Do Unemployment Benefit Extensions Explain the Emergence of Jobless Recoveries?*

Kurt Mitman† Stanislav Rabinovich‡

March, 2024

Abstract

Countercyclical unemployment benefit extensions in the United States act as a propagation mechanism, contributing to the high persistence of unemployment following recent recessions, as well as the weak correlation between unemployment and productivity. We show this by modifying an otherwise standard frictional model of the labor market to incorporate a stochastic and state-dependent process for unemployment insurance estimated on US data. Accounting for movements in both productivity and unemployment insurance, our calibrated model is consistent with post-war labor-market dynamics. It explains the emergence of jobless recoveries in the 1990s, the low correlation between unemployment and productivity, and the apparent shifts in the Beveridge curve following recessions.

Keywords: Business Cycles, Jobless Recoveries, Unemployment Insurance
JEL codes: E24, E32, J65

*We thank our discussants Tobias Broer and Moritz Kuhn as well as the editors for helpful feedback on the paper. We are grateful to colleagues at numerous institutions for constructive comments that significantly improved the paper.

†Corresponding author. Institute for International Economic Studies, Stockholm University, SE-106 91, Stockholm, Sweden. Email: kurt.mitman@iies.su.se

‡UNC Chapel Hill. Email: srabinov@email.unc.edu.

1 Introduction

Unemployment is not only volatile but also sluggish in recovering in the aftermath of a recession, as evidenced by the jobless recoveries following the recessions of 1990-1991, 2001, and 2007-2009. Moreover, as documented by Graetz and Michaels (2017), the emergence of jobless recoveries in these decades is unique to the US. In this paper, we assess the extent to which the unique features of the US unemployment insurance system — namely countercyclical and unemployment-dependent extensions of unemployment benefits — can account for the observed unemployment dynamics.

Explaining sluggish employment recoveries is distinct from the agenda of explaining the volatility of unemployment, which has been a focus of the macro-labor literature since at least Andolfatto (1996) and Shimer (2005) and has met with arguably mixed success. Indeed, the literature on labor market dynamics faces tension when trying to match both the volatility of unemployment and its persistence following recessions. Matching the high volatility of unemployment requires unemployment to be very sensitive to labor productivity, which implies, counterfactually, that the recovery of the labor market would mimic the recovery of output per worker. In a similar vein, the latter would imply a counterfactually high correlation between unemployment and productivity. The data does not support this and suggests the presence of a countercyclical and persistent labor wedge (Hall (1997), Gali (1999), Chari *et al.* (2007), Chang and Kim (2007), Shimer (2009), Ohanian (2010)). One interpretation of such a labor wedge is as a countercyclical tax (implicit or explicit) on labor. Most of the literature trying to explain jobless recoveries has, however, dismissed the literal interpretation of the labor wedge as a policy wedge because such countercyclical policy distortions are not observed in practice (or are small). In this paper, we recognize that such a countercyclical policy wedge does exist in the form of unemployment insurance. We evaluate the importance of this wedge quantitatively in explaining post-war labor-market dynamics in the US.

The mechanism we propose is simple and based on the now-standard equilibrium search model, in conjunction with the unemployment insurance (UI) system in the United States. The latter features automatic triggers that increase the duration of unemployment benefits during periods of high unemployment. Moreover, in all but one of the previous eight recessions, the government has enacted discretionary policies that extended UI benefit duration further. Crucially, because unemployment benefit duration depends on the unemployment rate, which lags productivity, high benefit duration persists long after labor productivity begins to recover following a recession. The standard search model predicts that unemployment insurance increases unemployment through reduced firm vacancy creation—because unemployment benefits lower the surplus from forming a match between a worker and a firm.[1] Thus, unemployment benefit extensions further propagate high unemployment, forestalling the recovery in the labor market even at a time when productivity has recovered. This mechanism has the potential to simultaneously increase the post-recession persistence of unemployment and reduce its correlation with productivity. The challenge is to quantify the contribution of this mechanism. The empirical literature on cross-sectional effects of unemployment benefits on unemployment does not address this, since it says nothing about either the aggregate effects or, most importantly, the timing. We address this using a dynamic model of the labor market.

Introducing countercyclical unemployment insurance into the equilibrium search model requires taking a stand on expectations. While there is a systematic component to unemployment benefits written into law (that automatically extends benefits when unemployment is high), they by no means depend deterministically on unemployment. In fact, while unemployment benefits typically get extended in recessions, the size and duration of extensions varies, and became progressively higher over the last 50 years. Moreover, as evidenced by the Great Recession of 2007-2009, Congress may reauthorize unemployment benefit extensions in a discretionary manner when they are scheduled to expire or after they expired and then have them apply retroactively. We incorporate these observations in a rational-expectations framework by assuming a stochastic state-dependent process for unemployment benefits, which we discipline from the data on actual unemployment benefit extensions.

We quantitatively evaluate the importance of this channel in a calibrated search

[1] We abstract from worker search effort in this paper, but standard search theory predicts that search effort should also decrease in response to increases in unemployment benefit duration.

model by simulating the series of productivity shocks observed in the 1960-2016 period and sequentially introducing the unemployment benefit extensions enacted during this period. We find that the model accounts well for the observed time series of unemployment, explaining between 35-47% of unemployment fluctuations (as measured by the R^2) over our sample. In particular, the model-generated recoveries were not jobless prior to 1990 and became jobless thereafter (though, all structural parameters in the model remained unchanged across the two time periods). The key to generating this result is the fact that the UI benefit extensions enacted after the recessions of 1990-1991, 2001, and 2007-2009 were large *relative* to the productivity recovery following these recessions (see figure 1). We also conduct counterfactual experiments to quantify the importance of the extensions: specifically, we examine how the cyclical behavior of unemployment would have been different had the extensions not occurred under two different scenarios. First, we take our baseline calibration and turn off benefits extensions. The model predicts a much faster recovery of employment if the unemployment benefit extensions are not enacted. Without benefit extensions, the model only explains 3-10% of the fluctuations in unemployment (as measured by the R^2), thus we attribute roughly 32-37% of the fluctuations in unemployment over this time period to unemployment benefit extensions. Under this counterfactual the model also fails do deliver amplification—the volatility of unemployment and labor market tightness is only 30% of that in the data. To make clear that solving the “volatility puzzle” by matching standard deviations of labor market variables does not imply the emergence of jobless recoveries we perform a second counterfactual. We recalibrate the model without unemployment benefit extensions to match the volatility of unemployment in the data. Despite solving the ”volatility puzzle”, this version of the model can explain at most 7% of fluctuations in employment in the data (as measured by the R^2). The poor performance of the two counterfactual models can be accounted for by the post-1985 period: the R^2 for both counterfactual models is negative over that period because they fail to account for the persistent unemployment and jobless recoveries in that time period.

In addition to matching the unemployment dynamics, we find that the model accounts for the dynamics of the Beveridge curve during recessions, including the apparent shift in the Beveridge curve observed following the 2007-2009 recession. The Beveridge curve — the observed negative correlation between unemployment and vacancies — is a robust feature of the post-war labor market, but weakens in the aftermath of recessions, in particular the Great Recession. We show that our simulated model reproduces an unemployment-vacancy correlation very similar to the one observed in the data — including the 2007-2013 period, during which the model reproduces the perceived shift in the simulated Beveridge curve.[2] In other words, the large unemployment benefit extensions implemented during this period acted as shocks that induced a substantial departure from the theoretical Beveridge curve, making it appear as if the curve itself shifted, although all the parameters of our model, including the matching function, have remained the same. Through the same mechanism, our model reproduces the perceived shifts in the Beveridge curve in the previous recessions as well.[3] In summary, we argue that countercyclical unemployment benefit extensions are a powerful endogenous propagation mechanism, capable of reconciling the standard search model with salient evidence on sluggish labor market recoveries.

1.1 Related Literature

This paper contributes to the literature rationalizing labor market dynamics in the US, most concretely the sluggish recoveries following recessions. We argue that incorporating time-varying and state-dependent unemployment insurance remarkably improves the model's ability to match observed dynamics; most notably this helps the model generate high unemployment persistence and low employment-productivity correlation without sacrificing on volatility. This sets our analysis apart from the large body of research that tries to explain the high volatility of unemployment, following the Shimer (2005) puzzle. As emphasized by Ljungqvist and Sargent (2017),

[2]We are not the first to suggest that unemployment benefit extensions contributed to the dynamics of the Beveridge curve during the Great Recession. Hobijn and Şahin (2013) and Daly *et al.* (2012) partially attribute the such shift to UI policy.

[3]Such shifts, and in particular the fact that they are not unique to the Great Recession, are well documented; see e.g. Diamond and Şahin (2015).

the size of the “fundamental surplus” from an employment match is crucial for unemployment volatility: unemployment is sensitive to productivity, and hence volatile, when this surplus is small. Achieving this sensitivity of unemployment to productivity, however, usually comes at the cost of a high unemployment-productivity correlation, and correspondingly low unemployment persistence that mimics that of productivity. In fact, nearly all of the theories put forth by the literature to resolve the Shimer puzzle (e.g. wage rigidities in Hall (2005), small surplus in Hagedorn and Manovskii (2008), Mitman and Rabinovich (2015), and Ljungqvist and Sargent (2017), or marginal worker-firm matches in Costain and Reiter (2008) and Menzio and Shi (2011)) feature a counterfactual correlation between labor productivity and unemployment close to unity. In contrast, our paper correctly predicts a much lower correlation between productivity and unemployment. We achieve this by introducing state-dependent unemployment insurance, which - as in the data - is a function of unemployment, a slow-moving variable that lags productivity. The quantitative success of the model is evidenced by the fact that it accounts well for the entire time series of unemployment. In particular, it correctly predicts the timing, not just the volatility, of unemployment dynamics, specifically the sluggish recovery of employment in the aftermath of certain recessions.

Our analysis contributes to an important line of research attempting to account for jobless recoveries: e.g. labor market polarization as in Jaimovich and Siu (2018) or Gaggl and Kaufmann (2019); labor hoarding as in Bachmann (2011); or countercyclical restructuring behavior of firms as in Berger (2011). Our paper is particularly close in nature to the innovative work by Herkenhoff (2019), who argues that increased access to credit has led households to take on more debt and be pickier about finding jobs, leading to slower recoveries. Similarly to that paper, we share the view that changes in the value of non-employment are important for generating observed unemployment patterns; our paper is distinct in emphasizing unemployment benefit extensions as the driving mechanism.

To our knowledge, Nakajima (2012) is the first quantitative model-based evaluation of the role of unemployment insurance in the Great Recession. Our results are fully consistent with the results of Nakajima (2012), in that unemployment benefit extensions are, in fact, very important quantitatively. We also share with that paper the methodology of using a quantitative model, calibrated to match cross-sectional evidence on the effects of UI, to conduct policy experiments. However, our methods and results differ along a number of salient dimensions. First, the philosophy of the experiment is quite different: Nakajima (2012) infers the path of productivity as a residual to exactly match the empirical path of the unemployment. In other words, the experiment evaluates the contribution of UI to the unemployment path, when the inferred productivity series ensures that the model with UI matches this path exactly. The inferred productivity path is strongly negatively correlated with the unemployment rate, contrary the data. As such, his study is not a test of whether unemployment benefit extensions act as an endogenous time-varying labor wedge that propagates productivity shocks, amplifying and propagating the unemployment response. Second, Nakajima (2012) studies only the Great Recession episode, as opposed to the full post-War time series for unemployment. We see our paper as addressing a fundamentally different question — assessing the role of UI as an endogenous propagation mechanism of observed productivity shocks — but view this as quite complementary to his contribution.

