extensive demographic characteristics and labor market information, including information on current employment occupation, past occupation (for those who are unemployed), education, age, gender, race, marital status, and state.

One of the main differences between the CPS and the LEHD is that while in the LEHD, we only observe the spell of non-employment, the CPS collects individual information on unemployment duration. This feature of the data enables us to construct the available UI benefits at the individual level as our main explanatory variable for our CPS analysis. The available UI benefits are calculated as the difference between the statutory UI benefits weeks in a state at a point in time minus the duration of unemployment for an unemployed individual in the sample. We restrict our CPS analysis to the period 2008-2013.[19]

[19] We restrict the analysis only to the Great Recession and its aftermath for two reasons. First, the BLS increased the sample size of the CPS in the early 2000s. Second, as showed in the previous sections, most of the variation in UI benefits comes from the Great Recession and its aftermath.

We link individuals from one month to the next using household and individual identifiers following the procedure outlined by Shimer (2012).[20] We then construct two different samples for our analysis. The first sample is restricted to workers who are present in the CPS in months 3, 4, 7 and 8 so that we made a transition from unemployment to employment either from month 3 to 4 or from month 7 to 8, or both. We use this sample to study how UI benefits extensions impact the job skill requirements for workers who move from unemployment to employment.

We use two distinct measures that capture the skill requirements. The first measure consists of the difference in the educational requirements the occupations in the current and previous jobs. To construct the educational requirements, we collect data from the U.S. Labor Department's O*NET database. The O*NET program gathers data on requirements for entry-level jobs, work styles, and task content within occupations by surveying each occupation's working population. We construct educational requirements by relying on the following survey question: "If someone was being hired to perform this job, indicate the level of education that would be required." The answers to the previous question are collected from the current employees.[21] The responses are recorded as a categorical variable.[22] To obtain a numerical proxy for the educational requirements for each occupation, we convert the categorical responses into years of education by multiplying the shares of responses for each category by the number of years required to attain the corresponding education level. For example, if 80% of respondents in the O*NET survey respond that a Ph.D. is required to perform the job of an Economist while

[20] In order to rule out spurious matches based on household and individual identifiers we perform checks such as the sex and the age of the individual is consistent from one month to the next.

[21] The survey respondents are reminded that their answers do not have to refer to the level of education that an incumbent or current employee has achieved, but they have to refer to a new hire.

[22] The categorical variable can take the following options: less than high school, high school, some college, associate' some college, associate, and graduate degree.

20% say that a Masters' degree is required, then we assign 17.6 (0.8 x 18 + 0.2 x 16) years of education as the requirement for Economists. We use these educational requirements of the workers' current jobs to construct our measure of educational mismatch as the difference between the workers' educational attainments and the level of education required in their current jobs. Thus, the educational mismatch is measured in terms of the difference in years of education, and it can take either positive or negative values reflecting whether a person is over-qualified or under-qualified for the job.[23]

The second measure we construct is an indicator that takes the value of one if the number of years of education required by the new occupation is greater than the number of years of education required by the previous occupation and zero otherwise.[24]

The second sample we have constructed is used to measure the earnings of workers who exit the labor market at a particular point in time. This sample is constructed by using the CPS monthly data files from IPUMS. We create longitudinal histories of workers several months apart by using the personal identifiers provided by IPUMS. As we are interested in wage outcomes for longer transitions, we focus on workers who have participated in all the 4-months in the CPS before leaving the sample[25] and we restrict our sample to workers who transitioned from unemployment to employment any time during months 1 through 3 and analyze their wages in month 4. As individuals may transit from unemployment to employment in one of the months from 1 through 3, we assign the available UI benefits for the month in which the individual made the transition. We use the this sample to study the medium-term effects on hourly wages.[26]

[23] We have also implemented alternative definitions of the educational mismatch such as using the mode of the responses as the required level of education for each occupation rather than the mean. Although these robustness checks are not reported in the paper, they generate quantitatively similar results.

[24] Notice that this variable would take a value of 0 for workers who do not change their occupation upon re-employment.

[25] In the analysis, we use individual weights from the 8th month.

[26] The nominal wages reported in the CPS are deflated by the 2010 national CPI to convert them in real wages.

Table 2 reports the descriptive statistics for the CPS. Column (1) shows the descriptive statistics for the entire sample, while columns (2) and (3) show the descriptive statistics for those in states with UI benefits duration above and below the average benefit weeks, respectively.[27] The first row of Table 2 shows that 81.5% of people exiting unemployment move to occupations with higher education requirements than their previous job. The share of those moving to jobs with higher education requirements after being unemployed is the same for those in states with UI benefit weeks above the average and those in states with shorter UI benefits durations. The third row shows that, on average, workers have 0.13 more years of education than it is actually required on the job, but that this mismatch is lower in more generous states where UI benefits duration is longer (0.11 vs. 0.19). The average age for those in the sample is 37 years old. Also, 81% of those in the sample are white, 58% are male, and 47% have some college or more as the highest educational attainment. Workers living in states with more generous UI benefits are more educated, younger, less likely to be white and male. It is worth highlighting that these summary statistics are not representative of the entire U.S. population. The differences are explained by the fact that our sample is restricted only to individuals who moved from unemployment to employment. For example, as a lower share of more educated workers is unemployed, it is reasonable to obtain a lower share of more educated individuals in the sample than in the national statistics. The last row of Table 2 reports the average hourly wages in month 4 after exiting from unemployment. The average hourly wage for the entire sample of states is about $12.50 US Dollars. The average wage for workers in states with UI benefit weeks above the

[27] The states above the average available UI are Alabama, Alaska, Arizona, California, Colorado, Connecticut, Delaware, District of Columbia, Georgia, Idaho, Illinois, Indiana, Kentucky, Maine, Massachusetts, Michigan, Nevada, New Jersey, New York, North Carolina, Ohio, Oregon, Pennsylvania, Rhode Island, Tennessee, Texas, Washington, West Virginia, and Wisconsin. The states below the average available UI are Arkansas, Florida, Hawaii, Iowa, Kansas, Louisiana, Maryland, Minnesota, Mississippi, Missouri, Montana, Nebraska, New Hampshire, New Mexico, North Dakota, Oklahoma, South Carolina, South Dakota, Utah, Vermont, Virginia, and Wyoming.

average is slightly lower than for those in states with shorter UI benefits durations (\$12.17 vs. \$12.74).[28]

4.3. State Controls

The extension of the UI benefits is concurrent with weak economic conditions, and some of the "tiers" were activated at the realization of extreme adverse economic events. Those weak economic conditions may also affect the quality of the matches between employers and employees. To address this potential issue in the identification of the effect of UI on labor market sorting, we include in the empirical specification state-specific macroeconomic variables to capture the deterioration of the economic environment when UI extensions are activated.

We collect unemployment rates for every state over time from the Local Area Unemployment Statistics (LAUS) at the Bureau of Labor Statistics (BLS). The BLS reports local unemployment rates at a monthly frequency. Since the LEHD microdata has quarterly periodicity, we average the 3-month unemployment rates within each quarter to obtain a quarterly rate. In the case of the monthly CPS data, instead, we use the monthly unemployment rates. The unemployment rate has reached its peak in the aftermath of the Great Recession with a spike of nearly 10%, driven mostly by large drops in demand and increases in layoffs.

We also include a second macroeconomic control −the quarterly gross state product (GSP)− from the Bureau of Economic Analysis (BEA). The BEA computes quarterly GSP data starting from 2005. For years before 2005, GSP series are calculated only annually. We assume

[28] The states above the mean of available UI are: Alabama, Alaska, California, Colorado, Connecticut, Delaware, District of Columbia, Idaho, Illinois, Indiana, Kentucky, Maine, Nevada, New Jersey, North Carolina, Ohio, Oregon, Pennsylvania, Rhode Island, Tennessee, Texas, Washington, West Virginia, and Wisconsin. The states below the mean of available UI are: Arizona, Arkansas, Florida, Georgia, Hawaii, Iowa, Kansas, Louisiana, Maryland, Massachusetts, Michigan, Minnesota, Mississippi, Missouri, Montana, Nebraska, New Hampshire, New Mexico, New York, North Dakota, Oklahoma, South Carolina, South Dakota, Utah, Vermont, Virginia, and Wyoming.

that for years before 2005, all quarters have the same value, and this value is equal to the annual GSP.[29]

The last control variable we use in some of the empirical specifications is the unionization coverage rate. Unionization coverage rates include both active union members as well as non-affiliated workers whose jobs are covered by a union or an employee association contract. The data are collected from the BLS as part of the Current Population Survey (CPS). The BLS reports annual rates for the period 2000-2013 by state. The annual rates are calculated as averages over the calendar year. In the empirical analysis, we assume that all quarters have the same annual unionization coverage rate.

