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May 2024

Working Paper 2021-13

https://doi.org/10.24148/wp2021-13

Suggested citation:

Petrosky-Nadeau, Nicolas and Robert G. Valletta. 2024. "UI Generosity and Job Acceptance: Effects of the 2020 CARES Act," Federal Reserve Bank of San Francisco Working Paper 2021-13. https://doi.org/10.24148/wp2021-13

UI Generosity and Job Acceptance: Effects of the 2020 CARES Act

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April 12, 2024*

Abstract

We assess labor market effects of the CARES Act $600 weekly UI supplement. We analyze labor force transitions using monthly CPS microdata and imputed UI benefits. The results show moderate disincentive effects of the supplement on job finding. We rationalize this result in a dynamic model of job acceptance decisions that yields a reservation level of UI benefits at which a recipient is indifferent between unemployment and employment at their prior wage. Calculations based on the model are consistent with the empirical analysis in regard to the moderate fraction of UI recipients who were likely to reject job offers.

JEL Classification: J64, J65.
Keywords: Unemployment, unemployment insurance, job acceptance, COVID-19, CARES Act.

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1 Introduction

The Coronavirus Aid, Relief, and Economic Security (CARES) Act, through the Pandemic Unemployment Compensation (PUC) provision, provided an additional $600 per week to supplement regular unemployment insurance (UI) benefits. The supplement was available during the initial outbreak of COVID-19 from late March though the end of July 2020. This historically unprecedented increase in the level of UI benefit payments meant that most UI recipients received more weekly income via UI payments than they earned on their prior jobs–i.e., their UI replacement rates exceeded 100 percent (Ganong, Noel and Vavra 2020). The enhanced benefits prompted concerns that the labor market recovery from the pandemic would be delayed as many UI recipients rejected offers to return to work, reflecting the standard moral hazard effect of UI benefits on job search (Feldstein 1976, Baily 1978, Chetty 2008).

We assess the disincentive effects of expanded pandemic-era UI benefits on job search and acceptance via two complementary approaches: (i) empirical analyses of observed labor force transitions; (ii) quantitative assessment of a dynamic model of the job acceptance decision. Despite the large increase in benefits, results from both approaches suggest only moderate disincentive effects of the $600 weekly UI supplement on job search and acceptance decisions.

We first conduct direct empirical analyses of labor force transitions using monthly data from the Bureau of Labor Statistics' (BLS) Current Population Survey (CPS) combined with imputed UI benefits. Our direct empirical tests are based on a difference-in-differences regression framework. We use it to assess whether the change in job-finding rates and other labor market transitions between the pre-CARES and CARES periods is larger for individuals who have higher UI replacement rates as a result of the supplemental payments. Our value added relative to prior analyses of the potential disincentive effects of the CARES Act supplemental payments arises from two specific features of our analyses: (i) we exploit individual variation in UI replacement rates rather than geographic or solely temporal variation; (ii) we directly assess the individual labor market transitions, in particular job-finding rates (exits from unemployment to employment), that may be affected by the moral hazard effect of UI benefit generosity.

Our regression analyses rely on labor market transition data formed using data on individuals matched across consecutive monthly CPS files. We use data for early- to mid-2020 only, to focus on the impact of the extra $600/week of UI payments specified by the CARES Act. We estimate UI replacement rates for individuals in our sample by applying the calculator developed by Ganong, Noel and Vavra (2020) to annual earnings data from the CPS Annual Social and Economic Supplement (ASEC).

Our results show moderate disincentive effects on job finding from this very large increase in UI replacement rates: for the typical UI recipient, our estimates imply a 22 percent reduction in job-finding rates due to the $600/week supplement. As we discuss in more detail in Section 2.3.3, the resulting estimated elasticities of search duration with respect to UI replacement rates are toward the low end of the range based on earlier research using U.S. data (Schmieder and von Wachter 2016).

Our results also are broadly comparable to the findings from two other recent papers that examine the effects of the pandemic-period expansions using administrative micro-data from financial services companies (Coombs et al. 2022, Ganong et al. 2022); using similar identification strategies, our estimated elasticity is essentially identical to that reported by Ganong et al. (2022). By contrast, some other recent studies found little or no effect of the pandemic UI enhancements on labor market outcomes. Altonji et al. (2020), Bartik et al. (2020), and Finamor and Scott (2021) found that states with more generous UI systems did not experience weaker labor market rebounds during the initial phase of economic recovery from the pandemic. As we discuss in Section 2.3.3, it is likely that our somewhat larger estimates of the UI benefit effects on job search arise because we focus on search responses at the individual level, commonly referred to as micro effects. By contrast, the studies that find smaller effects likely combine micro responses with offsetting macro effects on aggregate outcomes. The latter include the aggregate stimulus effects of UI payments, which help sustain labor demand and hence may offset the job search disincentives for individual UI recipients (Boone et al. 2021, Kekre 2023). We discuss these issues further in Section 2.3.3.

To complement and reinforce the regression estimates, our second approach relies on the development of a dynamic model of job acceptance decisions that includes heterogeneity in the underlying wage distribution for job seekers. We use this model to derive the level of benefits necessary for workers to be indifferent between accepting a job offer at their previous wage and rejecting it to remain unemployed, taking into account the remaining number of weeks of unemployment benefits available to them. We call this the *reservation benefit*: a job offer at the previous wage is accepted if the current level of benefits is below this level. For a given job offer, the level of the reservation benefit is determined by: (i) the expected duration of the employment spell for an accepted job – longer lasting jobs have a greater value and are rejected only for commensurately more generous unemployment insurance payments; (ii) the rate of arrival of new job offers – in a depressed labor market, when job offers are few and far between, any job offer is costly to refuse, raising the reservation benefit amount, and; (iii) the duration of benefits remaining – an additional week of benefits raises the opportunity cost of accepting an offer and lowers the reservation benefit level. In the limit of unbounded UI duration the reservation benefit converges to the prior wage. Conversely, with one week remaining of UI payments, the reservation benefit is always above the prior wage, implying that many UI recipients will accept a job even if their benefits exceed the offered wage.

We apply the reservation benefit concept to the period covered by the provisions in the CARES Act, including the extension of benefit payments for up to 52 weeks with the Pandemic Emergency Unemployment Compensation (PEUC) and state emergency extensions. We derive the probability that individuals within a group of job seekers would reject an offer to return to work. As in the regression analysis, we use data from the CPS ASEC to impute reservation benefit levels for workers in different skill (education) groups and occupations. Our quantitative analysis suggests that a moderate fraction of UI recipients, approximately 22 percent in our preferred specification, would refuse an offer to return to work at their previous pay. This proportion of rejected job offers is essentially identical to the magnitude of the disincentive effects on job finding rates found in the analysis of individuals' transition rates out of unemployment (the 22 percent reduction in job-finding rates noted above).

These findings imply that the value of a sustained job, especially in a depressed labor market, usually outweighs the value of the temporary additional UI income. By way of example, a typical (median) worker earning about $717 per week in their previous job received $959 per week in UI payments under the CARES act. We calculate that an offer received with 8 weeks of CARES UI payments remaining would be accepted as long as this worker's current benefit payment was below $1419, a reservation benefit level that is around twice the previous wage. Further down the earnings distribution, however, the value of UI benefits under the CARES Act can dominate the value of a job at the prior wage, thus exceeding the reservation benefit level for these job seekers. While our model suggests that only a moderate share of the overall population of job seekers receiving the $600 UI supplement would reject an offer to return to work during this period, rejection rates vary substantially across levels of education and broad occupations.

Our findings are broadly consistent with prior research on the effects of UI enhancements during past recessions. Most notably, our finding of limited disincentive effects of enhanced UI generosity during the pandemic recession is consistent with other work that finds substantial cyclicality in such effects, with little to no impact when labor market conditions are weak (Kroft and Notowidigdo, 2016). Similarly, analyses of the impact of the historically large increase in potential UI benefit duration during the Great Recession found small effects on unemployment exit rates, with the main impact instead being an increase in labor force attachment. (Rothstein 2011, Farber and Valletta 2015, Chodorow-Reich, Coglianese and Karabarbounis 2019).[1]

The framework used to derive the reservation benefit statistic is similar to Mortensen (1977) in incorporating the realistic feature of UI benefits that are limited in duration. It is also broadly related to the concept of the after-tax reservation wage in Shimer and Werning (2007), which represents the take home pay required to make a worker indifferent between working and remaining unemployed.[2] Finally, Boar and Mongey (2020) derive a quantitative framework along similar lines to our reservation benefits analysis and also find a likely limited impact of temporarily increased UI payments on job acceptance decisions during the pandemic.[3] Our analysis does not address the optimality or welfare effects of the supplemental income under the CARES Act. This

[1] See Moffitt (1985) for an early empirical study of the effect of UI benefits on unemployment durations. Lalive, Landais and Zweimüller (2015) use Austrian data and find that search disincentive effects of UI benefit extensions are offset somewhat by improved search outcomes for individuals who are not eligible for the extensions. A related question not addressed here is the impact of UI provisions on the joint behavior of workers and firms, and in particular on the duration of employment spells (see, for instance, Feldstein 1976 and Baker and Rea 1998).

[2] Berg (1990) extends Mortensen's analysis to a non-stationary environment to study the dynamic evolution of a worker's reservation wage as economic conditions evolve (exogenously). Contrary to Mortensen's reservation wage our reservation benefit and the reservation wage in Shimer and Werning (2007) are not choice variables affecting the arrival rate of job offers.

[3] Marinescu and Skandalis (2021), using French administrative data, find evidence of declining reservation wages (measured as a desired target wage) as exhaustion of UI benefit payments nears.

is the focus of Mitman and Rabinovich (2021), who argue the $600 supplemental income approximated an optimal UI benefit given the large and transitory nature of the COVID-19 shock to the labor market.

The rest of this paper is organized as follows. Section 2 describes the empirical design and provides the results from regression analyses of labor market transitions using matched monthly CPS data. Section 3 develops a modeling framework that rationalizes our empirical results. Section 3.1 describes the decision problem and derives a reservation benefit as a function of the state of the labor market, the wage offer, and the number of weeks of UI payments remaining. Section 3.2 adapts the reservation benefit statistic to the details of the CARES Act and uses CPS data to calculate benefit amounts for different categories of workers. Section 4 concludes.

2 CARES Act UI expansion and labor market transitions

We begin with a direct empirical assessment of the effects of the increase in UI benefit payments during the pandemic on job-finding rates and other labor market flows. We rely on a before/after regression framework to assess whether the change in job-finding rates and other labor market transitions between the pre-CARES and CARES periods is larger for individuals who receive the largest UI replacement rates due to the supplemental payments. As described in Section 1, our value added relative to prior analyses of the pandemic UI supplements arises from our reliance on individual variation in UI replacement rates and direct measurement of job-finding rates.[4]

Our regression analyses rely on labor market transition data formed using data on individuals matched across consecutive monthly CPS files. We use data for early- to mid-2020 only, to focus on the impact of the extra $600/week of UI payments specified by the CARES Act and available from late March through the end of July. We combine the monthly CPS data with estimated UI replacement rates formed using the calculator developed by Ganong, Noel and Vavra (2020). Our specific calculations rely on annual earnings data from the CPSASEC for the individuals observed in our matched monthly CPS data. We discuss these steps in detail in the next two sub-sections, including a discussion of the distribution of replacement rates in our sample.

[4] By comparison, Bartik et al. (2020) rely on state-level variation in median replacement rates and employment/hours, and Altonji et al. (2020) and Finamor and Scott (2021) examine labor market status but not flows between labor market states. These papers reported little or no disincentive effects of the enhanced UI payment generosity on employment status. In research conducted in parallel with ours, Coombs et al. (2022) and Ganong et al. (2022) used administrative data from financial services companies and found effects of the CARES Act UI benefit increases on job finding that are similar to or smaller than our estimates. We discuss the comparison of our results to these studies in more detail in Section 2.3.3. In addition, Marinescu, Skandalis and Zhao (2021) examined job applications in local labor markets during the pandemic and found moderate reductions in application rates in areas with greater increases in UI benefit amounts.