Our paper is likewise very different from our own prior work, Mitman and Rabinovich (2015), which studies optimal unemployment insurance over the business cycle. The model used there did incorporate countercyclical unemployment benefits, but — crucially — assumes that benefits are an exogenous function of productivity, not unemployment. This difference is vital to the model's ability to generate unemployment persistence. In fact, the model of Mitman and Rabinovich (2015) cannot generate the jobless recoveries observed in the data for precisely this reason.[4] To use the language of Ljungqvist and Sargent (2017), it is the persistence of the fundamental surplus, not merely its magnitude or even cyclicality, that matters for the ability of the model to generate jobless recoveries; here, we achieve a persistently low fundamental surplus post-recessions by indexing unemployment benefits

[4]The calibration strategy of Mitman and Rabinovich (2015) is to target unemployment volatility, which, as described previously, is insufficient for matching the timing of unemployment dynamics.

to unemployment, as they are in the data.[5] Moreover, rather than simply match second moments, we show that the simulated model successfully reproduces the entire (non-targeted) time series of unemployment.

More broadly, we contribute to the literature assessing the ability of productivity-driven business cycle models to explain observed labor market dynamics. The sluggishness of employment recoveries and the weak employment-productivity correlation suggest that productivity-driven models struggle to do so, motivating the literature on the labor wedge (e.g. Chari *et al.* (2007), Chang and Kim (2007), Shimer (2009), among others). From this point of view, our paper can be thought of as proposing and quantitatively evaluating a particular source of an endogenous labor wedge - countercyclical unemployment insurance.

In Section 2 we describe the model environment with time-varying unemployment benefits. We discuss how we bring the model to the data in Section 3 and describe the results and counterfactuals in Section 4. Section 6 concludes.

2 Model

Time is discrete and the time horizon is infinite. The economy is populated by a unit measure of workers and a larger continuum of firms. The model is one of "perpetual youth", where workers leave the economy stochastically at rate $1-\mu$ and are replaced with an identical mass of newborn workers. In any given period, a worker can be either employed (matched with a firm) or unemployed. Newborns begin life unemployed. Workers are risk-neutral expected utility maximizers. A worker born

[5] Chodorow-Reich and Karabarbounis (2016) argue that the fundamental surplus is pro-cyclical, calling into question the ability of search and matching models to match the volatility and persistence of unemployment. In their exercise, they consider the value of unemployment benefits to be 0.041 of the marginal product of employment, far lower than the actual replacement rate conditional on take-up of 45%. Controlling for eligibility ($\approx 60\%$) and take-up ($\approx 75\%$), a more accurate value for the value of unemployment benefits would be $0.45 \times 0.6 \times 0.75 = 0.223$, which would reverse their results on the cyclicality of the opportunity cost of unemployment. With the relevant measure of unemployment benefits, their exercise is fully consistent with our paper.

in period τ has expected lifetime utility

$$U = \mathbb{E}_\tau \sum_{t=\tau}^{\infty} (\beta\mu)^{t-\tau} x_t,$$

where $\mathbb{E}_\tau$ is the period-τ expectation operator, $\beta \in (0,1)$ is the discount factor, and x_t denotes consumption in period t. An unemployed worker produces h, which stands for the combined value of leisure and home production. Firms are risk-neutral, maximize profits, and have the same discount factor β. A firm can be either matched to a worker or vacant. A firm posting a vacancy incurs a flow cost k.

Unemployed workers and vacancies match in pairs to produce output. The number of new matches in period t equals

$$M(u_t, v_t),$$

where u_t is the unemployment level in period t and v_t is the measure of vacancies posted in period t. The matching function M exhibits constant returns to scale, and is strictly increasing and strictly concave in both arguments. Define $\theta_t = v_t/u_t$ to be the market tightness in period t. A worker finds a job with probability $f(\theta_t)$, where $f(\theta_t) = M(1, \theta_t)$, and a firm fills its vacancy with probability $q(\theta_t)$, where $q(\theta_t) = M(1/\theta_t, 1)$. Existing matches are exogenously destroyed with a constant job separation probability δ. Thus, any of the $1 - u_t$ workers employed in period t has a probability δ of becoming unemployed in period $t + 1$.

All worker-firm matches are identical in terms of their productivity. Specifically, each matched worker-firm pair produces output z_t in period t, where z_t is aggregate labor productivity. We assume that $\ln z_t$ follows an AR(1) process

$$\ln z_t = \rho \ln z_{t-1} + \sigma_\varepsilon \varepsilon_t, \tag{1}$$

where $0 \leq \rho < 1$, $\sigma_\varepsilon > 0$, and ε_t are independent and identically distributed standard normal random variables.

The government provides unemployment benefits to unemployed workers, which are described by a level level b, a re-entitlement rate ϱ, and duration D_t. The level b and the re-entitlement rate ϱ are assumed fixed, while D_t is time-varying and state-dependent; specifically, the transition process for D_t is allowed to depend on both u_t and z_t. UI eligibility status of a worker evolves as follows. Every *employed* worker can, at each point in time, be of two types: eligible or ineligible for unemployment benefits. Every *unemployed* worker can, at each point in time, be of three types: an eligible claimant, an eligible non-claimant, or ineligible. When an eligible employed worker loses his job, the worker has some probability P of taking up unemployment benefits, in which case he becomes a claimant; otherwise, he remains an eligible non-claimant. Similarly, an unemployed worker who is already an eligible non-claimant takes up benefits every period with the same probability P; otherwise, he remains an eligible non-claimant the following period. On the other hand, a job loser who was ineligible for benefits remains ineligible for the duration of the unemployment spell. An ineligible employed worker becomes eligible with probability ϱ. Finally, all new entrants start out as unemployed and ineligible.

We next describe how unemployment benefits expire for an unemployed worker currently claiming them. Contrary to the previous literature, we explicitly model deterministic unemployment benefit expiration. Therefore, the number of months of unemployment benefits remaining is an individual state variable for an unemployed UI claimant. Suppose that an unemployed worker currently receiving UI has n_t months remaining, and the current duration policy is D_t. Then, if the worker does not find a job this period, next period he is entitled to n_{t+1} months of UI, where

$$n_{t+1} = \max\{0, \max\{0, n_t - 1\} + D_{t+1} - D_t\} \tag{2}$$

To understand this law of motion, let us suppose first that D_t is fixed over time (i.e. there are no benefit extensions). Then the law of motion for an individual's months remaining is simply $n_{t+1} = \max\{0, n_t - 1\}$, i.e. it falls by 1 every month until reaching zero. On the other hand, an extension of unemployment benefits (an event where $D_{t+1} > D_t$) raises the entitlement length by an additional $D_{t+1} - D_t$ months. Finally, a cut in economy-wide benefit duration (an event where $D_{t+1} < D_t$, which occurs in case a UI extension is terminated) may render an otherwise eligible worker ineligible. For example, if a worker had 2 months of UI left, his UI entitlement drops to zero if economy-wide UI duration falls from 9 months to 6 months.

2.1 Recursive formulation

We proceed to characterize the environment recursively. At each point in time, the aggregate state of the economy, denoted by Ω, consists of aggregate productivity z, the current duration policy D, and the entire distribution of workers across employment and UI eligibility states. Given this distribution, one can compute the unemployment rate, u, and the fraction of unemployed workers eligible for UI benefits, ϖ. The timing is as follows. The economy enters the period with an aggregate state Ω. Firms then post vacancies v. Job matching is then realized, simultaneously with job separations. At this stage, wages are bargained and production takes place. Unemployed non-claimant workers eligible for UI then receive idiosyncratic shocks that determine whether they start claiming this period. Afterwards, labor market exit shocks hit and workers find out if they remain in the labor market the following period. This determines the new unemployment level

$$u' = \underbrace{1-\mu}_{\text{new entrants}} + \mu[\underbrace{\delta(1-u)}_{\text{job losers}} + \underbrace{(1-f(\theta))u}_{\text{non-finders}}] \tag{3}$$

At this point, a new aggregate productivity z' is drawn, and D' is drawn conditional on u' and D. Each unemployed worker's eligibility status n' is then updated based on the new D'.

2.2 Firm value functions

Consider a firm employing a worker eligible for UI. Such a firm has the value

$$J^E(\Omega) = z - w^E(\Omega) + \beta\mu(1-\delta)\mathbb{E}_{z',D'|u'}J^E(\Omega') \tag{4}$$

where $w^E(\Omega)$ is the bargained wage of such a worker, described below. A firm employing an ineligible worker has the value

$$J^I(\Omega) = z - w^I(\Omega) + \beta\mu(1-\delta)\mathbb{E}_{z',D'|u'}\left[\varrho J^E(\Omega') + (1-\varrho)J^I(\Omega')\right] \tag{5}$$

where $w^I(\Omega)$ is the bargained wage of an ineligible worker. The match survives to the next period with probability $\mu(1-\delta)$, which is the probability that the worker does not retire and does not separate into unemployment. If the worker was ineligible this period and the match survives, the worker becomes eligible the next period with probability ϱ.

Next, the value of posting a vacancy equals

$$V(\Omega) = -k + q(\theta(\Omega))\left[\varpi J^E(\Omega) + (1-\varpi)J^I(\Omega)\right] + \beta(1-q(\theta(\Omega))\mathbb{E}_{z',D'|u'}V(\Omega') \tag{6}$$

where ϖ is the fraction of UI-eligible workers among the unemployed, and hence the probability that, conditional on filling the vacancy, the worker hired is UI-eligible. Free entry of vacancies implies that $V(\Omega) = 0$ always, so that

$$k = q(\theta(\Omega))\left[\varpi J^E(\Omega) + (1-\varpi)J^I(\Omega)\right] \tag{7}$$

2.3 Worker value functions

We now turn to the workers' value functions. Denote by W^E and W^I the value functions of an employed worker eligible and ineligible for benefits, respectively. An employed worker eligible for benefits has the value

$$W^E(\Omega) = w^E(\Omega) + \beta\mu(1-\delta)\mathbb{E}_{z',D'|u'}W^E(\Omega') + \beta\mu\delta\mathbb{E}_{z',D'|u'}U^E(\Omega') \tag{8}$$

where U^E is the value function of an unemployed worker eligible for benefits, determined below. Similarly, an employed worker ineligible for benefits has the value

$$\begin{aligned} W^I(\Omega) = & w^I(\Omega) + \beta\mu(1-\delta)\mathbb{E}_{z',D'|u'}\left[\varrho W^E(\Omega') + (1-\varrho)W^I(\Omega')\right] \\ & + \beta\mu\delta\mathbb{E}_{z',D'|u'}U^I(\Omega') \end{aligned} \tag{9}$$

where U^I is the value function of an unemployed worker eligible for benefits. Now, consider the value functions of the unemployed. An ineligible unemployed worker has the value

$$\begin{aligned} U^I(\Omega) =& h + \beta\mu\mathbb{E}_{z',D'|u'} f(\theta'(\Omega')) W^I(\Omega') \\ &+ \beta\mu\mathbb{E}_{z',D'|u'} (1 - f(\theta'(\Omega'))) U^I(\Omega') \end{aligned} \tag{10}$$

In other words, such a worker receives only h while unemployed, and remains ineligible the following period. On the other hand, the value of an eligible unemployed worker is

$$U^E(\Omega) = PU^C(D, \Omega) + (1 - P) U^{NC}(\Omega) \tag{11}$$

An eligible unemployed worker becomes a *claimant* with probability P. We denoted by $U^C(n, \Omega)$ the value of a claimant with n months of UI remaining. A new claimant starts with the maximum allowable duration of UI, which is D. We denote by $U^{NC}(\Omega)$ the value function of an eligible non-claimant. The latter is given by

$$\begin{aligned} U^{NC}(\Omega) =& h + \beta\mu\mathbb{E}_{z',D'|u'} f(\theta'(\Omega')) W^E(\Omega') \\ &+ \beta\mu\mathbb{E}_{z',D'|u'} (1 - f(\theta'(\Omega'))) U^E(\Omega') \end{aligned} \tag{12}$$

In other words, an eligible non-claimant does not *receive* unemployment benefits, but may claim the next period if remaining unemployed. If finding a job, he receives W^E, since his outside option in wage negotiations is being unemployed and eligible.

