Labor Market Trends and Unemployment Insurance Generosity During the Pandemic*

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Abstract

We test whether changes in unemployment insurance (UI) benefit generosity under the CARES Act in the US are associated with differential employment outcomes under the distinct conditions of the pandemic. While we observe a negative association between UI generosity and employment, we show that the relative employment gap arises before the Act was instituted, decreases in magnitude when the augmented benefits were in place, and does not change when the benefits expansion expires.

Keywords: Unemployment Insurance, Employment, COVID-19, CARES Act

JEL Codes: J65, J68

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1 Introduction

The Coronavirus Aid, Relief, and Economic Stimulus (CARES) Act instituted a variety of economic policy responses to the Covid-19 pandemic in the United States. One such policy was a large, temporary expansion of unemployment insurance (UI) benefits known as Federal Pandemic Unemployment Compensation (FPUC). The expansion provided a $600 weekly payment in addition to any state unemployment benefits for which a worker would have already been eligible. The payment was designed to replace 100% of the *mean* wage when combined with existing benefits. However, the extra benefit yields a total UI that is greater than weekly earnings when working for the *median* worker (Ganong *et al.*, 2020). It is natural to ask whether such high replacement rates affect employment levels *under the distinct conditions of the pandemic*.[1]

We test whether higher UI benefits are associated with differential employment trends. We use data from Homebase, a private firm that provides scheduling and time clock software to small businesses, to exploit high-frequency observations to understand how firms and workers respond to policy changes in real time. The richness of data allows us to observe week-to-week changes around the introduction and expiration of FPUC. Additionally, the longitudinal aspect of the data allows us to estimate UI benefits for each worker in our sample.

We group individuals according to their *ex-post* UI replacement rates and test whether there are differential changes in employment and hours worked across groups. We flexibly control for state-industry-week trends that account for variation in the pandemic's severity and business restrictions. We find that workers with more generous benefits did not experience differential declines in employment when FPUC was in place. While there is a negative association between replacement rates and employment, it is fully established *before* the benefits were augmented. Our study benefits from high-frequency data that allows us to isolate the timing of the introduction and expiration of FPUC from other drivers of reduced employment beginning in mid-March.[2] Workers with higher replacement rates are also no less likely to return to work, even conditioning on working at firms with observably increasing labor demand.

Our findings are consistent with the recent literature analyzing labor market effects of the pandemic.[3]

[1] Schmieder and von Wachter (2016) review the literature on moral hazard effects of unemployment insurance.

[2] Such factors include the pandemic-induced decline in labor demand, fear and concern about public health as argued by Goolsbee and Syverson (2021), and increased childcare costs.

[3] Altonji *et al.* (2020), Bartik *et al.* (2020), Cajner *et al.* (July 2020), Chetty *et al.* (2020), Fairlie *et al.* (December 2020), Gupta *et al.* (2020), and Montenovo *et al.* (2020) use various data sources – including the Homebase data – to document trends in employment and spending during the pandemic and verify that lower-wage workers, who have higher *ex-post* replacement rates, were the most affected in the early weeks of the pandemic.

Bartik *et al.* (2020) and Marinescu *et al.* (2020) use average and median state and industry-state replacement rates and do not find evidence that worse employment or job-search outcomes can be attributed to FPUC alone. We contribute to this literature by leveraging high-frequency data linking individuals and firms to measure workers' replacement rates and observe individual employment, hours of work, and re-hiring over time.

We emphasize the limitations of our study. First, our empirical strategy does not estimate the causal effect of replacement rates. We can, however, test the differential responses before, during and after FPUC. Second, the Homebase data over-represents small businesses and is concentrated in specific sectors, e.g. restaurants. Third, in order to precisely estimate workers' replacement rates we restrict our sample to workers with relatively high attachment to their jobs.