5. Empirical Methodology

In this section, we first describe how we compute the quality of employers and employees. We then present three alternative strategies to identify the causal effect of extensions in UI benefits duration on the quality of the match between employers and workers.

5.1. Employer and Employee Quality Estimation

We derive the quality for each employer and employee by implementing a variation of the individual fixed effects methodology proposed in Abowd et al. (1999), and more recently, in Card et al. (2018). We estimate the following model:

$$\log(w_{ijsyq}) = \alpha_i + \theta_j + \mu_s + \tau_y + \omega_q + \varepsilon_{ijsyq} \qquad (1)$$

where w_{ijsyq} is the wage for worker i hired by employer j in state s in year y and quarter q, α_i are the employee fixed effects that capture the time-invariant worker characteristics that affect earnings, θ_j are the employer fixed effects which capture firm-specific pay premium, μ_s are the state fixed effects, τ_y are the year fixed effects, ω_q are the seasonal fixed effects, and finally,

[29] While the GSP is at quarterly frequency, the CPS data are monthly. We assume that all months in a quarter have the same value and that this value is equal to the quarterly statistic.

ε_{ijsyq} is the error term. We include year and seasonal fixed effects separately to capture both the fact that real wages may be changing over time, for example, due to productivity and that real wages may change due to changes in demand during high and low seasons that vary similarly every year.

Individual fixed effects are computed by implementing a variation of the algorithms used in Guimarães and Portugal (2010) and Crane et al. (2018).[30] Appendix B includes a detailed description of the steps of the computational algorithm. We estimate the individual fixed effects via an iterative procedure. We initialize the algorithm by guessing the worker types. We use as our initial guess the average of the workers' log earnings after controlling for life cycle patterns.[31] Then, we compute the employer fixed effects by removing the individual fixed effects from log earnings. The algorithm, then, computes iteratively the state, year, and seasonal fixed effects by progressively removing the previously computed fixed effects from the log earnings. Finally, we update the worker and employer fixed effects and calculate the R^2. The algorithm stops when convergence in the goodness-of-fit criterion is achieved, which we set to be a difference of less than 0.001 between the R^2's in two consecutive iterations.[32]

Although the quantification of the effects of UI on job matches uses only observations for the period 2000-2013, the estimation of the employee and employer fixed effects is based on the longest time series available. Short time series may generate imprecise estimates of the individual fixed effects due to the "limited mobility" bias.[33] The use of the longest time series possibly attenuates this bias and returns more precise estimates of the fixed effects.

[30] We are grateful to Henry Hyatt for having shared with us the SAS codes for his computational algorithm.

[31] Specifically, we regress the log earnings against the workers' age and squared age to absorb life cycle patterns for workers belonging to different birth cohorts.

[32] Convergence is achieved in approximately 10 iterations.

[33] See Abowd et al. (2004) and Andrews et al. (2008) for a discussion of the "limited mobility" bias and its effect on the individual fixed effects estimates. The "limited mobility" bias in the estimation of the worker fixed effects

5.2. Fixed Effects Ranking Approach

The first approach we use to get at the quality of matching is to measure positive assortative matching between employers and employees directly.[34]

We compute the percentile of each employer and employee fixed effects derived from equation (1). For any employee i hired by employer j, we then calculate the absolute value of the difference between the percentile of employee i and the percentile of employer j. This measure captures the distance in the position of an employer and an employee relative to their peers.

The distance between an employer and an employee is bounded between zero and one. The zero distance means that an employer and an employee are in the same percentile of their distribution. The distance of one implies that either an employer with the highest quality matches to an employee with the lowest quality or the other way around. The larger is the gap between an employer and employee, the more dissimilar they are. For interpretational purposes, we write the measure of positive assortative matching as one minus the absolute value of the difference between the employer and employee percentile:

$$d_{ijst} = 1 - |p_i - p_j| \qquad (2)$$

where p_i is the percentile for employee i, and p_j is the percentile for employer j.

This measure of positive assortative matching between an employer and an employee in equation (2) will then become our dependent variable in the following equation:

$$d_{ijst} = \delta_1 UR_{st-1} + \delta_2 \log(GSP_{st-1}) + \beta UI_{st-1} + u_{ijst} \qquad (3)$$

where UR_{st-1} is the lagged unemployment rate and $\ln GSP_{st-1}$ is the lagged log of the gross state product in state s in the previous period. We include these controls to make sure that we capture

occurs because the moves of workers across firms that helps identify these fixed effects is not frequent enough. Thus, the bias is bigger when there are fewer movers.

[34] We also examine the impact of UI on the likelihood of being employed by an employer with firm-ffects above the mean firm-effect in the sample.

the state of the economy in each labor market. UI_{st-1} are the maximum UI benefits weeks mandated in the state s in the previous period. A positive value of β is, thus, interpreted as a positive effect of UI extensions on assortative matching, i.e., employers hire more similarly ranked employees.

Extensions in UI benefits occur when economic conditions are weak. Weak economic conditions themselves may affect employer-employee matches.[35] This is why in the empirical model, we control for observable state's aggregate macroeconomic variables, such as unemployment rate and gross product, to directly capture the variation in job quality matches linked to weak economic conditions. In particular, we control for these to ensure that we distinguish the effect of the cycle from the effect of the unemployment insurance benefit duration. This is important because UI benefits were extended nationwide for longer periods when the state of the economy worsened.

Moreover, additional tiers of UI benefits sometimes triggered in response to higher unemployment rates in a state. Since worse labor market conditions may be related to hiring into worse jobs but also coincides with extended benefits, this may bias the effects of UI generosity on the quality of matches downwards. Finally, we used lagged explanatory variables by a quarter to address the possibility that better matches between employers and employees may be more productive and may lead to an increase in gross product, a decline in the unemployment rate, and UI benefits duration. The identification assumption we make is that past state economic conditions affect the current quality of matches, but the current quality of matches does not affect the previous economic conditions.

[35] Mueller (2017) shows that there are compositional changes in the pool of unemployed workers over the U.S. business cycle. Specifically, during recessions the pool of unemployed workers shifts towards workers with high wages in their previous job. Sedláček and Sterk (2017) and Moreira (2017) show that businesses births during downturns differ from businesses births during upturns. In fact, firms that start up during downturns start on a smaller scale and remain smaller over their entire lifecycle.

Legislative changes in UI benefits voted by Congress occurred due to weak economic conditions and aimed to ensure that unemployed workers during recessions do not suffer significant drops in incomes. As the current project focuses on the role of UI benefits extensions in potentially improving the functioning of the U.S. labor market through better matches between employers and employees, it is reasonable to think that the legislative changes are exogenous to individual matches. Therefore, we believe these legislative changes are a suitable natural experiment to investigate the causal effect of UI benefits on job quality matches. While discontinuity or kinked designs provide a clean method to identify UI impacts, an advantage of our identification design is that it can capture general equilibrium effects that will not be captured when one is examining local average treatment effects around a discontinuity.

We also investigate whether there are heterogeneous effects of UI benefits duration on match quality for different groups of workers. Specifically, we estimate models with interactions of the UI benefits durations with indicators for the following groups: men (vs. women); white (vs. non-white) workers; more-educated (vs. less-educated) workers; younger (vs. older) workers.[36]

5.3. Residuals Approach

The second measure of the quality of employer-employee matches is the residuals from the log wage model in equation (1). The residual term captures unobserved variables and, among other things, the quality of the match between worker i and employer j in state s in year t and quarter q. The residual term can also contain other effects on wages, including the bargaining power of workers or positive shocks to worker human capital. We include controls for the unionization coverage rate in state s at time t to control for the bargaining power of workers.

[36] We define a young worker as a worker younger than 40 years old. The choice of the threshold is motivated by the average age of the population in our sample reported in Table 1.

