Can the Unemployed Borrow? Implications for Public Insurance

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We empirically establish that unemployed individuals maintain significant access to credit and that upon a layoff, the unconstrained borrow while the constrained default and delever. Motivated by these findings, we develop a theory of credit lines and labor income risk to analyze optimal transfers to the unemployed. Since credit lines offer fixed interest rates and limits, credit lines are unresponsive to layoffs and provide greater consumption insurance relative to when debt is repriced period by period. At US levels of credit lines, the government can optimally reduce transfers to the unemployed, whereas this is not true when debt is counterfactually repriced period by period.

In the 2019 Survey of Consumer Finances (SCF), more than 39% of respondents reported revolving their credit card balances from month to

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month, and over 70% of respondents reported access to credit cards.[1] By the first quarter of 2022, aggregate credit card limits totaled 17% of GDP (gross domestic product).[2] In this paper, we explore how the prevalence of credit cards in the United States affects optimal public insurance provision. In particular, we ask whether there is scope to substitute away from public insurance and to rely more on private self-insurance through credit markets.

To answer this question, we first establish that credit limits are unresponsive to job loss and that a significant share of individuals borrow or default during job loss. We then develop a tractable theory of labor income risk in which lenders issue long-term credit contracts with fixed interest rates and limits. We refer to these contracts as "credit lines," and we show that by modeling credit lines, our theory is capable of matching our new set of facts. We find that optimal transfers to the unemployed—expressed as a replacement rate of lost income—are 6.6 percentage points (pp) lower in an economy with US levels of credit lines than in a counterfactual economy with no credit market. Importantly, the degree of substitutability between public transfers and private credit hinges on the availability of credit lines. In an economy in which credit lines do not exist and debt is repriced each period, consumption upon job loss is more sensitive to public transfers, implying less substitutability between public transfers and private credit.

Our empirical contribution is to measure workers' borrowing behavior and borrowing ability upon job loss. Using newly linked administrative earnings and credit bureau data, we document four facts that suggest that credit markets play an important role in the way workers self-insure: (1) before displacement, workers who lose their jobs can replace a significant fraction of their prior income with unused credit; (2) credit limits and credit scores do not immediately respond to job loss and do not decline

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[1] These statistics correspond, respectively, to the weighted fraction of 2019 SCF respondents with positive values for variable X413, "After the last payment(s) (was/were) made, what was the total balance still owed on (this account/all these accounts)?" and positive values for variable X411, "How many [Visa, MasterCard, Discover, or American Express] cards do you have?"

[2] This is based on the Federal Reserve Bank of New York's Consumer Credit Panel and the Bureau of Economic Analysis for 2022-I. Including home equity lines of credit, credit limits total 20% of GDP.

in an economically significant manner within 5 years after job loss; (3) unconstrained individuals, those with unused-credit limits in the top two quintiles before job loss, borrow and replace a significant fraction of lost earnings with credit; and (4) constrained individuals, who have unused-credit limits in the bottom two quintiles before job loss, default and delever. Both borrowing and defaulting allow individuals to transfer resources across time and states of the world, allowing unemployed individuals to partially self-insure their losses.

Our theoretical contribution is to develop a tractable model of credit lines and labor income risk capable of matching these new empirical facts. To generate the credit access and borrowing patterns we observe in the data, our theory relies on two features of the US credit market: (i) the credit registry generates reputation concerns in the form of exclusion from credit markets in the event of default, and (ii) lenders issue long-term contracts in the form of revolving lines of credit, such as credit cards and home equity lines of credit (HELOCs), whose limits and interest rates are not contingent on subsequent income changes. Because the unemployed value future access to credit markets, most workers upon job loss repay, and therefore lenders offer credit contracts to individuals both before and after job loss. Conversely, in a model without credit lines, where debt is individually priced each period, unemployed agents would face a sudden change in borrowing capacity, which is inconsistent with the facts we establish. Finally, we render the credit market tractable by incorporating directed search for credit lines (e.g., Moen 1997; Burdett, Shi, and Wright 2001; Menzio and Shi 2011). We demonstrate that our framework can incorporate rich worker heterogeneity, and we argue that the model is fungible to other contexts, including corporate and sovereign settings.

After estimating our framework to match aggregate credit access and borrowing moments in the early 2000s, we show that our model successfully replicates the nontargeted responses of borrowing, credit limits, and defaults upon job loss. Like the data, the model economy's borrowing limits do not exhibit an economically meaningful response to job loss. Additionally, as in the data, the model generates heterogeneity in borrowing and defaults following job loss. We estimate the same reduced-form empirical specifications on our model-simulated data and show that the model successfully captures the cross-sectional heterogeneity of borrowing and default rates present in the data. We show that in the cross section, upon job loss, the model simultaneously generates (1) deleveraging and defaults among constrained workers and (2) borrowing and repayment among unconstrained workers. Both groups of individuals, borrowers and defaulters, smooth consumption using credit markets. In particular, when individuals borrow they pay a premium in the form of a spread over the risk-free rate, reflecting default risk. In bad states of the world, such as when a borrower loses their job, they may default to smooth consumption.

Using the calibrated model, our quantitative contribution is to measure the extent to which the government can substitute away from public transfers to the unemployed, given current US credit levels. We answer this question by computing optimal transfers to the unemployed in the baseline economy and in an economy with zero credit, where we express the optimal transfers as a replacement rate of lost earnings during unemployment. The difference in the optimal replacement rate across the two economies indicates the degree to which credit markets allow the government to substitute away from public insurance for the unemployed.

We evaluate policies using the welfare of newborn agents. In our model, newborn agents draw both human capital and their degree of patience (i.e., discount factor) upon entering the workforce. We focus on newborn welfare "behind the veil of ignorance," before human capital and patience are realized, and we assume that transfers to the unemployed are funded by distortionary labor income taxes. This generates a trade-off for the government: greater transfers to the unemployed mitigate consumption losses after a layoff but require more distortionary taxes. Importantly, this trade-off depends on the prevalence of credit lines.

As transfers are cut, job losers can use credit lines to smooth consumption, dampening the costs of such a policy. Given credit access observed between 2002 and 2012, the utilitarian government's optimal steady-state replacement rate is 34.8%. When credit markets are counterfactually shut down, the optimal steady-state replacement rate is 41.4%. This implies a 6.6-pp difference in optimal replacement rates between our benchmark 2002–12 economy, in which 78% of individuals have credit access, and one in which none do.

What is surprising about our findings is that policy makers might naturally think that extensive access to private credit lines might allow the government to significantly cut unemployment insurance (UI). Instead, our model yields several general equilibrium forces that limit the desire of the government to substitute out of public insurance. For low levels of public insurance, default rates rise, leading to higher interest rates, less borrowing, and lower credit-finding rates. As a result, consumption losses upon a layoff become more severe as credit becomes more expensive, limiting the government's willingness to cut public transfers.

We demonstrate the importance of credit lines by reestimating the substitutability between public transfers and private credit in an economy with 1-period debt contracts (e.g., see Chatterjee et al. 2007, Livshits, MacGee, and Tertilt 2007, and ensuing literature). When debt is repriced each period, we find a 0.5-pp difference between our 1-period-debt economy and one with zero credit. What drives this lesser substitutability is that with 1-period debt, the consumption of the unemployed is more responsive to transfers. The reason why credit lines provide more insurance relative to 1-period debt is that long-term credit lines are established (most often) when an individual is employed. These credit lines do not respond to income changes—or transfer changes—as much as 1-period debt. We demonstrate this by showing that borrowing limits decline substantially upon job loss in the 1-period-debt economy, while limits are stable in the credit-line economy (consistent with our empirical facts). Consequently, in an economy with 1-period debt, consumption upon job loss falls by more for any given reduction in transfers, resulting in less substitutability between public transfers and private credit.

It is important to note that even with credit lines, the welfare gains from reoptimizing public transfers are economically small. Across steady states, cutting replacement rates from the current US policy of 41.2% to 34.8% yields a welfare gain worth 0.01% of a newborn's lifetime consumption. This subjectively small welfare gain reflects offsetting gains and losses across households with differing human capital and patience levels. We find that patient agents—who save and rely less on transfers—gain moderately from this policy, while impatient agents—who borrow and rely heavily on transfers—lose significantly. The net gains and losses across these two groups are positive but approximately offsetting. We find similar magnitudes of welfare gains along the transition path.

As we discuss in the conclusion, the presence of business and credit cycles may further erode the ability of the government to substitute between public insurance and private credit. In this regard, we view our estimates of the substitutability between public insurance and private credit as an upper bound.

Related literature.—Our empirical results reconcile two literatures with seemingly conflicting results. Studies based on checking-account data suggest that there is roughly zero net borrowing, on average, by workers who lose their jobs (e.g., Ganong and Noel 2019; Gelman et al. 2020). On the other hand, direct questions about borrowing among workers who lose their jobs and other survey data imply that roughly 20% of the unemployed borrow and that roughly 30% become delinquent on debt obligations (e.g., Sullivan 2008; Hurd and Rohwedder 2010; Gerardi et al. 2018).[3] We reconcile these results by showing that upon job loss some workers borrow while other workers default and delever. While these offsetting forces yield zero net borrowing by the unemployed, both the borrowers and the defaulters are using credit to smooth consumption.

[3] Papers that show ex post borrowing following job loss include Sullivan (2008), Hurd and Rohwedder (2010), Collins, Edwards, and Schmeiser (2015), and Herkenhoff (2019). Papers that show ex post default include Hurd and Rohwedder (2010), Gerardi et al. (2018), Keys (2018), and Herkenhoff, Phillips, and Cohen-Cole (2023). Surveys of bankruptcy also cite job loss as a factor (e.g., Sullivan, Warren, and Westbrook 1999). Finally, Baker and Yannelis (2017) illustrate significant differences in consumption losses between constrained and unconstrained individuals (see also Crossley and Low 2014).

Our paper contributes to recent work that has integrated credit markets into models with labor markets (e.g., Athreya and Simpson 2006; Athreya et al. 2015; Bethune, Rocheteau, and Rupert 2015; Luo and Mongey 2016; Bethune 2017; Herkenhoff 2019; Ji 2021). The most closely related paper is by Athreya and Simpson (2006), who compute the responsiveness of bankruptcies to public insurance provision, showing that more generous UI may actually raise bankruptcies. We build on Athreya and Simpson (2006) in three key ways: (1) we model long-term credit contracts, which allows us to match the degree of self-insurance provided by the credit market; (2) we model the labor market in general equilibrium; and (3) we calculate the optimal provision of public insurance. We also note that Athreya, Tam, and Young (2009) show that with 1-period debt, income risk transmits fully into consumption risk, implying limited insurance from consumer credit markets. Our results complement Athreya, Tam, and Young (2009) by showing that long-term credit lines allow job losers to partially smooth idiosyncratic job loss shocks.

Our model adds to a small but growing literature on individual credit lines, credit scoring, and long-term relationships between borrowers and lenders.[4] Of particular note, work by Mateos-Planas and Ríos-Rull (2010) analyzes bankruptcy reform in an economy with credit lines and private information about endowments. We depart from Mateos-Planas and Ríos-Rull (2010) by modeling the labor market, and we obtain tractability via competitive search over credit contracts.

Our paper is related to studies that integrate UI into Bewley-Huggett-Aiyagari frameworks (e.g., Lentz and Tranæs 2005; Krusell, Mukoyama, and Şahin 2010; Nakajima 2012a, 2012b) as well as studies of optimal UI with assets (inter alia, Chetty 2008; Shimer and Werning 2008; Lentz 2009; Koehne and Kuhn 2015; Griffy 2021; Chaumont and Shi 2022).[5] Related papers by Shimer and Werning (2008) and Lentz (2009) compute optimal UI in models with savings. Relative to these studies, we make several contributions: (i) we empirically document the large income-replacement or self-insurance role that credit markets play in the US economy, (ii) we incorporate the institutions that allow this self-insurance to exist in our model (long-term contracts, reputation concerns, and defaultable debt), and (iii) we quantify the substitutability between private borrowing and public forms of insurance.

[4] See Mateos-Planas and Seccia (2006), Mateos-Planas and Ríos-Rull (2010), and Mateos-Planas (2013) on models of credit lines; Chatterjee, Corbae, and Ríos-Rull (2008), Chen and Zhao (2017), and Chatterjee et al. (2020) on models of credit scoring; and Corbae and Quintin (2015), Hedlund (2016), and Mitman (2016) for models of long-term relationships between borrowers and lenders.

[5] Our paper also complements studies on optimal UI over the business cycle (Mitman and Rabinovich 2015, Birinci and See 2017, and references therein).

Finally, our article is also related to the literature on private UI (e.g., Chiu and Karni 1998; Hendren 2017). We contribute to this literature in two ways: (i) we focus on private self-insurance or income replacement through credit markets, and (ii) we include reputation concerns and long-run interactions between credit and UI. While the two papers take very different approaches to the question of how substitutable private and public forms of insurance are, our results are consistent with Hendren (2017), in the sense that the scope for private self-insurance is limited, even with long-term contracts and strong dynamic reputation concerns.

The paper proceeds as follows. Section I describes our main empirical results, section II describes the model, section III describes the calibration, section IV computes optimal transfers to the unemployed, and section V concludes.

I. Empirical Results Using Administrative Data

To examine the degree to which the government can substitute away from public insurance because of credit markets, we first measure the degree of insurance provided by credit markets under current government policy, using a new database of administrative earnings records that have been linked to individual credit reports.

A. Data

Our main dataset is a randomly drawn panel of roughly 5 million TransUnion credit reports linked through a scrambled Social Security number to the Longitudinal Employer–Household Dynamics (LEHD) administrative records database. The TransUnion database contains information on the balance, credit score, limit, and status (delinquent, current, etc.) across different types of consumer debt held by individuals at an annual frequency from 2002 through 2012.[6] The LEHD database is a matched employer-employee dataset covering 95% of US private sector jobs. The LEHD includes quarterly data on earnings, worker demographic characteristics, firm size, firm age, and average wages. Our primary sample of employment records includes individuals with credit reports between 2002 and 2012 from the 11 states for which we have LEHD data.[7]

Since job dismissal and reason of dismissal are not recorded in the LEHD, we identify layoffs using mass-layoff episodes in the spirit of Jacobson, LaLonde, and Sullivan (1993). We define a *mass layoff* to have occurred

[6] Our underlying sample is comprised of a random sample as well as an oversample of bankruptcies, foreclosures, and delinquencies. We reweight our combined sample to match the aggregate bankruptcy, foreclosure, and delinquency rates in the relevant states.

[7] The 11 states for which we have LEHD data are Arizona, California, Colorado, Delaware, Iowa, Illinois, Indiana, Maryland, Nevada, Virginia, and Washington.

when a firm with at least 20 employees decreases its employment by at least 20% between two consecutive quarters.[8]

In terms of credit report data, our analysis focuses on revolving credit because it can be drawn down immediately after job loss, with no additional application or income verification, and it can be repaid slowly. The main components of revolving credit include bank (bank credit cards), retail (retail credit cards), finance (other personal finance loans with a revolving feature), and mortgage related (HELOCs).[9] We also study the response of default activity, as measured through debt charge-offs, foreclosures, bankruptcies, and derogatory public flags.