2 Institutional background

On March 19, Senate Republicans introduced an economic relief package that did not include supplemental unemployment insurance (Sullivan, 19 March 2020). Legislators agreed to include supplemental unemployment benefits on March 22 (Cochrane *et al.*, 22 March 2020). The structure of unemployment benefits continued to be contested throughout the week. The bill passed the Senate on March 25 and the House of Representatives on March 26, and was signed into law on March 27. FPUC expired on July 31 in the absence of legislation to continue federal funding for the UI expansion. Beginning in the third week of August, several states were approved to receive federal funding for a $300 UI expansion. However, since states needed to apply for funding for this expansion, an individual worker would not have known in the weeks immediately following the expiration whether she would eventually re-qualify for augmented benefits (Holzhauer, 19 October 2020).

To compute an individual's eligibility for unemployment benefits, all states use workers' earnings in the four most recent completed quarters (Ganong *et al.*, 2020). While the CARES Act expanded eligibility for UI under FPUC, several institutional features restrict eligibility. First, a worker who quits her job is ineligible for FPUC, unless she quits due to exceptional circumstances related to Covid-19. Second, once a worker receives a "suitable offer of employment" – such as an offer to return to her previous job – she is no longer eligible for UI even if she rejects the offer.

3 Data

Our dataset comes from Homebase, a private firm that provides scheduling and time clock software to small businesses, covering hundreds of thousands of workers across the U.S. and Canada. Homebase's clients are primarily small firms that require time clocks for their day-to-day operations, nearly half of which are in the food and drink industry.[4] Workers are predominantly hourly, not salaried, employees.[5]

We observe workers' daily shift data, including hours worked and wages. Each worker is linked to a firm, for which we observe state, metro area, and industry. We impose the following data restrictions: i) we keep firms that logged positive hours for at least 5 weeks between 2019-2020, ii) we keep workers who worked at least 20 hours per week in the baseline period (January 19 to February 8, 2020), and iii) we keep only workers who were employed in all quarters of 2019, with positive earnings and at least 300 hours worked in each quarter. The last requirement is the most restrictive. We impose it because we need to observe a worker's 2019 earnings history to accurately compute their unemployment benefits. We compute pre- and post-FPUC UI benefit replacement rates using the calculator developed by Ganong *et al.* (2020) using state identifiers and earnings in the four quarters of 2019. When computing workers' quarterly earnings, we floor their wages at the state minimum wage.[6] The resulting dataset has 29,005 workers. We present descriptive statistics in table 1.

We only observe workers' earnings at Homebase client firms, which imposes limitations on our findings. Since we do not observe earnings outside Homebase, this may lead to measurement error in the replacement rates which may bias against finding an effect of FPUC. We address this concern by considering only individuals with longer earnings histories and with full employment in a Homebase firm in 2019, which makes our computation of unemployment benefits more accurate and less prone to this measurement error.

[4] The basic version of Homebase is free to firms.

[5] More details on the data are discussed by Altonji *et al.* (2020) and Bartik *et al.* (2020).

[6] Many workers' posted wages are below the state minimum because they work for tips. U.S. labor law requires that employers must "top up" workers' wages to the state minimum wage.

Table 1: Descriptive Statistics

Variable	N	Mean	St. Dev.	Min	Pctl(25)	Median	Pctl(75)	Max
Panel A - Workers								
Weekly Hours in base period	29,005	36.968	8.409	20.000	31.253	36.737	41.360	100.163
Hourly Wage in base period	28,912	13.317	4.734	2.130	10.500	13.000	15.000	95.000
Weekly Earnings in base period	28,922	495.406	221.202	0.303	355.672	467.133	598.717	4,355.363
Weekly Earnings in 2019	29,005	479.495	203.115	50.914	353.512	452.581	572.102	3,989.656
Pre-FPUC Replacement Rate	29,005	0.551	0.064	0.093	0.520	0.547	0.584	0.886
Post-FPUC Replacement Rate	29,005	1.924	0.495	0.273	1.587	1.860	2.199	4.108
Panel B - post-FPUC Replacement Rate, for each quintile								
Q1	5,801	1.318	0.190	0.273	1.237	1.369	1.460	1.527
Q2	5,801	1.641	0.063	1.527	1.587	1.641	1.695	1.747
Q3	5,801	1.861	0.067	1.747	1.803	1.860	1.918	1.980
Q4	5,801	2.124	0.090	1.980	2.046	2.118	2.199	2.294
Q5	5,801	2.678	0.335	2.294	2.416	2.582	2.856	4.108