While the residual is a proxy for the employer-employee match quality, we use it as an alternative measure. We estimate the following model:

$$\hat{\varepsilon}_{ijst} = \delta_1 UR_{st-1} + \delta_2 \log(GSP_{st-1}) + \delta_3 Union\ Coverage\ Rate_{st-1} + \beta UI_{st-1} + u_{ijst} \quad (4)$$

where $\hat{\varepsilon}_{ijst}$ are the residuals from equation (1) where t is the combination of the year and quarter subscripts, UR_{st-1} is the state unemployment rate, $\ln GSP_{st-1}$ is the log of the gross state product, $Union\ Coverage\ Rate_{st-1}$ is the unionization rate, UI_{st-1} is the main variable and it measures the generosity of the UI benefits duration in state s, and u_{ijst} is the error term. We cluster the standard error at the state level to capture the correlation in the error terms. We expect β to be positive, meaning that more generous UI benefits improve the employer-employee matches. As before, we control for the lagged values of the state economic conditions, including the unemployment rate and the gross domestic product, to ensure we do not confound the effects of the unemployment benefit duration with the economic conditions in the state.

As before, we also estimate the differential effects of the UI benefits on different groups by the interaction of the UI benefit variable with a dummy for being a man, a dummy for being white, a dummy for having some college or more, or a dummy for being 40 years old or younger.

5.4. Effects on Skills Requirements

In our third empirical approach, we examine how UI extensions decrease the educational mismatch between worker education and the skills required for the job and increase the educational requirements in the new job relative to the previous job. The third approach enables us to study whether UI extensions allows workers to find better jobs where they are no longer over-educated.

Our estimation strategy is similar to the work of Rothstein (2011), Farber and Valletta

(2015), and Farber et al. (2015), who also use individual-level data from the CPS. We estimate the following model:

$$Y_{ist} = \psi_1 UR_{st-1} + \psi_2 \log(GSP_{st-1}) + \varphi\, AvailableUI_{st-1} \qquad (5)$$
$$+ \beta X_{it} + \mu_s + \tau_y + \omega_m + \varepsilon_{ist}$$

where Y_{ist} are the re-employment job quality or the education mismatch measures. $AvailableUI_{st-1}$ measures the available duration of benefits in each state during each month for a given individual i calculated as the total UI benefits weeks in the state s at a point in time minus the duration of unemployment for an unemployed individual in the sample. We control for state effects μ_s, year effects τ_y, and month effects ω_m to make our analysis consistent with the previous specifications, which also controlled for these fixed effects. In addition to the fixed effects, we include a set of individual demographic characteristics as controls. The set of individual controls X_{it} includes age, squared age, years of education, a dummy for race, a dummy for gender, a dummy for marital status.[37] As above, we also include the lagged unemployment rate and GSP to disentangle the impact of the policy changes from the effect of weak labor demand on match quality.

6. LEHD Results: Impacts on Firm-Worker Match Quality

Several studies proxy the impact of UI benefits on the quality of jobs by examining the impact on wages. Individual wages, however, capture many other effects aside from the quality of the match, including unobservable worker characteristics, quality of the employer, and bargaining power. To illustrate the differences between our approach and the previous analysis that simply examine the impact on wages, we first report the impact of UI benefit duration on wages in Table 3. Column (1) in Table 3 shows the impact of unemployment insurance duration

[37] In the LEHD analysis we do not include this set of individual demographic characteristics as controls because we control for individual fixed effects that absorb all these characteristics.

on the logarithm of earnings for the period from 2000 to 2013. UI benefits show no impact on earnings for the period for which we have data. However, Column (2) in Table 3 shows a positive effect of UI duration on earnings during the period of the Great Recession. The results show that an increase of 53 weeks, from the standard benefits of 26 weeks to 79 weeks (the maximum benefits at the end of 2009), increases wages by 2.6%. In these regressions, we already control for the unemployment rate in the state and gross state product in addition to state, year, and seasonal fixed effects. As expected, a higher unemployment rate in the past reduces earnings, though the effect is insignificant. By contrast, the effect of GSP on wages is positive and significant at the 5% level.

6.1. Impacts on Ranking Measure

Instead of looking at wages, which captures many other effects, we examine the impact of UI duration on the distance between the ranking of firms and workers. Table 4 shows the results on ranking for the full period from 2000 to 2013 as well as for the period of the Great Recession. Column (1) in Panel A in Table 4 shows the impact of UI duration on the similarity in the ranking. The effect indicates that an increase of 53 weeks in UI benefits increases the similarity between the worker and firm ranking by 1%. Column (1) in Panel B in Table 4 shows the effects of UI benefit extensions during the Great Recession. The effects are bigger during the Great Recession –an increase of 53 weeks increases the similarity between workers' and firms' rankings by 1.1%. To ensure we do not confound the effect of labor market conditions with UI benefit duration, we control for the lagged unemployment rate and gross state product in all of our estimations. The effect of the unemployment rate (UR) is negative but insignificant, and the effect of Gross Domestic Product (GSP) is positive and significant at the 5% level. These results suggest that if a worker can receive UI benefits for a longer period, she will be able to find a job with an employer that is closer to her in terms of quality. This worker then is likely to leave another job open for someone else who is also likely to be better matched, and in turn that other worker can also leave vacant another job and relieve it to someone else, generating a chain reaction that makes many other workers, beyond the one receiving the UI extension, match better in the labor market.

We examine models with interactions to check whether the impacts are bigger on workers by gender and race. Column (2) in Panels A and B in Table 4 show that the interaction term of the UI duration on the male dummy is negative and statistically significant. The effects of UI benefits are, thus, bigger on women than men. An increase of 53 weeks of UI benefits increases similarity in the ranking by 1.3% for women and by 0.9% for men. Similarly, the effects were greater for women than men during the Great Recession. Column (3) in Panels A and B in Table 4 show that the interaction term of UI duration with a dummy for white workers. The interaction term with the white dummy is negative and statistically significant. The results show that an increase of 53 weeks improves the match quality by 1.2% for minority workers and by 0.9% for whites for the entire period. During the Great Recession, minority workers benefit even more from extended benefits; indeed, a 53 weeks extension increases the ranking similarity by 1.3%.

Columns (4) and (5) in Panels A and B of Table 4 also show the differential effects of UI extensions on workers with different levels of education and different ages. Column (4) shows that the interaction with the dummy for those with some college or more education is negative and significant. Thus, the least educated benefit more from UI benefit extensions both during the whole period as well as during the Great Recession sub-period. For instance, for the full period, an increase of 53 weeks of UI increases the similarity in the ranking by 1.3% for less educated workers and by 1% for more educated workers. Column (5) in Panels A and B also show that the interaction with the under 40 years old group is negative and statistically significant. Older workers benefit more from UI benefits both during the Great Recession and during the whole period. The results show that an increase of 53 weeks improves the ranking similarity by 1.7% for those over 40 and by 0.5% for workers under 40.

The larger effects on women, minorities, and less educated workers suggest that workers with greater credit constraints probably benefit most from having access to unemployment insurance benefits for a longer period of time. The larger effects for older workers also likely reflects that workers with families may be more likely to benefit from receiving UI benefits for a longer period instead of having to accept the first job that they receive an offer from.

6.2. Impacts on Residual Measure

Table 5 shows results in the wage residuals are the outcome, controlling for the state UR, GSP, and unionization coverage rates in the state. Column (1) in Panels A and B show the results for the full sample. The results for the full period of the analysis show no effects, but the effects of UI benefits on improved residual wages are positive during the Great Recession sub-period. The magnitude shows that an increase of 53 weeks in the duration of UI benefits (the equivalent of increasing maximum benefits from 26 weeks to 79 weeks) increases unexplained wages by 1.1%, and an increase of 73 weeks in the duration of UI benefits (i.e., an increase in maximum benefits from 26 to 99 weeks) increases unexplained wages by 1.5%, which is much lower than the 2.6% increase in log wages when not controlling for other factors. These specifications show, negative and significant effects of the lagged unemployment rate and positive and significant effects of lagged GSP. However, the results show no effect of unionization coverage rates on residual wages suggesting that bargaining power through unionization does not explain much of the unexplained component of wages.

Columns (2)-(5) in Panels A and B of Table 5 show differential effects on different demographic groups. As before, these results show that the impact of UI benefit duration is bigger for women and non-whites both for the full period and for the Great Recession (columns (2) and (3) in Panels A and B). Column (4) in Table 5 also shows greater effects on the less educated for the Great Recession, though the results show the opposite for the entire period. By contrast, these results show no differential effects between older and younger workers. Thus, these results confirm that women, minorities and less educated workers, all of whom typically face greater credit constraints, benefit the most from more generous UI benefits.