B. Sample Descriptions and Summary Statistics

We split the sample of workers at a firm undergoing a mass-layoff episode into two subsamples.[10]

1. *Panel sample*. Our first sample includes 24–64-year-olds who were at a firm that underwent a mass-layoff episode, had at least 3 years of tenure at the time of the mass layoff, and made at least $1,000 in each quarter at the firm in the prior year.[11] Since individuals may move to states outside of our sample, we require individuals to return to our sample with positive earnings before 2012. We split this sample into a treatment group and a control group. Our treatment group includes 92,000 individuals who were displaced as part of the mass layoff. Our control group includes 126,000 individuals who were coworkers of those in the treatment group during the mass layoff but were not displaced. If an individual is involved in two or more mass layoffs, we use only the first event, and we require those in the control group to never be displaced as part of a mass-layoff episode.
2. *Cross-sectional sample*. Our second sample includes 56,000 displaced workers in the treatment group who had a decline in annual earnings

[8] When defining layoffs, researchers face a trade-off: imposing stricter conditions reduces noise by isolating true layoffs, but the cost is weaker external validity. We defend external validity in app. A.1 (apps. A–G are available online) by showing that we obtain similar results using a sample of displaced and nondisplaced workers with a looser layoff requirement, which we refer to as the generic layoff sample.

[9] Appendix A.2 includes an analysis of bank cards (e.g., credit cards) that exhibits patterns similar to revolving credit. However, it is important to note that not all types of credit balances affect the budget constraint in the same way. A first mortgage lowers liquid resources on hand (buying a house involves handing money to the bank), whereas an increase in revolving debt augments liquid resources on hand.

[10] All sample sizes are rounded to the nearest thousand, in compliance with Census Bureau disclosure rules.

[11] These restrictions on tenure and prior earnings are common in the literature (e.g., Davis and von Wachter 2011) and are used to mitigate issues associated with seasonal employment or weak labor force attachment.

from the year before to the year after layoff, and we require them to have a positive credit limit.[12]

Table 1 includes summary statistics for both samples. Panel A of table 1 provides summary statistics for the treatment and control groups in the panel sample in the year before the layoff event. Annual earnings, as well as credit limits and balances, are deflated by the consumer price index. Column 1 of table 1 summarizes the treatment group, while column 2 summarizes the control group. The treatment group earned \$51,000 in the year before displacement, while the control group earned \$53,000. In the empirical analysis, we include individual fixed effects, controls for age, and proxies for wealth to account for differences across treatment and control groups.

Individuals have substantial revolving-credit limits in the year before job loss, with an average of nearly \$30,000 for the treatment group. Individuals in the treatment group can replace, on average, 39% of their income with unused revolving debt in the year before job loss.[13] It is important to note that the distribution of unused credit is highly skewed. As we report in section III, median ratio of unused revolving credit to income is 8.2%.

Panel B of table 1 includes summary statistics for the cross-sectional sample in the year before mass layoff. In the analysis that follows, we define credit constraints using an individual's share of unused revolving credit ((revolving limit − revolving balance)/revolving limit), which is equivalent to 1 minus the revolving-credit utilization rate.[14] From this point forward, we refer to the unused revolving-credit share as simply "unused credit." Table 1 shows that in the year before mass layoff, the majority of individuals have substantial unused credit. Individuals in the highest unused-credit quintile have their entire credit limit available to draw down, while individuals in the third quintile have over half of their credit limit available to draw down.[15]

The summary statistics of table 1 indicate that individuals have, on average, a large stock of credit before layoff. We next examine how access to and use of credit evolves following job loss.

[12] Since we stratify this sample by the share of credit that is unused (unused revolving credit divided by revolving-credit limit), we must require these individuals to have a nonzero revolving-credit limit in the year before layoff. In a prior draft, we included these individuals in the sample by stratifying our analysis by credit scores, and we found broadly similar results.

[13] Note that the ratio of unused revolving credit to income is winsorized at the 1% level at the top and bottom of the distribution.

[14] Let L denote the limit and B denote the balance. We define the share of unused revolving credit as $(L - B)/L = 1 - B/L$, where B/L is the utilization rate.

[15] Across quintiles individuals have substantial revolving-credit limits. On average, individuals in the first quintile have limits of over \$18,000, while in the fifth quintile, limits are over \$33,000.

TABLE 1
Summary Statistics

	A. Panel Sample (Year before Mass Layoff)		B. Cross-Sectional Sample (Year before Mass Layoff)
	Treatment (1)	Control (2)	Unused Revolving-Credit Share (3)
Annual earnings ($)	51,340	52,710	
Age (years)	40.7	42.2	
Revolving-credit balance ($)	11,300	11,890	
Revolving-credit limit ($)	29,780	33,330	
Unused revolving credit/income	.394	.491	
Observations	92,000	126,000	
Unused-credit quintile 1			−.0027
Unused-credit quintile 2			.3113
Unused-credit quintile 3			.5773
Unused-credit quintile 4			.8313
Unused-credit quintile 5			.9833

Note.—Sample selection criteria are in sec. I.B. Annual earnings, revolving-credit balance, and revolving-credit limit are in 2008 dollars. “Unused revolving-credit share” is defined as 1 minus the utilization rate (limit − balance)/limit. “Unused-credit quintile 1” is the average unused revolving-credit share among those between the 1st and 20th percentiles. The remaining quintiles are defined similarly.

C. *Average Response of Earnings and Credit Following Job Loss*

To gauge how credit access and usage evolve around job loss, we first estimate the average response of credit variables following job loss, using a distributed lag framework as in Jacobson, LaLonde, and Sullivan (1993) around mass-layoff episodes. This empirical strategy compares displaced to nondisplaced individuals before and after the layoff episode to identify how individuals’ access to and use of credit evolves following job loss.

Let i index individuals and t index years. Let α_i denote a set of individual fixed effects and γ_t denote year dummies. Let $Y_{i,t}$ denote the outcome of interest (such as real earnings, real revolving-debt balance, etc.). Let $D_{x,i,t}$ be a dummy variable taking the value 1 when an individual is x periods before (if x is negative) or after (if x is positive) displacement. For example, $D_{-1,i,t}$ is a dummy variable indicating that an individual is 1 period before displacement. The vector $X_{i,t}$ contains control variables, including a quadratic in age, and deciles for lagged cumulative earnings. We include deciles for lagged cumulative earnings to proxy for an individual’s wealth before displacement.[16] The specification we use is of the following form:

[16] Since states enter the LEHD at different times, these deciles are computed within a state.

$$Y_{i,t} = \alpha_i + \gamma_t + \sum_{j=-4}^{5} \beta_j D_{j,i,t} + \Gamma X_{i,t} + \varepsilon_{i,t}. \quad (1)$$

The objects of interest are $\beta_0, \beta_1, \dots, \beta_5$, which summarize the impact of job loss on the outcome variable in the year of displacement and subsequent years. To examine the validity of the point estimates, we test that the treatment and control groups have parallel trends before displacement.

Figure 1 plots the coefficient estimates from the estimation of equation (1), along with 95% confidence intervals.[17] The coefficients in figure 1 correspond to $(\beta_{-4}, \beta_{-3}, \dots, \beta_4, \beta_5)$ in equation (1) and are interpreted as the difference in the outcome variable between displaced and nondisplaced individuals.

We first examine how earnings evolve around job loss. Figure 1*A* plots the differences in real annual earnings between displaced and nondisplaced individuals. The figure shows that earnings losses following job loss are large and persistent. In the year after job loss, a displaced individual makes over $9,000 less than a nondisplaced individual, representing a decline of over 17.6% relative to prelayoff earnings. Five years after job loss, a displaced individual still earns over $4,000 less than a nondisplaced individual, representing almost a 9% decline relative to prelayoff earnings. These large and persistent declines in earnings following job loss align with estimates from Jacobson, LaLonde, and Sullivan (1993), Davis and von Wachter (2011), Huckfeldt (2022), and Jarosch (2023).[18]

We next examine how an individual's access to credit evolves following job loss. Figure 1*B* demonstrates that, despite the decline in earnings, credit limits are largely unresponsive to job loss. One year after displacement, a displaced individual's credit limit decreases relative to a nondisplaced individual's by $187, on average. In the year before displacement, individuals in the treatment group had, on average, a revolving-credit limit of nearly $30,000. Thus, credit limits decline by 0.6% following job loss, a magnitude that we view as economically insignificant. Credit limits remain statistically indistinguishable from those of the control group 5 years after job loss, suggesting that laid-off individuals maintain substantial lines of credit.

We find similar results for a conceptually distinct measure of credit access, the credit score. While credit limits reflect the stock of existing credit, credit scores reflect the marginal cost of acquiring new credit. To ease

[17] Table A6 (tables A1–A21 are available online) reports the results of estimating eq. (1). Additionally, in app. A.5 we present the raw average of the outcome variables of interest for both the treatment and control groups.

[18] We note that the increase in earnings of the treatment group relative to the control group before displacement is also observed in Davis and von Wachter (2011) and Jarosch (2023).

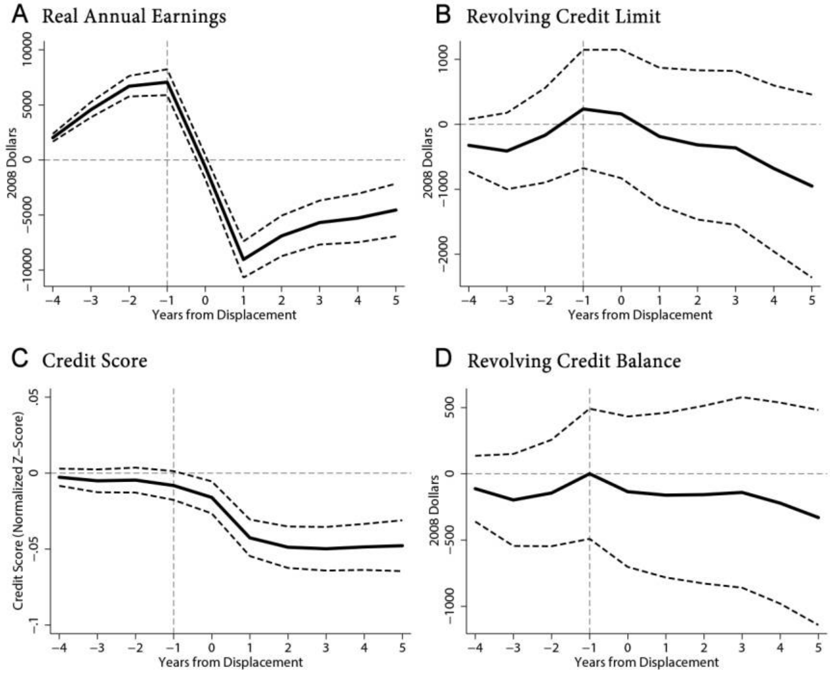

Fig. 1.—Average response of earnings and credit variables to displacement. Solid lines represent the difference in the outcome variable between displaced and nondisplaced individuals. Dashed lines represent 95% confidence intervals. Panels present coefficient estimates from table A6.

the interpretation of credit scores, we normalize the credit score to have a mean of zero and a standard deviation of 1 (i.e., a *Z*-score). Figure 1*C* shows that credit scores decline by less than 5% of a standard deviation in the year after a layoff. This very small decline suggests that the marginal cost of acquiring new credit does not decline in an economically meaningful way for workers upon job loss. The very small decline potentially reflects defaults, which we investigate in the next section.

We next examine the degree to which individuals borrow following job loss. We focus on revolving credit because it can be drawn down upon job loss without notice or further income verification. Figure 1*D* shows that, on average, displaced individuals do not borrow more than nondisplaced individuals. This zero response of borrowing following job loss is consistent with the recent work of Ganong and Noel (2019) and Gelman et al. (2020).[19] However, the cross-sectional analysis in section I.D reveals

[19] The results presented in fig. 1 and table A6 include all types of revolving credit (HELOCs, etc.) rather than just credit cards. In app. A.2, we present results for credit card (bank card) balances as well as limits. The pattern of the results for credit card balances is nearly identical to those for revolving balances.

that there is significant heterogeneity among workers who lose their jobs, as nearly one-third of laid-off workers default and/or delever, while another third of individuals borrow.

Default following job loss.—For individuals who cannot borrow, defaulting on scheduled debt repayments provides similar consumption-smoothing benefits. When a lender and a borrower enter into a debt contract, both sides know that there is potential for the borrower to not repay the loan. Lenders price contracts accordingly by charging a premium over the risk-free rate, and in bad states of the world, an indebted individual may default to self-insure. Figure 2 documents the propensity of individuals to default following job loss.[20]

We first examine how job loss affects debt charge-offs. A debt charge-off occurs when (1) an individual has skipped payments for a sufficient amount of time (typically 6 months) and (2) the creditor ceases collections, notifies the credit bureau to charge off the debt, and then potentially sells the account to a third-party collection agency. Figure 2*A* shows that in the year of job loss, the probability that a displaced individual has a debt charge-off is nearly 0.5 pp higher than that for a nondisplaced individual. One year after displacement, the difference is nearly 2 pp, which represents an increase of more than 15% in the flow rate of entry into charge-off relative to the year before a layoff. This result indicates that following job loss, individuals are skipping debt payments for upward of 6 months as a means to smooth consumption.

We see similarly elevated foreclosure, bankruptcy, and public derogatory flag rates around job loss. Figure 2*B* plots the effect of job loss on entering foreclosure within the past year. In the year after job loss, displaced individuals are 0.29 pp more likely to enter foreclosure relative to nondisplaced individuals. This represents a 33% increase compared to the prelayoff mean foreclosure rate of 0.9% per annum.

Figure 2*C* illustrates the effects of job loss on entering bankruptcy within the past year. In the year after job loss, the probability of entering bankruptcy increases by 0.13 pp. This represents an increase of more than 15% in the flow rate of entry into bankruptcy relative to the year before job loss. Individuals appear to be combining informal default through charge-offs with formal bankruptcy proceedings.

Finally, figure 2*D* illustrates the effects of job loss on new derogatory public flags within the past year. Derogatory public flags aggregate all relevant delinquency information: bankruptcy, foreclosure, tax liens, civil court judgments, and so on. We find that individuals are over 0.5 pp more likely to have a new derogatory public flag in the year after job loss. This

[20] Table A7 reports the results of estimating eq. (1) when the dependent variables are measures of default.

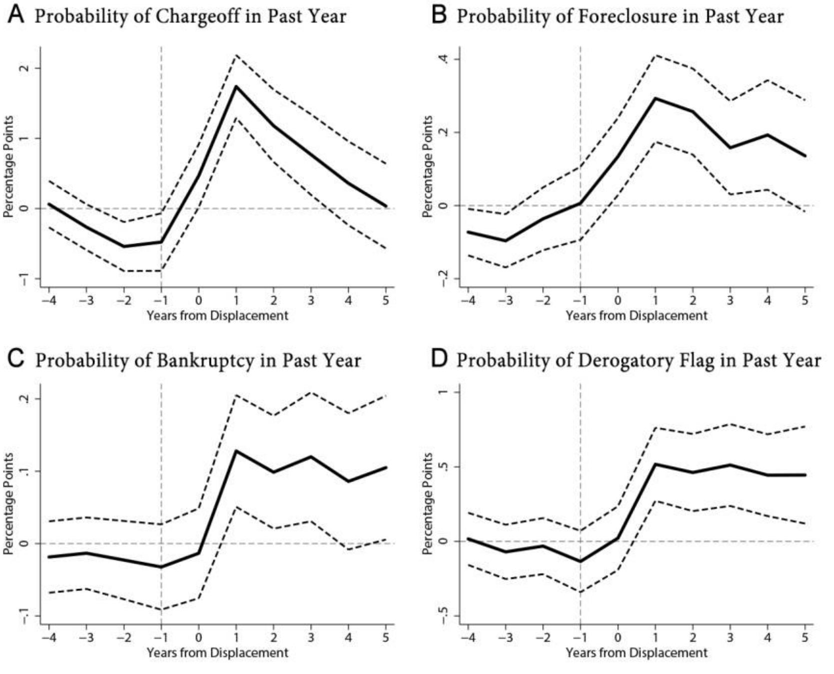

FIG. 2.—Average response of default variables to displacement. Solid lines represent the difference in the outcome variable between displaced and nondisplaced individuals. Dashed lines represent 95% confidence intervals. Panels present coefficient estimates from table A7.

represents a 17% increase in the flow rate of entry into a public derogatory flag relative to the year before job loss. Because of the way debt discharge is modeled in our framework—an endogenous exclusion period, with the ability to obtain credit access quickly after default—we view the model's notion of default as synonymous with derogatory public flags that reflect both the formal and informal default channels.