2.1 Matched CPS data on labor market flows

We use matched monthly data on individual labor force participants from the CPS (age 16 and over).[5] Because our empirical strategy requires linking monthly CPS files to annual earnings data from the CPS ASEC (see the next sub-section), our matched observations are limited to the months of February through July of 2020. This timeframe is narrow but enables us to focus on the period when the $600 supplement was in place (and the preceding two months of 2020, which are used as a pre-treatment comparison period).[6]

Due to the rotating sampling scheme used for the CPS, surveyed households and individuals are in the sample for two separate periods of 4 consecutive months (with an intervening 8-month period spent out of the sample). This enables consecutive month-to-month matching for about 70 percent of the sample.[7] The monthly match is based on household identifiers, which we validate by ensuring that the reported data on age, education, race, and gender do not conflict across matched observations. We identify labor market transitions by comparing an individual's labor force status in consecutive months. We focus primarily on transitions out of unemployment (U), to employment (E) or out of the labor force (N), denoting them as UE and UN transitions respectively. Given relaxed job search requirements under the CARES Act UI expansions, the behavior of individuals not actively searching for work may have been affected by the UI supplements. We therefore also examine transitions from out of the labor force to employment (NE).

A well-known concern regarding matched CPS data is the likelihood of spurious transitions in labor force status arising from inconsistent or error-ridden survey responses rather than meaningful changes (Abowd and Zellner 1985, Poterba and Summers 1986, 1995). Such spurious transitions could impart a downward bias to the estimated effects of UI payments on labor force transitions and reduce the precision of the estimates. We therefore follow past research by adjusting the data to minimize the incidence of spurious transitions (Rothstein 2011, Valletta 2014, Farber and Valletta 2015, Farber, Rothstein and Valletta 2015). In particular, for individuals identified as leaving unemployment one month, either through job finding or labor force exit, and then returning to unemployment the next month, we recode their records to show no transition (and retain the newly created observations). We refer to these as "two-month matches," although the resulting transitions are still measured on a consecutive monthly basis. This adjustment requires restriction of the final analysis sample to individuals who are observed to be in their first or second month of a consecutive four-month span in the sample, thereby reducing the matched sample count by approximately one-third and eliminating July 2020 observations from our analyses.[8]

[5] See Valletta (2014) for more details on construction of a similar sample for an earlier timeframe (in particular, Table 2 and the associated discussion in that paper). We exclude individuals who identify as serving in the armed forces.

[6] The ASEC is administered primarily in March, although some CPS respondents receive the supplement in other months. With the 4-month rotation in the monthly CPS, this enables us to use observations with ASEC information for the months of January through July of 2020. Forward matching yields observations on labor force transitions starting in February 2020.

[7] Most of the non-matched observations are from the "outgoing rotation groups" that are exiting the sample for eight months or permanently (one quarter of each monthly sample). In addition, a modest fraction of observations is lost because respondent households that move to different geographic locations are not followed.

[8] The adjustment reduces the monthly incidence of transitions out of unemployment by about 5 percentage points

The results for unemployment exits reported in subsequent sections generally are based on these adjusted transitions, although we also provide some comparison to specifications that do not make this adjustment. We do not apply this adjustment to our analysis of transitions from out of labor force to employed (NE), because the measurement distortion generally applies to transitions in and out of unemployment. As we will see in the results below, the adjustment for unemployment exits makes a substantial difference for the key results. This likely reflects the turbulence in labor market transitions and measurement during the time period used for our estimates, which corresponds to the early phase of the pandemic. Adjusting for short-term or spurious transitions is especially important in such circumstances.

2.2 UI replacement rates from CPS ASEC data

Our analysis relies on UI replacement rates calculated at the individual level, defined as the ratio of weekly UI payments to weekly earnings prior to the job loss that resulted in the UI claim. As discussed in Ganong, Noel and Vavra (2020), median UI replacement rates across all eligible workers typically are slightly below 0.5 in the United States (50 percent of prior earnings), absent benefit supplements. They estimated that the $600 CARES Act UI supplement raised the typical replacement rate substantially, to a median value of 1.34, implying that the majority of UI recipients were eligible for UI payments that exceeded their prior weekly earnings. As part of their research, Ganong et al. constructed a calculator for replacement rates based on individuals' recent prior earnings history, which they have made publicly available.[9]

We use the Ganong et al. calculator to form estimated UI replacement rates for the individuals in our data. Precise measurement requires individual employment and earnings data from prior quarters. We therefore restrict our matched monthly CPS sample to individuals who are included in the 2020 CPS ASEC sample. As noted above, this limits the sample to the months of January through July 2020. The ASEC includes information on weeks worked, hours, and earnings in the prior calendar year (2019 in this case, which largely contains the qualifying earnings period for potential UI recipients in our sample from early 2020).[10] Because no information is provided on the timing of employment and earnings across the four quarters of the year, we spread them out evenly across all four quarters for the purposes of applying the UI benefits calculator.[11]

One notable feature of the distribution of UI replacement rates is that because normal UI payments generally are determined as a fraction of prior earnings, the uniform $600 supplement increased replacement rates more for individuals with low versus high prior earnings. To illustrate this, we divided our sample of unemployed individuals during the months when the $600 supplement was available (April-July) into quintiles based on weekly earnings. The median replacement

on average (Valletta 2014).

[9] https://github.com/PSLmodels/ui_calculator

[10] Ganong et al. used pre-pandemic labor market data and 2018 as their base earnings year. Our use of observations on actual unemployed individuals in early 2020 combined with their 2019 earnings should yield relatively accurate measurement of UI replacement rates in our sample.

[11] The rules specifying which prior earnings quarters are used to determine UI eligibility and weekly payments vary across states.

rate ranged from about 2.5 in the lowest quintile down to about 0.6 in the highest quintile. This will also be reflected in the distribution of replacement rates across industries and occupations, given variation in typical skill levels and hence earnings across sectors.

This variation in UI replacement rates due to the $600 supplement raises potential concern about the identifying information that we use to estimate their effects on job-search behavior. In particular, job losses early in the pandemic were heavily weighted toward low-wage sectors and workers, notably individuals in high-contact services jobs concentrated in the retail, leisure and hospitality, and personal services sectors. Activity in these sectors remained disrupted well into the pandemic, curtailing job prospects for individuals laid off from these sectors. This correspondence between UI replacement rates and sectoral disruptions during the early pandemic period raises potential concern that our estimates of UI effects may be contaminated by sector-specific labor market conditions.

In our subsequent empirical analysis, we address this concern about possibly confounding effects from the sectoral pattern of labor market disruptions during the early pandemic period. To preview, Figure 1 illustrates the distribution of replacement rates across our complete sample of unemployed individuals and also within major industries and occupations (calculated for the months of April-July 2020, including the $600 supplement). The median replacement rate for the full sample is 1.40, slightly higher than the estimate of 1.34 from Ganong, Noel and Vavra (2020).[12] The figures show that median and mean replacement rates vary notably across industries and occupations, with higher replacement rates evident in low-wage sectors such as the leisure and hospitality industry and also for services occupations. However, replacement rates vary more within than between sectors, as reflected in standard deviation spreads within sectors that generally extend to or beyond the range of means and medians across sectors. This suggests that variation in replacement rates is not closely related to sector-specific effects of the pandemic. We explore this issue further when we present our regression results in the next section.

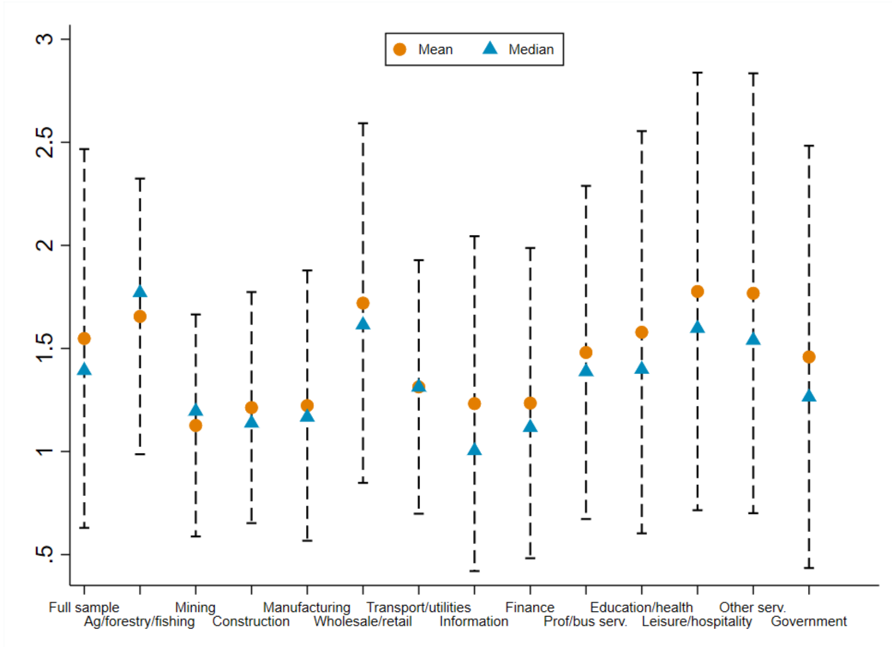
2.3 Regression specification and results

We begin our analyses with a conventional before/after regression design, essentially a differences-in-differences analysis that yields an estimate of the average treatment effect of variation in UI replacement rates. We include imputed UI replacement rates using the procedure described in the preceding section. Our specific regressions take the form:

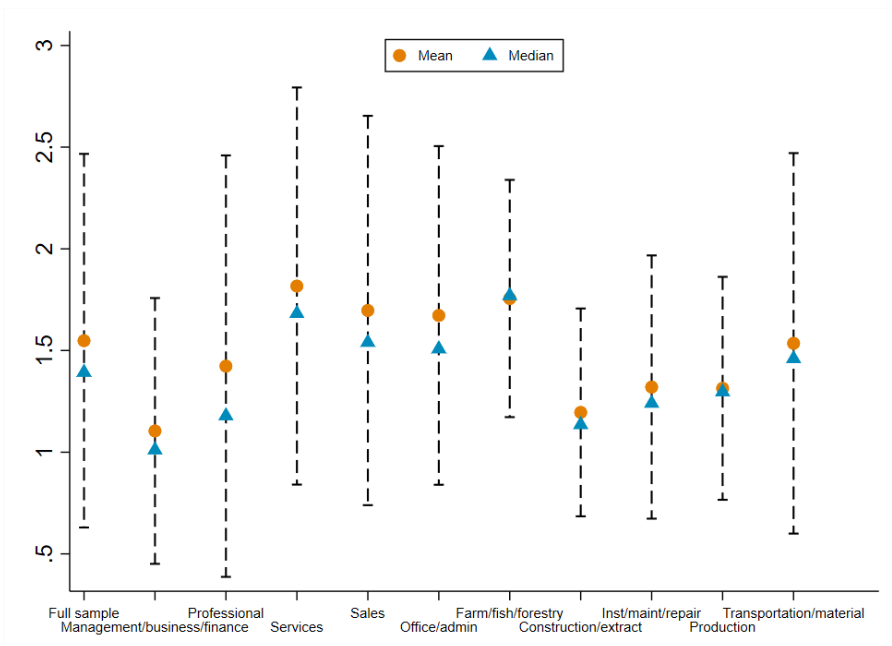
$$Pr(Y_{it} = 1) \quad = \quad \delta R_i + (\pi \times R_i \times (Apr - July)) + \gamma_t + \phi_s + \beta X_{i,t-1} + \lambda Z_{s,t-1} \tag{1}$$

In this equation, the dependent variable Y_{it} is an indicator for whether an individual i tran-

[12] Our higher estimated replacement rates are as expected: Ganong, Noel and Vavra (2020) noted that replacement rates estimated from individuals unemployed during the early pandemic period are likely to exceed their estimates based on pre-pandemic data, given the tilt toward low-wage individuals among job losers during this period. Our estimates are very similar between the full sample of unemployed individuals and our restricted two-month match sample.