Notes: Panel A presents the descriptive statistics for the workers in our Homebase sample, which encompasses individuals who (1) worked for at least 300 hours in each quarter of 2019 (2) worked at least 20 hours in the base period, defined as the three weeks from January 19 to February 1; (3) worked at the same firm throughout 2019 and in the base period. “Pre-FPUC replacement rate” and “Post-FPUC replacement rate” indicate the ratio of UI benefits for which the worker was eligible based on their 2019 earnings to their average weekly earnings in 2019, before and after the passage of the CARES Act, respectively. Workers with pre-FPUC replacement rates of zero are excluded from our analysis. Note that the minimum hourly wage in the base period reflects the minimum wage in some states for workers who receive tips. In our analysis we floor these wages at the state non-tipped minimum to reflect the provision in U.S. labor law that if a worker does not earn the state minimum wage in wages + tips, their employer must pay them the difference. Panel B shows the descriptive statistics for the Post-FPUC Replacement Rate for each quintile.

4 Empirical approach

We explore how workers facing different UI generosity under FPUC behave before, during and after the benefits were in place. We look into workers' weekly employment status and hours worked over time depending on their *ex-post* replacement rate.[7] We estimate the following specification:

$$Y_{ijst} = \alpha_0 + \sum_{\tau=0}^{T} \sum_{g=2}^{5} \beta_t^g R_{ijs}^g \mathbb{1}\{t = \tau\} + \eta_{jst} + \varepsilon_{ijst}, \tag{1}$$

where Y_{ijst} is the outcome for worker i associated with industry j in state s during week t. R_{ijs}^g is an indicator that worker i's with FPUC replacement rate places her in replacement rate quintile g. η_{jst} is a state-industry-week fixed effect, which subsumes all state-industry weekly variation, including the severity of the pandemic in each state and states' restrictions on business activities. Standard errors are clustered at the worker level.

This strategy allows us to test for differential labor market dynamics for individuals with different replacement rate levels, before, during and after the FPUC. Our strategy does not estimate a causal effect of the replacement rate on labor outcomes, as we do not estimate a counterfactual path of labor outcomes without the benefits expansion. Instead, we explore differences in the treatment intensity when the augmented benefits were in place.

The most important assumptions in our analysis are: i) that individuals did not anticipate the approval of FPUC and therefore did not time their labor market responses in advance, and (ii) there are no other factors that correlate with replacement rates, conditional on state-industry-week and are simultaneous to FPUC. We argue that anticipation before FPUC is unlikely for two reasons. First, the timeline of negotiations indicates that the $600 additional benefit was not agreed upon until at least March 24.[8] Second, workers will face any increased incentive to exit only after they are able to receive extra benefits. While workers may have anticipated the expiration of benefits, we would expect to observe at least one of two phenomena if benefits depressed labor supply: either at least some workers would time their return to work to coincide with the expiration of benefits, or we would observe a steady increase in relative employment in the weeks prior to

[7] The *ex-post* replacement rate R_{ijs} for individual i working in industry j in state s is determined by her pre-FPUC UI benefits (UI_{ijs}^{Pre}), the additional $600, and her weekly earnings in a chosen reference period (w_{ijs}). In our baseline specification we choose w_{ijs} to be her average weekly wage in 2019. Formally:

$$R_{ijs} = \frac{UI_{ijs}^{\text{Pre}} + 600}{w_{ijs}}.$$

[8] Even if workers stopped working the following day, they would be coded as employed in the entire pre-period.

expiration. We observe neither of these patterns. Moreover, once our exercise restricting the analysis to re-hired workers mitigates the concern because such workers have particularly low search frictions. In order to violate ii) a factor needs to correlate with replacement rates, within state-industry, and have a differential effect before and after the events on March 27 and July 31.