The results presented in figure 2 indicate that individuals miss debt repayments and default in response to job loss. A striking feature of these results is their persistence. Three years after job loss, individuals remain significantly more likely to have a new charge-off, foreclosure, bankruptcy, or derogatory public flag. The results in this section show that although individuals do not borrow on average, credit markets play a central role in their response to unemployment through the use of defaults. In the next section, we show that while there is zero borrowing on average, this result masks substantial heterogeneity in borrowing behavior following job loss.

D. *Heterogeneous Responses: Borrowing and Default*

In this section, we examine heterogeneous borrowing and default responses to job loss. Our primary metric for borrowing is the revolving-credit replacement rate (we refer to this as the “replacement rate” in this section). The *replacement rate* is the ratio of the change in an individual’s revolving-debt balance to the change in their earnings, where we measure the change in revolving-debt balance and earnings from the year before displacement to the year after displacement ($\mathrm{RR}_{it} = (\mathrm{debt}_{i,t+1} - \mathrm{debt}_{i,t-1})/-(\mathrm{earnings}_{i,t+1} - \mathrm{earnings}_{i,t-1}))$.[21] For the replacement rate to be well defined, we base our analysis on the cross-sectional sample, which isolates job losers with an earnings loss from $t-1$ to $t+1$.

Our theory, which we present in section II, as well as existing theories, predicts that credit constraints are an important determinant of the borrowing decision. To proxy for credit constraints, we separate individuals into unused-credit quintiles on the basis of their fraction of unused revolving-credit limits in the year before displacement ((revolving limit − revolving balance)/revolving limit).[22] Individuals in the first quintile have the lowest amount of unused credit, while individuals in the fifth quintile have the greatest amount of unused credit. Let $C_{y,i,t-1}$ be a dummy variable taking the value 1 when individual i is in unused-credit quintile y in year $t-1$ and will be displaced in year t. For example, $C_{3,i,t-1}$ is a dummy variable indicating that an individual is in the third unused-credit quintile 1 year before being displaced in year t.

In figure 3, we plot the kernel density of replacement rates in the cross-sectional sample (where 0.1 denotes a 10% replacement rate). We find that roughly one-third of workers who lose their jobs borrow, one-third delever or default, and roughly one-third do not alter their borrowing patterns. In appendix A.3, we also show that simple comparisons across unused-credit quintiles reveal significantly higher replacement rates among those in the fifth unused-credit quintile (the unconstrained) versus those in the first unused-credit quintile (the constrained). However, simple comparisons across unused-credit quintiles may capture selection into unused-credit quintiles on the basis of unobservables. We therefore adopt an empirical specification inspired by Sullivan (2008), and we exploit variation across the magnitude of earnings losses within each unused-credit quintile to gauge the heterogeneous use of credit in response to job loss.

[21] We measure the change in earnings and revolving-debt balances over a 2-year window because fig. 1*A* shows that the decline in earnings due to job loss is concentrated in the year after displacement. An earlier draft used a 1-year window (comparing t to $t-1$) and found similar results.

[22] Our requirement that those in the cross-sectional sample have positive credit limits implies that the unused-credit share is well defined for all individuals in our analysis.

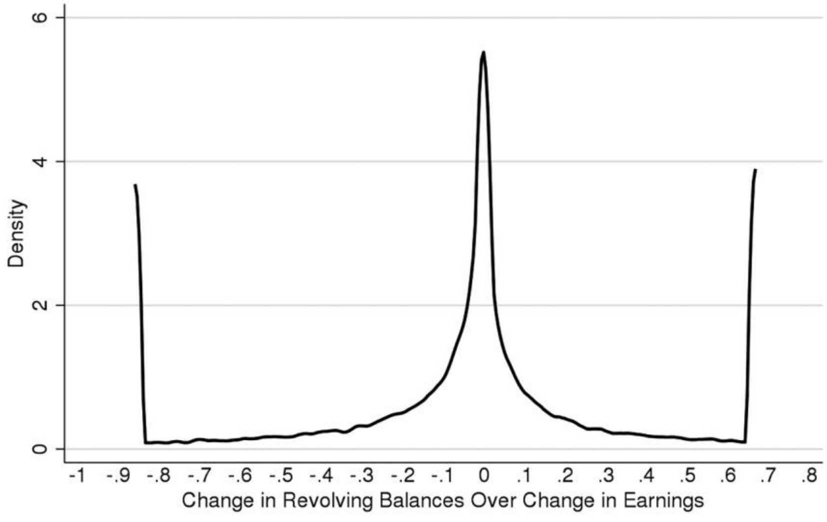

FIG. 3.—Replacement rate of lost earnings with revolving credit. Replacement rate is the negative of the change in revolving-credit balance over the change in earnings, where the change in earnings and the change in borrowing are measured from the year after displacement relative to the year before displacement. The replacement rate is defined for individuals who had a decline in earnings around displacement. A replacement rate of 0.2 indicates that an individual replaced 20% of their lost earnings with revolving credit.

Let $\Delta e_{i,t+1,t-1} = (e_{i,t+1} - e_{i,t-1})$ be the earnings loss from year $t-1$ to year $t+1$ for an individual i who was displaced in year t. The vector $X_{i,t}$ contains control variables, including a quadratic in age and deciles for lagged cumulative earnings. Let $Y_{i,t+1}$ be the outcome variable of interest (such as the change in real revolving-debt balances or an indicator variable for having a bankruptcy). Using our cross-sectional sample of displaced workers who had an earnings loss, we estimate regressions of the following form:

$$Y_{i,t+1} = \gamma_t + \eta + \mu\Delta e_{i,t+1,t-1} + \sum_{j=2}^{5}\left(\eta_j C_{j,i,t-1} + \mu_j C_{j,i,t-1} \times \Delta e_{i,t+1,t-1}\right) + \Psi X_{i,t} + \varepsilon_{i,t}. \tag{2}$$

The coefficient μ is the marginal change in the outcome variable for each dollar lost among individuals in the lowest unused-credit quintile, and the sum of the coefficients $\mu + \mu_j$ is the marginal effect for individuals in the jth unused-credit quintile. We relegate the corresponding tables to appendix A.6.

We first consider the heterogeneous responses of borrowing to changes in earnings. The dependent variable is the difference in revolving credit from $t-1$ to $t+1$ (i.e., $Y_{i,t+1} = \text{debt}_{i,t+1} - \text{debt}_{i,t-1}$), implying that $-1 \times (\mu + \mu_j)$ can be interpreted as a replacement rate of those in unused-credit quintile j. Figure 4*A* plots the earnings replacement rate from revolving-credit balances by unused-credit quintile. Individuals with the greatest amount of unused credit replace nearly 5% of lost earnings by borrowing. So for every \$10,000 of lost earnings, they borrow \$458 $(-10,000 \times (0.0506 - 0.0964))$. Individuals in the lowest unused-credit quintile reduce their credit balances by nearly 5% of lost earnings. For every \$10,000 of lost earnings, they reduce borrowing by \$506 $(-10,000 \times 0.0506)$. These results highlight that within unused-credit quintiles, the magnitude of the earnings loss is an important determinant of an individual's borrowing behavior following displacement.

We next consider the heterogeneous responses of default to changes in earnings. The dependent variable is now a default indicator in the year

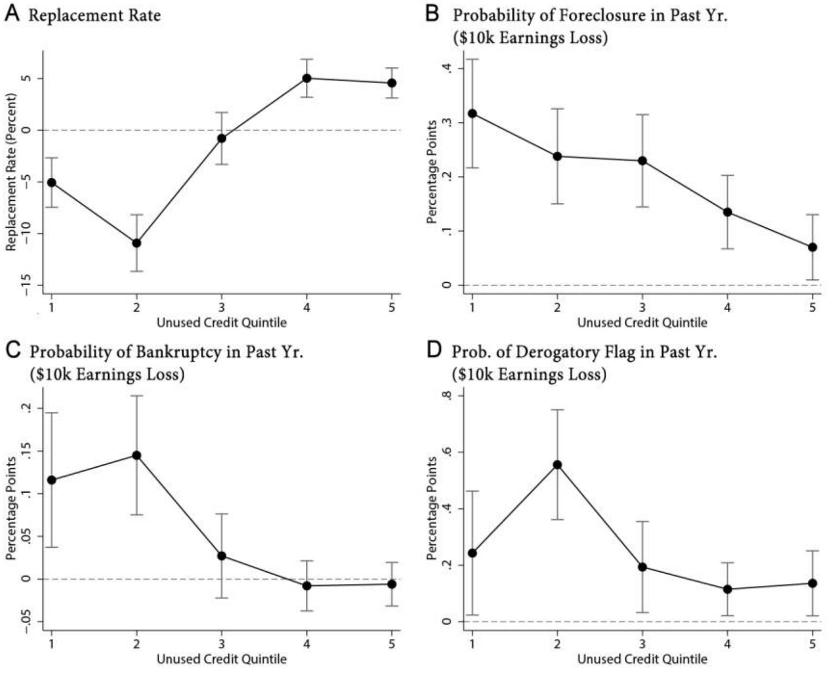

FIG. 4.—Marginal effect of earnings loss on borrowing and default activity. Circles represent the marginal effect of earnings loss on the variable of interest. Earnings loss is measured as the difference in real annual earnings in the year after displacement relative to the year before displacement. The estimates are taken from column 3 of tables A8 and A11. The coefficient for unused-credit quintile 1 corresponds to the coefficient "2 Yr. Chg. Earnings" from the table, while the coefficient for unused-credit quintile *k* corresponds to the sum of the coefficients "2 Yr. Chg. Earnings" and "2 Year Chg. Earnings Credit Quin *k*." Bars represent 95% confidence intervals.

after layoff. In order to more easily interpret the default propensities, we consider a \$10,000 earnings loss when interpreting the coefficients. Figure 4*B* plots the marginal effect of a \$10,000 earnings loss on the probability of foreclosure in the year after displacement. For individuals in the lowest unused-credit quintile, a \$10,000 decline in earnings increases the probability of a foreclosure by nearly 0.3 pp. Conversely, for individuals in the highest unused-credit quintile, a \$10,000 decline in earnings increases the probability of foreclosure by less than 0.1 pp. These results indicate that those who cannot borrow (i.e., those with very low amounts of unused credit) resort to skipping mortgage debt payments and defaulting in order to smooth consumption following job loss.

We find a similar pattern for the heterogeneous impact of earnings losses on default activity when looking at bankruptcy filings. Figure 4*C* plots the marginal effect of a \$10,000 earnings loss on the probability of bankruptcy in the year after displacement. For individuals in the two lowest unused-credit quintiles, a \$10,000 decline in earnings increases the probability of bankruptcy by between 0.10 and 0.15 pp, versus effectively 0 pp for those in the highest unused-credit quintile.

Figure 4*D* plots the marginal effect of a \$10,000 earnings loss on the probability of a derogatory public flag in the year after displacement. As discussed above, we view derogatory public flags as the closest proxy to our theoretic definition of informal and formal default. For individuals in the second-lowest unused-credit quintile, a \$10,000 decline in earnings increases the probability of having a derogatory public flag by over 0.55 pp. For those in the highest unused-credit quintile, a \$10,000 decline in earnings increases the probability of having a derogatory public flag by less than 0.15 pp.

Overall, the results of figure 4 indicate that unconstrained individuals replace a significant share—roughly 5%—of their income using credit markets. Individuals who are constrained and cannot borrow turn to default. By skipping debt payments and entering bankruptcy and collections, these individuals also use the credit market to smooth consumption.

E. Robustness

We briefly summarize two robustness exercises for our empirical analysis, the details of which are included in the appendix.

In appendix A.1, we relax the mass-layoff requirements and consider a broader "generic layoff" definition. We reestimate our empirical specifications using this weaker definition, and we find very similar results for both the distributed lag and cross-sectional specifications. While earnings losses are shallower in the generic layoff sample, the earnings elasticity of borrowing and default by quintile of credit access (e.g., $\mu + \mu_j$ in specification [2] in sec. I.D) are nearly identical to our benchmark results.

Second, in appendix A.2, we narrow our analysis to using bank cards only (e.g., credit cards) as our measure of credit. The results show that, on average, individuals maintain their access to bank cards following job loss. We additionally find that, similar to revolving credit, unconstrained individuals borrow using bank cards following job loss, while constrained individuals delever.

F. *Taking Stock*

We measure the degree to which the government can substitute away from public insurance when credit lines are prevalent using administrative earnings records linked to credit reports to examine displaced workers' ability to self-insure through credit markets. We document that, on average, individuals have substantial amounts of unused credit before job loss. We additionally show that upon job loss, on average, individuals maintain their access to credit and do not borrow, but relax their budget constraint by skipping debt payments and defaulting. However, these average responses to job loss mask substantial heterogeneity in the use of credit following job loss. Across unused-credit quintiles, individuals use credit markets to smooth consumption in very different ways. Unconstrained individuals in the highest unused-credit quintiles increase their revolving-credit balances in response to income losses. Conversely, constrained individuals in the bottom of the unused-credit distribution default in response to income losses. We conclude that both groups of individuals are using credit markets to smooth consumption. In subsequent sections, we develop a quantitative model to replicate these observations from the data, and then we use the framework to examine the substitutability of public insurance and private credit.

II. Model

In this section, we compute optimal transfers to the unemployed (which we also call "public insurance") in an environment that replicates the borrowing and default behavior documented in section I. We do so by integrating long-term credit lines (e.g., Mateos-Planas and Ríos-Rull 2010) into a model of labor search (e.g., Menzio and Shi 2011).

Time is discrete and runs forever. There is a unit measure of individuals, a continuum of potential risk-neutral lenders, and a continuum of potential entrant firms. There are $T \geq 2$ overlapping generations of risk-averse individuals who face idiosyncratic risk, similar to Menzio, Telyukova, and Visschers (2016). Each individual lives T periods. We assume that there are two types of individuals (indexed by $i \in \{1, 2\}$) that differ only by their permanent, observable discount factors, β_i. We set $0 < \beta_1 < \beta_2 < 1$; that is, type 1 individuals are less patient and generally more profitable to lenders than the more patient type 2 individuals. The share of type i individuals in the economy is π_i.

At the start of each period, individuals direct their search for jobs (e.g., Moen 1997; Burdett, Shi, and Wright 2001; Menzio and Shi 2011). Individuals then participate in an asset market where they make asset accumulation, borrowing, and default decisions. Let t denote age and t_0 denote birth cohort. We assume that individuals must apply (i.e., search) for credit contracts at utility cost κ_S. Let $S_{i,t,t+t_0}$ be a dummy that equals 1 if a type i, age t individual searches for credit in period $t + t_0$. Individuals may default on their loans $b_{i,t,t+t_0}$ at utility cost $\psi_D(b_{i,t,t+t_0})D_{i,t,t+t_0}$, where $D_{i,t,t+t_0}$ is a dummy that equals 1 in the event of default. The objective of an individual is to maximize the present discounted value of utility over nondurable consumption $(c_{i,t,t+t_0})$ net of any utility penalties of default and application costs:

$$\mathbb{E}_{t_0}\left[\sum_{t=1}^{T}\beta_i^t\left(u(c_{i,t,t+t_0}) - \psi_D(b_{i,t,t+t_0})D_{i,t,t+t_0} - \kappa_S S_{i,t,t+t_0}\right)\right].$$

For the remainder of the paper, we focus on a recursive representation of the problem, dropping the time subscript $t + t_0$.