(a) UI replacement rate, by major industry



(b) UI replacement rate, by major occupation

Figure 1: UI replacement rates with CARES Act $600 supplement
Notes: Calculated from authors' CPS monthly-ASEC match, using the UI benefits calculator from Ganong, Noel and Vavra (2020). Vertical bars extend one standard deviation above and below the mean.

sitions between the specified labor market states across consecutive months (observed in month t, based on status in months t and $t-1$). We focus primarily on job finding rates from unemployment (UE transitions) but also examine transitions that involve labor force exit or entry (UN and NE transitions), as discussed more below. Our preferred estimates rely on our two-month matched CPS data for the reasons described in Section 2.1, although we also examine results from the single-month matched sample. The underlying sample contains observations for transitions observed in the months of February through July of 2020, although the two-month match estimation samples end in June.

The key explanatory variables are the individual's imputed UI replacement rate (R_i) under the CARES Act and its interaction with an indicator for observations corresponding to the months of April through July, when the CARES Act \$600 supplement was available (with estimated coefficients δ and π). The replacement rate with the \$600 supplement included varies across individuals but not over time and hence is not the key source of variation in this equation. Instead, the treatment effect of the \$600 CARES supplement is captured by the impact of the replacement rate after the CARES Act was implemented and the supplemental payments were available. This period began in late March 2020, between the March and April CPS reference periods. These effects are estimated by the coefficient on the interaction between the replacement rate R_i and an indicator for observations in the months of April through July (or June for the two-month matched sample). This represents a conventional before/after estimation approach with regression controls (with the months of February and March combined used as the baseline control period).

We also estimate an expanded equation that allows the effects of the CARES Act supplement to vary across the months during which it was active:

$$\begin{aligned} Pr(Y_{it}=1) &= \delta_1 R_i + (\delta_2 \times R_i \times Mar) + (\pi_{1,2,3,4} \times R_i \times (Apr, May, Jun, Jul)) \\ &\quad + \gamma_t + \phi_s + \beta X_{i,t-1} + \lambda Z_{s,t-1} \end{aligned} \tag{2}$$

In this expanded equation, we separately identify each month in the sample, which enables us to examine the time pattern in the effects of the \$600 supplement. The months of February and March are once again used as baseline control periods. The February control period effect is identified as the omitted category in the regression (the first term on the right-hand side), and the March control period is separately identified (via the second term on the right hand side). The effects of the UI replacement rate (R_i) for each month when the supplement was available are represented by the coefficients $\pi_{1,2,3,4}$. As above, the sample period ends in June rather than July when we use our two-month matched sample.

The regression specifications also include indicators for calendar months (γ_t) and state of residence (ϕ_s). In addition, the vector $X_{i,t-1}$ consists of individual-level controls observed in the base (pre-transition) month: age (eight categories), education (five categories), race/ethnicity (five categories), gender by marital status, broad occupation (10 categories) and industry (13 categories) of prior employment, and duration to date of the individual's unemployment spell (10 categories, with the final category indicating duration of longer than one year).[13] The model also includes several state/month labor market controls (Z_{st}): cubics in the state unemployment rate and the three-month employment growth rate.

Estimation is via a logit model, with reported parameter estimates converted into average marginal effects. All estimates are weighted by the longitudinal weights that adjust the sample for the characteristics of the sequentially matched observations.[14] The analysis is restricted to individuals with non-zero estimated UI replacement rates under the CARES Act—i.e., individuals who are identified as eligible to receive UI payments based on their prior earnings history. This is a direct implication of our before/after design, since individuals who are not eligible to receive UI payments do not contribute any identifying variation to the estimation.

2.3.1 Main results

The results for the before/after regression specification for unemployment exits and other labor force transitions, based on equation 1 in the preceding section, are shown in Table 1 (estimated coefficients, with robust standard errors in parentheses below them).

Results from the preferred specification for unemployment exits are reported in the first column; this specification relies on the two-month match that corrects for temporary exits from unemployment. The estimated effect of the UI replacement rate measure during the combined months of April through June when the $600 supplement was available is negative and very precise in the first column, attaining significance at better than the 1 percent level.

The second column of Table 1 shows results from the alternative specification for job-finding rates (UE transitions), with the two-month match restriction removed: all consecutive monthly transitions are included and no correction is made for reported temporary exits from unemployment. This enables use of observed transitions through July. As expected, the prevailing labor market turbulence during our sample frame appears to introduce noise in the measurement of monthly transition rates: the estimated interaction coefficient is reduced somewhat in size, although it remains statistically significant at nearly the 1 percent level.

The estimated effects of UI replacement rates during the CARES Act period in columns 1 and 2 of Table 1 are economically meaningful, implying moderate disincentive effects of the enhanced UI payments on job acceptance decisions. In particular, because they are stated as average marginal effects, the interaction coefficients imply a reduction in job-finding rates ranging from 7.2 percentage points (column 1) down to 4.9 percentage points (column 2) when UI replacement rates increase from 0.5 to 1.5 (for example). We discuss interpretation of these magnitudes in more detail in in Section 2.3.3 below, where we compare our estimates to findings from the existing literature on UI benefit levels and job search.[15]

[13] For regressions in which the initial state is out of the labor force, the unemployment duration, industry, and occupation variables are excluded.

[14] The regression results are highly robust to use of different survey weights.

[15] The disincentive effects of UI generosity on job-finding might vary depending on the duration of unemployment. Given the massive job loss early in the pandemic, the typical duration of an in-progress unemployment spell in our

Table 1: Regression results: UI replacement rates and labor force transitions, before/after design

	(1) UE (2-month match)	(2) UE (1-month match)	(3) UN (2-month match)	(4) NE (1-month match)
UI rep rate	0.050**	0.026	0.004	-0.001
	(0.022)	(0.019)	(0.016)	(0.007)
UI rep rate*CARES months	-0.073***	-0.049**	0.013	-0.013
	(0.027)	(0.022)	(0.016)	(0.010)
Observations	2860	5449	2768	7124

*** p<0.01 ** p<0.05, * p<0.1 Notes: Logit regression model results (average marginal effects, with robust standard errors in parentheses) from matched CPS micro-data, Feb.-Jul. 2020, combined with 2020 CPS ASEC data to form individual UI replacment rates (including the $600 supplement from the CARES Act). Regression controls include: age (eight categories), education (five categories), race/ethnicity (five categories), gender by marital status, broad occupation (10 categories) and industry (13 categories) of prior employment, and duration to date of the individual's unemployment spell (10 categories, with the final category indicating duration of longer than one year); state/month economic conditions (cubics in the unemployment rate and log 3-month employment growth); and complete vectors of calendar month and state dummies. The duration, occupation, and industry controls are excluded from column 4.

One notable element of the results is that the pre-CARES (February and March) exit rates are higher for individuals with the highest UI replacement rates under the CARES Act enhancements, with a statistically significant estimate evident for this baseline effect in the first column. This higher unemployment exit rate for individuals who will later receive high UI replacement rates suggests possible violation of the conventional parallel trends assumption for the validity of difference-in-differences estimates. While we cannot reject this interpretation, we view this baseline difference as reflecting systematic unobserved differences between individuals with high and low replacement rates–e.g., it is likely that individuals with high replacement rates under the CARES Act supplement were employed prior to the pandemic in low-wage labor markets with high turnover and job-finding rates. Moreover, this baseline difference is not evident in the second column, which is based on the full set of monthly matches.

We also examined whether UI generosity affects exits from unemployment to out of labor force (UN), with the results displayed in column 3 of the table.[16] This follows earlier empirical results suggesting that UI benefits may increase labor force attachment, because active job search generally is a requirement for UI eligibility in the United States (e.g., Farber, Rothstein and Valletta 2015, Card, Chetty and Weber 2007). The results in column 3 show no meaningful effect of UI

sample is quite short: the mean and median duration are about 8.7 and 5 weeks, and only about 10 percent of the sample has been unemployed for more than 6 months (27 weeks or more). This short duration distribution precludes reliable estimation of our model for individuals with prolonged spells of unemployment. Estimates with the sample divided between those above and below the mean duration suggests somewhat larger and more consistent effects for those with shorter durations.

[16]Relative to the column 2 sample, a small number of observations are lost due to exact collinearity in the column 3 regression.

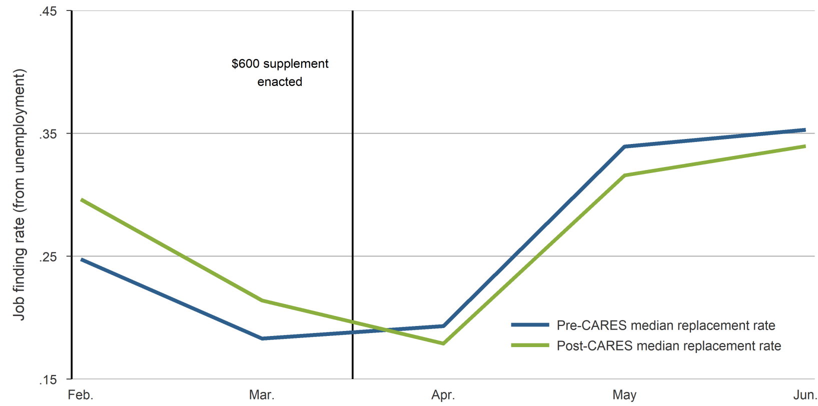


Figure 2: Job finding rates (from unemployment), by UI replacement rates (pre/post CARES Act)
Notes: Calculated from logit regression results (column 1 of Table 2).

replacement rates on reported labor force exits from unemployment. This contrasts with the earlier empirical findings of enhanced labor force attachment due to extended UI durations, likely in part because the job search requirements for UI eligibility were relaxed during the initial phase of the COVID-19 pandemic in the first half of 2020. Finally, column 4 presents results for job-finding rates from out of the labor force (NE). We include this analysis because the relaxation of job-search requirements implies that the $600 supplement may have altered the job search and acceptance decisions of individuals who self-report as not actively searching and hence out of the labor force. However, the results provide no evidence that these transition rates were affected by the increase in UI generosity due to the CARES Act.

The regression results based on equation 2 in section 2.3, which allows the effects of the UI replacement rates to vary across months, are displayed in Table 2. In this specification, the omitted month is February, so the UI replacement rate variable and its interaction with the month indicator for March both reflect the pre-CARES baseline comparison period. The interaction effects for subsequent months represent the impact of the higher replacement rates generated by the extra $600/week UI benefits available through the CARES Act. The table is otherwise structured identically to Table 1, with the results for our preferred two-month match specification reported in the first column.

The results show that the estimated negative effect of UI benefit generosity from column 1 of table 1 vary somewhat across the months when the $600 supplement was available. The estimates are the same size and attain similar statistical significance for April and May, while the estimate for June is slightly smaller and much less precise. In other words, the estimated effects are largest early in the implementation period and then decline somewhat over time, although the monthly estimates are too imprecise for the differences between them to be statistically significant. Declining effects of the $600 supplemental payments as their expiration date approaches is one implication of our model of reservation benefits discussed subsequently, in Section 3.

Figure 2 shows the time pattern of UI generosity effects on job-finding rates based on the

Table 2: Regression results: UI replacement rates and labor force transitions, monthly effects

	(1) UE (2-month match)	(2) UE (1-month match)	(3) UN (2-month match)	(4) NE (1-month match)
UI rep rate	0.057*	0.022	0.016	0.002
	(0.033)	(0.026)	(0.015)	(0.010)
UI rep*Mar	-0.013	0.007	-0.029	-0.005
	(0.043)	(0.037)	(0.034)	(0.013)
UI rep*Apr	-0.080*	-0.025	0.003	-0.011
	(0.043)	(0.033)	(0.018)	(0.017)
UI rep*May	-0.080**	-0.046	-0.001	-0.021
	(0.038)	(0.030)	(0.017)	(0.013)
UI rep*June	-0.069	-0.064**	0.011	-0.020
	(0.046)	(0.032)	(0.028)	(0.019)
UI rep*July	–	-0.004	–	0.008
		(0.046)		(0.037)
Observations	2860	5449	2768	7124

*** p<0.01 ** p<0.05, * p<0.1 Notes: See Table 1 notes. Pre/post-CARES period replaced by individual month indicators.

column 1 results from Table 2, comparing exit rates for individuals at the pre- and post-CARES Act average levels of replacement rates (holding both groups characteristics at the full sample averages). A drop in relative job-finding rates for those with higher replacement rates is evident in April. In subsequent months, job-finding rates increase for both groups, but the job-finding rates for those with higher post-CARES replacement rates remain somewhat lower than for those with lower replacement rates.