5 Results

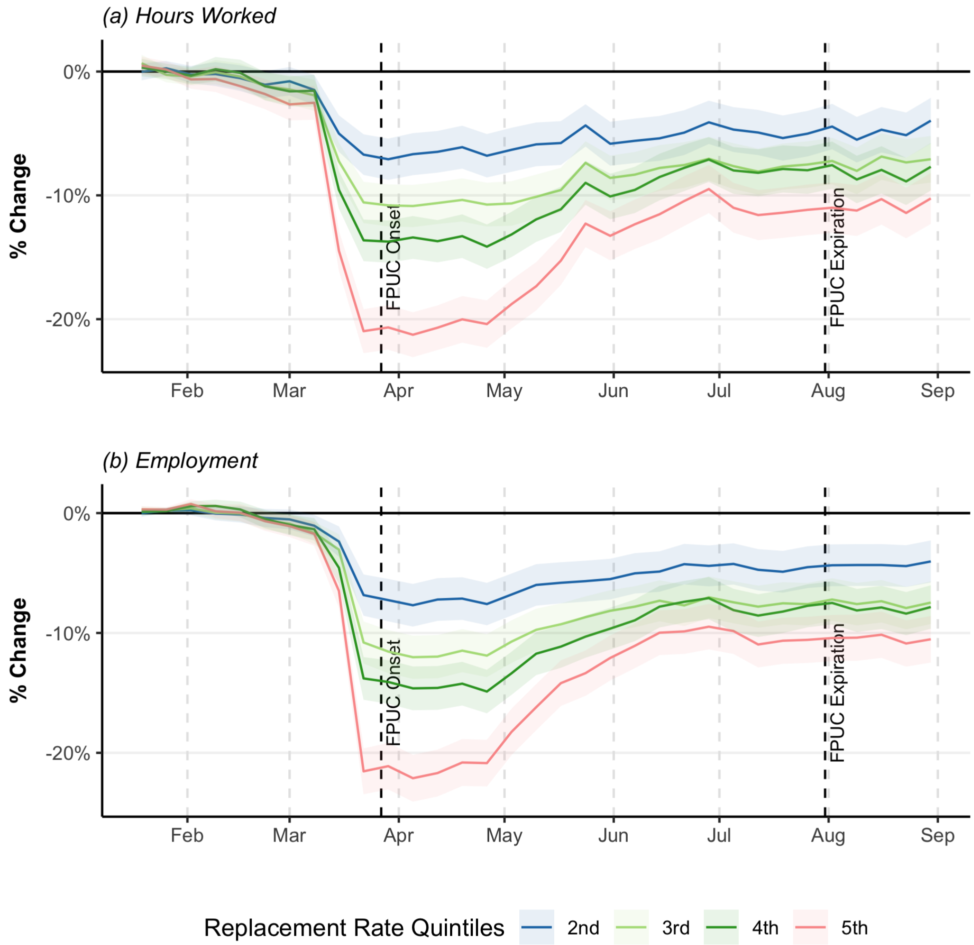
Figure 1 plots the β_t^g coefficients on the interaction between the week dummies t and the replacement rate quintiles g in equation 1. The coefficients represent the percentage change in hours (panel a) and probability of employment (panel b) relative to the first replacement rate quintile in a given week.