Worker heterogeneity.—In addition to types, individuals are heterogeneous along multiple dimensions. Individuals are either employed or unemployed. Workers differ with respect to their piece rate $\omega \in [0, 1]$, which denotes the share of their per-period match output received as a wage. We let $\omega = 0$ correspond to unemployment and $\omega > 0$ correspond to employment. Thus, ω simultaneously encodes all relevant information about wages and employment status. Let $\vec{h} \in \mathcal{H} \equiv [\underline{h}, \bar{h}] \times [\underline{\epsilon}, \bar{\epsilon}] \subset \mathbb{R}^2$ be a tuple representing an individual's human capital. Human capital is comprised of two components, a persistent component $(\tilde{h})$ and a transitory component (ϵ), and we assume that human capital follows a Markov chain that depends on an individual's employment status. Let $b \in \mathcal{B} \equiv [\underline{B}, \bar{B}] \subset \mathbb{R}$ denote the net asset position of the individual, where $b > 0$ indicates saving and $b < 0$ indicates borrowing. Individuals are also heterogeneous with respect to their borrowing limit $\underline{b} \in \underline{\mathcal{B}} \equiv [\underline{B}, 0] \subset \mathbb{R}_-$ as well as their interest rate $r \in \mathcal{R} \equiv \{0\} \cup [\underline{r}, \bar{r}] \subset \mathbb{R}_+$. Those with credit access have nonzero limits and interest rates $(\underline{b}, r) \neq (0, 0)$. Those without credit access $((\underline{b}, r) = (0, 0))$ may save but cannot borrow.

Transfers and home production.—We assume that unemployed individuals $(\omega = 0)$ receive government transfers $z > 0$ and home production $g(\vec{h}) > 0$, whereas employed individuals $(\omega > 0)$ do not. Government transfers are financed by a proportional labor income tax τ. To economize on notation, we treat government transfers z as constant in our exposition of the model; however, when we map the model to the data, we follow

Mitman and Rabinovich (2015) and assume that benefits expire stochastically (see sec. III for details).

Labor market.—Unemployed individuals direct their search for employment across vacancies that specify a fixed piece rate ω for the duration of the employment match. Let $M(u, v)$ denote the labor market matching function, and define labor market tightness to be the ratio of vacancies (v) to unemployed workers (u). Since search is directed, there is a separate labor market tightness for each submarket defined by an agent's age (t), requested piece rate (ω), and human capital $(\vec{h})$. Although individuals differ along other dimensions, an agent's age, human capital, and requested piece rate are the only characteristics that matter for firm profitability. In each submarket, the job-finding rate for individuals, $p(\cdot)$, is a function of labor market tightness $\theta_t(\omega, \vec{h})$, such that $p(\theta_t(\omega, \vec{h})) = M(u_t(\omega, \vec{h}), v_t(\omega, \vec{h}))/u_t(\omega, \vec{h})$.

On the other side of the market, the hiring rate for firms $p_f(\cdot)$ is also a function of labor market tightness and is given by $p_f(\theta_t(\omega, \vec{h})) = M(u_t(\omega, \vec{h}), v_t(\omega, \vec{h}))/v_t(\omega, \vec{h})$. Once matched with a firm, a worker produces $f(\vec{h}) : \mathcal{H} \to \mathbb{R}_+$ and keeps a share ω of this production as their wage. Matches end exogenously each period with probability δ. It is important to note that because we model piece-rate contracts, workers' wages grow over time with their human capital. The prospect of higher future earnings gives workers a motive to borrow while employed as a means to smooth consumption. Generating borrowing among the employed is essential to match deleveraging upon job loss, since only individuals with preexisting debts can delever following job loss.

Credit market.—Individuals who do not default and are not hit by the credit separation shock (described below) choose whether to apply for a new credit line. Applying for a new credit line entails a common cost κ_S. If the agent decides to apply for credit, they direct their search over the menu of credit lines.[23] Each credit line specifies a borrowing limit $\underline{b}$ and interest rate r. Let $M_C(u_C, v_C)$ denote the credit market matching function, and define the credit market tightness to be the ratio of vacant credit contracts (v_C) to individuals searching for a credit contract (u_C). As in the labor market, since search is directed, credit market tightness is specific to each submarket. A submarket is defined by an agent's age (t), type (i), piece-rate wage (ω), prior debt (b), human capital $(\vec{h})$, and the requested contract $(\underline{b}, r)$. In each submarket, the credit-finding rate for individuals, $p^C(\cdot)$, is a function of the credit market tightness, where credit market tightness is given by $\theta^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$.[24]

[23] Note that while directed search is not necessary for generating interest rate and credit limit dispersion, directed search allows us to find a block-recursive solution and thus tractably compute transition dynamics without having to resort to bounded rationality.

[24] In particular, we define the credit-finding rate as, $p^C(\theta^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r)) = M_C(u_{C,i,t}(\omega, b, \vec{h}; \underline{b}, r), v_{C,i,t}(\omega, b, \vec{h}; \underline{b}, r))/u_{C,i,t}(\omega, b, \vec{h}; \underline{b}, r)$.

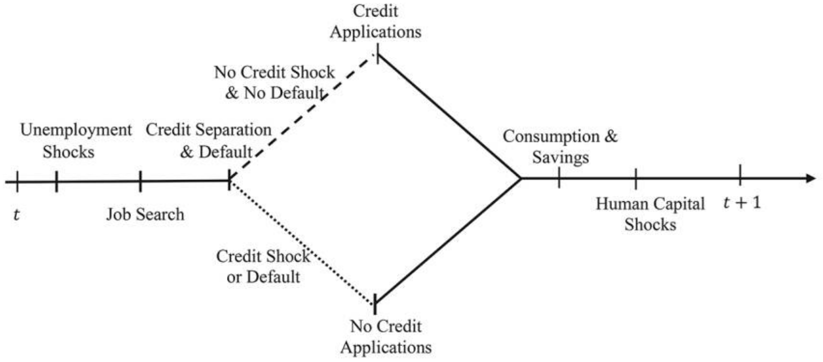

FIG. 5.—Model timeline.

On the other side of the market, the probability a lender matches with a borrower, denoted $p_f^{\mathrm{C}}(\cdot)$, is also a function of credit market tightness and is given by $p_f^{\mathrm{C}}(\theta_{i,t}^{\mathrm{C}}(\omega, b, \vec{h}; \underline{b}, r))$.[25] An individual remains matched with a lender until the individual successfully applies and obtains a new credit line, defaults, or is hit by the exogenous credit separation shock (δ_{C}). We assume that individuals are unable to search for new credit lines in periods when the individual defaults or exogenously separates from the lender.

Timing.—We assume that unemployment shocks are realized at the start of the period. Unemployed individuals then enter the labor market and apply for jobs. After the labor market closes, the agent may endogenously separate from lenders by defaulting or may receive an exogenous credit separation shock. Individuals who did not default and were not hit by the credit separation shock then enter into the credit application stage.[26] After the credit application stage, individuals make borrowing, saving, and consumption decisions. Idiosyncratic human capital risk is then realized, and the next period begins. We illustrate the timing of the model in figure 5, and in the subsequent sections we lay out the problem solved by agents in each stage of the period.

A. *Bellman Equations*

This section presents the Bellman equations that govern the behavior of workers, firms, and lenders in equilibrium.

Consumption/savings decision.—We first detail the consumption and savings problem that each agent faces. Let $V_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$ denote the value

[25] The credit-finding rate for lenders is defined as, $p_f^{\mathrm{C}}(\theta_{i,t}^{\mathrm{C}}(\omega, b, \vec{h}; \underline{b}, r)) = M_{\mathrm{C}}(u_{\mathrm{C},i,t}(\omega, b, \vec{h}; \underline{b}, r), v_{\mathrm{C},i,t}(\omega, b, \vec{h}; \underline{b}, r))/v_{\mathrm{C},i,t}(\omega, b, \vec{h}; \underline{b}, r)$.

[26] Note that individuals without credit access at the start of the period cannot default and are not subjected to the credit separation shock.

of entering the consumption-savings stage for an age t, type i individual with wage rate ω, net assets b, human capital $\vec{h}$, and credit contract $(\underline{b}, r)$. Note that the unemployed have piece rate $\omega = 0$ and that agents without credit have a contract $(\underline{b}, r) = (0, 0)$. Upon entering this stage, the agent makes their consumption/savings decision, where their asset decision (b') is constrained by their borrowing limit $\underline{b}$. After the individual makes their consumption/savings decision, shocks to human capital are realized and the period ends. At the start of the next period, the agent enters into the labor market, where $V^{\mathrm{L}}_{i,t+1}(\omega, b', \vec{h}'; \underline{b}, r)$ denotes the value to an individual of entering into the labor market. The value to an agent of entering the consumption-savings stage is given by

$$V_{i,t}(\omega, b, \vec{h}; \underline{b}, r) = \max_{b' \geq \underline{b}} u(c) + \beta_i \mathbb{E}\left[V^{\mathrm{L}}_{i,t+1}(\omega, b', \vec{h}'; \underline{b}, r) \right] \quad \forall\, t \leq T,$$

$$V_{i,T+1}(\omega, b, \vec{h}; \underline{b}, r) = 0,$$

subject to the budget constraint,

$$c + q(b', r)b' \leq w(\omega, \vec{h}) + b,$$

where the bond price $q(b', r)$ includes both the discount on the face value of loans and the savings rate,

$$q(b', r) = \mathbb{I}\{b' < 0\} \frac{1}{1 + r} + \mathbb{I}\{b' \geq 0\} \frac{1}{1 + r_f},$$

the function $w(\omega, \vec{h})$ governs how an individual's employment status, wage rate, and human capital translate into a wage and is given by

$$w(\omega, \vec{h}) = \begin{cases} z + g(\vec{h}) & \text{if } \omega = 0, \\ (1 - \tau)\omega f(\vec{h}) & \text{if } \omega \neq 0, \end{cases}$$

and the law of motion for human capital is indexed by the individual's employment status,

$$\vec{h}' = \begin{cases} H_{-}(\vec{h}) \text{ if } \omega = 0, \\ H_{+}(\vec{h}) \text{ if } \omega \neq 0. \end{cases} \tag{3}$$

Human capital evolves so that, on average, unemployed agents see their human capital decline, while employed agents experience an increase in human capital. Unemployed agents ($\omega = 0$) receive a public insurance transfer z, which is provided by the government and funded through taxes on employed agents. We model the public insurance transfer to encapsulate all forms of assistance that unemployed workers receive, which can include unemployment compensation and emergency unemployment assistance as well as general transfer programs, such as welfare and food stamps, that unemployed individuals may use. Additionally, unemployed individuals receive the value of home production $g(\vec{h})$, which is assumed to be a function of a worker's human capital $\vec{h}$. In the model, home production proxies for other resources that individuals access during unemployment, such as transfers from friends and family or changes in spousal labor supply.

Employed agents ($\omega \neq 0$) receive a wage that is a piece rate $\omega \in (0, 1]$ of their per-period production $f(\vec{h})$. They then pay a proportional tax τ on labor earnings to finance public insurance transfers.

After the agents make their consumption/savings choice, shocks to human capital are realized and the period ends. At the start of the next period, individuals enter into the labor market.

Labor market.—Let $V^{\mathrm{L}}_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$ denote the value of entering the labor market for a type i, age t individual. In the labor market, unemployed workers ($\omega = 0$) search for jobs across potential wage piece rates $\tilde{\omega}$. In choosing where to apply, the worker faces a trade-off, since jobs with higher wage piece rates have lower job-finding rates. With probability $p(\theta_t(\tilde{\omega}, \vec{h}))$, an individual matches with a job that pays wage piece rate $\tilde{\omega}$ and becomes an employed worker. With probability $1 - p(\theta_t(\tilde{\omega}, \vec{h}))$, the worker does not match with the job and continues as an unemployed worker. After the labor market closes, agents enter the default stage, where $V^{\mathrm{D}}_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$ denotes the value to an individual of entering the default stage.

For agents who enter the labor market as employed ($\omega \neq 0$), with probability δ they become unemployed. For agents who become unemployed, with probability $\lambda_{\mathrm{S}} \in [0, 1]$ the agent is able to search immediately. With probability $1 - \lambda_{\mathrm{S}}$, the agent is unable to search and thus enters the default stage as an unemployed worker. Allowing some agents to search immediately upon job loss allows us to discipline the size of earnings losses after layoff. The value to entering the labor market is given by

$$V^{\mathrm{L}}_{i,t}(\omega, b, \vec{h}; \underline{b}, r) = \begin{cases} \max\limits_{\tilde{\omega}} p(\theta_t(\tilde{\omega}, \vec{h}))V^{\mathrm{D}}_{i,t}(\tilde{\omega}, b, \vec{h}; \underline{b}, r) + (1 - p(\theta_t(\tilde{\omega}, \vec{h}))V^{\mathrm{D}}_{i,t}(0, b, \vec{h}; \underline{b}, r)) & \text{if } \omega = 0, \\ (1 - \delta)V^{\mathrm{D}}_{i,t}(\omega, b, \vec{h}; \underline{b}, r) + \delta[\lambda_{\mathrm{S}} V^{\mathrm{L}}_{i,t}(0, b, \vec{h}; \underline{b}, r) + (1 - \lambda_{\mathrm{S}})V^{\mathrm{D}}_{i,t}(0, b, \vec{h}; \underline{b}, r)] & \text{if } \omega \neq 0. \end{cases}$$

Credit separations and default.—Let $V^{\mathrm{D}}_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$ denote the value of entering the default stage for a type i, age t individual. At the start of the default stage, individuals with credit (i.e., $(\underline{b}, r) \neq (0, 0)$) are exogenously separated from their lenders with probability δ_C. After the credit separation shock, agents decide whether to default. If an agent defaults, (i) they incur a utility penalty ($\psi_D(b)$) that is increasing in the amount of assets defaulted upon, (ii) their assets are set to zero, and (iii) they are excluded from searching for credit in the current period.[27] Nondefaulters who avoid the credit separation shock may search for new credit lines, where $V^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$ denotes the value of entering the credit-search stage. This formulation allows agents to engage in "in-the-contract" credit search and move to credit contracts that have more generous borrowing limits and lower interest rates.

The value of entering the default stage is given by

$$\begin{aligned} V^D_{i,t}(\omega, b, \vec{h}; \underline{b}, r) = {} & \delta_C \max\{V_{i,t}(\omega, 0, \vec{h}; 0, 0) - \psi_D(b); V_{i,t}(\omega, b, \vec{h}; 0, 0)\} \\ & + (1 - \delta_C) \max\{V_{i,t}(\omega, 0, \vec{h}; 0, 0) - \psi_D(b); V^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r)\}. \end{aligned} \tag{4}$$

For individuals without credit (i.e., $(\underline{b}, r) = (0, 0)$) there is no default decision, and they simply proceed to the credit-search stage.

Credit application and search.—Let $V^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r)$ denote the value of entering the credit-search stage for a type i, age t individual. In the credit-search stage, the agent decides whether to apply for credit at utility cost κ_S. When an agent chooses to apply for credit, they direct their search across a menu of credit contracts $(\underline{b}, r)$. The value of entering the credit-search stage is given by

$$V^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r) = \max\{V^A_{i,t}(\omega, b, \vec{h}; \underline{b}, r) - \kappa_S, V_{i,t}(\omega, b, \vec{h}; \underline{b}, r)\}$$

where $V^A(\omega, b, \vec{h}; \underline{b}, r)$ denotes the value of applying for credit,

$$\begin{aligned} V^A_{i,t}(\omega, b, \vec{h}; \underline{b}, r) = {} & \max_{(\tilde{\underline{b}}, \tilde{r})} p(\theta^C_{i,t}(\omega, b, \vec{h}; \tilde{\underline{b}}, \tilde{r})) V_{i,t}(\omega, b, \vec{h}; \tilde{\underline{b}}, \tilde{r}) \\ & + \left[1 - p(\theta^C_{i,t}(\omega, b, \vec{h}; \tilde{\underline{b}}, \tilde{r}))\right] V_{i,t}(\omega, b, \vec{h}; \underline{b}, r). \end{aligned}$$

After the credit market closes, agents enter into the consumption/savings problem.