The second column of Table 2 show results for the alternative specification for job-finding rates, similar to the same column in Table 1, with the two-month match restriction removed. As in Table 1, this reduces the magnitude and precision of the estimated coefficients. Only the June estimate is statistically significant at the conventional 5 percent level (despite the reduced standard errors afforded by the larger sample size compared with the first column).

Finally, results for the UN and NE transitions in columns (3) and (4) confirm no effect of UI replacement rates on these transitions during any of the months of our sample.

2.3.2 Sensitivity to industry and occupation

As noted in Section 2.2, the concentration of early pandemic job losses among low-wage services sectors raises potential concern that our estimates of UI benefit effects may be contaminated by sector-specific labor market conditions. More precisely, rather than solely reflecting effects of expanded UI generosity, our estimates may instead reflect reduced job prospects for workers previously employed in sectors in which activity was directly constrained by pandemic effects. As already illustrated in that earlier section, the wide distribution of UI replacement rates across and within occupations and industries in our sample partly alleviates this concern. This concern is also partly addressed by the inclusion of broad occupation and industry dummies in the regressions reported in the preceding section.

We further explore the sensitivity of our results to sector-specific effects by restricting the sample used for our regressions based on occupation and industry. The small sample sizes for individual industry and occupation groups preclude reliable estimation of UI replacement rate effects within those groups.[17] However, we can explore whether our main results are driven by key sectors by sequentially excluding each occupation and industry from the estimation sample.

The results are shown in panels A and B of Tables 3. For straightforward interpretation, we once again rely on the simplified "before/after" design, as in Table 1 above. The specification is identical to that from column 1 of Table 1, but with the indicated occupations and industries sequentially excluded from the full sample. Across all columns of both panels, the estimated effect of the UI replacement rate during the CARES Act $600 supplement period (April-July) is tightly distributed around the full sample estimate of -0.072 (from column 1 of Table 1) and in all cases is precisely estimated. One notable and perhaps surprising exception is the leisure and hospitality industry (column 11 of Table 3, panel A). When this industry is excluded, the estimated effect of UI replacement rates under the CARES Act rises substantially, suggesting that the effect is small in this sector. Overall, these results bolster the case for interpreting our findings as reflecting variation in the UI replacement rates associated with the $600 supplement rather than sector-specific differences in job prospects for individuals who lost jobs early in the pandemic.

[17] When estimated for subsets of grouped occupations and industries, the estimated effects of the UI replacement rate in our before/after design varied widely across sectors and generally were statistically imprecise.

Table 3: Regression results: UI replacement rates and job finding (UE rates, two-month match), before/after design, occupations and industries eliminated

Panel A: occupations eliminated

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Manag	Prof/Tech	Services	Sales	Admin	Farm/ Fish	Const/ Extract	Inst/ Maint/Rep	Prod.	Trans/ Materials
UI rep rate	0.057**	0.040	0.049*	0.045*	0.062**	0.050**	0.045*	0.054**	0.050**	0.048**
	(0.024)	(0.025)	(0.026)	(0.024)	(0.026)	(0.023)	(0.024)	(0.023)	(0.022)	(0.023)
UI rep* CARES months	-0.074***	-0.074**	-0.073**	-0.070**	-0.076***	-0.072***	-0.062**	-0.074***	-0.068**	-0.074***
	(0.028)	(0.030)	(0.029)	(0.029)	(0.029)	(0.027)	(0.027)	(0.027)	(0.026)	(0.028)
Observations	2568	2409	2191	2547	2611	2833	2582	2774	2645	2580

Panel B: industries eliminated

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Agric	Mining	Const	Manuf	Whl/Ret	Trans/ Util	Info	Financial	Prof/ Bus	Educ/ Health
UI rep rate	0.050**	0.051**	0.043*	0.059**	0.059**	0.054**	0.051**	0.045**	0.051**	0.048*
	(0.023)	(0.023)	(0.023)	(0.024)	(0.024)	(0.023)	(0.023)	(0.023)	(0.025)	(0.025)
UI rep* CARES months	-0.072***	-0.071***	-0.058**	-0.076***	-0.079***	-0.076***	-0.073***	-0.067**	-0.067**	-0.078**
	(0.027)	(0.027)	(0.027)	(0.028)	(0.029)	(0.027)	(0.027)	(0.027)	(0.029)	(0.031)
Observations	2825	2828	2547	2548	2487	2730	2817	2746	2567	2362

	(11)	(12)	(13)
	Leis/Hosp	Oth Serv	Gov
UI rep rate	0.046*	0.048**	0.051**
	(0.026)	(0.022)	(0.023)
UI rep* CARES months	-0.091***	-0.062**	-0.072***
	(0.029)	(0.026)	(0.027)
Observations	2336	2699	2811

Notes: *** p<0.01 ** p<0.05, * p<0.1 Notes: Occupations and industries excluded as indicated in column titles. Specification is from column 1 of Table 2.

2.3.3 Assessing the magnitude of the UI effect on job finding

We assess the magnitude of the estimated impact of the CARES supplement based on the results from the first column of Table 1, which applies the before-after design to our two-month match specification. We focus on this specification because the two-month match correction for spurious transitions bolsters the precision of the results, and the before/after framework provides a straightforward averaging of the UI replacement rate effects across the months when the $600 supplement was available.

Given the wide span of post-CARES UI replacement rates observed in our data, various metrics could be used to interpret the size of the estimated effect.[18] Interpretation of the coefficients is straightforward, however: the replacement rate is measured relative to a value of 1.0 (UI payments equal to prior earnings), and the coefficients are average marginal effects that represent the effect of an increase in the UI replacement rate of 1 (100 percentage points) on the probability of observing the relevant transition.

We conduct straightforward calculations based on these considerations. The $600/week additional payments raised the median replacement rate from 0.5 to 1.39 in our two-month matched sample of unemployed individuals.[19] This represents an increase in the typical replacement rate of 0.89. Based on the coefficient on the UI replacement rate interaction with the CARES Act months in column 1 of Table 1, this implies that the job-finding rate for the typical recipient of enhanced UI benefits during those months was reduced by about 6.5 percentage points, or 0.065.[20] This is of moderate size relative to job-finding rates averaging about 0.29 during the months when the $600 supplement was available. Specifically, this represents a 22 percent reduction in job-finding rates due to the availability of the $600 weekly UI benefit supplement.

We can also translate our estimate into an elasticity of job-finding with respect to variation in UI replacement rates. This enables us to put our findings further into context and compare them directly with results from other research. We use mean rather than median changes for consistency with conventional elasticity calculations. The $600 supplement raised average UI replacement rates from 0.44 to 1.53 in our sample from column 1 of Table 1 (a 245 percent increase). This increase of 1.09 in the average replacement rate lowers job-finding rates by 7.8 percentage points (1.09*(-0.073)=-0.080), or 28 percent relative to a base rate of about 0.29. The calculated job-finding elasticity is -0.11, with a standard error of 0.042 and a 95 percent confidence interval of -0.045 to -0.18.[21]

[18] As noted earlier, the $600 supplement substantially raised the typical replacement rates. It also widened the dispersion substantially, with the standard deviation of replacement rates across UI-eligible individuals rising by nearly a factor of seven.

[19] We limit this calculation to individuals observed during the period when the supplement was available, from April through July. The median replacement rate of 1.39 in this two-month matched sample is very close to the value of 1.40 for the full sample of unemployed individuals discussed in section 2.2.

[20] The specific calculation is 0.89·(-0.073)=-0.065.

[21] The exact calculation is is [(-0.078)/(0.286)]/[(1.090)/(0.444)]=-0.111. The standard error and confidence intervals for this point estimate are obtained via a bootstrap, using 500 iterations of random re-sampling with replacement from the regression sample. Our estimated elasticity range is somewhat lower if we instead use the interaction coefficient from the sample of one-month transitions in column 2 of Table 1 (point estimate of -0.077, with a standard error of 0.031

This estimated elasticity of -0.11 and corresponding 95 percent confidence interval are at the low end of the range of elasticity estimates using U.S. data summarized in Schmieder and von Wachter (2016). In particular, the low end of the range summarized in their Table 2 is -0.10 to -0.15, with about half of the reported estimates exceeding -0.5.[22]

Our estimated elasticity is close to two other careful estimates of pandemic UI effects, both using very different data than ours (administrative microdata from financial services companies, from two separate sources). Ganong et al. (2022) rely on two identification strategies, including one that is very similar to ours, based on a difference-in-differences framework using individual variation in UI replacement rates. With this approach they obtain an estimate of the elasticity of the UI benefit effect on job finding that is essentially identical to ours, -0.11.[23] Coombs et al. (2022) obtain a somewhat higher estimated elasticity range of -0.13 to -0.22; this is largely within the confidence interval for our estimate and like ours is toward the lower end of the range from past estimates (Schmieder and von Wachter 2016).

As noted in Section 1, other studies found much smaller labor market effects of the $600/week UI benefit enhancement. In particular, Bartik et al. (2020), Altonji et al. (2020), and Finamor and Scott (2021) found that states that experienced the largest increases in UI payments due to the $600/week supplement did not experience weaker labor market rebounds during the initial phase of economic recovery from the pandemic.

It is likely that our somewhat larger estimates of the pandemic UI benefit effects on job finding, along with those in Ganong et al. (2022) and Coombs et al. (2022), arise because we focus on search responses at the individual level, commonly referred to as micro effects. Our regressions rely on variation in UI replacement rates across individuals before and during the periods when the CARES Act $600 supplement was available, with no direct channel for broader macro effects to influence the estimates. Reinforcing this point, although we include measures of state labor market conditions in our regressions, our results for the UI replacement rates are largely invariant to their inclusion or exclusion.

By contrast, the studies that find smaller effects likely combine responses along the micro or individual margin with more general macro effects on aggregate outcomes. These macro effects can take various forms. The most important one in regard to the pandemic UI payments is their aggregate stimulus effects via consumption spending. Recent research has found these stimulus effects to be substantial during the pandemic, with a large marginal propensity to consume out of the UI benefits paid (Ganong et al. 2022). This added household spending likely helped sustain labor demand and hence offset the job search disincentives for individual UI recipients (Boone et al. 2021, Kekre 2023). Analyses that rely on outcomes aggregated above the individual level (i.e., state and local labor markets) are likely to capture a combination of the micro responses at

and 95 percent confidence interval of -0.015 to -0.117)

[22]Schmieder and von Wachter (2016) focus on the elasticity of search duration, which takes on positive values, in their literature summary. Under the assumption of constant exit rates from unemployment, the elasticities of search duration and job-finding probabilities are nearly identical in absolute value.

[23]Ganong et al. (2022) report their estimated elasticities in duration rather than transition terms, hence with opposite sign to ours.

the individual level and the offsetting stimulus effects at the macro level. Our estimates instead focus on the narrow micro or individual responses.[24]

3 UI income and job acceptance decisions: a reservation benefits framework

Our direct assessment of individual job-finding rates in the preceding section uncovered moderate disincentive effects of the $600 weekly UI supplement. In order to further understand this result, this section develops a dynamic framework of job acceptance decisions and applies it to the specific features of the CARES Act UI benefits expansion. Importantly, our modeling framework incorporates heterogeneity in the distribution of prior earnings within groups of job seekers (defined primarily by educational attainment or prior occupation). The model sheds light on why many UI recipients will accept job offers even when their benefits exceed the offered wage. As in the preceding section, we use data from the CPS ASEC to infer the underlying wage distribution for job seekers and CARES Act replacement rates. Our quantitative analysis suggests that a moderate fraction of UI recipients would have refused an offer to return to work at their previous pay during the period when the $600 weekly supplement was available.