Finally, we consider the value of a claimant with n months left (including the current month). This is given by

$$\begin{aligned} U^C(n, \Omega) &= h + \mathcal{I}_{n>0} b \\ &+ \beta\mu\mathbb{E}_{z',D'|u'} f(\theta'(\Omega')) \left[\mathcal{I}_{n'>0} W^E(\Omega') + (1 - \mathcal{I}_{n'>0}) W^I(\Omega')\right] \\ &+ \beta\mu\mathbb{E}_{z',D'|u'} (1 - f(\theta'(\Omega'))) U^C(n', \Omega') \end{aligned} \tag{13}$$

where $\mathcal{I}$ denotes the indicator function and

$$n' = n'(n, D, D') = \max\{0, \max\{0, n - 1\} + D' - D\} \tag{14}$$

A claimant with n months left receives unemployment benefits this period if $n > 0$. The evolution of his own UI status, n', then depends both on his current status n and on the evolution of the policy, D'. If the claimant finds a job, his outside option is that of an eligible worker if $n' > 0$, and that of an ineligible worker if $n' = 0$.

2.4 Wage bargaining

Wages are determined by Nash bargaining. Specifically, $w^E(\Omega)$ and $w^I(\Omega)$ are set so that the the firm in each match receives a share $1-\xi$ of the current match surplus:

$$J^E(\Omega) = \frac{1-\xi}{\xi}\left(W^E(\Omega) - U^E(\Omega)\right) \tag{15}$$

$$J^I(\Omega) = \frac{1-\xi}{\xi}\left(W^I(\Omega) - U^I(\Omega)\right) \tag{16}$$

2.5 Laws of motion

We next describe the law of motion for the endogenous aggregate state, namely the distribution of workers across employment and UI eligibility states. Denote by u^I the measure of ineligible unemployed workers, by u^E the measure of eligible unemployed workers not yet claiming, and by u_n the measure of eligible claimants with n months left. Similarly, denote by ℓ^E the measure of eligible employed and by ℓ^I the measure of ineligible employed. Finally, denote by u and ℓ the total measures of unemployed and employed, respectively. The law of motion for the aggregate unemployment rate is straightforward:

$$u' = 1 - \mu + \mu\left[\delta\ell + (1 - f(\theta))\,u\right] \tag{17}$$

The laws of motion for the measures of employed workers of each type are

$$\ell^{E\prime} = \mu\left\{(1-\delta)\,\ell^E + \varrho\,(1-\delta)\,\ell^I + f(\theta)\left[u - u^I - u_0\right]\right\} \tag{18}$$

and

$$\ell^{I\prime} = \mu\left[(1-\varrho)\,(1-\delta)\,\ell^I + f(\theta)\,(u^I + u_0)\right] \tag{19}$$

Next, the law of motion for the measure of ineligible unemployed workers is given by

$$u^{I\prime} = 1 - \mu + \mu \left[\delta \ell^I + (1 - f(\theta)) u^I\right], \tag{20}$$

and the measure of eligible, but not yet claiming, unemployed workers evolves according to

$$u^{E\prime} = \mu (1 - P) \left[\delta \ell^E + (1 - f(\theta)) u^E\right] \tag{21}$$

The measure of claimants with the highest possible duration remaining is

$$u'_{D'-1} = \mu P \left[\delta \ell^E + (1 - f(\theta)) u^E\right] + \mu (1 - f(\theta)) \sum_{n=1}^{D} u(n) \mathcal{I}_{n'(n,D,D')=D'-1} \tag{22}$$

This is composed of eligible employed workers who just lost their jobs and started claiming, eligible non-claimants who did not find a job and started claiming, and legacy claimants whose next-period remaining duration happens to be $D'-1$. Finally, the measure of claimants with m months remaining, for $0 \leq m < D' - 1$, evolves according to

$$u'_m = \mu (1 - f(\theta)) \sum_{n=1}^{D} u_n \mathcal{I}_{n'(n,D,D')=m} \tag{23}$$

The above also yields the fraction of unemployed workers eligible for benefits the following period, as

$$\varpi' = 1 - \frac{u^{I\prime} + u'_0}{u'} \tag{24}$$

2.6 Equilibrium

An equilibrium consists of value functions

$$\{J^E(\Omega), J^I(\Omega), W^E(\Omega), W^I(\Omega), U^E(\Omega), U^I(\Omega), U^{NC}(\Omega), U^C(n, \Omega)\}$$

together with market tightness $\theta(\Omega)$, wages $w^E(\Omega)$, $w^I(\Omega)$, and a transition rule $\Omega' = H(\Omega)$ satisfying all the firm and worker Bellman equations, Nash bargaining, and the laws of motion.[6]

3 Calibration

Table 1: Internally Calibrated Parameters

	Parameter	Value	Target	Value
h	Value of non-market activity	0.74	Steady state u	0.055
ξ	Bargaining power	0.28	Increase in unemp. duration after 1 month extension	0.15
λ	Matching parameter	0.9	Steady state v/u	0.6
1-μ	Labor force exit rate	0.0055	Fraction u eligible for benefits	0.6
P	Benefit take-up rate	0.333	Fraction u claiming benefits	0.34

We calibrate the steady state of the model to match average aggregate US data targets over the 1960-2015 period and then assess the model's fit with respect to the time series of unemployment. As we explain below, we target aggregate moments of the key data series over this period. We also use, as calibration targets, the literature's estimates on the elasticity of unemployment with respect to UI benefits. We *do not* target the correlation between unemployment and productivity, nor any other moment that directly affects the timing of unemployment or the speed of the recovery. The success of the model can then be assessed by how well it matches the time series of unemployment and its correlation with productivity.

The model period is taken to be 1 month. We normalize mean monthly productivity to one. We use the seasonally adjusted unemployment series constructed by the BLS, and measure vacancies using the composite Help-Wanted Index of Barnichon (2010).

We set the discount factor $\beta = 0.99^{1/3}$, implying a yearly discount rate of 4%. The job separation parameter δ is set to 0.014 to match the average monthly job

[6] The aggregate state includes the entire distribution of unemployed workers over different unemployment durations. The model is solved globally and non-linearly using a similar computational technique to Krusell and Smith (1998). Details of the computational procedure are in Appendix D.

separation rate. We set $k = 0.5$ following Hagedorn and Manovskii (2008), who estimate the combined capital and labor costs of vacancy creation.

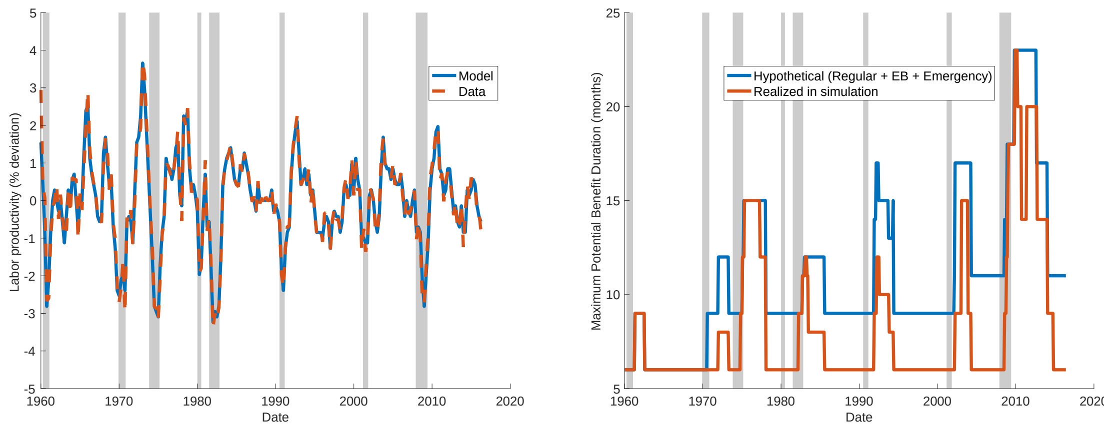
Following den Haan *et al.* (2000), we assume the functional form of the matching function to be

$$M(u,v) = \frac{u \cdot v}{\left[u^{\lambda} + v^{\lambda}\right]^{1/\lambda}}$$

We set $b = 0.45$ to be consistent with the average replacement rate in the U.S. We set the re-entitlement rate to $\varrho = 1/12$ to account for the fact that it takes about one year to gain eligibility for unemployment benefits. The baseline duration of unemployment insurance in normal times, $\overline{D}$, is set to equal 6 months (and is constant in steady state).

This leaves five parameters to be internally calibrated: the value h of non-market activity; the worker's bargaining weight ξ; the matching function parameter λ; and the labor market exit and benefits take-up rates, $1-\mu$ and P. We calibrate these five parameters jointly to match five data targets, chosen to capture relevant statistics from the US labor market. The first three of these statistics are aggregate targets. We target the average unemployment rate of 0.055 and the average vacancy-unemployment ratio of 0.634. We also target the fraction of unemployed eligible for benefits equal to 0.6, and the fraction of unemployed claiming UI benefits equal to 0.34 (see Auray *et al.* (2019) and Birinci and See (2021)). Finally, we target the effect of a benefit duration increase on unemployment duration. Research by Moffitt and Nicholson (1982), Moffitt (1985), and Katz and Meyer (1990), among others, reached consensus estimates that a one week increase in benefit duration increases the average duration of unemployment spells by 0.1 to 0.2 weeks. We target an elasticity in the middle of this range: a one month increase in benefit duration increases average duration of unemployment by 0.15 months.[7]

[7] In a previous version of this paper, we conduct our own empirical analysis, extending the methodology of Hagedorn *et al.* (2013), to show that the elasticity of unemployment with respect to benefit duration remained stable between the pre- and post-jobless recoveries periods. These results are available in Mitman and Rabinovich (2019).



(a) Labor productivity (deviations from trend)

(b) Maximum potential benefit duration

Figure 1: The productivity and policy series in the simulated model. Left panel: simulated and actual labor productivity, in deviations from HP-filtered trend with parameter 1600. Right panel: maximum statutory potential benefit duration (regular program + extended benefits + emergency programs) if "high unemployment," and realized maximum potential benefit duration in the simulation. NBER dated recessions are shaded.

4 Simulation and Results

Using our calibrated model, we quantify the extent to which the model can replicate observed labor-market dynamics, given the realized path for productivity and unemployment benefit legislation in the data. To do so requires solving for the stochastic version of the model with aggregate shocks to productivity. Following Shimer (2005), labor productivity z_t is taken to mean real output per worker in the non-farm business sector. This measure of productivity is taken from the quarterly data constructed by the BLS. The parameters for the productivity shock process are estimated, at the monthly level, to be $\rho = 0.9587$ and $\sigma_\varepsilon = 0.006$.