1. Lenders

There is a continuum of potential lenders who are risk neutral and can obtain funds without constraint at the risk-free rate r_f. Lenders discount

[27] Those who are exogenously separated from their lenders are excluded from immediately searching for credit.

their stream of future profits at rate $\beta_{lf} \in (0, 1)$. Lenders offer credit contracts which specify a borrowing limit $\underline{b} < 0$ and an interest rate r. Let $\Pi_{i,t}(\vec{s})$ denote the present value of profits to a lender of being matched with a type i, age t, individual where an individual's state is given by $\vec{s} = (\omega, b, \vec{h}; \underline{b}, r)$.[28] We first derive lender flow profits. We then derive the present value of flow profits and detail the lender entry decision.

Lender flow profits.—At the end of the period, an age t agent makes their asset decision, $b'_{i,t}(\vec{s})$. If the individual is borrowing, $b'_{i,t}(\vec{s}) < 0$, then in the next period the lender earns the spread between the interest rate r and the risk-free rate r_f. However, the lender faces default risk on the outstanding loan $b'_{i,t}(\vec{s})$. Let $\hat{D}_{i,t+1}(\vec{s})$ denote the expected probability of default for an agent with state $\vec{s}$. The expected probability of default incorporates the probability of the credit separation shock, as well as shocks to human capital and the individual's job search decision.[29] With the expected probability of default defined, we can write the flow profits to the lender as

$$m_{i,t}(\omega, b, \vec{h}; \underline{b}, r) = \beta_{lf} b'_{i,t}(\vec{s}) \left(\frac{r_f - r}{1 + r} + \hat{D}_{i,t+1}(\vec{s}) \right) \times \mathbb{I}\{b'_{i,t}(\vec{s}) < 0\}. \tag{5}$$

Present value of lender flow profits and free entry.—Lenders make entry decisions on the basis of present value of flow profits, and the present value of flow profits crucially depends the match duration. A lender's match continues if (1) the match is not hit by the credit separation shock, (2) there is no default, and (3) the individual does not move to another lender via in-the-contract search. Let $\Gamma(\omega', b', \vec{h}'; \underline{b}, r)$ denote the probability that the match between the lender and the agent continues to the next period.[30] The present value of profits to the lender are then given by

$$\begin{aligned}\Pi_{i,t}(\omega, b, \vec{h}; \underline{b}, r) &= m_{i,t}(\omega, b, \vec{h}; \underline{b}, r) + \beta_{lf}\mathbb{E}\left[\Gamma_{i,t+1}(\omega', b', \vec{h}'; \underline{b}, r)\Pi_{i,t+1}(\omega', b', \vec{h}'; \underline{b}, r)\right] \\ &\quad \forall\ t \leq T, \\ \Pi_{i,T+1}(\omega, b, \vec{h}; \underline{b}, r) &= 0.\end{aligned}$$

Free entry determines the number of lenders who enter each submarket in equilibrium. The free-entry condition is

$$\kappa_C \geq p_f^C\left(\theta^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r)\right)\Pi_{i,t}(\omega, b, \vec{h}; \underline{b}, r). \tag{6}$$

The free-entry condition binds for all submarkets such that $\theta^C_{i,t}(\omega, b, \vec{h}; \underline{b}, r) > 0$.

[28] Let $\vec{s}'$ denote the state space of the individual in the next period.

[29] See app. B.1.1 for the derivation of the expected probability of default.

[30] See app. B.1.2 for the derivation of the probability that the match between the lender and agent proceeds to the next period.

2. Firms

Firms use a linear production technology $f(\vec{h})$, and they exogenously separate from their workers at rate δ. Firms have the same discount factor β_{lf} as lenders. The continuation value of a firm that has committed to pay piece rate ω to their age t employee with human capital $\vec{h}$ is

$$\begin{aligned} J_t(\omega, \vec{h}) &= (1-\omega)f(\vec{h}) + \beta_{lf}\mathbb{E}\big[(1-\delta)J_{t+1}(\omega, \vec{h}')\big] \ \ \forall\, t \le T, \\ J_{T+1}(\omega, \vec{h}) &= 0, \end{aligned}$$

subject to the law of motion for human capital for employed individuals,

$$\vec{h}' = H_+(\vec{h}).$$

Firms must pay cost κ to post a vacancy. A vacancy specifies a wage piece rate ω as well as a human capital requirement $\vec{h}$ and age t. Free entry requires that

$$\kappa \ge p_f\big(\theta_t(\omega, \vec{h})\big)J_t(\omega, \vec{h}). \tag{7}$$

The free-entry condition binds for all submarkets such that $\theta_t(\omega, \vec{h}) > 0$.

3. Government

The government provides public transfers z to the unemployed. Public transfers are paid for by a proportional labor income tax, τ, which is levied on all employed individuals to yield period-by-period budget balance,

$$z\sum_{(i,t)}\sum_{\vec{s}}\hat{u}_{i,t}(\vec{s}) = \sum_{(i,t)}\sum_{\vec{s}}\tau\omega f(\vec{h})\hat{e}_{i,t}(\vec{s}), \tag{8}$$

where $\hat{u}_{i,t}(\vec{s})$ is the share of individuals with state $\vec{s}$ that are type i and age t who are unemployed, and $\hat{e}_{i,t}(\vec{s}) = 1 - \hat{u}_{i,t}(\vec{s})$ is the share who are employed.

B. *Equilibrium*

In equilibrium, individual decision rules are optimal, free entry holds in both the credit and labor markets, the government balances its budget, and the distribution of individuals across states is consistent with the decision rules. The formal definition of equilibrium is given in appendix B.3.

In appendix B.3, we prove that if the government budget constraint is ignored and τ is exogenously given, then the model is block recursive (e.g., Menzio and Shi 2011). Given an exogenous τ, block recursivity means that the individual, lender, and firm problems can be solved independently of the distribution of individuals across states.

The equilibrium tax rate that balances the government budget constraint will ultimately depend on the distribution of individuals across states and, in the case of transition dynamics, the path of tax rates will also depend on the path of the distribution of individuals across states. However, the fact that equilibrium prices and the distribution of individuals across states are linked only by τ greatly simplifies our computation of the transition path. We refer to this property of our model as "conditional block recursivity."

III. Calibration

Because of the computationally demanding nature of the model, our calibration strategy is to assign values from the literature to standard parameters wherever possible and then estimate the remaining nonstandard parameters to match moments from the data.[31] We estimate our steady state to match moments from 2002 to 2012. Everywhere possible, we calibrate the model to match moments from our linked LEHD-TransUnion sample.[32] However, several of our moments are available only at different points in time or from other sources. While the calibration of model parameters is performed jointly, we discuss the moments that are most informative for each model parameter.

Preferences and demographics.—The period is 1 quarter. A worker's life span is set to $T = 120$ quarters (30 years). Newly born individuals enter as unemployed workers, with zero assets and without a credit contract. Their initial persistent human capital is drawn from an exponential distribution with parameter λ_H. We calibrate the parameter $\lambda_H = 2.37$ to match the P75-P25 ratio of residualized log earnings among 25–29-year-olds.[33] We estimate this ratio to be 0.662 in our LEHD-TransUnion sample.

Individual preferences over nondurable consumption are given by

$$u(c) = \frac{c^{1-\sigma} - 1}{1 - \sigma}.$$

We set the risk aversion parameter to a standard value, $\sigma = 2$. We set the annualized risk-free rate to 4%, and the corresponding quarterly discount

[31] Appendix C describes our solution algorithm in detail.

[32] When calibrating parameters that are not specific to layoffs, we use the full sample of linked LEHD-TransUnion data to estimate moments, and not just the sample of individuals who experience a layoff or coworkers of an individual who experienced a layoff. To align with the 30-year working careers of agents in the model, we limit our linked LEHD-TransUnion sample to individuals between the ages of 25 and 54.

[33] We compute a Mincer-style log earnings regression for workers between the ages of 25 and 29. We residualize earnings by removing age, as well as year, industry, race, and gender fixed effects. We estimate our Mincer-style regression on all individuals in our LEHD-TransUnion sample with earnings over $5,000.

factor for firms and lenders is $\beta_{lf} = 0.99$. The patient worker type also discounts the future at the same rate, $\beta_2 = 0.99$. The parameters that govern the impatient type are determined by using cross-sectional moments on credit usage. We estimate the discount factor of the impatient type, $\beta_1 = 0.832$, to match that 30.6% of individuals increase their revolving-credit balance in the year after layoff relative to the year before layoff. A lower discount factor makes individuals more constrained upon job loss and reduces the share who borrow after displacement.

We calibrate the fraction of individuals who are impatient, denoted $\pi_1 = 1 - \pi_2 = 0.387$, to match the average unused revolving-credit share for individuals in the second unused-credit quintile, which in section I we measured to be 0.311.[34] A larger share of impatient agents increases the share of individuals with little unused credit upon job loss. We discuss the role of impatient agents for optimal policy in appendix F.1.[35]

Labor market.—We set the job destruction rate to a constant 6.87% per quarter, $\delta = 0.0687$.[36] For the labor market matching function, we use a constant-returns-to-scale matching function that yields well-defined job-finding probabilities:

$$M(u, v) = \frac{u \cdot v}{(u^{\zeta} + v^{\zeta})^{1/\zeta}} \in [0, 1).$$

The matching elasticity parameter is chosen to be $\zeta = 1.6$, as measured in Schaal (2017). The labor vacancy posting cost $\kappa = 0.512$ is estimated to target an unemployment rate of 5.7% among 24–54-years-olds in the BLS (Bureau of Labor Statistics) from 2002–12. When an individual is hit by the job separation shock, with probability λ_S they are able to search for a job immediately. We calibrate the parameter λ_S to match the size of earnings losses around job loss. In particular, we calibrate λ_S to match the trough earnings loss following displacement, which in section I we measured to be 17.6% of prior earnings.

Human capital evolves following a Markov chain with a persistent and transitory component. Let $\vec{h} = (\tilde{h}, \epsilon)$, denote the human capital of an agent, where $\tilde{h}$ denotes the individual's persistent human capital and ϵ denotes

[34] We calibrate to the second unused-credit quintile, since the first unused-credit quintile is zero for a large range of parameters, yielding a flat objective function and weak identification.

[35] In app. F.3 we restrict the share of impatient agents to be half of what it is in the baseline (e.g., type 1's comprise 20% of agents) and find a moderately weaker degree of substitution between private and public insurance of 5 pp. In earlier calibrations with lower discount factors of the impatient types and alternate shares of the impatient types, we found moderately weaker degrees of substitution between private and public insurance of 3–5 pp. From this perspective, we view 6 pp as an upper bound on the degree of substitution.

[36] Using the method of Shimer (2005) and data from the Current Population Survey (CPS) for the years 2002–12, we estimate a quarterly job separation rate of 6.87%.

the transitory component. We assume that the production function is linear and additive in the human capital of the worker, $f(\vec{h}) = \tilde{h} + \epsilon$. The process for the persistent component of human capital is governed by two parameters $p_{\tilde{h},\mathrm{L}}$ and $p_{\tilde{h},\mathrm{H}}$:

$$H_{\mathrm{P},-}(\vec{h}) = \tilde{h}' = \begin{cases} \tilde{h} - \Delta & \text{with probability } p_{\tilde{h},\mathrm{L}} \text{ if unemployed,} \\ \tilde{h} & \text{with probability } 1 - p_{\tilde{h},\mathrm{L}} \text{ if unemployed.} \end{cases}$$

$$H_{\mathrm{P},+}(\vec{h}) = \tilde{h}' = \begin{cases} \tilde{h} + \Delta & \text{with probability } p_{\tilde{h},\mathrm{H}} \text{ if employed,} \\ \tilde{h} & \text{with probability } 1 - p_{\tilde{h},\mathrm{H}} \text{ if employed.} \end{cases}$$

The grid for the persistent component of human capital $\tilde{h} \in [0.6, 0.7, \dots, 1.2, 1.3]$ and the step size $\Delta = 0.1$ between grid points are taken as given. To estimate the probability that the persistent component of a worker's human capital increases while employed, $p_{\tilde{h},\mathrm{H}} = 0.062$, we target the 0.95% semielasticity of earnings with respect to age in the LEHD-TransUnion sample.[37] To estimate the probability that a worker's productivity decreases while unemployed, $p_{\tilde{h},\mathrm{L}} = 0.737$, we target the 8.9% decline in earnings 5 years following job loss, as measured in section I.C. Note that with probability $\lambda_{\mathrm{S}} = 0.586$ we allow for immediate job search (within the quarter) if an agent receives an exogenous separation shock δ. Therefore, human capital depreciation applies only to a much smaller subset of agents who do not immediately find a job.

The process for the transitory component of human capital is governed by the parameters $p_{\epsilon,\mathrm{L}}$ and $p_{\epsilon,\mathrm{H}}$:

$$H_{\mathrm{T},+}(\tilde{h}') = \epsilon' = \begin{cases} \Delta_\epsilon(\tilde{h}') & \text{with probability } p_{\epsilon,\mathrm{H}}, \\ 0 & \text{with probability } 1 - p_{\epsilon,\mathrm{L}} - p_{\epsilon,\mathrm{H}}, \\ -\Delta_\epsilon(\tilde{h}') & \text{with probability } p_{\epsilon,\mathrm{L}}. \end{cases} \tag{9}$$

The step size $\Delta_\epsilon(\tilde{h}') = 0.095\tilde{h}'$ is taken as given, and we estimate the parameters $p_{\epsilon,\mathrm{H}} = 0.064$ and $p_{\epsilon,\mathrm{L}} = 0.058$ to target the share of employed workers who experience a 9.5% wage increase and decrease over a given

[37] We estimate the earnings gain associated with an increase in age using the following regression of age on earnings in period t: $\ln(Y_{i,t}) = \alpha + \beta_{\mathrm{age}}\mathrm{Age}_{i,t} + \varepsilon_{i,t}$, where $Y_{i,t}$ denotes the earnings of individual i in year t, and $\mathrm{Age}_{i,t}$ denotes the age of individual i in year t. The coefficient β_{age} estimates the average increase in log earnings associated with an increase in age. We estimate the regression for all individuals for whom we can link a TransUnion credit report to the LEHD and the individual had annual earnings greater than \$5,000 in year t. For this sample, we estimate a relative gain in earnings with a 1-year increase in age of 0.95%. We additionally include dummies for year, industry, sex, and race in the estimation.

year, respectively, as reported in Kurmann and McEntarfer (2017).[38] Given the processes for the transitory and persistent components of human capital, the evolution of human capital proceeds as

$$\begin{aligned} H_{+}(\vec{h}) &= (H_{\mathrm{P},+}(\vec{h}), H_{\mathrm{T}}(H_{\mathrm{P},+}(\vec{h}))), \\ H_{-}(\vec{h}) &= (H_{\mathrm{P},-}(\vec{h})). \end{aligned}$$

Transfers and home production.—The public transfer to unemployed workers $z = 0.470$ is estimated to match the 41.2% public transfer replacement rate (change in public transfers divided by change in annual income) among laid-off workers observed in the PSID (Panel Study of Income Dynamics) between 2001 and 2013.[39]

To further connect the model to the UI literature, we extend the model to include benefit expiration. To maintain tractability, we model benefit expiration as in Mitman and Rabinovich (2015), where UI benefits expire stochastically.[40] Let φ denote the probability than an individual's unemployment benefits expire. Standard unemployment benefits expire after 26 weeks, and, given the quarterly timing of the model, we set $\phi = 1/3$ to account for individuals who transition to unemployment and immediately lose unemployment benefits.[41] When an individual's UI benefits expire, they receive a transfer αz, where $\alpha < 1$ reflects the non-UI component of transfers. We calibrate α using the decline in consumption after benefit

[38] Kurmann and McEntarfer (2017) use the LEHD for the state of Washington, where both hours and earnings are reported, which allows for measuring wages. Kurmann and McEntarfer (2017) report that between 2009 and 2010, 7.65% of job stayers (individuals who report being at the same establishment [state employer ID number, or SEIN] for 10 consecutive quarters) experienced a wage decline of at least 9.5% during that year. They report that 19% of job stayers experienced a wage increase of 9.5% or higher during that year.