3.1 Reservation benefits

We study the problem of a risk neutral insured job seeker considering a job offer to return to work at the previous wage, w_j, characterizing the level of UI benefits that leave a job seeker indifferent between accepting and rejecting the offer.[25] The framework highlights the circumstances under which a job seeker may accept a job offer for a wage below the value of their current weekly UI payments. Finally, note that the reservation benefit statistic developed here does not take into account risk aversion, which would increase the value of a long stream of earned income on the job compared with temporary UI payments.

Consider a worker comparing the present value of the job, $W_E(w_j)$ to that of remaining unemployed with UI benefits b_j, which depend on prior earnings w_j, and t remaining weeks of eligibility, $W_U(b_j, t)$. The decision takes into account the likely duration of the job and that of finding an alternative offer – through the probabilities of losing and finding a job s and f, respectively – and the discounting of time at rate r:

$$W_E(w_j) \quad = \quad w_j + \frac{1}{1+r}\left[(1-s)\,W_E(w_j) + sW_U\left(b_j, T\right)\right] \tag{3}$$

[24]Other channels for macro effects that may offset the individual search response include job rationing and spillovers to individuals not covered by the UI enhancements (Michaillat 2012, Lalive, Landais and Zweimüller 2015, Landais, Michaillat and Saez 2018). These channels may have affected recent assessments of expanded UI effects, although the unusually widespread availability of enhanced UI benefits during the pandemic likely limited the spillovers to ineligible individuals.

[25]Although there is little evidence of significant wage cuts during the recession triggered by the COVID-19 pandemic, the approach developed here is straightforward to adapt to any wage offer.

$$W_U\left(b_j,t\right) = b_j + \frac{1}{1+r}\left[(1-f)\,W_U(b_j,t-1) + f\max\left[W_E(w_j),W_U(b_j,t-1)\right]\right] \quad \text{for } 1 < t \leq T \tag{4}$$

$$W_U\left(b_j,1\right) = b_j + \frac{1}{1+r}\left[(1-f)\,W_U(0) + f\max\left[W_E(w_j),W_U(0)\right]\right] \tag{5}$$

$$W_U(0) = 0 + \frac{1}{1+r}\left[(1-f)\,W_U(0) + f\max\left[W_E(w_j),W_U(0)\right]\right] \tag{6}$$

where T is the maximum duration of UI, $W_U(0)$ is the value of unemployment after exhaustion of unemployment benefits, $W_U(b_j,T)$ is the value of unemployment at the start of a new unemployment spell following a job loss and, for a positive wage, $\max\left[W_E(w_j),W_U(0)\right] = W_E(w_j)$.[26]

If employment if preferred to remaining unemployed at a date $t+1$ then, from the value functions above, the value of unemployment up to the maximum duration of UI of T weeks can be re-expressed as:

$$W_U\left(b_j,t\right) = B_j(t) + \left(\frac{f}{r+f}\right)W_E(w_j) \text{ for } 1 < t \leq T \tag{7}$$

which highlights that unemployment is valued for the discounted present value of expected UI payments with t weeks of eligibility remaining, $B_j(t) = \sum_{i=0}^{t-1} b_j \left(\frac{1-f}{1+r}\right)^i$, and the discounted value of finding a job and moving into employment.

Since the value of unemployment in (7) is increasing in the weekly benefit amount, there exists a reservation benefit $b_j^r(t,w_j)$ to be paid out for the remaining weeks of eligibility t such that an individual is indifferent between remaining unemployed and receiving that amount or accepting a job offering pay w_j. That is, a job offer with pay w_j will be turned down if the current level of weekly benefit payments b_j is greater than this reservation level $b_j^r(t,w_j)$. Formally:

Proposition 1. *The reservation benefit for an unemployed individual with t weeks of UI eligibility remaining and considering a job offer at wage w_j solves:*

$$W_U\left(b_j^r(t,w_j),t\right) = W_E(w_j) \tag{8}$$

Given the value functions for employment and unemployment (3) and (7) the reservation benefit is

$$b_j^r(t,w_j) = \frac{b_j^r(1,w_j)}{\sum_{i=0}^{t-1}\left(\frac{1-f}{1+r}\right)^i} \quad \textit{for} \quad 0 < t \leq T \tag{9}$$

where

$$b_j^r(1,w_j) = \left(\frac{r}{r+f}\right)W_E(w_j) = \left(\frac{r}{r+f}\right)\left(\frac{(1+r)\,w_j + sW_U(b_j,T)}{r+s}\right) > w_j \tag{10}$$

[26] We assume that employment immediately affords eligibility to full UI whereas state UI systems have different work and earnings requirements to establish UI eligibility. Moreover, we make the simplifying assumption of considering job offers at the previous wage w_j and an arrival rate of job offers that is independent of prior earnings. Detailed derivations for all results are provided in the appendix.

Job seekers will accept an offer to return to work at their previous wage if weekly income from UI benefits is lower than their reservation level of benefits with t weeks of payments remaining, $b_j < b^r(t, w_j)$.

For a given wage offered, the level of reservation benefit leading to a job offer being rejected is determined by: (i) the duration of benefits remaining (t); (ii) the expected duration of the employment spell ($\approx 1/s$), and; (iii) the rate of arrival of new job offers (f). With an unbounded duration of UI payments ($T \to \infty$) the reservation benefit is equal to the wage, $b_j^r(\infty) = w_j$. In this limit, a replacement rate above 100 percent will induce workers to reject a job offer at their previous wage rate. With one week remaining, the reservation benefit $b^r(1, w_j)$ is the annuity value of the present discounted value of the job offered. It is always the case that, with a week remaining, the reservation benefit is greater than the wage offer ($b^r(1, w_j) > w_j$). In other words, replacement ratios above 100 percent do not necessarily lower job offer acceptance rates. More generally, for UI benefit payments of finite duration, the reservation benefit $b^r(t, w_j)$ is declining with weeks remaining of UI benefits, trading off an additional week of benefits at the reservation level against the forgone employment value.

The level of the reservation benefit depends crucially on the expected duration of the employment spell and the rate of arrival of new job offers, over and above the considerations from the duration of remaining weeks of UI eligibility. Longer lasting employment spells (lower s) are of greater value and rejected only for commensurately generous unemployment insurance payments. In a depressed labor market, when job offers are few and far between (low f), any job offer is costly to refuse as new offers are hard to find. This can be seen in the discounting terms in equations (9) and (10).

3.2 Reservation benefits during the pandemic

This section provides estimates of reservation benefits for different categories of workers during the COVID-19 recession. We adapt the general problem to reflect institutional details from the CARES Act and then use micro data from the CPS to obtain the relevant moments entering the definition of a reservation benefit level. The main set of results are based on the experience during the recovery out of the Great Recession of 2007-09, especially with respect to the expected hazard rates out of unemployment. Additional results, obtained by varying the assumptions on the expected durations of unemployment and employment spells, are provided and are meant to capture the ranges of reservation benefits and offer rejection rates under alternative worker assessments of the state of the labor market.

3.2.1 CARES Act specific formulation

The temporary nature of the supplemental PUC income relative to the duration of payments of baseline UI requires a small modification to the unemployment Bellman equations above. Let t_c denote the weeks of expanded UI eligibility, and t_p the weeks of supplemental UI income under the PUC remaining for a given unemployment spell. For simplicity it is assumed that $t_p < t_c$ for all unemployed. In addition, let $\bar{b}_j$ denote baseline UI payments and the additional income provided through the PUC by b_p. Note the baseline payments $\bar{b}_j$ are conditional on past earnings while the PUC payments b_p are not. The value of unemployment under the CARES Act is:

$$\begin{aligned} W_U\left(\bar{b}_j, t_c, b_p, t_p\right) &= \bar{b}_j + b_p + \frac{1}{1+r}\left[(1-f)\, W_U(\bar{b}_j, t_c-1, b_p, t_p-1)\right. \\ &\quad \left. + f \max\left[W_E(w_j), W_U(\bar{b}_j, t_c-1, b_p, t_p-1)\right]\right] \text{ for } t_c, t_p > 1 \qquad (11) \\ W_E(w_j) &= w_j + \frac{1}{1+r}\left[(1-s)\, W_E(w_j) + s W_U\left(\bar{b}_j, T_c\right)\right] \qquad (12) \end{aligned}$$

Following similar steps as in the previous section, the value of unemployment under the CARES Act with t_c weeks of regular UI payments and t_p weeks of PUC payments may be expressed as:

$$W_U\left(\bar{b}_j, t_c, b_p, t_p\right) = \overline{B}_j(t_c) + B_p(t_p) + \frac{f}{r+f} W_E(w_j)$$

where $\overline{B}_j(t) = \sum_{i=0}^{t-1} \bar{b}_j \left(\frac{1-f}{1+r}\right)^i$ is the present discounted value of expected baseline UI payments and $B_p(t) = \sum_{i=0}^{t-1} b_p \left(\frac{1-f}{1+r}\right)^i$ the present discounted value of expected supplemental UI payments.

We focus on the the level of *supplemental* UI payments leading to indifference to job offers at the previous wage w_j, denoted $b^r_{p,j}(t, t_c, w_j)$. The reservation benefit for a job seeker during the period of the CARES Act is the sum of regular benefit payments $\bar{b}_j$ and this supplemental reservation benefit payments: $b^r_j(t, t_c, w_j) = \bar{b}_j + b^r_{p,j}(t, t_c, w_j)$. This level of reservation for the supplemental benefit depends on the number of weeks of regular benefit payments remaining, t_c, and, for 1 or t weeks remaining in PUC payments, is given by:

$$b^r_{p,j}(1, t_c, w_j) = \frac{r}{r+f} W_E(w_j) - \overline{B}_j(t_c) \qquad (13)$$

$$b^r_{p,j}(t, t_c, w_j) = \frac{b^r_{p,j}(1, t_c, w_j)}{\sum_{i=0}^{t-1} \left(\frac{1-f}{1+r}\right)^i} \qquad (14)$$

3.2.2 Calculating reservation benefits during the pandemic

As previously mentioned, the reservation benefits during the pandemic calculated below is the sum of regular and supplemental reservation benefit payments for an individual considering a of job offer at their previous wage w_j, $b^r_j(t, t_c, w_j) = \bar{b}_j + b^r_{p,j}(t, t_c, w_j)$. We specify the baseline UI program as a weekly payment $\bar{b}_j = \min\left[\bar{\tau} \times w_j, b_{cap}\right]$ for a maximum duration of $\bar{T} = 26$ weeks, where $\bar{\tau} \in (0, 1)$ is a replacement rate set to 50 percent and b_{cap} a cap on weekly payments of \$500.[27] The PEUC extended the duration of UI payments an additional 13 weeks for a total of 39

[27]This assumption for regular UI compensation is slightly more generous than the typical U.S. state program. See Department of Labor (2019) for a review of the heterogeneity in eligibility requirements and benefit levels and duration across US states. Note also the discount rate r is set to an annualized rate of 5 percent.

weeks, but in some states emergency extensions provide an additional 13 weeks for a maximum of 52 weeks. We set T^c to 52 weeks. The additional income provided through the PUC is denoted by $b_p = \$600$ per week. Payments first began the week ending April 4, 2020 and the last went out the week ending July 25, 2020, for a total of $T_p = 17$ weeks. Finally, the CARES Act provision of additional UI income is assumed to no longer be available at the end of the employment spell of any job offer under consideration.[28]

The remaining moments for weekly earnings, and job finding and job separation rates required to calculate reservation benefits are obtained from the monthly CPS. These are reported for the overall population, prime aged workers, by level of education, and by occupation of prior job in Table 4. As in the regression analysis, individual weekly earnings from 2019 are obtained from the 2020 CPS ASEC file, for individuals recorded as unemployed in the months of January through July 2020; this information captures the likely distribution of earnings for workers considering job offers during the period of the CARES Act provisions. More precisely, we assume that the distributions of weekly earnings for job seekers within a group are drawn randomly from a log-normal distribution $G(\cdot)$, with mean μ_w and standard deviation σ_w; these parameters are estimated directly from the CPS ASEC individual earnings data just referenced. For example, the distribution of log weekly earnings for our overall sample of unemployed individuals has a mean of 6.58 and a standard deviation of 0.73, while the distribution of log weekly earnings for unemployed individuals who were in managerial occupations in their prior job has a mean of 7.01 and a standard deviation of 0.69.