7. CPS Results: Impacts on Educational Requirement Matches

In this section, we examine the impact of UI benefits on wages and educational mismatch using CPS data. We first examine the impact of UI benefits on wages and then turn to the impacts of UI benefits on match quality as measured by the disparity between a workers' educational attainments and the educational requirements of their jobs and on getting access to jobs with higher educational requirements.

Column (3) in Table 3 shows similar effects of UI on wages with the CPS data, as we found with the LEHD in Columns (1) and (2) in Table 3. The results in Column (3) show that 53 additional weeks of available unemployment benefits increase wages by 4.4%. The relation of UI to wages is, thus, even bigger than the 2.6% increase in wages we found with LEHD data. However, in this case, the lagged unemployment rates and lagged GSP do not have a significant impact on wages.

Table 6 shows results of the impact of UI on whether an individual gets a better quality job and a job better matched to their skills.

We start by analyzing whether access to longer UI benefits leads to any decreases in the mismatch between worker education and the skills required for the job in Panel A of Table 6. Column (1) in Panel A of Table 6 shows that higher available weeks of benefits reduce mismatch, but the effect is not statistically significant. Columns (2)-(5) show models with interactions. The main effect is significant in all of these cases except for in the model in Column (5) which includes the young vs. old interaction. Also, the only interaction term which is significant, but in the opposite direction, is the one with more educated workers. These estimates, thus, show that the effects are bigger for women, non-whites, and less educated workers. The findings imply that an increase in the availability of UI benefits reduces the mismatch between the educational attainment of the person and the educational requirement of the job. An increase of 53 weeks of available UI benefits reduces the mismatch by 60% for women and for non-whites, and by about 100% for those with less than a college degree.

Panel B in Table 6 shows the effects of UI benefits on the difference in education requirements of the job upon exiting unemployment compared to the worker's last job. We find that workers who have access to more weeks of UI benefits can find jobs that have higher education requirements than the ones they were doing before. Thus, we find evidence that having access to a safety net that allows workers to search longer can lead to occupational upgrading. The estimate from Column (1) implies that an additional 53 weeks of UI benefits increases the likelihood that workers end up in jobs with higher education requirement by 11.7 percentage points or by 14.4%. Columns (2) through (5) show models with interactions with a male dummy, a white dummy, a college dummy, and a dummy if the person is younger than 40 years old. None of these interactions are significant, but the main effect is significant in all of these cases, meaning that the effects are bigger for women, non-whites, non-college graduates, and older workers. Given the evidence in Marinescu (2017) showing that firms do not change skills requirements in response to UI during the Great Recession, we interpret this as reflecting workers' response to UI by finding better jobs in which they are not over-educated. Note that this is consistent with the lack of finding of UI improving the quality of firms workers move to, since workers could be moving to better jobs in new occupations even if they stay in firms of the same quality.

Thus, like the evidence from the LEHD, the results from the CPS show that more generous unemployment insurance during the Great Recession increased wages. Moreover, our CPS results show that more generous UI improved the quality of matches and the quality of jobs obtained after unemployment, so as to reduce over-education.

8. Conclusion

This paper provides new evidence that more generous UI benefits improve sorting in the labor market. Most previous papers in this area focus on the adverse effects of unemployment insurance benefits in terms of prolonging unemployment spells. By contrast, there is a much slimmer body of literature focusing on the potential positive effects of unemployment benefits. There are a small number of convincing papers showing the consumption smoothing effects of UI and an equally small but less conclusive number of studies examining the impact of UI on earnings in Europe.

Here, we provide new evidence on the positive impact of UI extensions on wages in the U.S. during the last two recessions using both LEHD and CPS data. Moreover, we examine if this positive impact of UI on wages is due to improved employer-employee matches or movements to higher paying firms. Using employer-employee data from the LEHD, we find that increasing the weeks of UI benefits reduces the difference in the quality rankings between workers and employers. The fact that match quality improves weighs against an interpretation that UI increases the bargaining power of workers vis-à-vis employers. If that was the case, there would be more (not less) disparity between the worker and the firm ranking. By contrast, we find that higher UI is not associated with higher wage premiums paid by firms during the Great Recession. We similarly find that the unexplained part of earnings increases when UI benefits durations are lengthened even after controlling for bargaining power from unionization, and we find that the impact on wages over-estimates the impact on match quality. We find that the effect of UI benefits duration on these two measures of match quality is greater for less-educated workers, women, and minority workers, all of whom are more likely to be credit-constrained.

We also find positive impacts of UI extensions during the Great Recession using CPS data. We find evidence that greater duration of UI benefits reduces the mismatch between the actual educational attainment of workers and educational requirements for jobs and, thus, reduces the phenomenon of over-education for women, non-whites and less educated workers. We also find that a greater number of weeks of UI benefits increases the likelihood of finding a job with higher educational requirements after unemployment, which we interpret as being due to the workers' ability to find better jobs where they are no longer over-educated even if they remain in a firm of the same quality.

Our paper provides evidence using different measures and different data sets that more generous UI improves the functioning of the labor market by allowing better sorting of workers and employers. In particular, we find that UI benefits those who are more likely to be liquidity constrained by allowing them to search for a longer period of time until they find jobs better fit to their skills. Our evidence is, thus, consistent with the "liquidity effect" dominating the "moral hazard effect" of UI extensions and UI benefits enhancing welfare.

At the same time, UI benefits may also have distributional effects. Our findings imply that not only will better workers be matched to better firms, but also worse workers may be matched with worse employers. Thus, lower performing workers may now earn less because they are no longer matched with higher productivity firms. While more even matches may be efficient, more balanced matches may reduce earnings for lower performing workers who may had previously been able to find a job with a better employer. In this sense, the welfare enhancing role of UI may be smaller when UI affects employee-employer assignments rather than the creation of better jobs.

There is, however, another distributional effect of UI which likely enhances welfare. UI can improve the likelihood that lower performing workers are now able to get a job rather than remain unemployed because other jobs are now freed up by UI recipients moving to better jobs. The change in search behavior by UI recipients is likely to have positive externalities on many other workers in their labor market. When a worker waits to get another job that is better suited to their skills, they turn down other jobs that may also be better suited for others in the same labor market. These externalities enhance the welfare effects of UI extensions like those introduced during the past few recessions.

Our paper suggests that UI extensions during the Great Recession probably helped to mitigate the big wage drops experienced by displaced workers due the loss of valuable specific worker-employer matches found by Lachowska, Mas and Woodbury (forthcoming). The UI extensions were also likely important in boosting workers' wages given the inability to move to higher paying firms due to the collapse in firm-wage ladders during this time documented by Haltiwanger et al. (2018). Similarly, the Pandemic Unemployment Assistance (PUA) introduced recently as part of the CARES Act or future legislation in response to the COVID-19 Recession will potentially play a role in increasing wages after workers are able to get back to work following the economic fallout related to the pandemic.

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Table 1: Descriptive Statistics of Workers' Characteristics, LEHD 2000-2013

	2000-2013			2008-2013		
	All States	States Above UI Average	States Below UI Average	All States	States Above UI Average	States Below UI Average
Share of Men	0.5175	0.5192	0.5142	0.5139	0.5146	0.5125
Share of Whites	0.7964	0.7935	0.8024	0.7911	0.7879	0.7973
Share with Some College or More	0.5677	0.5788	0.5453	0.5619	0.5743	0.5377
Age (years)	39.08	39.20	38.84	39.62	39.75	39.37
Quarterly earnings	4,614 (211.4)	4,628 (203.3)	4,584 (233.6)	4,693 (211.7)	4,688 (209.4)	4,704 (218.5)

Notes: The average number of weeks of UI is computed over the period 2000-2013, and it is equal to 46 weeks. The states with UI benefits above the average are Arizona, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Nevada, Pennsylvania, Tennessee, and Washington. The states with UI benefits below the average are Arkansas, Iowa, Kansas, Maine, Maryland, Oklahoma, Texas, and Wisconsin.