[39] Our measure of income from the PSID is household income less transfers, which is the sum across household members of (1) wage and salary income, (2) business income, and (3) interest dividend income. Transfers are also measured at the household level. We measure the public transfer replacement rate (change in transfers over the change in household income less transfers), for households where either the head of household or spouse has an involuntary unemployment spell with a duration of greater than 1 quarter. We additionally require an income decline of at least $1,000, and we winsorize the replacement rate at the 1% level. We focus on involuntary layoffs to avoid unemployment spells due to quits and because involuntary layoffs are more consistent with the notion of a layoff in the model. We similarly use individuals with an unemployment duration of at least 3 months, given the quarterly timing of the model where unemployed individuals are out of work for at least a full quarter. Using the Survey of Income and Program Participation (SIPP), Rothstein and Valletta (2017) estimate a replacement rate (changes in transfers over changes in earnings) of 43.6%.

[40] Stochastic benefit expiration allows us to avoid keeping an individual's unemployment duration as a state variable.

[41] We allow for individuals to transition to unemployment and immediately lose benefit eligibility to align with the estimates of Chodorow-Reich and Karabarbounis (2016) that 35% of unemployed individuals do not take up UI.

expiration from Ganong and Noel (2019). In appendix B.2, we present the Bellman equations that incorporate benefit expiration.[42]

In addition to public insurance transfers, the unemployed also receive home production, which proxies for other resources that the unemployed have, such as transfers from friends and family or changes in spousal labor supply. We define home production to be a function of human capital such that

$$g(\vec{h}) = g - \eta(\bar{h} - \tilde{h}),$$

where $\bar{h}$ is the highest value on the grid of persistent human capital. The parameter g governs the base level of home production and is calibrated to match the decline in consumption upon job loss. Using the PSID, we estimate that, on average, individuals who experience at least 1 quarter of unemployment have annual consumption that is 94.7% of their consumption level before layoff.[43] The parameter η governs the heterogeneity in home production across human capital, with $\eta = 0$ yielding homogeneous home production and $\eta > 0$ yielding home production that increases with persistent human capital ($\tilde{h}$). Additionally, given the law of motion for persistent human capital, if $\eta > 0$, then workers expect home production to decrease with the length of their unemployment spell. Home production also influences job search behavior, with flatter profiles ($\eta \approx 0$) implying greater unemployment durations among low-human-capital (low-earning) individuals. Accordingly, we calibrate η to match the relationship between prior earnings and unemployment rates.[44]

Credit market.—We calibrate the exogenous credit separation rate to 2.4% per quarter, $\delta_C = 0.024$, to match the time-aggregated annual credit separation rate we observe in our TransUnion sample. In both model and data, we define a credit separation to be a 90% (or more) reduction in

[42] We obtain an estimate for the non-UI share of transfers (25%) similar to the estimate in Nakajima (2012b), 33.4%. In app. F.4, we show that our welfare results are robust to using the estimate from Nakajima (2012b) for the non-UI share of transfers.

[43] In the PSID, we measure the change in family consumption across survey waves for families where the head of household had an unemployment spell with a duration of at least 1 quarter between 2001 and 2013. Additionally, we require that the household have at least $1,000 of consumption both before and after layoff and that the head of household was employed in the prior wave of the PSID. We winsorize the change in consumption among this sample at the 5% level.

[44] Using the Annual Social and Economic Supplement (ASEC) of the CPS, we place individuals into income deciles based on their annual earnings in the prior year. We then measure the share of individuals who report that they are unemployed by decile at the time of the survey in March and take the difference between the first decile (individuals with the lowest prior earnings) and the tenth decile (individuals with the highest prior earnings). Using data from 2002–12 for workers between the ages of 25 and 54, we measure this difference to be 8.7 pp. We refer to this differential as the "unemployment by prior earnings slope."

credit limits across two consecutive years. In the TransUnion data, 5.3% of individuals experience a credit separation in a given year.

For the credit market matching function, we again use a constant-returns-to-scale matching function that yields well-defined credit-finding probabilities in discrete time:

$$M_{\mathrm{C}}(u_{\mathrm{C}}, v_{\mathrm{C}}) = \frac{u_{\mathrm{C}} \cdot v_{\mathrm{C}}}{\left(u_{\mathrm{C}}^{\zeta_{\mathrm{C}}} + v_{\mathrm{C}}^{\zeta_{\mathrm{C}}}\right)^{1/\zeta_{\mathrm{C}}}} \in [0, 1).$$

The matching elasticity parameter is chosen to be $\zeta_{\mathrm{C}} = 0.37$ as measured in Herkenhoff (2019).[45]

There is an exogenously given grid of interest rates for credit contracts over the interval $[\underline{r}, \bar{r}]$. We set the minimum annual interest rate ($\underline{r}$) to be 4.4%, which is the 10th percentile of the real credit card interest rate distribution in the SCF.[46] We set the maximum interest rate ($\bar{r}$) to be 18.6%, which is the 90th percentile of the real credit card interest rate distribution in the SCF.

Credit contracts also specify a borrowing limit that must lie in the interval $[\underline{B}, 0)$, where $\underline{B} < 0$ is the minimum value of the asset grid. We estimate $\underline{B} = -0.97$, so that the median unused-credit-to-income ratio is 8.2%, as measured in the LEHD-TransUnion data. The credit posting cost $\kappa_{\mathrm{C}} = 2.18 \times 10^{-5}$ is estimated so that the credit-finding rate in the model matches the new-borrower credit approval rate of 51.4% in the LEHD-TransUnion data.[47]

The utility cost of searching for a credit contract $\kappa_{\mathrm{S}} = 8.41 \times 10^{-4}$ is calibrated to match the fact that 77.9% of the population in our LEHD-TransUnion database has a positive revolving-credit limit. The utility penalty of default is assumed to be linear in the amount of assets defaulted upon:

$$\psi_{\mathrm{D}}(b) = -b \cdot \psi.$$

We set the default penalty $\psi = 18.1$ to match the peak probability of a new derogatory public flag after job loss.[48] In section I, we estimated that

[45] Using data from Synovate on direct-mail credit card offers and credit applications from the SCF, Herkenhoff (2019) estimates the matching elasticity in the credit market to be 0.37 via nonlinear least squares.

[46] We use the SCF to define the grid on interest rates because interest rates are not reported on TransUnion credit reports.

[47] To measure the new-borrower credit-finding rate, we take the ratio of the number of individuals who have a credit inquiry and take out their first revolving-credit line in a year t over the number of individuals who did not have a revolving-credit line in year $t-1$ and had a credit inquiry in year t.

[48] We calibrate the default penalty to the response of derogatory public flags to job loss in order to isolate defaults arising as a result of job loss. We note that the consumption equivalent of our default costs implies a $934 cost of bankruptcy, which is very comparable to Albanesi and Nosal's (2015) estimated costs of Chapter 7 bankruptcy of $697–$975.

TABLE 2
Model Parameters

Variable	Value	Description
Nonestimated:		
r_f (%)	4	Risk-free rate
β_{lf}	.99	Discount factor: lenders and firm
β_2	.99	Discount factor low worker type
δ	.0687	Exogenous job destruction rate
φ	1/3	Benefit expiration probability
ζ	1.6	Labor match elasticity
ζ_{C}	.37	Credit match elasticity
$\underline{r}$ (%)	4.4	Minimum (annualized) interest rate
$\bar{r}$ (%)	18.6	Maximum (annualized) interest rate
σ	2	Risk aversion
T (quarters)	120	Life span
Jointly estimated:		
z	.470	Public insurance transfer to unemployed
κ	.512	Firm entry cost
κ_{C}	2.18×10^{-5}	Lender entry cost
κ_{S}	8.41×10^{-4}	Utility penalty of searching for credit
ψ_{D}	18.1	Utility penalty of default
$p_{\tilde{h},\mathrm{L}}$	.737	Probability of persistent human capital decrease
$p_{\tilde{h},\mathrm{H}}$	.062	Probability of persistent human capital increase
$p_{\epsilon,\mathrm{L}}$	.058	Probability of transitory human capital low
$p_{\epsilon,\mathrm{H}}$	.064	Probability of transitory human capital high
λ_{H}	2.37	Exponential parameter initial persistent human capital
α	.250	Public insurance transfer share after expiration
g	.345	Home production
$\underline{B}$	−.97	Lower bound for borrowing limit
β_1	.832	Discount factor: impatient worker type
π_1	.387	Share of impatient agents
δ_{C}	.024	Exogenous credit destruction rate
λ_{S}	.586	Probability of searching immediately after job loss
η	.049	Slope of home production function

in response to job loss, the peak probability of a new derogatory public flag is 0.517 pp higher and occurs 1 year after layoff. We calibrate the default penalty to match this moment.

Table 2 contains a summary of the model parameters, and table 3 displays the calibrated parameters and their calibration targets. The estimated model matches the targeted moments very well. We discuss nontargeted moments in the next section.

A. *Model Estimates of Credit Access and Usage*

In this section, we compare the model's estimates of credit access and usage to the data.[49] We first examine how credit access and usage respond

[49] In app. B.4, we examine additional nontargeted moments, in particular, the distribution of unused credit to income as well as gross debt positions.

TABLE 3
Model Calibration

Variable	Value	Target Used to Match Variable	Model	Data	Source
z	.470	Transfer replacement rate (%)	41.1	41.2	PSID 2001–13
κ	.512	Unemployment rate (%)	5.7	5.7	BLS, 24–54-year-olds, 2002–12
κ_C	2.18×10^{-5}	New-borrower credit-finding rate (%)	51.3	51.4	LEHD-TU 2002–12
κ_S	8.41×10^{-4}	Share of individuals with credit access (%)	77.8	77.9	LEHD-TU 2002–12
ψ_D	18.1	Peak derogatory flag rate	.003	.005	LEHD-TU 2002–12
$p_{\bar{h},L}$	.737	Earnings loss 5 years after layoff (%)	−3.6	−8.9	LEHD-TU 2002–12
$p_{\bar{h},H}$	.062	Earnings gain with age (%)	.61	.95	LEHD-TU 2002–12
$p_{\epsilon,L}$	.058	Share of individuals with 9.5% wage decline (%)	4.5	7.7	KM 2017
$p_{\epsilon,H}$	.064	Share of individuals with 9.5% wage increase (%)	14.7	19.0	KM 2017
λ_H	2.37	P75-P25 residual log wage ratio, 25–29-year-olds	.491	.662	LEHD-TU 2002–12
α	.250	Consumption after benefit expiration (%)	86.5	88.0	GN 2019
g	.345	Consumption after layoff (%)	94.8	94.7	PSID 2001–13
$\underline{B}$	−.97	P50 unused credit to income (%)	8.3	8.2	LEHD-TU 2002–12
π_1	.387	Unused-credit quintile 2 share (%)	34.1	31.1	LEHD-TU 2002–12
β_1	.832	Share of individuals borrowing around job loss (%)	23.6	30.6	LEHD-TU 2002–012
δ_C	.024	Credit separation rate (%)	5.1	5.3	TU 2002–12
λ_S	.586	Trough percent earnings loss (%)	−18.0	−17.6	LEHD-TU 2002–12
η	.049	Unemployment by prior earnings slope (%)	6.46	8.70	CPS-ASEC 2002–12

Note.—TU = TransUnion; KM = Kurmann and McEntarfer (2017); Px = xth percentile; GN = Ganong and Noel (2019).

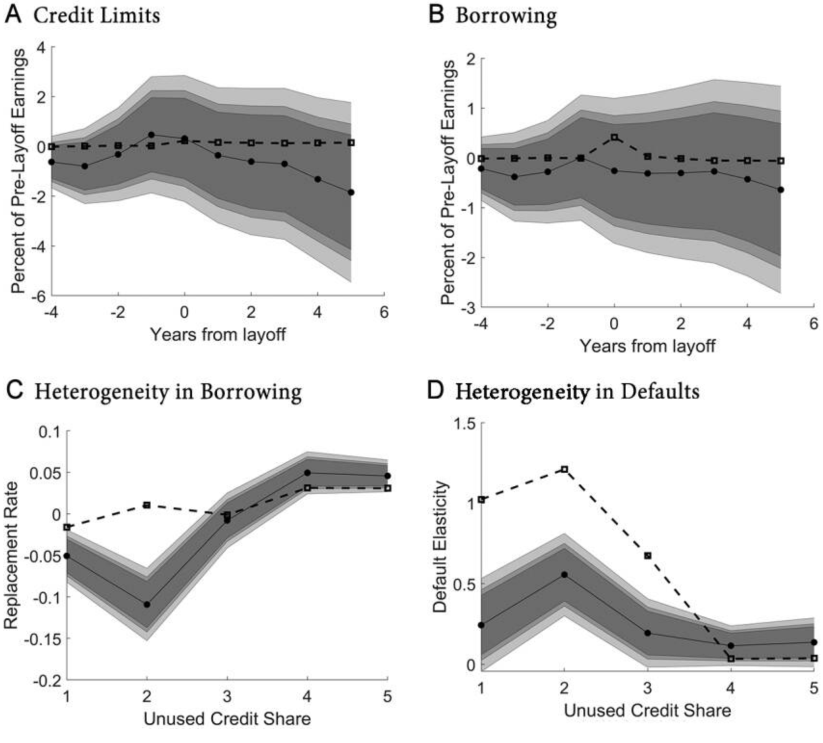

FIG. 6.—Model predictions of credit access and usage around displacement, comparing estimates from the data (solid line with circles) to estimates from the model (dashed line with squares). The darkest shaded region represents the 90% confidence interval, the middle shaded region represents the 95% confidence interval, and the lightest shaded region represents the 99% confidence interval.

to job loss. To make our analysis comparable with our empirical analysis in section I, we estimate the distributed lag regression model given by equation (1) on model-simulated data. We impose the same sampling requirements in the simulation as in the data. In particular, we require individuals to have 3 years of tenure at a firm in order to be in either the treatment or control group.[50]

Figures 6*A* and 6*B* plot the estimated coefficients. To facilitate the comparison between model estimates and data, we normalize reported

[50] We define an individual to be in the treatment group if they are hit by the job separation shock (δ) and satisfy the job tenure requirement. Individuals are defined to be in the control group if they were not hit by the job separation shock and satisfy the job tenure requirement.

coefficients by predisplacement earnings.[51] Figure 6*A* plots credit limits following job loss. Despite the large and persistent decline in earnings, the figure shows that borrowing limits are largely unaffected by job loss. In the model, a small fraction of job losers take out new credit lines to smooth consumption, causing a modest increase in limits. However, the path of borrowing limits is within the 90% confidence interval throughout the 5-year window following job loss.

We next examine the path of borrowing after job loss. Figure 6*B* reveals that debt is largely unresponsive to job loss in both the model and data. Borrowing increases marginally upon job loss but quickly reverts lower than its prelayoff value. In all years following job loss, the path of borrowing is within the 90% confidence interval of the data.