In our baseline calculations, the measure of expected durations of unemployment and employment spells draws on the experience from early recovery phase following the Great Recession (the full calendar year 2010). This period is chosen as a reasonable reference point for a job seeker's expectation of job offer arrival rates coming out of the initial phase of the COVID-19 recession. In addition, below we present results under an alternative assumption regarding job offer arrival rates, to assess model properties and robustness. The job offer arrival rate $f_t = UE_t/U_{t-1}$ is the sum of transitions from unemployment to employment over the previous period's stock of unemployed individuals. The separation rate out of employment $s_t = (EU_t + EN_t)/E_{t-1}$ is the sum of transitions out of employment into either unemployment or non-employment over the preceding period's stock of employed individuals. It is worth noting that durations of unemployment spells based on outflow rates are significantly shorter than the average durations reported by CPS respondents. In other words, our chosen measure of job offer arrival rates will imply lower levels of reservation benefits compared to using self-reported durations of unemployment spells.[29]

Transition rates into employment in a specific occupation are not easily defined due to the

[28] Allowing for the additional UI income to be available upon reemployment, at least partially, would increase the value of a job offer. The levels of the reservation benefit would be somewhat higher due to strong discounting over the duration of a typical employment spell.

[29] Table A1 provides durations of unemployment spells as self-reported in the CPS for comparison to the durations implied by the finding rate f. In particular, it reports the average duration of the unemployment spell preceding a transition into employment, which can be compared to the imputed finding rate based on durations by occupation. Table A4 of the appendix reports the equivalent moments for 2015. See also the discussion in Farber and Valletta (2015).

ambiguity of identifying the pool of potential job seekers within each occupation. Our solution to this measurement challenge follows Hall and Schulhofer-Wohl (2018). We estimate a logit on the outcome of a potential transition from unemployment into employment into a specific occupation, $f = \exp(\beta_f X) / \left[1 + \exp(\beta_f X)\right]$, based on a set of demographic characteristics in the vector X that includes age, education, race/ethnicity, sex, and marital status. The regressions, using all months of 2010, are then used to predict the average transition rate by occupation (see appendix B for further details).

3.2.3 Results: Overall, by education, and by occupation

The PUC benefit expired July 31st 2020. As such, we focus on reservation benefit levels, and the corresponding UI benefit amounts, for individuals considering an offer to return to work at the previous wage between the end of April and the end of June 2020. This is the core period of concern over the disincentive effects of the supplemental UI payments, during which job seekers were considering leaving behind 4 to 12 weeks of increased UI payments by returning to work. This timeframe largely overlaps with that in the analysis of individual transition rates in the preceding section. The second to last column of Table 4 reports the value of reservation benefits at a midpoint in this period, the end of May, for the median worker within each group.

The final column of Table 4 displays our main results. It reports the proportion of individuals within each group who will reject a job offer at their prior wage based on the comparison between their imputed UI benefits and their calculated reservation benefit level. That is, let $o_j^r(t)$ be an indicator function equal to one if an individual with prior wage w_j would reject an offer to return to work with t weeks of PUC payments remaining: $o_j^r(t) = 1$ if $b_j^r(t) \leq \bar{b}_j + b_p$. The share of individuals rejecting an offer to return to work with t weeks remaining is then calculated as $\int o_j^r(t) dG(w_j)$. For example, for our overall sample, the results in the first row, column 7, show that the fraction of individuals who would reject job offers based on their UI benefits is about 22 percent; we discuss this result in more detail below.

The findings in column (7) of Table 4 are best understood by reviewing the components in the prior columns that contribute to the calculation. Focusing on the overall sample in the first row, the table shows that a typical (median) unemployed worker earning about \$717 per week in their previous job received \$959 per week in UI payments under the CARES Act (columns 1 and 5). Considering an offer at the previous wage takes into account that the proposed employment spell is expected to last just under two years and, if rejected, unemployment can be expected to last 22 weeks (column 2). We calculate that an offer received with 8 weeks of CARES UI payments remaining would be accepted as long as this worker's current benefit payment was below $b_j^r(8, w_j = 717)$ of around \$1419 (column 6), a reservation benefit level that is about twice the previous wage.[30] This is well above weekly UI payments under the CARES Act, implying that the

[30]The following example illustrates the calculation of the reservation benefit level with 8 weeks of PUC payments remaining for the median earner facing a weekly job arrival rate of $f = 0.047$, separation rate $s = 0.012$. The first step is to calculate supplemental benefit level with 1 week of PUC and $t_c = T_c - T_p + 8$ weeks of regular payments remaining

Table 4: Reservation benefits, replacement rates, and offer rejection rates

	(1) Earnings	(2) Duration of:	(3)	(4) Weekly UI comp.	(5)	(6)	(7) Rejection rate (%)
	w (wkly)	U (wks)	E (yrs)	$\bar{b}$	b^C	$b^r(8)$	Average[a]
Overall	717	22	1.7	359	959	1419	21.5
Age 25 to 54 years	792	21	2.3	396	996	1598	15.6
Education:							
Less then HS	514	23	0.74	257	857	908	43.1
High School	650	22	1.6	325	925	1290	22.0
College and above	955	19	2.8	477	1077	1853	15.2
Occupation:							
Management	1103	20	2.3	500	1100	2196	8.4
Prof. and tech.	850	20	2.3	425	1025	1639	19.0
Services	541	22	1.2	271	871	1026	34.6
Sales	637	21	1.4	318	918	1206	29.5
Admin.	608	22	1.6	304	904	1215	26.4
Farm., fish.	608	20	1.0	304	904	1084	30.5
Constr., extract.	929	20	1.5	465	1065	1717	13.1
Inst., maint., rep.	831	20	1.7	415	1015	1572	15.3
Production	798	21	1.5	399	999	1545	12.6
Transp., materials	688	21	1.4	344	644	1314	24.2

Notes: Earnings data calculated using the 2019 CPS ASEC for indidivuals unemployed in any of the months of March through July 2020. Durations of unemployment and employment in columns 2 and 3 are calculated using the Dec. 2009 to Dec. 2010 CPS. w: median weekly earnings. Weekly job finding and separation rates entering the resevation benefits are obtained by converting the monthly flow rates to a weekly frequency (see appendix for details); $\bar{b}$: median regular weekly unempmloyment benefits; b^C: median weekly benefits under the CARES Act, $\bar{b} + 600\$$; $b^r(t_p)$ median reservation benefit level with t_p weeks left under the CARES Act. a: Average rate of rejecting an offer arriving between 12 and 4 weeks of PUC payments remaining.

median worker would not reject an offer to return to work at the prior wage. However, there is significant dispersion in weekly earnings such that many workers at the lower end of the distribution might be dissuaded from accepting a job offer at their prior wage. As noted above, the final column reports the fraction of workers whose weekly UI payments under the CARES Act exceed their reservation benefit level, prompting them to reject job offers at their prior wage. We find that about 22 percent would reject an offer to return to work, calculated as an average for the period from the end of April through the end of June 2020. This proportion is essentially identical to the magnitude of the disincentive effects on job finding rates found in the previous section's analysis.[31] These model-based calculations are not exactly comparable to the regression results, which are based on an empirical sample and a before/after research design. However, the close correspondence between the estimated effect of the \$600 UI supplement in the two settings is reassuring, given the independent design of the two approaches.

As shown in the second row of Table 4, the results are similar when the analysis is restricted to the prime age workforce, aged 25 to 54 years old, though the share who reject job offers is somewhat lower than for the full sample (16 percent versus 22 percent). The next three rows of Table 4 present the results for workers in three educational attainment groups (less than high school, high school, and college and above). College educated workers were less likely than those with lower educational attainment to turn down a job offer at the previous wage during the period of supplemental payments under the CARES Act: their earnings are higher relative to UI benefits and their employment spells are longer compared with individuals with less education. Conversely, high school educated unemployed workers, with median earnings of \$650 per week and expected durations of employment and job-finding rates close to the overall average, were moderately deterred by a high replacement rate under the CARES Act.

The subsequent rows of Table 4 present reservation benefit calculations for workers within 10 major occupations. Median weekly earnings among the unemployed within these occupations range from just under \$541 per week (Services) to over \$1100 per week (Managers), with average durations of employment spells from under a year to about two and a half years. The median reservation benefits levels with 8 weeks remaining in PUC payments for each occupation are reported in the second to last set of columns in Table 4, while Figure 3 plots an occupation's median weekly earnings against the median reservation benefit within that occupation with 8 weeks of PUC payments remaining. A 100 percent replacement rate (black line) separates the graph in two regions, shaded in blue for replacement rates below 100 percent. Regular UI payment rates are represented by the bottom line (red), increasing at a rate of 50 percent of the prior wage until hitting a cap at \$1000 in weekly earnings for a maximum benefit payment of \$500 per week. The

as in equation (13), resulting in $b^r_{p,j}(1, t_c, w_j = 717)$ where we have used the definition of the value for employment in (12) as detailed in appendix section A.2in equation (A.5). The supplemental benefit level with 8 weeks reamaining is then obtained as $b^r_{p,j}(8, t_c, w_j = 717) = b^r_{p,j}(1, t_c, w_j = 717) / \sum_{i=0}^{8-1} \left(\frac{1-f}{1+r}\right)^i = 1060$, to which we add regular weekly UI payments of $\bar{b}_j = 359$, yielding the value reported in the first row of Table 4.

[31] As noted in Section 2.3.3: "Specifically, this represents a 22 percent reduction in job-finding rates due to the availability of the \$600 weekly UI benefit supplement."

UI payment schedule under the CARES Act is shifted up by $600 (green line), and any individual with earnings below $1100 per week receives more on UI with the PUC payments than on the previous job. Each occupation's median weekly earnings and reservation benefit level with 8 weeks of PUC supplemental payments remaining are plotted as blue dots. Only median earners employed in Services have UI payments under the CARES Act close to their respective median reservation benefit levels.[32] Individuals in these occupations were more likely than others to reject a job offer at the previous wage during the period analyzed.

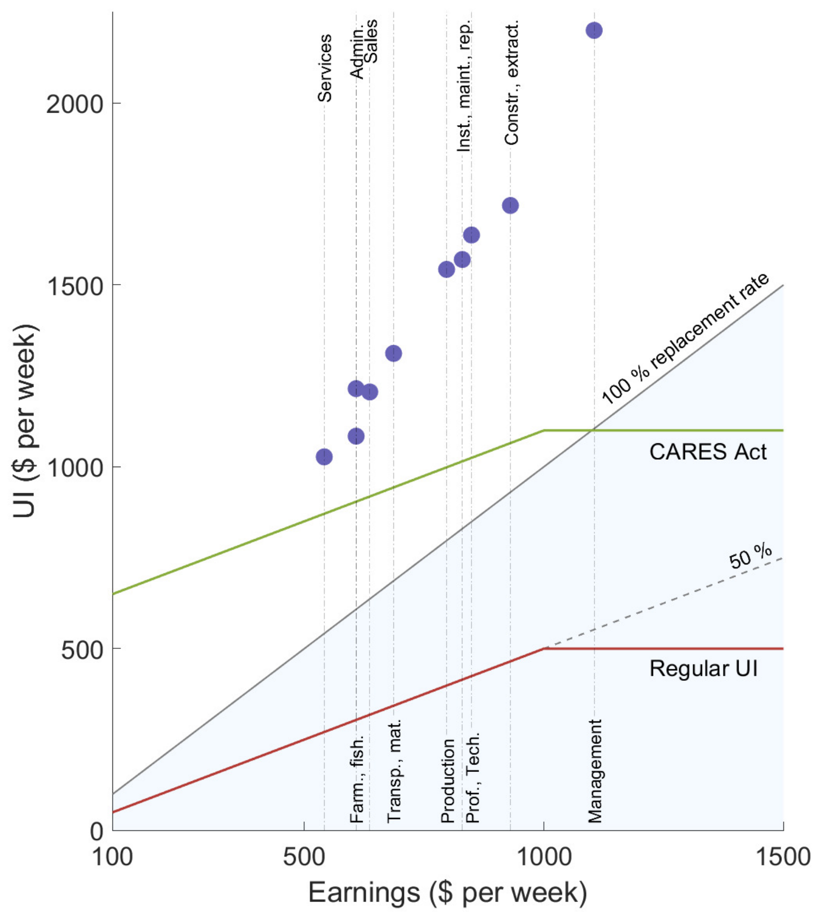


Figure 3: Regular, CARES Act and reservation level UI benefit payments for median earner by occupation

Notes: Each dot corresponds to the reservation benefit for a with median earnings within each occupation calculated according to equation (14) with 8 weeks of PUC payments remaining.