While the workers with the highest replacement rates experience the largest declines in employment relative to the January baseline, the differential decline occurs entirely in the weeks prior to the passage of FPUC. Furthermore, the figure suggests that workers with larger increases in benefit generosity are no slower to return to work than others with more modest UI increases. Even if many states experienced implementation delays of several weeks, we could expect at least some drop in the first full week and a significant drop relative to the pre-period once all states had implemented the expanded UI benefits, when controlling for state-industry-week effects. We observe no such pattern.

Additionally if workers who had experienced larger increases in benefit generosity were choosing not to return to work because of the benefits expansion, we would expect to see a differential increase in employment after the benefits expired at the end of July. However, there appears to be no differential change in employment around the expiration for workers with more generous benefits under FPUC. We should interpret this result with caution. It is plausible that the ability to return to one's job decreased from March to July – either because firms that did not rehire workers shut down, or because firms looking to rehire workers filled jobs with new workers who would not appear in our sample. Thus, frictions may have prevented an immediate return-to-work response to the expiration even if workers chose not to return to work due to the benefits expansion. We address this concern with evidence from firms with observably increasing labor demand.

Since many firms have ceased operating entirely during the pandemic, our null result could be driven by the total lack of labor demand. We aim to address this concern by testing whether firms that have *increasing* labor demand experience difficulty in rehiring workers. This allows us to exclude the effect of depressed

Figure 1: Changes in hours and employment, by replacement rate quintile



Notes: These figures show the specification from equation 1 showing the estimated β_t^g coefficients for each quintile of post-FPUC replacement rate. The omitted category is the first quintile — i.e., those with lowest replacement rates. The regression was estimated in the weekly data and the specification includes state-industry-week fixed effects. The outcomes are weekly hours worked compared to the baseline (January 19 to February 8) and probability of employment, where individuals were coded as being employed (employment = 1) if they worked any positive hours in the week. The vertical line indicates the day CARES act was passed (March 27) and the date of FPUC expiration (July 31). Standard errors were estimated using cluster at the worker level. The shaded areas represent 95% confidence intervals.

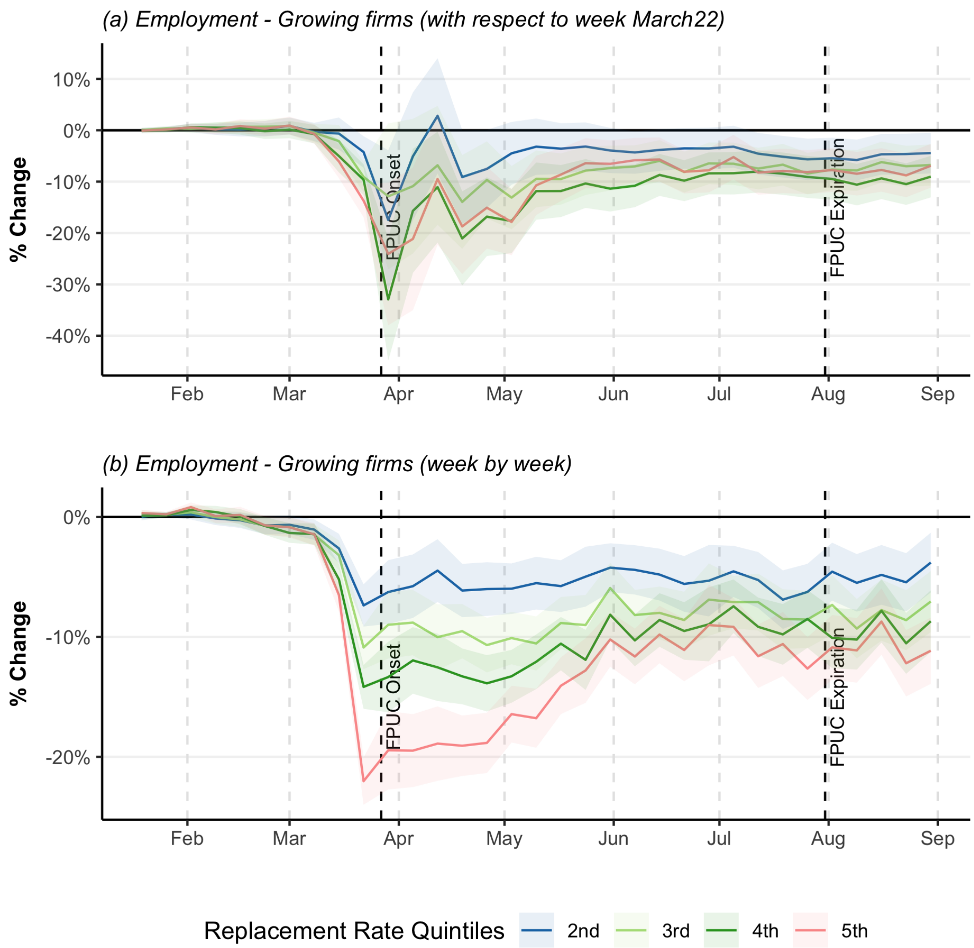
labor demand across the economy. We test whether workers at firms that are growing have differentially lower probabilities of employment if they have higher UI replacement rates. We define a firm as "growing" based on a leave-out measure of growth in hours worked.[9]