Figures 6*A* and 6*B* show that in our calibrated model, on average, individuals maintain their access to credit (e.g., credit limits do not respond to job loss). As we show below, this muted response of borrowing is masking the fact that some agents borrow significantly in response to job loss, while other agents delever and default following job loss.

Our next exercise measures the heterogeneous response of credit usage following job loss. In this exercise, we define a cross-sectional sample of model-simulated agents exactly as in the data. We require individuals to have 3 years of tenure, a nonzero borrowing limit in the year before job loss, and an earnings loss. We stratify this sample into quintiles on the basis of unused-credit share in the year before job loss. With this sample of simulated agents, we estimate equation (2) using model-simulated data. Figures 6*C* and 6*D* present the results.

Figure 6*C* plots the credit replacement rate in both the model (dashed line with squares) and the data (solid line with circles). Qualitatively, the model replicates the untargeted feature that constrained individuals delever in response to job loss, while the unconstrained borrow. In the data, we estimate that displaced workers in the first unused-credit quintile delever 5.1 cents per dollar of lost income, with a lower bound on the 99% confidence interval of 1.9 cents per dollar. In the model, agents with the lowest amount of unused credit delever and decrease their borrowing by 1.6 cents per dollar of lost income, falling marginally outside the 99% confidence interval. Conversely, agents in the top two quintiles, those with the greatest amount of unused credit, increase their borrowing in response to greater earnings losses. The model predicts that those in the top quintile borrow 3.1 cents per dollar of lost income, whereas the point estimate in the data is 4.6 cents per dollar of lost income. Quantitatively, the model

[51] In app. B.4, we present the path of earnings and default around job loss. As part of the calibration exercise, we target the size of earnings losses upon impact as well as the 5-year earnings loss. Additionally, we target the increase in default propensity upon job loss.

TABLE 4
Percent of Displaced Workers Who Borrow and Delever around Layoff

	Model (%)	Data (%)
Displaced workers with $1,000 decline in revolving-credit balances (untargeted)	17.8	37.3
Displaced workers with $1,000 increase in revolving-credit balances (targeted)	23.6	30.6

Note.—Data are from table A3. Model statistics are computed with identical $1,000 balance change thresholds.

underpredicts deleveraging through default but does well at matching borrowing rates.

Despite understating the elasticity of deleveraging to lost income among constrained agents, the model succeeds at generating roughly half of the unconditional deleveraging observed among workers upon job loss in the data. Table 4 reports that the fraction of agents who delever upon job loss is 17.8% in the model, versus 37.3% in the data. While the model generates a reasonable share of deleveragers and defaulters (shown next), job losers deleverage on relatively small debts. The net result is that the model understates the overall deleveraging elasticity of constrained households.

We next examine the heterogeneous response of defaults to greater earnings losses in the model. Figure 6*D* plots the marginal increase in the probability of default in response to a $10,000 decline in earnings following job loss (where the choice of the simulated earnings loss is solely to facilitate exposition).[52] In the model, agents in the first two quintiles have a significantly higher probability of defaulting, compared to agents in the top three quintiles. Constrained individuals who lose their jobs default in order to deleverage, whereas unconstrained individuals borrow and avoid default.

In the model, the default propensity among those in the lowest quintile of unused credit increases by 1.02% per $10,000 dollars of lost income, versus 0.243% in the data (with an upper bound on the 99% confidence interval of 0.532%). Without additional sources of default risk in the model such as expense, health, or divorce shocks, the only factor driving defaults is earnings losses. Thus, our estimated elasticities are predictably larger than those in the data. However, our response of defaults to job loss (which is targeted) is 0.003 in the model, versus 0.005 in the data.[53] It is only the elasticity of defaults with respect to earnings that is overstated. Nonetheless,

[52] Note that in the data, a $10,000 earnings loss corresponds to a 19% =$10,000/$51,340 decline in earnings relative to the year before layoff. We appropriately scale the model earnings loss so that it corresponds to the same share of prelayoff earnings in the model.

[53] That is, upon job loss the probability of defaulting in the model increases by 0.3 pp, compared with 0.5 pp in the data.

the pattern of declining default sensitivity for those with greater credit access is qualitatively in line with the data.

Overall, we view figure 6 as evidence that the calibrated model is able to match the responsiveness of credit access and usage to job loss both on average and in the cross section. We view the model's ability to reproduce unresponsive borrowing among workers with job loss, despite featuring strong precautionary motives and rising defaults, as providing a validation of the model.[54]

B. 1-Period Debt versus Credit Lines

We next compare our model of credit lines to the standard model of 1-period debt (e.g., Chatterjee et al. 2007; Livshits, MacGee, and Tertilt 2007). In appendix D, we show that for certain parameter values, our framework nests the 1-period-debt model.[55]

There are two important features of the 1-period-debt model: (1) the price of borrowing between a borrower and lender is renegotiated each period, and (2) all nondefaulting individuals have credit access. To accurately measure how the repricing of debt changes credit access following job loss, we perform a ceteris paribus counterfactual. We isolate the set of job losers who have zero assets in our credit-line economy, and we compare their implicit credit limit (defined below) and borrowing paths under the assumption of sole access to (1) credit lines and (2) 1-period debt.[56] In other words, conditional on a fixed state vector, we compare borrowing limits and usage after job loss in the credit-limit and 1-period-debt economies.

We first analyze implicit credit limits after job loss. We define an individual's effective credit limit to be the point at which the interest rate in the 1-period-debt model equals the 90th percentile of real interest rates in the SCF. Figure 7*A* shows that with 1-period debt contracts, credit limits sharply decrease after layoff. As a fraction of annual income, credit limits fall by over 8%.[57] The dramatic reduction in limits reflects greater

[54] In app. B.5, we explore the mechanisms that allow our model to replicate these features of the data.

[55] We thank an anonymous referee for pointing out how our framework connects to the previous literature on 1-period-debt models with default.

[56] We assume that 1-period debt contracts arrive unexpectedly, but from that point forward, agents understand that the only contract available is 1-period debt. We recalibrate the default penalty so that defaults upon job loss are similar across the two economies—in particular, we recalibrate the default cost because the likelihood of default is what determines the implicit amount that an individual is able to borrow via the bond-pricing equation. See app. D.1.3 for more details. We isolate zero-asset workers with job loss to remove mechanical rollover risk effects (i.e., among already-indebted agents, moving to 1-period debt mechanically generates defaults, as agents cannot roll over their debts).

[57] In the baseline model, a small fraction of job losers take out new credit lines to smooth consumption, causing a modest increase in limits around job loss. In app. D.1.1, we examine

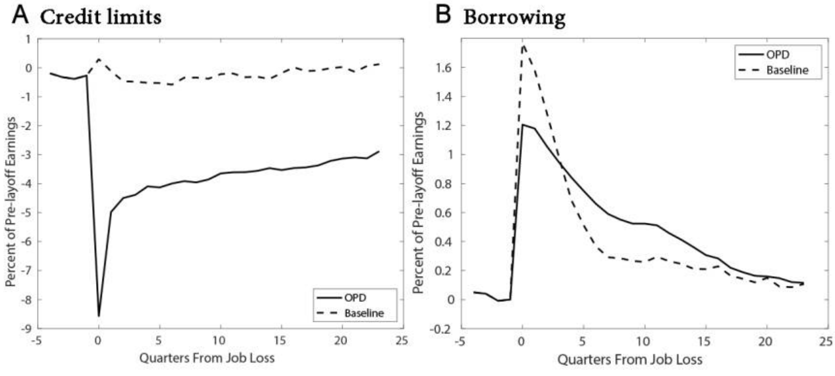

FIG. 7.—Borrowing limits and borrowing behavior in 1-period-debt economy (OPD): counterfactual estimates of borrowing limits (*A*) and borrowing around job loss (*B*) in the baseline economy (dashed line) and in a counterfactual economy where individuals are given a 1-period debt contract in the quarter of job loss (solid line).

default risk and is counterfactual to the stable path of postlayoff limits documented in section I. In the data and the credit-line model, borrowing opportunities are effectively unchanged by job loss, whereas this is not true in the 1-period-debt model. Since the 1-period-debt economy is nested within our credit-line framework, we view these results as providing a strong rejection of the 1-period-debt economy vis-à-vis the data.

Second, we study the borrowing response to job loss in figure 7*B*. Two opposing forces are present in the 1-period-debt economy. The price of credit rises sharply in the period of layoff, discouraging borrowing; on the other hand, all nonbankrupt agents have access to credit, significantly expanding the set of individuals who can (and do) borrow in response to job loss. Figure 7*B* shows that the higher cost of credit is strong enough that there is a weaker borrowing response in the 1-period-debt economy relative to the credit-line economy. Therefore, ceteris paribus, long-term credit lines provide greater self-insurance against job loss than 1-period debt. In the next section, we examine the implications of greater self-insurance via credit lines for the provision of public insurance to the unemployed.

IV. Optimal Public Insurance to the Unemployed

In this section, we compute optimal public transfers to the unemployed under various levels of credit access.

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the heterogeneity in borrowing limits in the counterfactual by human capital and public insurance transfers.

A. *Optimal Policy in Steady State*

We first compute optimal transfers to the unemployed in steady state. We compute optimal policy on the basis of the welfare of a newborn who is "behind the veil of ignorance" and has not yet realized their type (patient or impatient) or human capital draw. Let $F(\vec{h})$ denote the distribution of newborn human capital.[58] Social welfare is given by

$$\mathcal{W} = \int_{\vec{h}} \sum_i \pi_i V_{i,1}^{\mathrm{L}}(0, 0, \vec{h}; 0, 0)\, dF(\vec{h}). \tag{10}$$

We define the optimal replacement rate to be the level of transfers z that maximizes social welfare $\mathcal{W}$ when all parameters except for z are held fixed at their values in table 2. When reporting optimal policy, instead of reporting the level of transfers z, we report the replacement rate of public transfers, which is the average fraction of lost earnings replaced by a given level of transfers z.

As is the case in most optimal UI problems (e.g., Baily 1978 and Chetty 2006), when determining the optimal policy, the government faces a trade-off between the consumption-smoothing benefits of increasing UI and the distortionary effects of payroll taxes and moral hazard of job search.[59]

We first explore our model's ability to replicate the consumption-smoothing benefits of UI. Figure 8*A* plots the consumption of the unemployed as the replacement rate of government transfers varies. We refer to the slope of this profile as the "consumption elasticity." In response to a 10% increase in the public insurance transfer, relative consumption of the unemployed increases by nearly 2.55%.[60] Column 2 of table 5 shows that our model's consumption elasticity is similar to recent estimates from Ganong and Noel (2019).[61]

[58] Note that $\vec{h} = (\tilde{h}, \epsilon)$ and since the newborns are unemployed, $\epsilon = 0$ and thus $F(\vec{h}) = e^{-\lambda_H \tilde{h}}$. We also note that this welfare criterion includes intertemporal considerations generated by the life-cycle profile of human capital.

[59] In app. F.7, we discuss the relationship between our exercise and standard Baily-Chetty formulas. We note that Baily-Chetty formulas allow researchers to assess whether current policy is optimal using a limited set of sufficient statistics; however, these sufficient statistics can be computed only at current levels of US insurance, and thus our structural approach is necessary to compute the optimal replacement rate, which, in our case, differs significantly from current US policy. We also note that Baily-Chetty formulas can be modified to include general equilibrium effects, although Landais, Michaillat, and Saez (2018) show that such modifications are difficult even in simple equilibrium settings.

[60] Note that while our calibration directly targets consumption losses in the PSID using 2-year windows, we compare our model to the on-impact loss of consumption reported in Chodorow-Reich and Karabarbounis (2016) and find that our model produces a decline in consumption within their reported range.

[61] Using webplotdigitalizer, we estimate a 95% confidence interval of 1.95 and 5.31, with a coefficient of 3.63, from fig. 5 of Ganong and Noel (2019).

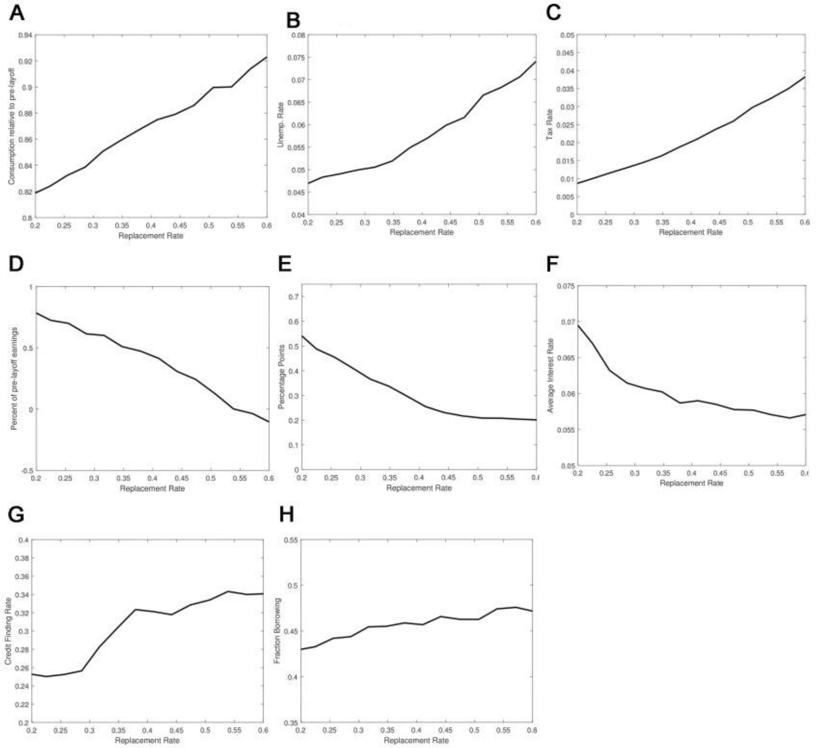

FIG. 8.—Steady-state welfare experiment: steady-state values of model output when all parameters except for z are held fixed at their values in table 2. The panels plot consumption after job loss (c_t/c_{t-1} for those who transit from employment at $t-1$ to unemployment at t; A), the unemployment rate (B), the budget-balancing distortionary labor tax rate (τ; C), borrowing upon job loss, defined to be the date 0 coefficient on the year of layoff from regression equation (1) (D), default upon job loss, defined to be the date 0 coefficient on the year of layoff from regression equation (1) (E), average interest rate r on open credit contracts (F), the credit-finding rate $p(\theta^C)$ averaged over those applying for credit contracts (G), and the fraction borrowing, $b < 0$ (H).

We next explore the strength of moral hazard in our model. As the public insurance transfer increases, the unemployed start to search for jobs that pay a higher wage piece rate (ω), but have a lower job-finding rate. Thus, the model features moral hazard in the labor market because search is directed. In response to an increase in the public insurance transfer, both unemployment (fig. 8B) and unemployment duration increase. As a consequence, as transfers increase, so does the tax rate that balances the government's budget (fig. 8C). Column 2 of table 5 shows that our model's duration elasticity of 0.326 is well within the range of existing estimates (see Krueger and Meyer 2002 for a survey). Moreover, our estimates align well with recent work by Card et al. (2015), who estimate an

TABLE 5
OPTIMAL PUBLIC INSURANCE: CONSUMPTION INSURANCE VERSUS MORAL HAZARD

	Data (1)	Baseline (2)	No Credit (3)	1-Period Debt (4)
Consumption elasticity	.34[a]	.255	.404	.398
Relative consumption unemployed	[.72, .91][b]	.875	.842	.841
Duration elasticity	[.1, 1.0][c]	.326	.334	.330

NOTE.—The current US replacement rate is 41.2% ($z = .470$). "Consumption elasticity" is the elasticity of the unemployed-to-employed consumption ratio with respect to the replacement rate, $d(c_u/c_e)/dz$. "Relative consumption unemployed" refers to the unemployed-to-employed consumption ratio, c_u/c_e. "Duration elasticity" is the elasticity of unemployment duration with respect to the replacement rate, $\epsilon_{D,b}(z)$.