[32]These results by occupation are related to the robustness exercise in the earlier section evaluating whether specific occupations or industries drive the regression results. However, our sample sizes for the regression analysis are too small to enable regressions for each occupation.

In order to provide a sense of how reservation benefits and rejections of job offers are affected by varying expectations for labor market conditions, we performed the same calculations as summarized in Table 4 under an alternative assumption for job offer arrival rates and durations of employment spells. This alternative uses data from 2015 rather than 2010 to obtain transition rates, thereby incorporating stronger labor conditions and better job-finding prospects than in our initial calculations. Under this scenario, the arrival rate of job offers is raised by nearly 50 percent. The results are reported in the appendix, in Table A4, which is structured identically to Table 4. The increased arrival rate for job offers lowers the level of reservation benefits in all occupations. For the full sample, about 34 percent would reject an offer to return to work at their previous wage, notably higher than the 22 percent obtained in our initial scenario and also in the regression analysis. However, this scenario is meant purely to illustrate the sensitivity of our results to labor market conditions. We regard the results from our primary scenario displayed in Table 4 as a better reflection of the prevailing labor conditions and job acceptance decision by recipients of the $600 UI supplement in early 2020.[33]

4 Conclusion

This paper examines the impacts of the $600 weekly UI benefits supplement provided by the CARES Act on job search and job acceptance decisions in early 2020. We first conduct direct empirical analyses of labor force transitions using matched CPS data, linked to annual earning records from the CPS income supplement to form UI replacement rates. The results show moderate disincentive effects of the UI supplemental payments on job finding rates. Our estimated elasticities of job-finding rates with respect to UI replacement rates are at the low end of the range of prior estimates using U.S. data (Schmieder and von Wachter 2016). Our estimates are broadly comparable to results from two other recent papers that use administrative micro-data from financial services companies to assess the impacts of the pandemic UI supplements (Coombs et al. 2022, Ganong et al. 2022). In fact, our point estimate of the UI benefit elasticity for job finding is essentially identical to a similarly obtained estimate from different data in Ganong et al. (2022). Our estimates are aimed at identifying the micro effects of the UI benefit expansions on job search, whereas other recent analyses of the pandemic UI expansions likely combine micro effects at the individual level with more general macro effects that alter aggregate labor market conditions (Altonji et al. 2020, Bartik et al. 2020, Finamor and Scott 2021).

In the second part of our analysis, we derive a level of UI benefit payments over the duration of remaining UI eligibility at which unemployed individuals are indifferent between accepting a job paying their previous wage and remaining unemployed. With finite UI benefit duration, this

[33] All else equal, the frequency of rejected job offers in the model increases with the expected job offer arrival rate. Intuitively, UI recipients are more willing to continue receiving benefits if they know that additional job offers will come sooner rather than later. Similarly, if UI recipients expect the availability of benefits to be extended, as they were in the aftermath of the recession of 2007-09, this will increase the job offer rejection rate in the context of our model. We interpret the close alignment of our regression results and model calculations in Table 4 as indicating that the latter are based on a reasonably accurate representation of labor market conditions and job seeker expectations in early 2020.

reservation benefit is always above the previous wage. In a depressed labor market with lower job offer arrival rates, the gap between the previous wage and the reservation benefit widens, leaving room for replacement ratios above 100 percent without large negative effects on job acceptance rates and the resulting speed of the labor market recovery. Our model-based calculations using CPS monthly and ASEC microdata show that only a moderate faction of unemployed individuals receiving the temporary CARES Act $600 weekly UI supplement would refuse an offer to return to work at their previous wage. Under our preferred, most likely scenario, the model-based calculations yield reductions in job acceptance rates that are essentially identical to the estimated impact on job finding in our preceding regression analysis.

It is worth noting a few considerations that may have a meaningful impact on an individual's job acceptance decision in the context of our job-search model. First, our analysis does not incorporate job-related health risks during the pandemic that would reduce the value of a job offer and the corresponding value of the reservation benefit. We focus more narrowly on the financial disincentive of the supplemental UI income on the job acceptance decision, and the calculations based on the microdata suggest that the additional UI income alone was not likely to have deterred a large nmber of workers from returning to work. Second, our model does not incorporate human capital depreciation or other factors that would result in a declining job arrival rate over the duration of the unemployment spell. This consideration would act to increase the reservation benefit level, especially as individuals experience longer unemployment spells during a protracted slowdown. Finally, these are partial equilibrium exercises, which do not take into account general equilibrium effects of expanding UI policies on job offer arrival and separation rates, and are meant to be the model counterparts of the individual level estimates uncovered in the first part of the analysis. Consideration of such general equilibrium effects is left to future work.

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Online appendix

A Detailed derivations

A.1 Main derivations

Recall the Bellman equations:

$$\begin{aligned}
W_E &= w_j + \frac{1}{1+r}\left[(1-s)\,W_E + sW_U\left(b_j, T\right)\right] && \text{(A.1)}\\
W_U\left(b_j, t\right) &= b_j + \frac{1}{1+r}\left[(1-f)\,W_U(b_j, t-1) + fW_E\right] \text{ for } T \geq t > 1 && \text{(A.2)}\\
W_U\left(b_j, 1\right) &= b_j + \frac{1}{1+r}\left[(1-f)\,W_U(0) + fW_E\right] && \text{(A.3)}\\
W_U(0) &= 0 + \frac{1}{1+r}\left[(1-f)\,W_U(0) + fW_E\right] && \text{(A.4)}
\end{aligned}$$

From the last line we have $W_U(0) = \frac{f}{r+f}W_E$, then:

$$\begin{aligned}
W_U\left(b_j, 1\right) &= b_j + \frac{1}{1+r}\left[(1-f)\,\frac{f}{r+f}W_E + fW_E\right] = b_j + \frac{f}{r+f}W_E\\
W_U\left(b_j, 2\right) &= b_j + \frac{1}{1+r}\left[(1-f)\,W_U\left(b_j, 1\right) + fW_E\right]\\
&= b_j + b_j\left(\frac{1-f}{1+r}\right) + \frac{1}{1+r}\left[(1-f)\,\frac{f}{r+f} + f\right]W_E\\
&= b_j + b_j\left(\frac{1-f}{1+r}\right) + \frac{f}{r+f}W_E
\end{aligned}$$

and finally:

$$W_U\left(b_j, t\right) = \sum_{i=0}^{t-1} b_j\left(\frac{1-f}{1+r}\right)^i + \left(\frac{f}{r+f}\right)W_E$$

Let $b_j^r(t, w_j)$ denote the value of unemployment benefit with t weeks of eligibility remaining such that an individual is just indifferent between a job offer and remaining unemployed. With one week of benefits remaining:

$$\begin{aligned}
W_U\left(b_j^r(1, w_j), 1\right) &= W_E\\
b_j^r(1, w_j) + \frac{f}{r+f}W_E &= W_E\\
b_j^r(1, w_j) &= \left(\frac{r}{r+f}\right)W_E
\end{aligned}$$

With two weeks remaining:

$$
\begin{aligned}
W_U\left(b_j^r(2,w_j),2\right) &= W_E \\
b_j^r(2,w_j)\left[1+\left(\frac{1-f}{1+r}\right)\right]+\frac{f}{r+f}W_E &= W_E \\
b_j^r(2,w_j) &= \frac{\left(\frac{r}{r+f}\right)W_E}{\left[1+\left(\frac{1-f}{1+r}\right)\right]}=\frac{b^r(1,w_j)}{\left[1+\left(\frac{1-f}{1+r}\right)\right]}
\end{aligned}
$$

such that $b_j^r(2,w_j) < b_j^r(1,w_j)$. More generally: for $T > t > 1$

$$
b_j^r(t,w_j) = \frac{b_j^r(1,w_j)}{\sum_{i=0}^{t-1}(\frac{1-f}{1+r})^i}
$$

Finally, we can re-express the value of employment as:

$$
\begin{aligned}
W_E &= \frac{w_j+\frac{s}{1+r}W_U(b_j,T)}{1-(\frac{1-s}{1+r})} \\
W_E &= \left(\frac{1+r}{r+s}\right)w_j+\left(\frac{s}{r+s}\right)W_U(b_j,T)=\left(\frac{1+r}{r+s}\right)w_j+\left(\frac{s}{r+s}\right)B_j(T)+\left(\frac{s}{r+s}\right)\left(\frac{f}{r+f}\right)W_E \\
rW_E &= \frac{r+f}{r+f+s}\left[(1+r)w_j+sB_j(T)\right]
\end{aligned}
$$

such that

$$
b_j^r(1,w_j) = \frac{(1+r)\,w_j+sB_j(T)}{r+s+f}
$$

A.2 Application to the 2020 CARES Act

The value of unemployment under the CARES Act is:

$$
\begin{aligned}
W_U\left(\bar{b}_j,t_c,b_p,t_p\right) &= \bar{b}_j+b_p+\frac{1}{1+r}\left[(1-f)\,W_U(\bar{b}_j,t_c-1,b_p,t_p-1)\right. \\
&\quad \left. +f\max\left[W_E(w_j),W_U(\bar{b}_j,t_c-1,b_p,t_p-1)\right]\right] \text{ for } t_c,t_p>1 \\
W_U\left(\bar{b}_j,t_c,b_p,1\right) &= \bar{b}_j+b_p+\frac{1}{1+r}\left[(1-f)\,W_U(\bar{b}_j,t_c-1,0,0)+f\max\left[W_E(w_j),W_U(\bar{b}_j,t_c-1,0,0)\right]\right] \\
W_U\left(\bar{b}_j,t_c,0,0\right) &= \bar{b}_j+\frac{1}{1+r}\left[(1-f)\,W_U(\bar{b}_j,t_c-1,0,0)+f\max\left[W_E(w_j),W_U(\bar{b}_j,t_c-1,0,0)\right]\right] \\
W_U\left(\bar{b}_j,1,0,0\right) &= \bar{b}_j+\frac{1}{1+r}\left[(1-f)\,W_U(0)+f\max\left[W_E(w_j),W_U(0)\right]\right] \\
W_U(0) &= \frac{f}{r+f}W_E(w_j) \\
W_E(w_j) &= w_j+\frac{1}{1+r}\left[(1-s)\,W_E(w_j)+sW_U\left(\bar{b}_j,T_c\right)\right]
\end{aligned}
$$

With one week and t_c weeks of regular UI remaining and exhaustion of PUC benefits:

$$\begin{aligned} W_U\left(\bar{b}_j, 1, 0, 0\right) &= \bar{b}_j + \frac{f}{r+r} W_E(w_j) \\ W_U\left(\bar{b}_j, t_c, 0, 0\right) &= \bar{b}_j \sum_{i=0}^{t-1} \left(\frac{1-f}{1+r}\right)^i + \frac{f}{r+r} W_E(w_j) = \overline{B}_j(t_c) + \frac{f}{r+r} W_E(w_j) \end{aligned}$$

With t_c weeks of regular UI payments and one week of PUC payments:

$$\begin{aligned} W_U\left(\bar{b}_j, t_c, b_p, 1\right) &= \bar{b}_j + b_p + \frac{1}{1+r}\left[(1-f)\, W_U(\bar{b}_j, t_c - 1, 0, 0) + f W_E(w_j)\right] \\ W_U\left(\bar{b}_j, t_c, b_p, 1\right) &= \overline{B}_j(t_c) + b_p + \frac{f}{r+r} W_E(w_j) \end{aligned}$$