Table 2: Descriptive Statistics of Workers' Characteristics, CPS

	2008-2013		
	All States	States Above UI Average	States Below UI Average
Share of Workers that Move to a Job with Higher Education Requirement	0.815	0.815	0.815
Difference in Education Requirement between New and Old Job	-0.0405 (1.353)	-0.0183 (1.334)	-0.0510 (1.362)
Education Mismatch (Years)	0.131 (1.895)	0.185 (1.846)	0.105 (1.918)
Years of Education	12.99 (2.126)	12.97 (2.175)	13.01 (2.080)
Share of Men	0.583	0.573	0.587
Share of White	0.809	0.789	0.818
Share with Some College or More	0.471	0.476	0.469
Age (years)	36.83	36.61	36.93
Hourly Wage upon Exit from Unemployment (USD)	12.49 (1.782)	12.17 (1.795)	12.74 (1.770)

Notes: The statistics in all rows but the last are computed using the Monthly Data files from the Current Population Survey for the period 2008-2013. The last row uses the CPS monthly data files from IPUMS to construct longitudinal histories of workers. The states are classified as above or below the mean based on the "short sample" averages. The states above the average UI are Alabama, Alaska, Arizona, California, Colorado, Connecticut, Delaware, District of Columbia, Georgia, Idaho, Illinois, Indiana, Kentucky, Maine, Massachusetts, Michigan, Nevada, New Jersey, New York, North Carolina, Ohio, Oregon, Pennsylvania, Rhode Island, Tennessee, Texas, Washington, West Virginia, and Wisconsin. The states below the average UI are Arkansas, Florida, Hawaii, Iowa, Kansas, Louisiana, Maryland, Minnesota, Mississippi, Missouri, Montana, Nebraska, New Hampshire, New Mexico, North Dakota, Oklahoma, South Carolina, South Dakota, Utah, Vermont, Virginia, and Wyoming.

Table 3: Effects on UI Duration on log wages

	LEHD		CPS
	(1)	(2)	(3)
	2000-2013	2008-2013	2008-2013
Unemployment Insurance Duration	0.0001397	0.0004995***	0.000831***
	(0.0002679)	(0.0001752)	(0.000280)
Unemployment Rate	-0.007230	-0.002270	-0.0116
	(0.005166)	(0.002706)	(0.00808)
Gross State Product	0.1903**	0.2574***	-2.05e-09
	(0.06964)	(0.04148)	(0.000000330)
No. Observations	3,419,000,000	1,485,000,000	9824
R^2	0.002661	0.002773	0.023
State FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Seasonal FE	Yes	Yes	Yes

Notes: Results in columns 1 and 2 are based on the LEHD data and include the following states: Arizona, Arkansas, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Iowa, Kansas, Maine, Maryland, Nevada, Oklahoma, Pennsylvania, Tennessee, Texas, Washington, and Wisconsin. Earnings in Columns 1 and 2 are deflated by using the CPI Index adjusted for the Regional Price Parities Index to account for differences in purchasing power between states. Results in column 3 use the "longitudinal" sample from the monthly CPS data. The CPS sample includes all 50+1 states. While in columns 1 and 2, the frequency of the seasonal FE is quarterly; in column 3, the seasonal frequency is monthly. Column 3 includes as controls a set of individual demographic characteristics such as age, squared age, number of years of education, and dummies for the race, gender, and marital status. The wages in Column 3 are deflated by the 2010 national CPI. Standard errors are included in parenthesis and are clustered at the state level. The asterisks *, **, and *** reflect the 10%, 5%, and 1% significance levels, respectively.

Table 4: Effects of UI Duration on Ranking Difference between Workers and Firms, LEHD 2000-2013

	(1) All Workers	(2) Men vs. Women	(3) Whites vs. Non-whites	(4) More-educated vs. Less-educated	(5) Young vs. Old
	Panel A: Period 2000-2013				
UI Duration	0.0001**	0.0002**	0.0002**	0.0002**	0.0002***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Group Dummy		0.0339***	0.0187***	0.0146***	0.0307***
		(0.0009)	(0.0034)	(0.0009)	(0.0017)
UI Duration × Group Dummy		-0.0001***	-0.00004**	-0.00004***	-0.0002***
		(0.0000)	(0.0000)	(0.0000)	(0.0000)
Unemployment Rate	-0.0008	-0.0008	-0.0006	-0.0009	-0.0007
	(0.0011)	(0.0011)	(0.0011)	(0.0011)	(0.0011)
Gross State Product	0.0040**	0.0039*	0.0045**	0.0040*	0.0039*
	(0.0019)	(0.0019)	(0.0019)	(0.0020)	(0.0019)
No. Observations	3,419,000,000	3,419,000,000	3,419,000,000	3,419,000,000	3,419,000,000
R^2	0.0005	0.0062	0.0015	0.0014	0.0037
	Panel B: Period 2008-2013				
UI Duration	0.0002***	0.0002***	0.0002***	0.0002***	0.0002***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Group Dummy		0.0305***	0.0194***	0.0120***	0.0260***
		(0.0008)	(0.0033)	(0.0009)	(0.0019)
UI Duration × Group Dummy		-0.00002**	-0.00005**	-0.00001*	-0.0001***
		(0.0000)	(0.0000)	(0.0000)	(0.0000)
Unemployment Rate	-0.0018*	-0.0018*	-0.0017*	-0.0020*	-0.0018**
	(0.0010)	(0.0009)	(0.0010)	(0.0010)	(0.0009)
Gross State Product	0.0036*	0.0035*	0.0041**	0.0037*	0.0036*
	(0.0018)	(0.0018)	(0.0018)	(0.0018)	(0.0018)
No. Observations	1,485,000,000	1,485,000,000	1,485,000,000	1,485,000,000	1,485,000,000
R^2	0.0003	0.0054	0.0012	0.0010	0.0024

Notes: The states included in the analysis are Arizona, Arkansas, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Iowa, Kansas, Maine, Maryland, Nevada, Oklahoma, Pennsylvania, Tennessee, Texas, Washington, and Wisconsin. The variable "Group Dummy" changes from columns 2 to 5. In Column (2), it takes the value of 1 if an employee is male and 0 otherwise. In Column (3), it takes the value of 1 if an employee is white and 0 otherwise. In Column (4), it takes the value of 1 if an employee's highest education is some college or more and 0 otherwise. In Column (5), it takes the value of 1 if an employee is younger than or equal to 40 years old and 0 otherwise. Standard errors are included in parenthesis and are clustered at the state level. The asterisks *, **, and *** reflect the 10%, 5%, and 1% significance levels, respectively.

Table 5: Effects of UI Duration on Wage Residuals, LEHD 2000-2013

	(1) All Workers	(2) Men vs. Women	(3) Whites vs. Non-whites	(4) More-educated vs. Less-educated	(5) Young vs. Old
	Panel A: Period 2000-2013				
UI Duration	0.0002	0.0004***	0.0004***	0.0002	0.0002
	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)
Group Dummy		0. 0098***	0.0068**	0.0061***	0.0008
		(0.0013)	(0.0032)	(0.0012)	(0.0016)
UI Duration × Group Dummy		-0.0003***	-0.0002***	0.0001***	0.0000
		(0.00003)	(0.00005)	(0.00003)	(0.00004)
Unemployment Rate	-0.0031**	-0.0031**	-0.0032***	-0.0031**	-0.0031**
	(0.0012)	(0.0012)	(0.0011)	(0.0012)	(0.0012)
Gross State Product	0.0026**	0.0026**	0.0025**	0.0026**	0.0026**
	(0.0009)	(0.0009)	(0.0010)	(0.0009)	(0.0009)
Union Representation	0.0003	0.0003	0.0003	0.0003	0.0003
	(0.0003)	(0.0003)	(0.0003)	(0.0003)	(0.0002)
No. Observations	3,419,000,000	3,419,000,000	3,419,000,000	3,419,000,000	3,419,000,000
R^2	0.00001	0.00005	0.00003	0.00005	0.00001
	Panel B: Period 2008-2013				
UI Duration	0.0002**	0.0003**	0.0004***	0.0003**	0.0002**
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
Group Dummy		-0.0025	0.0025	0.0224***	0.0009
		(0.0043)	(0.0026)	(0.0019)	(0.0039)
UI Duration × Group Dummy		-0.0002***	-0.0002***	-0.0001***	0.0000
		(0.0000)	(0.0000)	(0.0000)	(0.0000)
Unemployment Rate	-0.0030***	-0.0030***	-0.0032***	-0.0030***	-0.0030***
	(0.0009)	(0.0009)	(0.0009)	(0.0009)	(0.0009)
Gross State Product	0.0054***	0.0054***	0.0051***	0.0054***	0.0054***
	(0.0013)	(0.0013)	(0.0012)	(0.0013)	(0.0013)
Union Representation	-0.0001	-0.0001	-0.0010	-0.0002	-0.0001
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
No. Observations	1,485,000,000	1,485,000,000	1,485,000,000	1,485,000,000	1,485,000,000
R^2	0.00004	0.0001	0.0001	0.0001	0.00004

Notes: The states included in the analysis are Arizona, Arkansas, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Iowa, Kansas, Maine, Maryland, Nevada, Oklahoma, Pennsylvania, Tennessee, Texas, Washington, and Wisconsin. The variable "Group Dummy" changes from Columns (2) to (5). In Column (2), it takes the value of 1 if an employee is male and 0 otherwise. In Column (3), it takes the value of 1 if an employee is white and 0 otherwise. In Column (4), it takes the value of 1 if an employee's highest education is some college or more and 0 otherwise. In Column (5), it takes the value of 1 if an employee is younger than or equal to 40 years old and 0 otherwise. Standard errors are in parenthesis and are clustered at the state level. The asterisks *, **, and *** reflect the 10%, 5%, and 1% significance levels, respectively.