4.1 Expectations about UI extensions

Unemployment benefit programs changed significantly throughout our sample, as we describe in Appendix A. Given the forward-looking nature of vacancy creation decisions, we have to take a stand on how to model expectations about future benefit policies. In particular, households need to form expectations over the probability that an emergency benefit extension will be passed (if one is not currently in place), or that it will be allowed to expire (if a program is currently in place). Further, nearly all discretionary extensions entail a state-contingent menu of benefit durations (providing longer duration when the unemployment rate is higher). In reality, such decisions arise endogenously from government decisions. We abstract from modeling the government's decisions here and parameterize the expectations to be consistent with the behavior of the U.S. government over the sample.[8]

We make the following assumptions on the duration of unemployment benefits. D_t is the months of benefits that are available to the currently unemployed and consists of three components:

$$D_t = \overline{D} + I_t^{EB} D^{EB} + I_t^{EUC} D_t^{EUC}. \tag{25}$$

First, there is a baseline duration of unemployment insurance in normal times, $\overline{D} = 6$. Second, there are automatic triggers for extending unemployment benefits when unemployment is high. We denote these triggers "extended benefits" or EB after the Extended Benefits program in Federal-State Extended Unemployment Compensation Act of 1970. $D^{EB} = 3$ denotes the added months of unemployment insurance when the trigger is on. I_t^{EB} is an indicator variable, equal to 1 when the trigger is on, and 0 when it is off:

$$I_t^{EB} = \begin{cases} 1 & \text{if } u_t \geq \hat{u} \\ 0 & \text{otherwise,} \end{cases} \tag{26}$$

where $\hat{u}$ is a pre-specified threshold that we set $\hat{u} = 6.5\%$.[9] Third, there are discretionary extensions, which we call "emergency unemployment compensation," or EUC. $D_t^{EUC}(u_{t-1})$ denotes the added weeks of unemployment insurance when the

[8] Recent work by Bardóczy and Guerreiro (2023) and Broer *et al.* (2021) explore the impact of unemployment benefit policies in models that depart from full-information rational expectations.

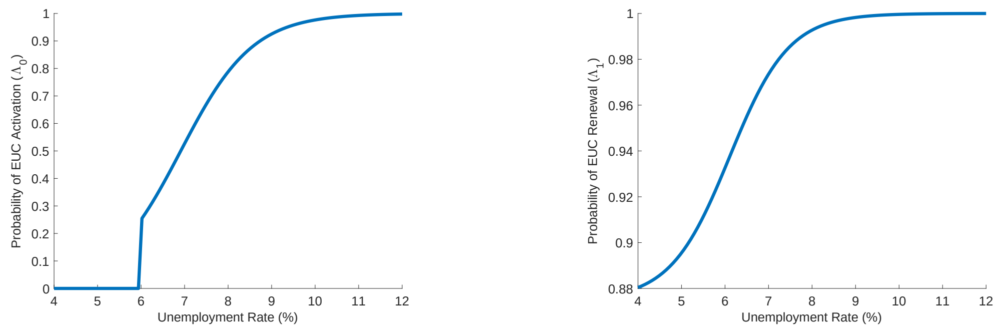
[9] The triggers for EB are more complicated in practice. States have the option of setting triggers based on the insured unemployment rate or the total unemployment rate. Further, prior to 1982 a national trigger also existed.

discretionary extension is on given last month's unemployment rate. I_t^{EUC} is an indicator variable, equal to 1 when such an extension is in effect, and 0 when it is not. It is assumed to follow a Markov process

$$Prob\left(I_t^{EUC} = 1 | I_{t-1}^{EUC} = i\right) = \Lambda_i\left(u_t\right), \tag{27}$$

where Λ_0 is the (activation) probability that the EUC program gets turned on conditional on being off in the previous period, and Λ_1 is the (renewal) probability that EUC continues being on conditional on being on in the previous period. In estimating that process, we seek to account for the fact that, while these extensions and reauthorizations do not follow any pre-set rule (and are therefore not perfectly predictable, even given the unemployment rate), they are nevertheless persistent and correlated with unemployment. To this end, we estimate transition probabilities for I_t^{EUC} from the data on actual unemployment benefit extensions and reauthorizations, with transition probabilities between t and $t+1$ dependent on the unemployment rate in period t. Specifically, we estimate separate logit models for the probability that a discretionary extension is passed and for the probability that a discretionary extension is reauthorized conditional on being in place, as a function of the unemployment rate. For the authorization regression, we find a constant of 10.99, and a coefficient on the unemployment rate of 0.99. For the re-authorization regression, we find a constant of 8.89 and a coefficient on the unemployment rate of 1.47. We plot the estimated functions for Λ_i in Figure 2. Appendix B reports additional details on the construction of the figure.

We keep the process for Λ_i fixed over the simulation, but $D_t^{EUC}(u_{t-1})$ changes over time using the actual extensions described in Appendix A. We assume that households believe that the extensions of the most recent recession will be the extensions enacted in all future recessions. For example, in the 1975 recession the government provided up to 26 weeks of benefits during the recession as part of the discretionary extensions. Going into the 1982 recession, agents' expectations were for discretionary extensions up to 26 weeks (that occur stochastically with the estimated transition probabilities that are kept constant). In September 1982 agents



(a) Probability EUC program enacted

(b) Probability EUC program renewed

Figure 2: Estimated activation (left panel) and renewal (right panel) probabilities as a function of the unemployment rate estimated from post-war EUC programs. Probabilities are monthly.

are surprised when the government instead only enacts up to 10 weeks of extensions (i.e. it is a probability 0 event). But going forward, they assume that this is the expected discretionary response of the government until they are "surprised again." We found this to be a parsimonious and plausible way to handle the beliefs about future discretionary actions. To be clear, throughout all time periods agents beliefs about the probability that an extension will occur or expire are constant functions of the unemployment rate and current extension status, with the actual realization taken from the data. Thus, the simulation forward of the model is by a sequence of MIT shocks whenever the D^{EB} and D^{EUC} policies change, but agents always have full rational expectations over z_t, I_t^{EB}, and I_t^{EUC}.

To illustrate that this procedure generates plausible belief dynamics, we perform two exercises. First, we perform a Monte Carlo study. We perform 1000 simulations of the model as in the benchmark, but instead of taking the realization of extensions, I_t^{EUC} from the data, we draw them according to the beliefs of the agents $\Lambda_i(u)$. Thus, we can evaluate how "atypical" the actual realization of policy in the data was compared to the distribution implied by the expectations. We should note that we keep changes in the EUC schedule (D_t^{EUC}) the same (since those are treated as "MIT shocks" as discussed above). Figure 3 plots the realization of benefits in our benchmark, the mean realization across the 1000 simulations, and then the 95% CI across the simulations (± 1.96 standard deviations across simulations). The expected benefits from the estimated belief process align remarkably well with the realized benefit path. Some notable "misses" where agents predicted a benefit extension include 1970 (that didn't arrive until 1972) and the 1980 recession—the only post-war recession that didn't feature a discretionary program.

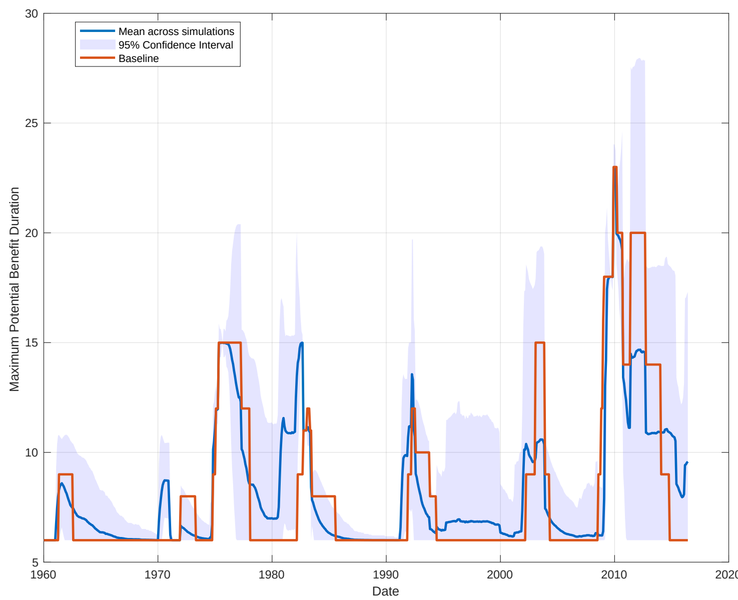


Figure 3: Monte Carlo Study to validate the belief process. Blue line is the mean benefit duration across 1000 monte carlo simulations. Red line is realized benefit duration in the simulation taking I_t^{EUC} from the data. Shaded areas represent 95% CI from the simulations.

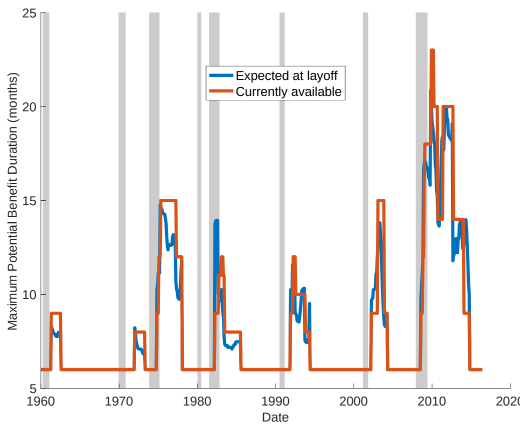
Second, we plot the expected months of benefits that a newly unemployed agent would expect to receive over the subsequent 24 months in the left panel of Figure 4. The line labeled "expected at layoff" shows the expected future duration of benefits of a newly unemployed, where the expectation is taken with respect to the uncertainty about future z_t, I_t^{EB}, and I_t^{EUC} (and in particular future unemployment rates) but is conditional on current beliefs about D^{EB} and D^{EUC} policies. The line labeled "currently available" plots the actual D_t, the months of benefits currently available. Note that the actual D_t is a cross-sectional measure at a point in time, i.e. at time t what's the longest duration unemployed that could be collecting benefits. The "expected at layoff" measure, by contrast, is forward-looking and measures what durations of benefits will be available in the future as that individual's unemployment duration increases. The "expected at layoff" measure turns out to be quite accurate. To illustrate this, in the right panel of Figure 4 we plot the "perfect foresight" measure of benefit duration, that is how many months of benefits an agent at time t would actually collect (if they remained unemployed) in our simulation.[10] As is evident from the figure, the expectations and perfect foresight measures are quite close, providing additional reassurance that our process for beliefs is reasonable.

At the beginning of recessions, newly unemployed (correctly) expect to receive relatively long durations, but the expectation falls over time as agents forecast that the probability of reauthorization, Λ_1, is falling. The "surprise" in 1982 mentioned in the previous paragraph is evident from the figure—in the model agents expected the emergency compensation to mimic that of from 1975 (explaining why the blue line is above the red), but then after the passage of the Federal Supplemental Compensation (FSC) program in 1982, they revise their beliefs downwards. As the end of the Great Recession, expect benefits to last longer than occurred in reality. This can be explained by the unexpected nature of the end of the EUC08 program at the end of 2013, as discussed in Hagedorn *et al.* (2015).

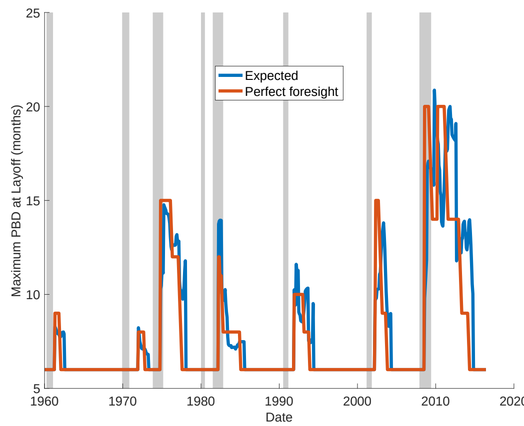
4.2 Main Results

To generate our baseline results we then simulate the stochastic model, selecting the sequence of shocks to match the time series for productivity and benefits from Figure 1. The left panel of Figure 1 displays the time series of productivity shocks that serve as inputs into the model. The right panel displays the time series of

[10] When no extensions are in place we set the perfect foresight measure to six, otherwise it would include anticipation of recessions.