With t_c weeks of regular UI payments and t_p weeks of PUC payments:

$$W_U\left(\bar{b}_j, t_c, b_p, t_p\right) = \overline{B}_j(t_c) + B_p(t_p) + \frac{f}{r+r} W_E(w_j)$$

Reservation supplemental benefit with one week of PUC remaining $b_j^r(t_c, t_p = 1, w_j)$:

$$\begin{aligned} W_U(\bar{b}_j, t_c, b_p^r(1), 1) &= W_E(w_j) \\ \overline{B}_j(t_c) + b_{p,j}^r(1, t_c, w_j) + \frac{f}{r+r} W_E(w_j) &= W_E(w_j) \\ b_{p,j}^r(1, t_c, w_j) &= \frac{r}{r+f} W_E(w_j) - \overline{B}_j(t_c) \end{aligned}$$

Using similar steps as in the previous section, this can be expressed as

$$b_{p,j}^r(1, t_c, w_j) = \frac{(1+r)\, w_j + s\bar{B}_j(T_c)}{r+s+f} - \frac{b_j}{\sum_{i=0}^{t_c} \left(\frac{1-f}{1+r}\right)^i} \tag{A.5}$$

Reservation supplemental benefit with two weeks of PUC remaining $b_j^r(t_c, t_p = 2, w_j)$:

$$\begin{aligned} W_U(\bar{b}_j, t_c, b_{p,j}^r(2), 2) &= W_E(w_j) \\ \overline{B}_j(t_c) + B_p(2) + \frac{f}{r+r} W_E(w_j) &= W_E(w_j) \\ b_{p,j}^r(2, t_c) &= \frac{\frac{r}{r+f} W_E(w_j) - \overline{B}_j(t_c)}{\sum_{i=0}^{1} \left(\frac{1-f}{1+r}\right)^i} \end{aligned}$$

Reservation supplemental benefit with t weeks of PUC remaining $b_j^r(t_c, t_p = t, w_j)$:

$$W_U(\bar{b}_j, t_c, b_{p,j}^r(t), t) = W_E(w_j)$$

$$\overline{B}_j(t_c) + B_p(t) + \frac{f}{r+r} W_E(w_j) = W_E(w_j)$$

$$b^r_{p,j}(t,t_c) = \frac{\frac{r}{r+f} W_E(w_j) - \overline{B}_j(t_c)}{\sum_{i=0}^{t-1} \left(\frac{1-f}{1+r}\right)^i}$$

B Data

Unemployment duration is the inverse of the weekly job finding rate calculated by converting the monthly flow rate $f_m = UE_t/U_{t-1}$, to a weekly frequency as $f_w = 1 - (1-f_m)^{1/4}$; The duration of an employment spell is the inverse of the weekly job separation rate calculated from the monthly flow rate $s_m = (EU_t + EN_t)/E_{t-1}$, converted to a weekly rate by solving
$s = s_w \left\{ [(1-f_w) + (1-s_w)] \left(2 s_w f_w + (1-f_w)^2 + (1-s_w)^2 \right) \right\}$.

Table A1: Measures of weekly earnings, unemployment and employment duration

	Log weekly earnings dist.		Duration of: unemployment[a]			employment[b]
				Reported	Flow	Flow
	μ_w	σ_w	mean	cond. on U-E	$1/f_w$	$1/s_w$
Overall	6.58	0.73	31.7	20.5	22	1.7
Age 25 to 54 years	6.67	0.68	33.73	22.1	21	2.3
Education:						
Less then HS	6.24	0.70	28.6	18.4	23	0.74
High School	6.48	0.61	32.5	21.1	22	1.6
College and above	6.86	0.81	32.8	21.1	19	2.8
Occupation:						
Management	7.01	0.69	–	18.9	20	2.3
Prof. and tech.	6.74	0.81	–	16.9	20	2.3
Services	6.29	0.64	–	20.6	22	1.2
Sales	6.46	0.77	–	22.9	21	1.4
Admin.	6.41	0.67	–	23.9	22	1.6
Farm., fish.	6.41	0.53	–	16.7	20	1.0
Constr., extract.	6.84	0.64	–	16.4	20	1.5
Inst., maint., rep.	6.72	0.62	–	21.1	20	1.7
Production	6.68	0.53	–	17.9	21	1.5
Transp., materials	6.53	0.70	–	20.3	21	1.4

Notes: (a) weeks; (b) years. Earnings data calculated using the Jan. to Dec. 2019 CPS. Durations calculated using Dec. 2009 to Dec. 2010 CPS. w: weekly earnings; Weekly job finding f_w and separation s_w rates calculated by converting the monthly flow rates to a weekly frequency.

Job finding rates by major occupation are obtain from a logit on the outcome of a transition from unemployment into employment, $f = \exp(\beta_f X) / \left[1 + \exp(\beta_f X)\right]$, based on a set of demographic characteristics in the vector X that includes age, education, race/ethnicity, sex and marital status. The regression results are reported in Table A2.

Table A2: Predicting Finding and Separation Rates for 2010

	UE		*EU + EN*	
Age				
25-34	0.0128	-0.0539	-0.953	-0.833
	(0.0341)	(0.0362)	(0.0215)	(0.0227)
35-44	-0.0316	-0.135	-1.166	-0.976
	(0.0356)	(0.0408)	(0.0222)	(0.0257)
45-54	-0.195	-0.310	-1.274	-1.070
	(0.0363)	(0.0430)	(0.0220)	(0.0263)
55-64	-0.333	-0.460	-0.970	-0.757
	(0.0437)	(0.0504)	(0.0230)	(0.0275)
65-79	-0.468	-0.604	-0.0557	0.159
	(0.0759)	(0.0812)	(0.0268)	(0.0315)
Education				
H.S. Diploma	0.0721	0.0755	-0.536	-0.529
	(0.0336)	(0.0336)	(0.0211)	(0.0211)
Some College	0.149	0.170	-0.672	-0.672
	(0.0355)	(0.0356)	(0.0214)	(0.0215)
College Degree & Above	0.287	0.309	-1.020	-1.014
	(0.0408)	(0.0410)	(0.0236)	(0.0236)
Race/Ethnicity				
Black	-0.373	-0.343	0.408	0.356
	(0.0353)	(0.0357)	(0.0221)	(0.0224)
Hispanic	0.147	0.137	0.269	0.268
	(0.0322)	(0.0323)	(0.0209)	(0.0209)
Asian/Pacific Islander	-0.248	-0.260	0.147	0.141
	(0.0635)	(0.0637)	(0.0338)	(0.0338)
Other	-0.0771	-0.0627	0.291	0.267
	(0.0623)	(0.0624)	(0.0403)	(0.0404)
Sex				
Female		-0.169		0.0984
		(0.0238)		(0.0141)
Marital Status				
Married (Spouse Absent)		0.243		0.221
		(0.0866)		(0.0572)
Widowed		-0.0420		0.109
		(0.0962)		(0.0465)
Divorced		-0.133		0.0810
		(0.0393)		(0.0254)
Separated		0.00183		0.213
		(0.0669)		(0.0477)
Never Married		-0.185		0.291
		(0.0323)		(0.0195)
Constant	-1.540	-1.314	-1.761	-2.070
	(0.0323)	(0.0450)	(0.0210)	(0.0277)
Observations	52442	52442	536849	536849

Note: Groups "16-24", "Less than H.S. Diploma", "White", "Male", and "Married (Spouse Present)" are included as reference categories, respectively.

C Additional tables

Table A3: Reservation benefits, replacement rates, and offer rejection rates

	Earnings	Duration of:		Weekly UI compensation					Rejection Rate (%)		
	w (wkly)	U (wks)	E (yrs)	$\bar{b}$	b^C	$b^r(12)$	$b^r(8)$	$b^r(4)$	12 wks	8 wks	4 wks
Overall	717	22	1.7	359	959	1105	1419	2373	38.3	21.7	5.8
Age 25 to 54 years	792	21	2.3	396	996	1245	1598	2668	30.3	15.2	2.4
Education:											
Less then HS	514	23	0.74	257	857	708	908	1514	65.8	45.4	14.6
High School	650	22	1.6	325	925	1003	1290	2161	42.0	21.8	3.3
College and above	955	19	2.8	477	1077	1461	1853	3060	27.0	15.1	3.0
Occupation:											
Management	1103	20	2.3	500	1100	1712	2196	3672	17.7	7.7	1.0
Prof. and tech.	850	20	2.3	425	1025	1287	1639	2702	32.6	19.2	5.0
Services	541	22	1.2	271	871	801	1026	1714	57.8	35.9	8.4
Sales	637	21	1.4	318	918	944	1206	2002	47.8	30.5	8.9
Admin.	608	22	1.6	304	904	942	1215	2038	46.3	26.8	5.7
Farm., fish.	608	20	1.0	304	904	854	1084	1785	56.5	30.9	4.2
Constr., extract.	929	20	1.5	465	1065	1355	1717	2828	26.8	12.4	1.6
Inst., maint., rep.	831	20	1.7	415	1015	1237	1572	2597	30.7	14.8	1.9
Production	798	21	1.5	399	999	1207	1545	2580	28.7	11.0	0.7
Transp., materials	688	21	1.4	344	644	1026	1314	2186	42.7	24.5	5.3

Notes: Earnings data calculated using the 2019 CPS ASEC for indidivuals unemployed in any of the months of March through July 2020. Durations of unemployment and employment in columns 2 and 3 are calculated using the Dec. 2009 to Dec. 2010 CPS. w: median weekly earnings. Weekly job finding and separation rates entering the resevation benefits are obtained by converting the monthly flow rates to a weekly frequency (see appendix for details); $\bar{b}$: median regular weekly unempmloyment benefits; b^C: median weekly benefits under CARES act, $\bar{b} + 600\$$; $b^r(t_p)$ median reservation benefit level with t_p weeks left under the CARES act.

Table A4: Reservation benefits, replacement rates, and offer rejection rates - stronger recovery scenario

	(1) Earnings	(2) Duration of:	(3)	(4) Weekly UI compensation	(5)	(6)	(7) Rejection rate (%)
	w (wkly)	U (wks)	E (yrs)	$\bar{b}$	b^C	$b^r(8)$	Average[a]
Overall	717	22	1.7	359	959	1141	34.2
Age 25 to 54 years	792	21	2.3	396	996	1291	26.9
Education:							
Less then HS	514	23	0.74	257	857	780	54.3
High School	650	22	1.6	325	925	1028	37.7
College and above	955	19	2.8	477	1077	1485	25.1
Occupation:							
Management	1103	20	2.3	500	1100	1700	18.1
Prof. and tech.	850	20	2.3	425	1025	1262	32.5
Services	541	22	1.2	271	871	811	52.9
Sales	637	21	1.4	318	918	953	44.6
Admin.	608	22	1.6	304	904	941	43.8
Farm., fish.	608	20	1.0	304	904	842	53.8
Constr., extract.	929	20	1.5	465	1065	1354	26.0
Inst., maint., rep.	831	20	1.7	415	1015	1252	28.8
Production	798	21	1.5	399	999	1239	25.8
Transp., materials	688	21	1.4	344	644	1044	39.0

Earnings data calculated using the 2019 CPS ASEC for indidivuals unemployed in any of the months of March through July 2020. Durations of unemployment and employment in columns 2 and 3 are calculated using the Dec. 2014 to Dec. 2015 CPS. w: median weekly earnings. Weekly job finding and separation rates entering the resevation benefits are obtained by converting the monthly flow rates to a weekly frequency (see appendix for details); $\bar{b}$: median regular weekly unempmloyment benefits; b^C: median weekly benefits under CARES act, $\bar{b} + 600\$$; $b^r(t_p)$ median reservation benefit level with t_p weeks left under the CARES act. a: Average rate of rejecting an offer arriving between 12 and 4 weeks of PUC payments remaining.