Table 6: Effects of UI Duration on Educational Requirement Matches, CPR 2008-2013

	(1) All Workers	(2) Men vs. Women	(3) Whites vs. Non-whites	(4) More-educated vs. Less-educated	(5) Young vs. Old
	Panel A: Mismatch in Years of Education				
Available UI Duration	-0.000665	-0.00142**	-0.00140**	-0.00270***	-0.000438
	(0.000494)	(0.000651)	(0.000594)	(0.000509)	(0.000638)
Group Dummy		-0.176***	-0.229***		
		(0.0508)	(0.0613)		
Available UI Duration × Group Dummy		0.00125	0.000970	0.00434***	-0.000396
		(0.000794)	(0.000850)	(0.000551)	(0.000678)
Unemployment Rate	0.0444**	0.0444**	0.0441**	0.0447**	0.0445**
	(0.0206)	(0.0206)	(0.0206)	(0.0205)	(0.0206)
Gross State Product	0.00000158***	0.00000159***	0.00000158***	0.00000151***	0.00000158***
	(0.000000302)	(0.000000303)	(0.000000302)	(0.000000302)	(0.000000303)
No. Observations	14994	14994	14994	14994	14994
R^2	0.518	0.518	0.518	0.519	0.518
	Panel B: Probability of Higher Educational Requirements				
Available UI Duration	0.00221***	0.00239***	0.00214**	0.00234***	0.00223***
	(0.000520)	(0.000615)	(0.000864)	(0.000541)	(0.000511)
Group Dummy		0.0387	0.0208		
		(0.0579)	(0.0637)		
Available UI Duration × Group Dummy		-0.000291	0.0000918	-0.000280	-0.000318
		(0.000865)	(0.00110)	(0.000508)	(0.000520)
Unemployment Rate	-0.0543**	-0.0543**	-0.0543**	-0.0543**	-0.0543**
	(0.0206)	(0.0206)	(0.0206)	(0.0206)	(0.0206)
Gross State Product	-0.000000413	-0.000000416	-0.000000413	-0.000000409	-0.000000413
	(0.000000537)	(0.000000537)	(0.000000537)	(0.000000533)	(0.000000537)
No. Observations	13759	13759	13759	13759	13759
R^2	0.007	0.007	0.007	0.007	0.007

Notes: All specifications include state, year, and month fixed effects, and a set of individual demographic characteristics such as age, squared age, the number of years of education, and dummies for the race, gender, and marital status. The variable "Group Dummy" changes from columns 2 to 5. In Column (2), it takes the value of 1 if an employee is male and 0 otherwise. In Column (3), it takes the value of 1 if an employee is white and 0 otherwise. In Column (4) and Column (5), we do not include any dummy variable because we include continuous variables for the age and the number of years of education. Standard errors are included in parenthesis and are clustered at the state level. The asterisks *, **, and *** reflect the 10%, 5%, and 1% significance levels, respectively.

Figure 1: UI Duration during and in the aftermath of the Great Recession

Panel A: One-Tiered EUC Program, July 2008 - March 2009

Regular UI	EUC Tier 1
Up to 26 weeks	Up to 13 weeks Up to 20 weeks from November 2008

Panel B: Two-Tiered EUC Program, March 2009 - December 2009

Regular UI	EUC Tier 1	EUC Tier 2
Up to 26 weeks	Up to 20 weeks	Up to 13 weeks

Panel C: Four-Tiered EUC Program, December 2009 – May 2012

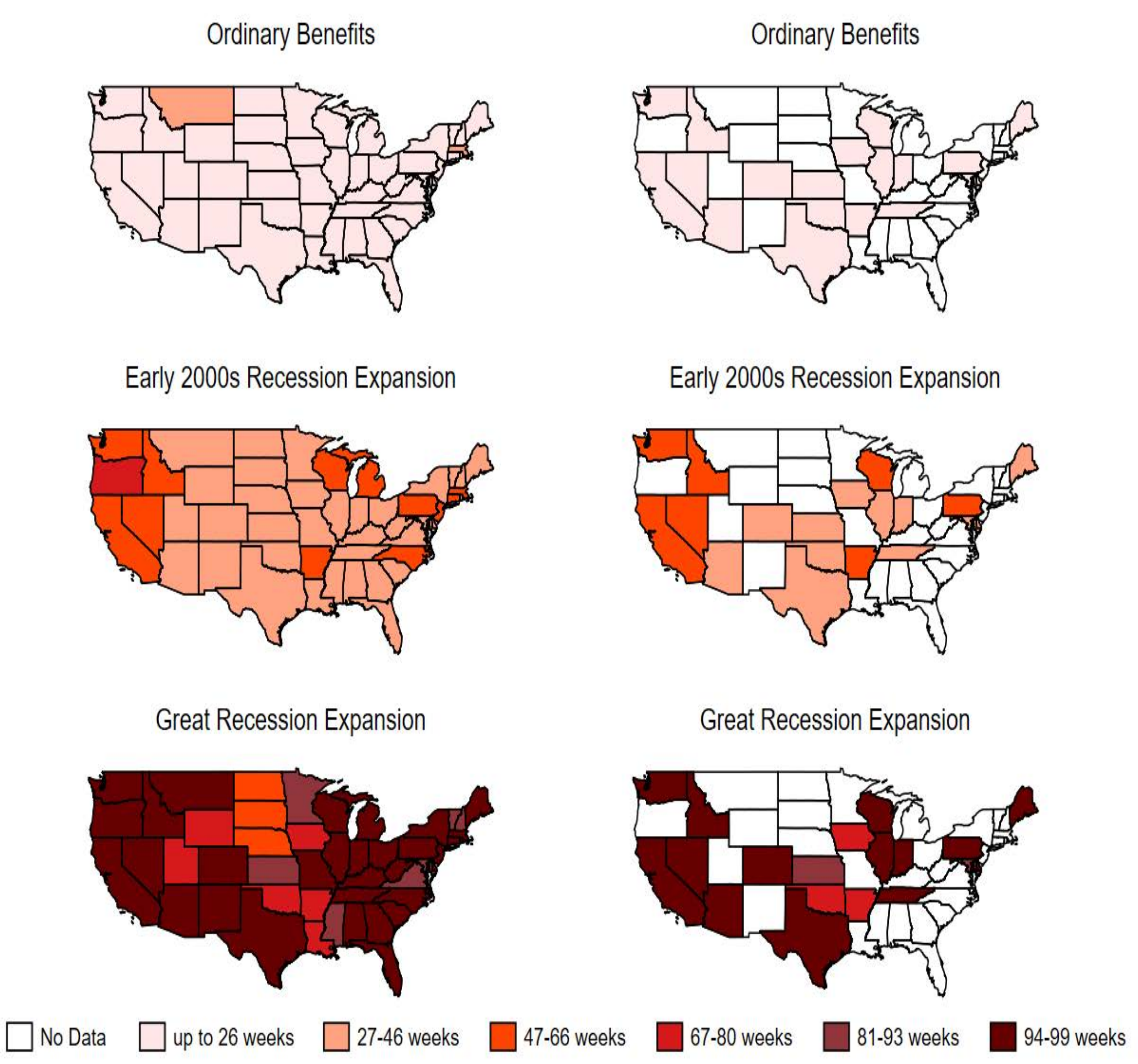
Regular UI	EUC Tier 1	EUC Tier 2	EUC Tier 3	EUC Tier 4
Up to 26 weeks	Up to 20 weeks	Up to 14 weeks	Up to 13 weeks	Up to 6 weeks