[9] Specifically, for worker i at firm j in week t, we define j to be growing if the number of hours worked by workers other than i is higher in week t than in a chosen reference week t^*. The hours growth rate, HG_{ijt}, is given by

$$HG_{ijt} = \frac{\sum_{k \neq i} h_{kj,t}}{\sum_{k \neq i} h_{kj,t^*}} - 1. \tag{2}$$

Figure 2 shows the results for the sampling of growing firms where the reference period is fixed at the week before the CARES Act (panel a) or on a rolling basis, comparing each week to the previous week (panel b). For both definitions, individuals with higher replacement rates do not have worse employment outcomes that could be fully attributable to FPUC.

Figure 2: Changes in hours and employment, by replacement rate quintile — Subset of growing firms



Notes: The figure shows the specification from equation 1 showing the estimated β_t^g coefficients for each quintile of the replacement rate ratio (where the omitted category is the first quintile — i.e., those with lowest replacement rates). The regression was estimated in the weekly data and the specification includes state-industry-week fixed effects. The top figure estimates in the sub-sample of firms that were growing (leaving individual i's hours) with respect to the week of March 22th. The bottom figure estimates in the subsample of firms that were growing after the CARES act, in a week by week comparison. The outcome is employment probability, where individuals were coded as being employed (employment = 1) if they worked any positive hours in the week. The vertical line indicates the day CARES act was signed (March 27th) and the FPUC expiration (July 31). Standard errors were estimated using cluster at the worker level and the shaded areas represent 95% confidence intervals.

6 Conclusion

We test whether individuals with ex-post higher UI replacement rates experienced negative labor market effects while FPUC was in place. While we do find this negative association, we show that it is explained by changes that occur before the benefit expansion — indicating that it was driven by the pandemic itself and not by the policy response to it. Furthermore, individuals with higher replacement rates under FPUC were no more likely to return to work after benefits expired.

Our results do not speak to the disemployment effects of UI generosity during more normal times. The severity of the decline in labor demand and the health risks to workers make the current pandemic different. We conjecture that the unique conditions of the pandemic may explain this null result. First, there has been a broad-based decline in labor demand. Second, several other factors may also drive workers to choose to stay home: fear, risk, and generalized concern about the pandemic could play a large role, as argued by Goolsbee and Syverson (2021), and childcare costs have risen substantially. Third, for many workers health insurance and other benefits are tied to their jobs. Fourth, the extra benefit is limited in duration. These factors limit the perceived value of the expanded UI, and increase the value of having a job (Boar and Mongey (2020), Petrosky-Nadeau (2020)).

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