[a] Ganong and Noel (2019), fig. 5*A*.

[b] Chodorow-Reich and Karabarbounis (2016), fig. 4.

[c] Krueger and Meyer (2002).

elasticity of 0.35 for the 2003–7 period and an elasticity in the range of 0.65–0.90 for the 2008–13 period.

Weighing these trade-offs, column 1 of table 6 reports that the optimal replacement rate in our baseline US economy is 34.8%. Patient individuals are willing to give up 0.04% of lifetime consumption to be born in an economy with a 34.8% replacement rate, relative to an economy with a 41.2% replacement rate. On the other hand, impatient individuals prefer current US policy, and their welfare loss from a 34.8% replacement rate equals 0.55% of lifetime consumption. While type-specific consumption-equivalent variation reveals a disproportionate loss by the impatient types, this does not necessarily imply that society is worse off from the proposed policy according to $\mathcal{W}$. In fact, behind the veil of ignorance (i.e., before types are realized), the welfare gain of the patient types marginally outweighs the

TABLE 6
OPTIMAL PUBLIC INSURANCE TO THE UNEMPLOYED

	Baseline Optimal (%) (1)	No Credit (%) (2)	1-Period Debt (%) (3)
Optimal replacement rate	34.8	41.4	40.9
Patient consumption-equivalent welfare	.042	.003	0
Impatient consumption-equivalent welfare	−.548	.049	0
Consumption-equivalent welfare ("behind the veil")	.010	.006	0

NOTE.—"Welfare" is the consumption equivalent of leaving an economy with the US policy of a 41.2% replacement rate to an economy with an alternate replacement rate. For example, in col. 1, the consumption-equivalent welfare change of 0.01% indicates that a newborn individual "behind the veil" would give up 0.01% of lifetime consumption to have a 34.8% replacement rate, as opposed to a 41.2% replacement rate in the baseline model.

welfare loss of the impatient types, resulting in a small positive consumption-equivalent welfare gain worth 0.010% of lifetime consumption.[62]

As we explore in more detail below, the model's consumption elasticity depends crucially on the credit market response to public transfers. Figure 8 shows that as public insurance transfers are cut, credit becomes more costly and the consumption insurance provided by credit diminishes. In our benchmark setting with credit lines, figures 8*D*–8*H* report how the cost and usage of credit vary with transfers. Figure 8*D* shows that as transfers are cut, individuals who lose their jobs increase their borrowing.[63] However, the unemployed also become more likely to default following job loss (fig. 8*E*).[64] As a result, figure 8*F*, shows that the average interest rate on open credit contracts rises steadily as transfers are cut. Compared to the current US replacement of 41.2%, the interest rate increases by roughly 1 pp at a replacement rate of 20%. With lower transfers, the increase in default risk causes lenders to post fewer credit contracts, and the ability of agents to obtain credit declines. Figure 8*G* shows that the credit-finding rate (for both new and existing credit customers) steadily declines as the transfer is cut from the optimum.[65] Finally, figure 8*H* shows that the fraction of individuals who borrow is positively correlated to the transfer rate. With low transfers, credit becomes more costly, individuals build greater asset buffers, and individuals borrow less before job loss.[66]

B. *Credit and Optimal Policy*

In this section, we first explore how the extensive margin of credit access shapes optimal policy. We then show that modeling credit lines—as opposed to 1-period debt—is essential for optimal policy.

[62] In app. F.1, we discuss type-specific policies.

[63] In fig. 8*D*, we present the coefficient on the year of layoff from estimating eq. (1) separately for each level of the public insurance transfer, where the dependent variable is borrowing. This coefficient estimates the amount of borrowing upon job loss by level of public insurance transfer. To ease interpretation, we present this coefficient as a share of prelayoff earnings.

[64] In fig. 8*E*, we present the coefficient on the year of layoff from estimating eq. (1) separately for each level of the public insurance transfer, where the dependent variable is defaults. This coefficient estimates the probability of defaulting upon job loss by the level of the public insurance transfer. Hsu, Matsa, and Melzer (2018) provide empirical support for this mechanism, by exploiting geographic variation in UI generosity and showing that in response to increases in UI generosity individuals become less likely to default on mortgage payments.

[65] In calibrating the model economy, we use the credit-finding rate for individuals who did not previously have a credit line, since it can be cleanly measured in the TransUnion data. However, to examine credit access in the model economy, the overall credit-finding rate, which incorporates applications from individuals without a credit line as well as individuals engaging in-the-contract credit search, is the appropriate measure.

[66] Recent work by Bornstein and Indarte (2022) documents complementarity between medicaid expansions and borrowing.

1. Optimal Policy with Zero Credit

We now counterfactually close credit markets (i.e., no borrowing, $\underline{\mathcal{B}} = \{0\}$, and thus $\underline{b} = 0$ for all contracts) and recompute optimal steady-state transfers. This exercise allows us to study how optimal policy interacts with the presence of credit markets and the degree to which credit lines allow the government to substitute away from public insurance. Column 2 of table 6 reports our results. The optimal replacement rate increases to 41.4% of lost earnings when credit markets are shut down. Therefore, public insurance replacement rates can be cut by 6.6 pp (41.4% − 34.8%) as we move from a steady state in which 0% of individuals have access to credit (col. 2 of table 6) to a steady state in which 77.5% of individuals have access to credit (col. 1 of table 6).[67]

What drives the higher optimal replacement rate in the economy with no credit? Column 3 of table 5 reports the consumption and duration elasticities in the "no-credit" model at the current US level of transfers. Without access to credit, the consumption elasticity with respect to transfers is 0.404, which is over 50% greater than that in our baseline economy. Furthermore, upon job loss, consumption declines by over 3 pp more when credit markets are turned off.

Figure 9 illustrates the relationship between the relative consumption of the unemployed and transfers for the baseline economy (solid line) and the no-credit economy (dashed line). Without credit access, the consumption of the unemployed is more sensitive to transfers and lower, relative to the baseline economy at all replacement rates. The utilitarian government is less able to substitute away from public insurance in the absence of credit because consumption insurance "dries up" faster as public insurance is cut.

On the other hand, column 3 of table 5 shows that the duration elasticity remains approximately unchanged. Thus, consumption insurance deteriorates in the absence of credit, while the moral hazard effects of insurance are unaltered. These forces yield higher optimal public transfers.[68]

[67] In app. F.6, we show that cutting the replacement rate of public insurance transfers to 34.8% (the optimal policy in the baseline economy) results in a welfare loss equivalent to nearly 0.07% of lifetime consumption in the no-credit economy. This welfare loss is nearly an order of magnitude larger than the oppositely signed gains from moving to the optimal policy in the no-credit economy.

[68] In app. F.3, we show that the size of the gap between optimal policies in the baseline credit-line economy and the no-credit economy is only moderately weaker with a lower share of impatient agents. In particular, we recalibrate our baseline economy, imposing that only 20% of agents are impatient—i.e., $\pi_1 = 0.20$—and reperform our welfare exercise. While the level of the optimal replacement rate differs from our main-text calibration, we find that the optimal policy is a 5-pp-lower replacement rate in the economy in which 78% of individuals have credit access, relative to an economy in which 0% of individuals have credit access. Thus, the degree of substitutability between public insurance and private credit is moderately weaker. Additionally, in app. F.5, we show that we obtain similar results for the size of the gap between optimal policies in the credit-line economy and the no-credit economy with a lower value of the risk-free rate.

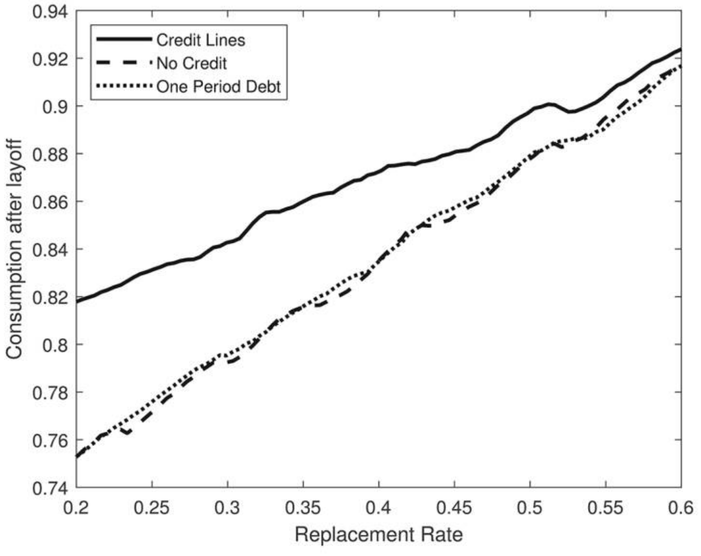

FIG. 9.—Relative consumption upon job loss: consumption after a layoff (y-axis) by replacement rate of public insurance transfers (x-axis) in the baseline credit-lines version of the mode (solid line), the no-credit version of the model (dashed line), and the 1-period-debt version of the model (dotted line).

2. Optimal Policy with 1-Period Debt

To further understand the role of credit lines in shaping the optimal transfer to the unemployed, we repeat our optimal policy exercise with 1-period debt contracts. Importantly, with 1-period debt, the terms of the credit contract are renegotiated every period on the basis of an agent's state.[69]

Column 3 of table 6 shows that the optimal public insurance transfer to the unemployed in an environment with 1-period debt replaces 40.9% of lost earnings.[70] The optimal replacement rate with 1-period debt is only 0.5 pp lower than the optimal policy when credit markets are counterfactually shut off. Alternatively, in the baseline model of credit lines, the optimal replacement rate is 6.6 pp lower than the no-credit benchmark. Therefore, with 1-period debt, there is less ability to substitute away from public insurance to the unemployed, relative to the credit-line economy.

[69] As we discuss in sec. III.B, our framework nests a 1-period-debt model.

[70] As we discuss in sec. III.B, in the 1-period-debt economy we recalibrate the default penalty so that defaults upon job loss are similar to those in the baseline model of credit lines. We recalibrate the default cost because the likelihood of default is what determines the implicit amount that an individual can borrow via the bond-pricing equation. See app. D.1.3 for more details. Additionally, note that our baseline transfer corresponds to a 40.9% replacement rate in the 1-period-debt economy.

To understand what yields such limited substitutability, figure 9 compares consumption after a layoff as public transfers vary in a 1-period-credit version of the model (dotted line) and the baseline credit-lines version of the model (solid line). The steeper slope of the line for the 1-period-debt model indicates that there is less consumption insurance available to households in the 1-period-debt model as public insurance transfers are cut. As transfers are cut, consumption insurance "dries up" at a rate very similar to that in the economy without credit. Consequently, there is very little substitutability between credit and public transfers in the economy with 1-period debt.

The reason why credit lines provide more insurance relative to 1-period debt is that long-term credit lines are established (most often) when an individual is employed. These credit lines do not respond to income changes (or transfer changes) as much as 1-period debt. Recall that in figure 7*A*, we showed that borrowing limits decline substantially upon job loss in the 1-period-debt economy, whereas limits are stable in the credit-line economy, which is consistent with our empirical findings in figure 1. As a result, the consumption elasticity with respect to transfers is greater in the economy with 1-period debt (see col. 4 of table 5).

On the other hand, column 4 of table 5 reveals that moral hazard effects in the 1-period-debt economy are similar to the credit-line economy. Since the trade-off between consumption insurance and distortionary taxes guides the optimal choice of transfers, we find that there is a lower optimal replacement rate with credit lines than with 1-period loans.[71]

C. Transition Path

Our final exercise computes welfare gains along the transition path when public insurance is unexpectedly and permanently cut from the current US replacement rate of 41.2% to 34.8%. We allow taxes to adjust each quarter after the transition begins in order to balance the government budget, and agents have perfect foresight over the path of taxes and benefits once the transition begins. We measure welfare along the transition path using a utilitarian welfare criterion for individuals alive at the time of the policy reform. We provide details of the transition experiment in appendix E.

We find that those alive at the time of the transition experience a welfare gain worth 0.03% of remaining lifetime consumption. When we stratify by age, we find that the largest welfare gains accrue to middle-aged individuals. The extremes of the age distribution are less likely to be

[71] In app. F.2, we further examine how the characteristics of credit lines shape optimal public insurance transfers to the unemployed, by varying the cost of searching in the credit market and the size of credit limits.

employed and have lower stocks of precautionary savings, generating weaker welfare gains for the young and welfare losses for the old. The welfare gains along the transition path are larger than those computed “behind the veil” across steady states; however, the gains from implementing a 34.8% replacement rate remain economically small.

V. Conclusions

In this paper, we empirically and theoretically examine the extent to which personal credit can be used as a substitute for public UI to insure workers’ consumption upon job loss. Empirically, we link TransUnion credit reports to administrative employment records to examine borrowing by workers upon job loss. We build a model and match our model’s parameters to key moments from our empirical work to measure the gains from reoptimizing public insurance in the presence of private credit markets.

We contribute novel empirical results showing that workers who lose their jobs maintain access to credit and that unconstrained workers who lose their jobs borrow, while constrained workers who lose their jobs default and delever. We thus show that there is important heterogeneity in borrowing by displaced workers. While displaced workers do not borrow on average, we show that roughly one-third of displaced workers default and delever and that roughly one-third of displaced workers borrow more. Thus, credit markets are important for both sets of workers in their borrowing and consumption decisions. These results reconcile previous conflicting results, as studies based on checking-account data suggest that there is roughly zero net borrowing, on average, by workers who lose their jobs, while direct questions about borrowing among workers who lose their jobs and other survey data imply that roughly 20% of the unemployed borrow and roughly 30% become delinquent on debt obligations.

We use these moments to estimate our theoretic framework that integrates credit lines (e.g., Mateos-Planas and Ríos-Rull 2010) into a competitive labor search model with employment risk (e.g., Moen 1997; Burdett, Shi, and Wright 2001; Menzio and Shi 2011). We show that the model simultaneously generates the main patterns of borrowing and default observed among displaced workers in our data: (1) credit limits unresponsive to job loss, (2) deleveraging and defaults among constrained workers, and (3) borrowing and repayment among unconstrained workers.

We find that under US levels of credit access observed between 2002 and 2012, the optimal UI replacement rate is 34.8%, versus the current US replacement rate of 41.2%. We find that the welfare gains from reoptimizing public transfers are economically small but that further reductions in UI generate increases in the cost of credit, leading to significantly weaker consumption insurance for the unemployed. When credit markets are shut down, we find that replacement rates are 6.6 pp higher. We then show that modeling credit lines is essential for optimal policy by reestimating the substitutability between public transfers and private credit in an economy with 1-period debt contracts. With 1-period debt, we show that consumption is more sensitive to changes in transfers, implying less substitutability between public insurance and private credit.

Several avenues for future research exist. First, the presence of business cycles may further limit the ability of the US government to substitute between credit and UI. The block-recursive structure of our paper makes it possible to tackle such questions. Second, we believe that our theory may explain the lack of private credit markets in developing countries in which safety nets are limited or nonexistent. The long-term credit model developed in this paper is flexible enough to study a variety of safety net programs, allowing future researchers to model a variety of institutional details while accurately capturing the way credit and labor markets interact. In concurrent work, we are using credit bureau data and modifying the model framework to (i) identify permanent and transitory income processes (Braxton et al. 2022) and (ii) study the impact of credit access on earnings mobility (Braxton et al. 2024).

Data Availability

The steps to access the proprietary databases used in this paper, as well as the codes necessary to replicate all tables and figures, are available in Braxton, Herkenhoff, and Phillips (2023), in the Harvard Dataverse, https://doi.org/10.7910/DVN/TGULMG.

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