Panel D: Four-Tiered EUC Program, May 2012 – September 2012

Regular UI	EUC Tier 1	EUC Tier 2	EUC Tier 3	EUC Tier 4
Up to 26 weeks	Up to 14 weeks	Up to 14 weeks	Up to 9 weeks	Up to 10 weeks

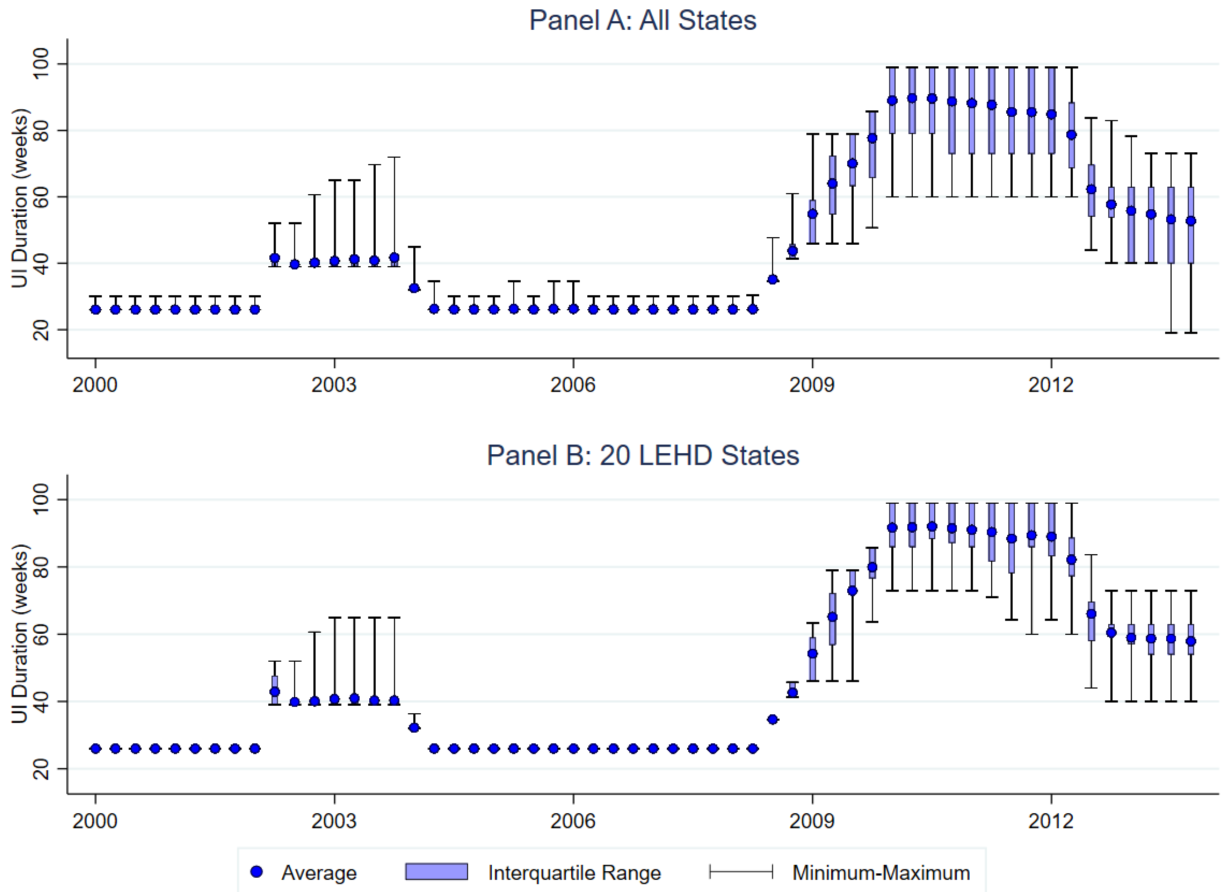
Figure 2: Maximum Weeks of UI Benefits across States and over Time

Panel A: All States **Panel B: 20 LEHD States**



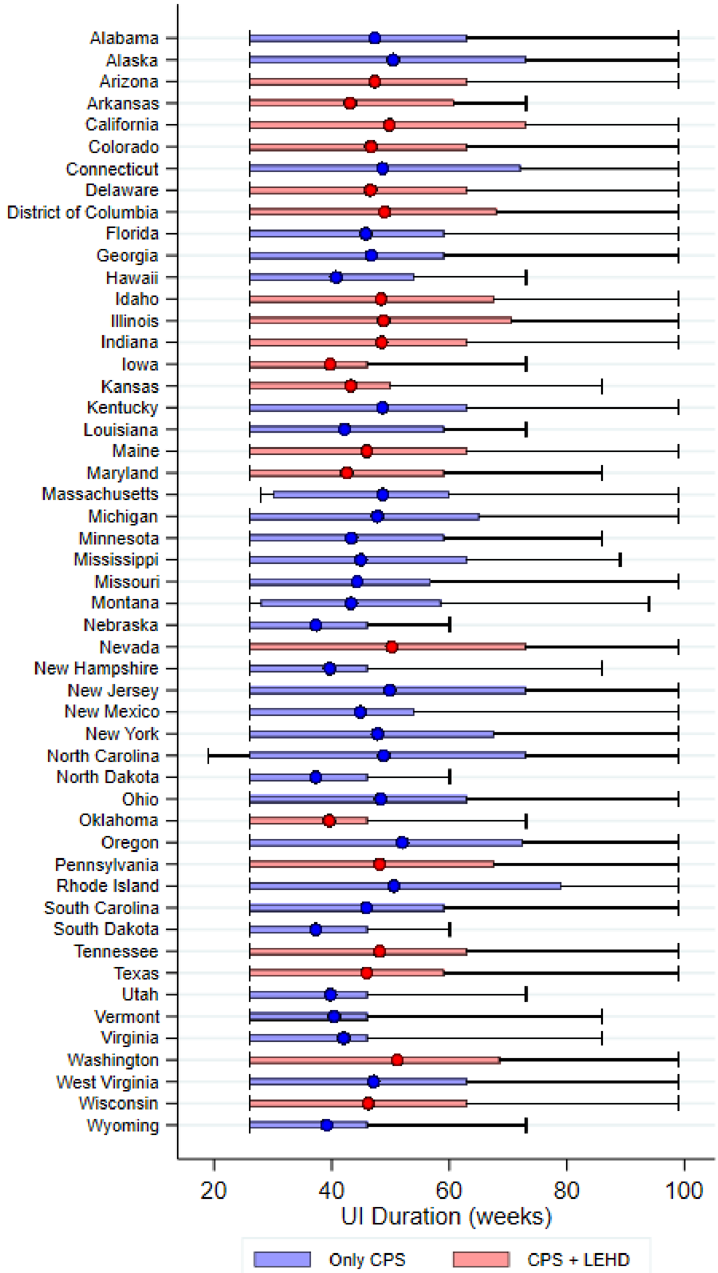
Notes: The states included in Panel B are: Arizona, Arkansas, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Iowa, Kansas, Maine, Maryland, Nevada, Oklahoma, Pennsylvania, Tennessee, Texas, Washington, and Wisconsin.