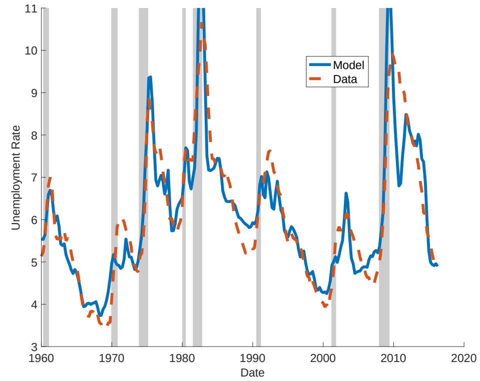


(a) Expected and currently available

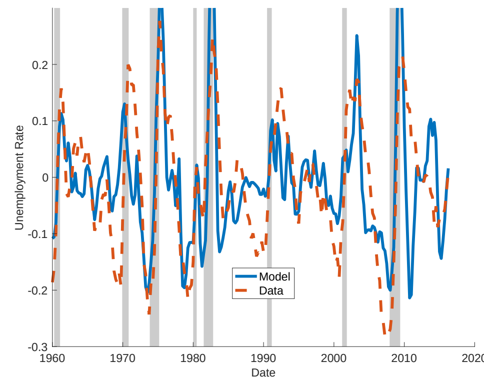


(b) Expected and perfect foresight

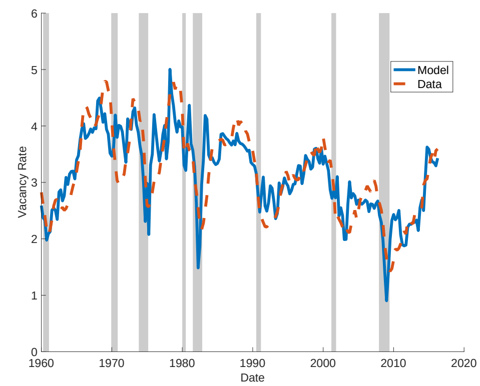
Figure 4: Measures of benefit duration in the simulation. In the left panel, the blue line is the expected maximum potential benefit duration (PBD) at layoff, using the estimated beliefs and law of motion from the model, whereas the red line is the maximum PBD currently available. The right panel reproduces the expected maximum PBD and adds the perfect foresight measure of PDB, the realized maximum months of benefits an unemployed person could claim.



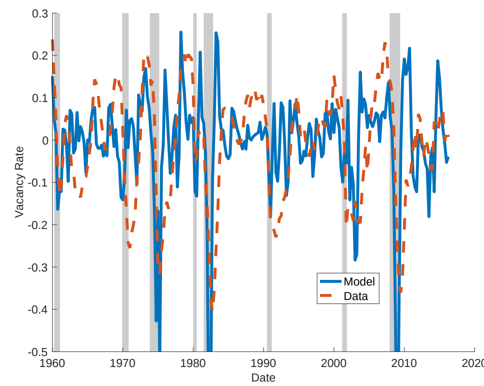
(a) Unemployment (levels)



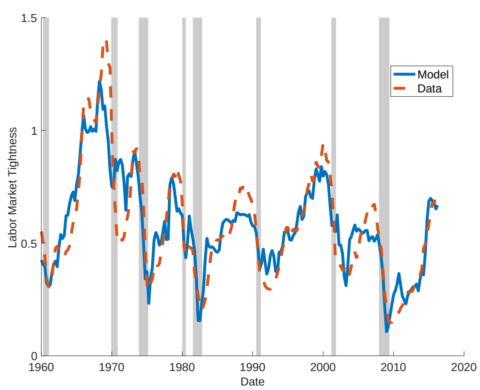
(b) Unemployment (deviations from trend)



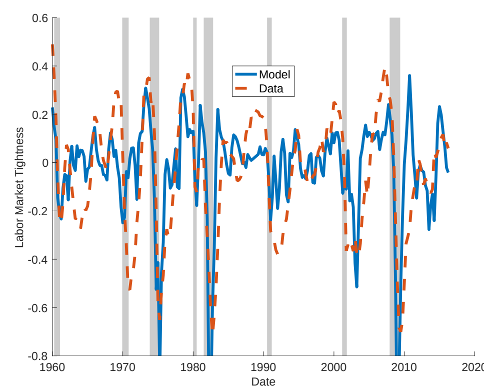
(c) Vacancy rate (levels)



(d) Vacancy rate (deviations from trend)



(e) Vacancy/unemp. ratio (levels)



(f) Vacancy/unemp. ratio (deviations from trend)

Figure 5: Simulated and actual labor market variables (levels left column, deviations from HP-filtered trend right column) from 1960:I through 2016:II. NBER dated recessions are shaded.

unemployment benefit extensions. The "hypothetical" series shows the maximum statutory UI benefit duration at each point in time, given the UI schedule in place at that moment. The "realized" series shows the UI extensions that actually realize in the simulation, which are lower than the "hypothetical" whenever the unemployment rate does not reach a sufficiently high level for all the extensions to be activated.

The model can account for key features of the post-war labor market. In Figure 5a, we plot the unemployment rate generated from the model and that observed in the data. The model with the implemented US unemployment benefit policy generates a time series of unemployment in levels that closely matches what is seen in the data ($R^2 = 0.47$). In addition, in Figure 5b we plot the log deviations from trend both in the data and in the model. Again, notice that the model does an excellent job of matching the data ($R^2 = 0.35$). As shown in Figures 5b-5f, the model also does an excellent job of capturing the time-series of the vacancy rate and the vacancy-unemployment ration. Next, we confirm the model's ability to match key business cycle statistics. Tables 2 and 3 report the summary statistics from US data and from the model. The model slightly under-predicts the volatility of the labor market.

		u	v	v/u	f	w	z
Standard Deviation		0.1169	0.1327	0.2491	0.0827	0.0106	0.0122
	u	1	-0.9222	-0.9572	-0.9296	-0.3837	-0.2141
Correlation	v		1	0.9852	0.8949	0.4222	0.3880
Matrix	v/u			1	0.9243	0.4140	0.3289
	f				1	0.3408	0.2036
	w					1	0.5713
	z						1

Table 2: Summary statistics, quarterly US data, 1960:I to 2016:II. Standard deviations and correlations are reported in logs as quarterly deviations from an HP-filtered trend with smoothing parameter 1600.

		u	v	v/u	f	w	z
Standard Deviation		0.1101	0.1353	0.2100	0.1480	0.0109	0.0118
	u	1	-0.5706	-0.8857	-0.8791	-0.6446	-0.5123
Correlation	v		1	0.8749	0.8806	0.3657	0.6605
Matrix	v/u			1	0.9977	0.5720	0.6542
	f				1	0.5460	0.6377
	w					1	0.8381
	z						1

Table 3: Results from the calibrated model. Standard deviations and correlations are reported in logs as quarterly deviations from an HP-filtered trend with smoothing parameter 1600.

4.3 Role of UI extensions

We perform two counterfactual experiments to quantify the contribution of unemployment benefit extensions in generating the observed labor-market dynamics. First, we simulate the model with all unemployment benefit extensions turned off. Table 4 reports the same summary statistics from the simulated model with no benefit extensions. These results show that while the calibrated model performs well in matching the cyclical behavior of the labor market, shutting down time-varying unemployment benefit extensions substantially worsens the model's ability to match the observed dynamics. In particular, without extensions unemployment is far less volatile, and all labor market variables are more strongly corrected with labor productivity. Note that the inclusion of time-varying unemployment benefit extensions was not guaranteed to improve the model's fit, since what matters for the latter is the timing of the extensions relative to productivity, which was not targeted in the calibration. The R^2 between model and data falls to 0.03 (levels) and 0.10 (deviations) in the model without extensions, thus we infer that unemployment benefit extensions explain roughly 32-37% of observed unemployment fluctuations.

We perform a second counterfactual where we recalibrate a version of the model without benefit extensions to match the volatility of unemployment in the data. The purpose of this experiment is to make clear that solving the "volatility puzzle" does not imply the emergence of jobless recoveries and the weak correlation between labor market variables and productivity. Table 5 reports the labor-market statistics from the recalibrated model. While the model is successful in matching the standard deviations of labor market variables, it fails to match the low correlation between productivity and labor-market variables. The R^2 between model and data falls to 0.07 (levels) and -0.15 (deviations) in the model without extensions, thus we infer that unemployment benefit extensions explain roughly 40-50% of observed unemployment fluctuations.

		u	v	v/u	f	w	z
Standard Deviation		0.0355	0.0421	0.0672	0.0418	0.0101	0.0118
	u	1	-0.6717	-0.9363	-0.9365	-0.9879	-0.9017
Correlation	v		1	0.8882	0.8881	0.7667	0.8415
Matrix	v/u			1	0.9998	0.9742	0.9571
	f				1	0.9739	0.9564
	w					1	0.9457
	z						1

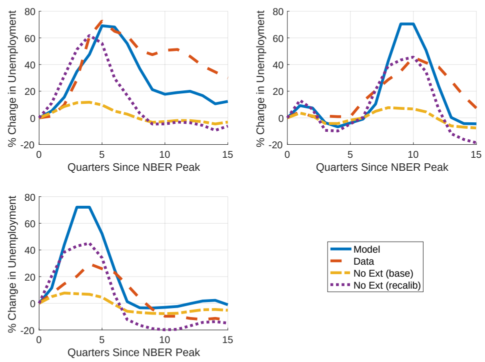
Table 4: Results from the model with no benefit extensions (benchmark calibration). Standard deviations and correlations are reported in logs as quarterly deviations from an HP-filtered trend with smoothing parameter 1600.

We next investigate whether the model is consistent with the emergence of jobless recoveries. In Figure 6a, we plot the change in unemployment - actual and predicted by the model - relative to the NBER peak before the 1973-1975, 1980 and 1981-1982 recessions. The model replicates the response of employment over those periods quite well. Next, in Figure 6b, we similarly plot the change in employment for the 1990-1991, 2001 and 2007-2009 recessions. The model is able to replicate the observation that, unlike the previous three recessions, the recovery of productivity was not matched in this case by a rapid rise in employment. To understand the role of unemployment benefit extensions in generating jobless recoveries, we reproduce the same plots for the two aforementioned counterfactual no extension models (baseline and recalibrated). The results are shown as the yellow dot-dashed and purple

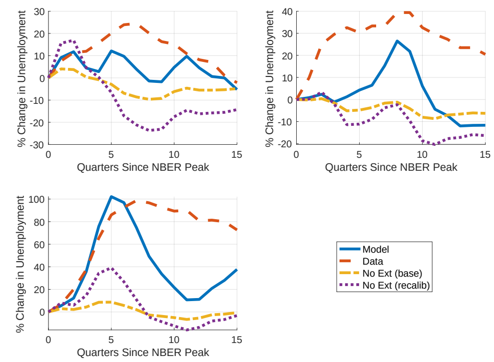
		u	v	v/u	f	w	z
Standard Deviation		0.1161	0.1378	0.2292	0.1362	0.0107	0.0118
	u	1	-0.7386	-0.9434	-0.9438	-0.9970	-0.8619
Correlation	v		1	0.9094	0.9102	0.7424	0.9099
Matrix	v/u			1	0.9975	0.9494	0.9457
	f				1	0.9455	0.9346
	w					1	0.8727
	z						1

Table 5: Results from the model with no benefit extensions (recalibrated). Standard deviations and correlations are reported in logs as quarterly deviations from an HP-filtered trend with smoothing parameter 1600.

dotted lines in Figures 6a and 6b. The recalibrated model without the extensions does a good job matching unemployment dynamics in the early recessions, but fails to generate jobless recoveries: unemployment recovers much faster in the model than it does in the data. Unemployment benefit extensions are thus quantitatively important for explaining the cyclical behavior of employment. While the recalibrated model is able to deliver *amplification* of productivity shocks, both counterfactual models fail to deliver *propagation* of those shocks. We reinforce this point by reporting the correlation of labor market variables with lagged productivity in Table 6. The reason for these patterns is that, since productivity is the driving exogenous shock, unemployment tracks productivity very closely in the model with no extensions. Introducing state-dependent UI extensions breaks this dependence, despite the fact that productivity is still the only exogenous shock. A negative productivity shock leads to a rise in unemployment, which in turn triggers UI extensions and hence a rise in unemployment in the future. UI extensions effectively make unemployment (a stock variable) a state, which now affects vacancy posting and hence future unemployment.