Figure 3: UI Benefits Variations within Quarters



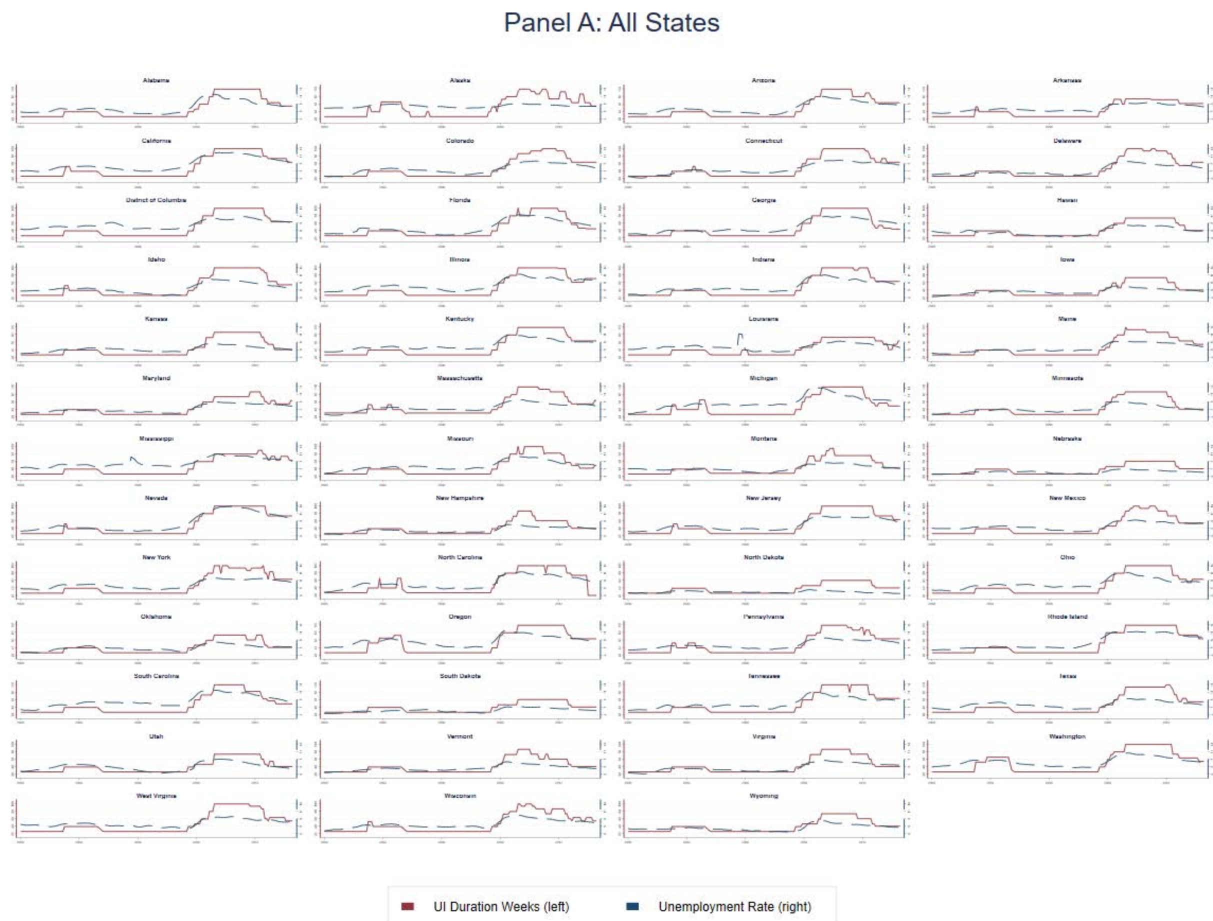
Note**:** The states included in Panel B are: Arizona, Arkansas, California, Colorado, Delaware, District of Columbia, Idaho, Illinois, Indiana, Iowa, Kansas, Maine, Maryland, Nevada, Oklahoma, Pennsylvania, Tennessee, Texas, Washington, and Wisconsin.

Figure 4: UI Benefits Variations within States

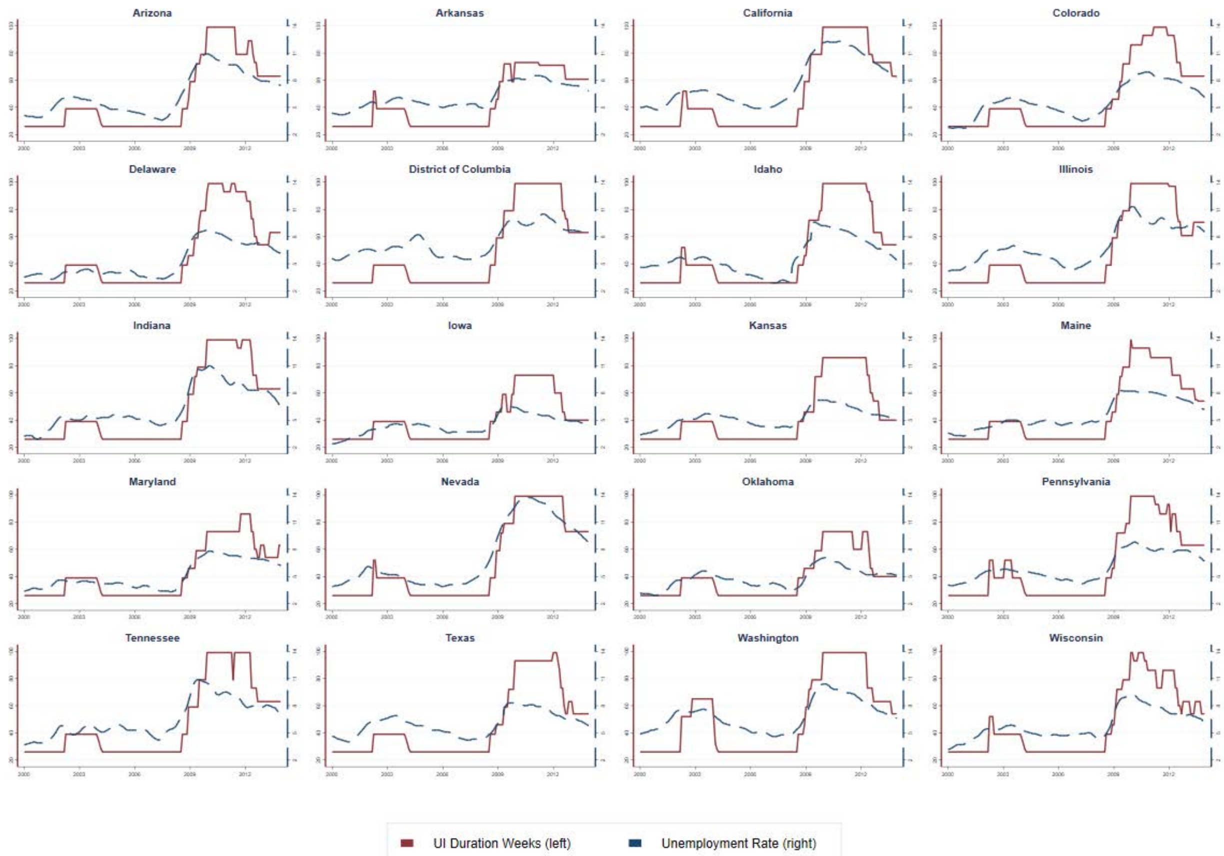


Appendix A

Figure A1: UI Benefits Duration and Unemployment Rate by States



Panel B: 20 LEHD States



Appendix B

We estimate worker and firm fixed effects via an iterative algorithm that is a modified version of the algorithm used by Crane et al. (2018). Our goal is to obtain the employee and employer fixed effects that determine earnings $\ln w_{ijsyq}$ for worker i employed at firm j in state s from year y and quarter q which are defined via the following formula:

$$\ln w_{ijsyq} = \alpha_i + \theta_j + \mu_s + \tau_y + \omega_q$$

where α_i is the worker effect, θ_j is the firm effect, μ_s is the state effect, τ_y is the year effect, and ω_q is the quarter effect. We solve for α_i, θ_j, μ_s, τ_y, and ω_q for the universe of our 20 states of matched employer-employee data. We solve using the following iterative algorithm:

1) Compute initial guess for the worker effects α_i as the average log earnings of each worker
2) Estimate the firm effects θ_j as the average by firms of the difference $\ln w_{ijsyq} - \hat{\alpha}_i$
3) Calculate the goodness of fit (R_o^2)
4) Estimate the state effects μ_s as the average by states of the difference $\ln w_{ijsyq} - \hat{\alpha}_i - \hat{\theta}_j$
5) Estimate the year effects τ_y as the average by year of the difference $\ln w_{ijsyq} - \hat{\alpha}_i - \hat{\theta}_j - \hat{\mu}_s$
6) Estimate the quarter effects ω_q as the average by quarters of the difference $\ln w_{ijsyq} - \hat{\alpha}_i - \hat{\theta}_j - \hat{\mu}_s - \hat{\tau}_y$
7) Update the worker effects α_i as the average by individuals of the difference as $\ln w_{ijsyq} - \hat{\theta}_j - \hat{\mu}_s - \hat{\tau}_y - \hat{\omega}_q$
8) Update the firm effects θ_j as the average by firms of the difference $\ln w_{ijsyq} - \hat{\alpha}_i - \hat{\mu}_s - \hat{\tau}_y - \hat{\omega}_q$
9) Recalculate the goodness of (R_n^2)
10) Check for convergence as the difference between (R_n^2 - R_o^2). If convergence is achieved, terminate the algorithm. If convergence is not achieved, define $R_o^2 = R_n^2$ and proceed back to step 4