(a) 1973-75, 1980 and 1981-82 recessions.



(b) 1990-91, 2001 and 2007-09 recessions.

Figure 6: Simulated and actual percentage change in unemployment from NBER peak before NBER dated recessions. The blue line is the model, dashed red line is the data, the yellow dot-dashed line is the model without extensions under the baseline calibration, and the purple dotted line is the model without extensions recalibrated. Data and model are not filtered. Data is from CPS, unemployment.

4.4 The Beveridge Curve

As shown in Figure 7 the model is also consistent with apparent "shifts" in the Beveridge curve in the Great Recession. Theoretically, the Beveridge curve is a steady-state relationship between vacancies and unemployment, and movements along it in the standard DMP model are generated by changing labor productivity. It is important to note that tightness (and vacancies) adjust immediately (they are jump variables), but unemployment takes time to adjust. Thus, at a monthly frequency, a drop in productivity would imply an immediate drop in tightness (and vacancies) but a fixed unemployment rate. This would be a downward departure from the theoretical Beveridge curve. But, the model would transit in the upward-right direction along a path of constant tightness until it returned to the Beveridge curve. This is inconsequential for small shocks: when aggregated to quarterly frequency, this movement would be masked and it would appear as if the economy remained on the theoretical Beveridge curve. However, for large shocks - such as unemployment benefit extensions in recessions - it would take the economy time to return to the Beveridge curve and what appear to be "departures" from the "true" Beveridge

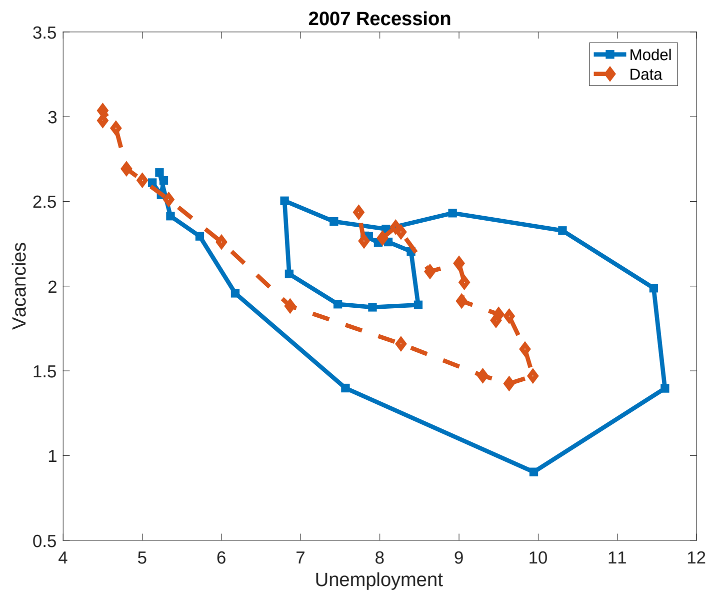


Figure 7: Beveridge curve in the Great Recession (levels).

curve would emerge. In Appendix C we show that this logic holds in all previous recessions as well, and that both the model and data exhibit counterclockwise shifts in the Beveridge curve during previous recessions.

4.5 Robustness to Changing Expectations

As described in Appendix A, unemployment benefit extensions became more generous in recent recessions. This motivates us to consider an alternative estimation of beliefs, in which we estimate activation and renewal probabilities separately for the pre-1985 and post-1985 periods. The results of this estimation and its implications for the simulations are reported in Appendix B.1. Figure 8 illustrates the estimated activation and renewal probabilities for the split-sample approach. As can be seen in the figure, only the estimated activation probabilities, rather than renewal probabilities, are substantially different before and after 1985. We confirm this statistically by running a Chow test between the models. Because it is the renewal probabilities

Variable	Data	Model	Model w/o Extensions (baseline)	Model w/o Extensions (recalibrated)
u_t	-0.3935	-0.6973	-0.9486	-0.9475
v_t	0.5433	0.5199	0.4593	0.6240
v_t/u_t	0.4985	0.6950	0.8078	0.8732
f_t	0.3862	0.6801	0.8071	0.8621
w_t	0.6006	0.4624	0.8731	0.8830

Table 6: Correlation with lagged productivity, z_{t-1}. Correlations are reported in logs as quarterly deviations from an HP-filtered trend with smoothing parameter 1600.

that are instrumental to the persistence of a UI extension, this suggests that our main results will be robust to splitting the sample this way. This conjecture turns out to be correct. Figure 9 compares the baseline simulation to the simulation under this alternative expectation process, showing that the model fit is not materially different. Figure 10 confirms that the model similarly reproduces the individual recessions, both in the pre-jobless recoveries period of the 1970's and 1980's, and the jobless recoveries period of the 1990's and 2000's.

5 Discussion of the Results

In this section, we discuss the implications of abstracting from endogenous separations and the broader applicability of our results to countries other than the US. These two questions are related because the relative importance of separations vs. job-finding rates in driving unemployment dynamics has been shown to differ across countries. In section 5.1, we discuss the assumption of a constant exogenous job separation rate. Section 5.2 discusses the relevance of our findings for other advanced economies.

5.1 Exogenous separations

Our model assumes exogenous separations so that unemployment benefits affect unemployment entirely through the job-finding rate. Despite this, the calibrated model generates unemployment dynamics that are very close to the data. Since separations are countercyclical in US data and account for a non-negligible part of US unemployment dynamics (see, e.g., Fujita (2010), Fujita and Ramey (2012)), this suggests that our model, while accounting well for unemployment dynamics overall, may attribute too much to job-finding rather than separations. Incorporating endogenous separations would reduce the measured role of the vacancy posting channel. Still, it would most likely affect the spike in unemployment at the start of a recession rather than the persistence of the subsequent recovery. In an earlier draft, Mitman and Rabinovich (2019), we extend the model to include endogenous separations and show that the propagation of shocks induced by unemployment benefit extensions still occurs mostly through vacancy posting. The role of separations, however, may be larger in other economies, which we discuss below.

5.2 External validity

Our paper has focused on unemployment dynamics in the US. In what follows, we discuss the unique features of the US labor market and whether our results apply to other developed economies. We emphasize three observations here. First, the US is an outlier regarding the emergence of jobless recoveries post-1985. Second, the US is also an outlier in terms of countercyclical UI extensions. Finally, the US is distinct from European economies in terms of the role played by job-finding rather than job separations in driving unemployment volatility. Below, we discuss the broader implications of our results in light of these observations.

Post-1985 jobless recoveries are a phenomenon specific to the US, as documented by Graetz and Michaels (2017). While recoveries of GDP have been slower in the latter period across OECD countries, there is no evidence of the kind observed for the US that recoveries of employment slowed down relative to recoveries of GDP. This observation motivates a search for a US-specific mechanism behind jobless recoveries, which makes policy a natural candidate. Countercyclical unemployment-dependent extensions of unemployment benefits are likewise a phenomenon specific to the US. While unemployment benefits are lower in the US on average than in many European countries, the US is distinct in extending unemployment benefit durations in response to downturns. From this standpoint, our paper shows that these two observations about the US labor market are (causally) related.

The unique feature of US policy notwithstanding, our results beg the question of whether unemployment benefit extensions in recessions would have a similar propagation effect in other economies. The existing literature suggests they would, though the exact channel may differ. One important difference between the US and other OECD countries, in addition to policy, has to do with the relative role of job separations in driving unemployment dynamics. In our model, the persistence of unemployment — and the effect of unemployment benefits — is affected entirely through the job-finding channel. However, as described above, endogenous separations into unemployment also play a role. In fact, Elsby *et al.* (2013) and Jung and Kuhn (2014) document that the dynamics of separations account for a much larger fraction of unemployment volatility in continental European economies. Following up on this, Hartung *et al.* (2018) show that cuts in unemployment benefits duration have large effects on unemployment in Germany, though it operates largely through endogenous job separations rather than job finding.[11] As shown in Jung and Kuhn (2014) and Hartung *et al.* (2018), both the larger role of separations for unemployment volatility and its larger role in driving the effects of unemployment insurance stem from the lower job finding rate in Germany compared to the US. Thus, we believe that the broader result on the importance of unemployment benefits for unemployment persistence would still manifest itself in other economies, but whether this occurs through the separation or job-finding margin would likely depend on the individual country.

[11]The evidence in the US, by comparison, suggests cuts in benefits have little effect on separation rates (Karahan *et al.*, 2019).

6 Conclusion

We have shown that unemployment benefit extensions act as an important propagation mechanism, contributing to unemployment's post-recession persistence and its weak correlation with productivity. More generally, our analysis implies that unemployment benefit extensions are a natural and compelling candidate for the endogenous labor wedge needed to reconcile an apparently weak productivity-labor market correlation with a theory of business cycles driven by productivity shocks.

Our analysis has been positive in nature. An important future direction for research is studying the optimal provision of unemployment benefits over the business cycle. Mitman and Rabinovich (2015, 2021) make progress in this dimension by solving Ramsey and Markov-Perfect problems in a similar framework. A full quantitative evaluation would require performing this analysis in an extended model that incorporates more frictions, explicit heterogeneity, and incomplete markets (e.g. Braxton *et al.* (2020), Hagedorn *et al.* (2019), and Lentz (2009)). We leave this for future research.

References

Andolfatto, D. (1996). Business cycles and labor-market search. *The american economic review*, pp. 112–132.

Auray, S., Fuller, D. L. and Lkhagvasuren, D. (2019). Unemployment insurance take-up rates in an equilibrium search model. *European Economic Review*, **112**, 1–31.

Bachmann, R. (2011). *Understanding the Jobless Recoveries After 1991 and 2001.* working paper.

Bardóczy, B. and Guerreiro, J. (2023). *Unemployment Insurance in Macroeconomic Stabilization with Imperfect Expectations*. Tech. rep., mimeo.

Barnichon, R. (2010). Building a composite help-wanted index. *Economics Letters*, **109** (3), 175–178.

Berger, D. (2011). *Countercyclical Restructuring and Jobless Recoveries.* working paper.

Birinci, S. and See, K. (2021). *Labor Market Responses to Unemployment Insurance: The Role of Heterogeneity*. Tech. rep., Federal Reserve Bank of St. Louis.

Braxton, J. C., Herkenhoff, K. F. and Phillips, G. M. (2020). *Can the unemployed borrow? implications for public insurance*. Tech. rep., National Bureau of Economic Research.

Broer, T., Kohlhas, A., Mitman, K. and Schlafmann, K. (2021). *Information and wealth heterogeneity in the macroeconomy*. CEPR, Discussion Paper No. 15934.

Chang, Y. and Kim, S.-B. (2007). Heterogeneity and aggregation: Implications for labor-market fluctuations. *American Economic Review*, **97** (5), 1939–1956.

Chari, V. V., Kehoe, P. J. and McGrattan, E. R. (2007). Business cycle accounting. *Econometrica*, **75** (3), 781–836.

Costain, J. S. and Reiter, M. (2008). Business cycles, unemployment insurance, and the calibration of matching models. *Journal of Economic Dynamics and control*, **32** (4), 1120–1155.

DALY, M. C., HOBIJN, B. and ŞAHIN, A. (2012). A search and matching approach to labor markets: Did the natural rate of unemployment rise? *Journal of Economic Perspectives*, **26** (3), 3–26.

DEN HAAN, W. J., RAMEY, G. and WATSON, J. (2000). Job destruction and propagation of shocks. *American Economic Review*, **90** (3), 482–498.

DIAMOND, P. A. and ŞAHIN, A. (2015). Shifts in the beveridge curve. *Research in Economics*, **69** (1), 18–25.

ELSBY, M. W., HOBIJN, B. and ŞAHIN, A. (2013). Unemployment dynamics in the oecd. *Review of Economics and Statistics*, **95** (2), 530–548.

FUJITA, S. (2010). *Effects of the UI Benefit Extensions: Evidence from the CPS*. working paper 10-35, Federal Reserve Bank of Philadelphia.

— and RAMEY, G. (2012). Exogenous versus endogenous separation. *American Economic Journal: Macroeconomics*, **4** (4), 68–93.

GAGGL, P. and KAUFMANN, S. (2019). The cyclical component of labor market polarization and jobless recoveries in the us. *Journal of Monetary Economics*.

GALI, J. (1999). Technology, employment, and the business cycle: do technology shocks explain aggregate fluctuations? *American economic review*, **89** (1), 249–271.

GRAETZ, G. and MICHAELS, G. (2017). Is modern technology responsible for jobless recoveries? *American Economic Review*, **107** (5), 168–73.

HAGEDORN, M., KARAHAN, F., MANOVSKII, I. and MITMAN, K. (2013). *Unemployment benefits and unemployment in the great recession: the role of macro effects*. Tech. rep., National Bureau of Economic Research.

— and MANOVSKII, I. (2008). The cyclical behavior of equilibrium unemployment and vacancies revisited. *American Economic Review*, **98** (4), 1692–1706.

—, — and MITMAN, K. (2015). *The impact of unemployment benefit extensions on employment: the 2014 employment miracle?* Tech. rep., National Bureau of Economic Research.

—, — and — (2019). *The fiscal multiplier*. Tech. rep., National Bureau of Economic Research.

Hall, R. E. (1997). Macroeconomic fluctuations and the allocation of time. *Journal of labor Economics*, **15** (1, Part 2), S223–S250.

— (2005). Employment fluctuations with equilibrium wage stickiness. *American Economic Review*, **95** (1), 50–65.

Hartung, B., Jung, P. and Kuhn, M. (2018). *What Hides Behind the German Labor Market Miracle? Unemployment Insurance Reforms and Labor Market Dynamics*. CEPR, Discussion Paper No. 13328.

Herkenhoff, K. F. (2019). The Impact of Consumer Credit Access on Unemployment. *The Review of Economic Studies*, **86** (6), 2605–2642.

Hobijn, B. and Şahin, A. (2013). Beveridge curve shifts across countries since the great recession. *IMF Economic Review*, **61** (4), 566–600.

Jaimovich, N. and Siu, H. E. (2018). Job polarization and jobless recoveries. *Review of Economics and Statistics*, (18334), 1–20.

Jung, P. and Kuhn, M. (2014). Labour market institutions and worker flows: comparing germany and the us. *The Economic Journal*, **124** (581), 1317–1342.

Karahan, F., Mitman, K. and Moore, B. (2019). *Micro and Macro effects of UI policies: Evidence from Missouri*. CEPR, Discussion Paper No. 14158.

Katz, L. F. and Meyer, B. D. (1990). The impact of the potential duration of unemployment benefits on the duration of unemployment. *Journal of Public Economics*, **41** (1), 45–72.

Krusell, P. and Smith, A. A., Jr (1998). Income and wealth heterogeneity in the macroeconomy. *Journal of political Economy*, **106** (5), 867–896.

Lentz, R. (2009). Optimal unemployment insurance in an estimated job search model with savings. *Review of Economic Dynamics*, **12** (1), 37–57.

Ljungqvist, L. and Sargent, T. J. (2017). The fundamental surplus. *American Economic Review*, **107** (9), 1630–1665.

Menzio, G. and Shi, S. (2011). Efficient Search on the Job and the Business Cycle. *Journal of Political Economy*, **119** (3), 468 – 510.

Mitman, K. and Rabinovich, S. (2015). Optimal unemployment insurance in an equilibrium business-cycle model. *Journal of Monetary Economics*, **71**, 99–118.

— and — (2019). *Do Unemployment Benefit Extensions Explain the Emergence of Jobless Recoveries?* Tech. Rep. 12365, Institute of Labor Economics (IZA).

— and — (2021). Whether, when and how to extend unemployment benefits: Theory and application to covid-19. *Journal of Public Economics*, **200**, 104447.

Moffitt, R. (1985). Unemployment insurance and the distribution of unemployment spells. *Journal of Econometrics*, **28** (1), 85–101.

— and Nicholson, W. (1982). The Effect of Unemployment Insurance on Unemployment: The Case of Federal Supplemental Benefits. *The Review of Economics and Statistics*, **64** (1), 1–11.

Nakajima, M. (2012). A quantitative analysis of unemployment benefit extensions. *Journal of Monetary Economics*, **59** (7), 686–702.

Ohanian, L. E. (2010). The Economic Crisis from a Neoclassical Perspective. *Journal of Economic Perspectives*, **24** (4), 45–66.

Shimer, R. (2005). The cyclical behavior of equilibrium unemployment and vacancies. *American Economic Review*, **95** (1), 25–49.

— (2009). Convergence in macroeconomics: The labor wedge. *American Economic Journal: Macroeconomics*, **1** (1), 280–97.

APPENDIX

A The Post-War US Unemployment Insurance System: An Overview

By the late 1950s, most unemployment insurance systems in U.S. states offered 26 weeks of benefits to newly displaced workers. The deep recession of 1957-58, however, prompted the federal government to lengthen the duration of benefits available. Under the Temporary Unemployment Compensation Act (TUC), the federal government offered interest free loans to states in order to provide up to 13 additional weeks of benefits. Seventeen states opted to participate in the program, which lasted from June of 1958 until June of 1959.

The first federally financed extension of unemployment benefits occurred during the 1960-1961 recession. The federal government passed the Temporary Extended Unemployment Compensation Act (TEUC). Whereas TUC was a voluntary program, TEUC was mandatory for all states and provided up to 13 weeks of additional benefits to unemployed workers from April 1961 until June 1962. The extra weeks of benefits were entirely financed by the federal government (which raised the Federal Unemployment Tax to offset the extensions).

Guided by TUC and TEUC, the federal government sought to develop an automatic system of extending unemployment benefits during recessions. In 1970 the Employment Security Amendments developed the Extended Benefits (EB) program, which would provide additional weeks of benefits to states experiencing high unemployment. The EB program is a state-federal partnership, with the costs of the extended benefits shared equally between the state and federal government. The EB program provided up to 13 weeks of additional benefits. The extended benefits can be "triggered" nationally when the unemployment rate crosses certain thresholds, or triggered within individual states when the state-level unemployment crosses certain thresholds.

Following the recession of 1969-1970, in addition to additional benefits provided by the EB program, the federal government passed the Emergency Unemployment Compensation Act of 1971 (EUCA) which provided for an additional 13 weeks of benefits to states with high unemployment financed fully by the federal government. Thus, unemployed workers could receive up to 52 weeks of benefits under the regular, EB and EUCA programs[12]. The EUCA provided benefits from January 1972 through March 1973.

[12] The triggers under EUCA were different than under the EB program. Thus some states only qualified for EB, others only for EUCA, and others for both EB and EUCA.

During the 1973-1975 recession, the federal government passed the Federal Supplemental Benefits (FSB) program, which was in effect from January 1975 through October 1977. The program initially provided for 13 weeks of additional benefits financed from the federal government, but was amended to provide 26 weeks of benefits in March 1975. The EB program triggered on nationwide from February 1975 through December 1977. Thus, from March 1975 through October 1977 displaced workers could receive a total of 65 weeks of benefits (26 state + 13 EB + 26 FSB).

In 1980 and 1981, through the Omnibus Reconciliation Acts of those years, the federal government altered the EB program. It eliminated the national trigger for EB and raised the thresholds for the state level triggers. In addition, it imposed stricter eligibility requirements for unemployed workers to receive benefits under the EB program.

During the 1981-1982 recession, the federal government established the Federal Supplemental Compensation (FSC) program in September of 1982. The tightening of the EB program under the OBRA legislation made roughly half of states ineligible to additional benefits under that program. FSC was amended several times from 1982 through early 1985. For the majority of the program duration, it provided up to 14 additional weeks of benefits financed by the federal government. Thus, the maximum weeks of benefits that could be received were 53 (26 state + 13 EB + 14 FSC).

After the 1990-1991 recession, the federal government passed the Emergency Unemployment Compensation (EUC) Act of 1991. The extension was amended several times from 1991 through 1994 providing at various times an additional 20, 26, 33 or 15 additional weeks of benefits. The benefits were financed entirely by the federal government. The maximum weeks of benefits that an individual could have received was 72 (26 state + 13 EB + 33 EUC in states with unemployment over 9%)), from February to June 1992. After that the maximum weeks of benefits declined to 65, where it remained for the majority of the EUC program. In addition, the EB program was amended to increase the maximum number of weeks payable. States with unemployment rates above 8% would now receive 20 weeks of benefits instead of 13.

In March 2002, after the 2001 recession, the federal government passed the Temporary Extended Unemployment Compensation (TEUC) act. The act provided up to 26 additional weeks of federally financed unemployment benefits through March of 2004. The maximum weeks of benefits that an individual could have received was 65 (26 state + 13 EB + 26 EUC).

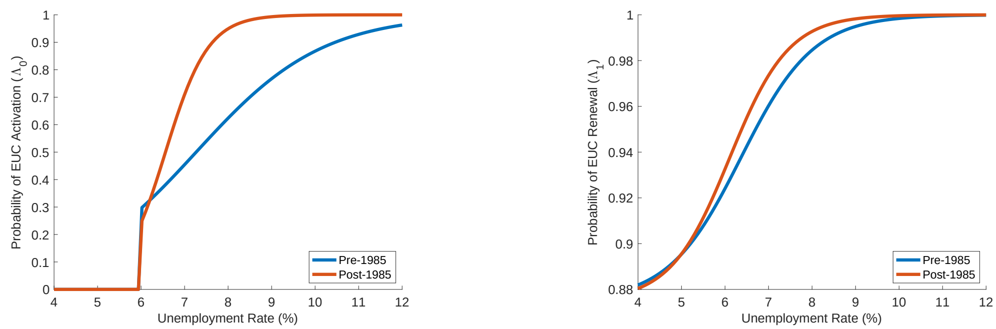
During the 2007-2009, the federal government passed the Emergency Unemployment Compensation (EUC08) Act of 2008. The program initially provided up to 13 weeks of additional benefits financed by the federal government. The EUC08 has been amended or reauthorized 12 times to day, gradually raising the maximum additional benefits provided by the federal government to 53 weeks, and thereby making the total compensation that an unemployed worker could receive 99 weeks (26 state + 20 EB + 53 EUC08). The program expired on January 1, 2014.

B Details on Estimation of Expectations

In order to map the realized history of emergency programs to beliefs about the discretionary programs we proceed as follows. First, we estimate the probability using a logit at a monthly frequency that an EUC program is passed as a function of the unemployment rate. We set the probability to 0 whenever the unemployment rate is less than six percent since all EUC programs passed have required unemployment above six percent to activate. We plot the estimated function in Figure 2a. For the renewal probabilities, $\Lambda_1(u) = \Pr(\text{not expire}) + \Pr(\text{expire}) \times \Pr(\text{renew}|\text{expire})$. We first compute that the average EUC program lasts for eight months, as legislated, and set $\Pr(\text{not expire}) = 7/8$. Next, we estimate $\Pr(\text{renew}|\text{expire})$ using a logit model in the data as a function of the unemployment rate when EUC programs lapses. Thus, Λ_1 is a combination of the expected duration of the program along with the probability that it gets renewed, conditional on expiring. We plot the estimated function in Figure 2b.

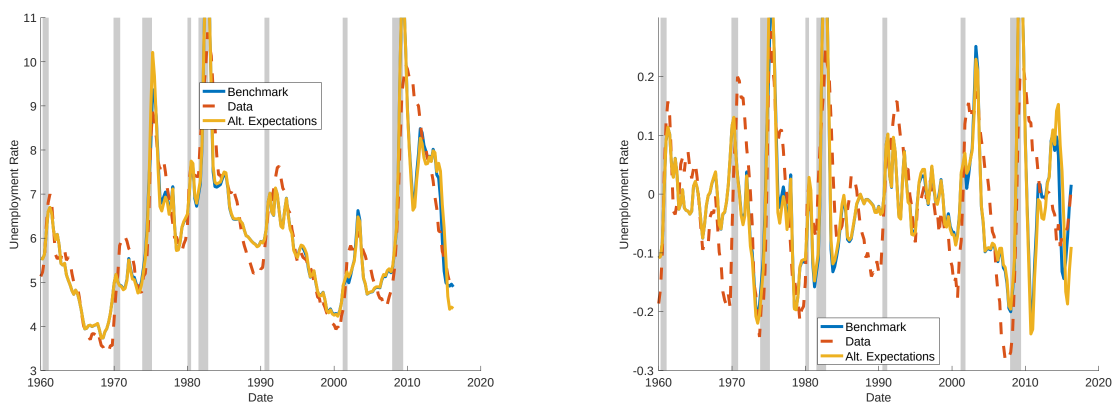
B.1 Alternative Process for Expectations

As a robustness check, we split the sample into pre-1985 and post-1985 periods, and estimate activation and renewal probabilities separately for the two periods. The results of this estimation are reported in Figure 8, which illustrates that only the estimated activation probabilities (rather than renewal probabilities) are substantially different before and after 1985. As described in the main text, we confirm that the alternative expectations process does not significantly change the simulation results. Figure 9 compares the baseline simulation to the simulation under this alternative expectation process, showing that the model fit is not materially different. Figure 10 confirms that the model similarly reproduces the individual recessions, both in the pre-jobless recoveries period of the 1970's and 1980's, and the jobless recoveries period of the 1990's and 2000's.



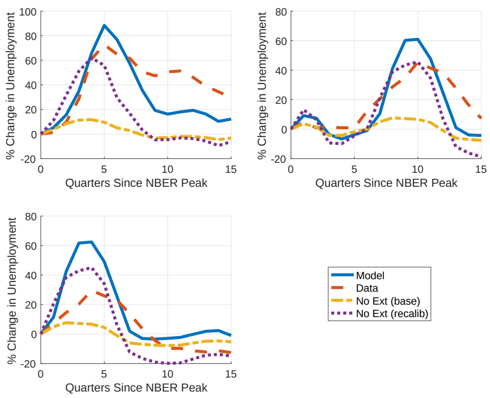
(a) Probability EUC program enacted (b) Probability EUC program renewed

Figure 8: Estimated activation (left panel) and renewal (right panel) probabilities as a function of the unemployment rate estimated from post-war EUC programs. Probabilities are monthly.

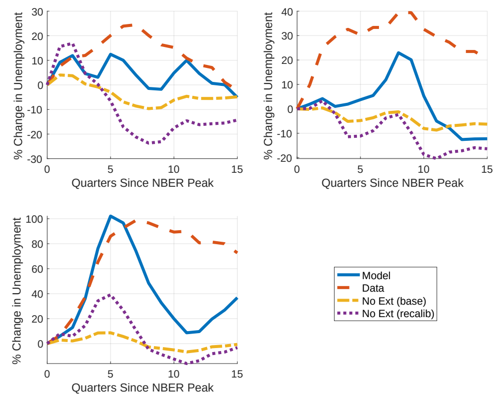


(a) Levels. (b) Deviations from trend

Figure 9: Simulated and actual labor market variables (levels left panel, deviations from HP-filtered trend right panel) from 1960:I through 2016:II. The benchmark model simulations are in blue, simulations from the model with the version of the model with alternate expectations are in yellow. The dashed red line is the data. NBER dated recessions are shaded.



(a) 1973-75, 1980 and 1981-82 recessions.

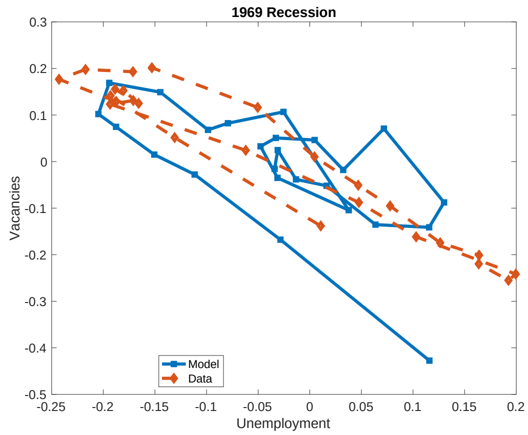


(b) 1990-91, 2001 and 2007-09 recessions.

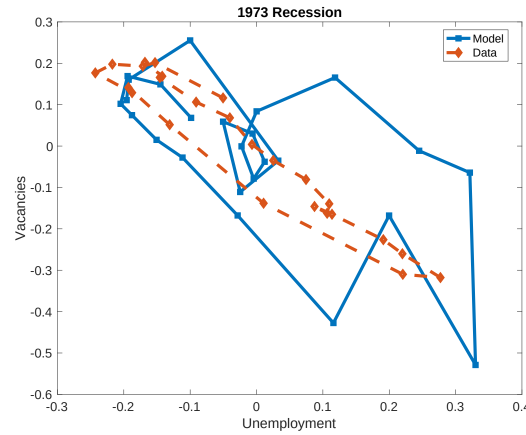
Figure 10: Simulated and actual percentage change in unemployment from NBER peak before NBER dated recessions. The blue line is the alternate expectations model, the dashed red line is the data, the yellow dot-dashed line is the model without extensions under the baseline calibration, and the purple dotted line is the model without extensions recalibrated. Data and model are not filtered. Data is from CPS, unemployment.

C Beveridge Curve in Earlier Recessions

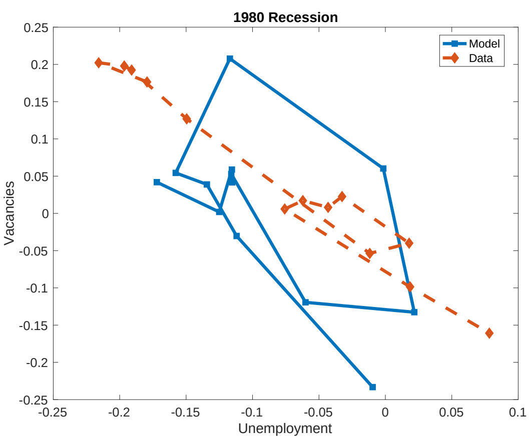
In this section we show that the model generates counterclockwise shifts in the Beveridge curve in all previous recessions, consistent with the data on recessions from 1969 onwards. In figure 11 we plot the Beveridge curve in deviations from trend and in figure 12 we plot both unemployment and vacancy rates in levels.



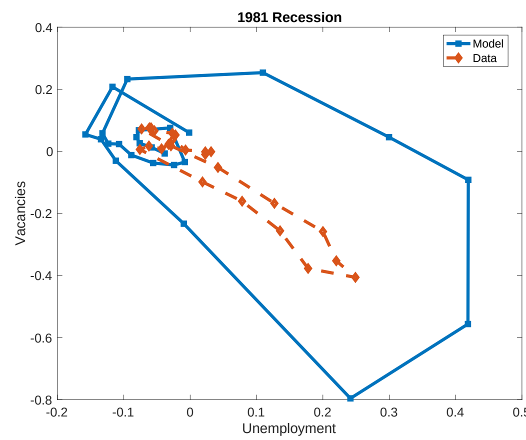
(a) 1969 Recession



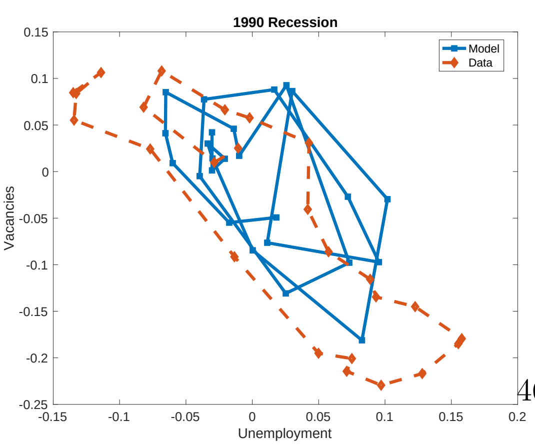
(b) 1973 Recession



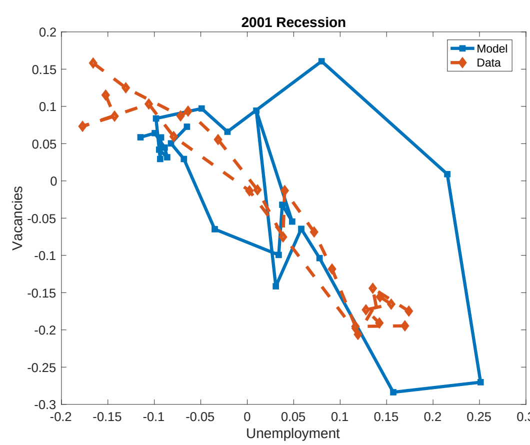
(c) 1980 Recession



(d) 1981 Recession

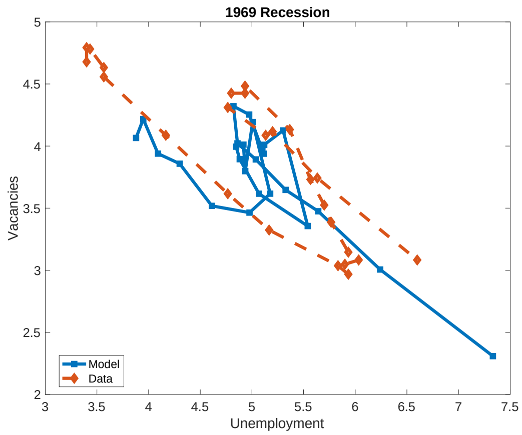


(e) 1990 Recession

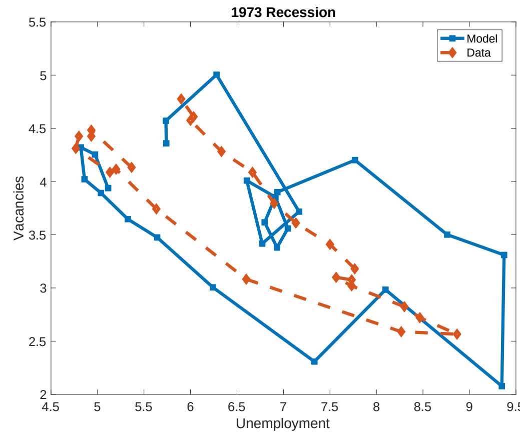


(f) 2001 Recession

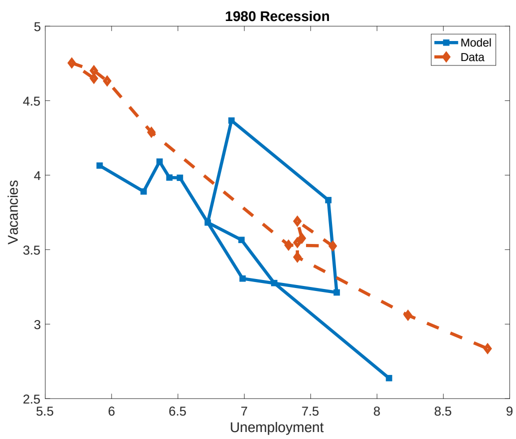
Figure 11: Beveridge Curve in Recessions (deviations from trend)



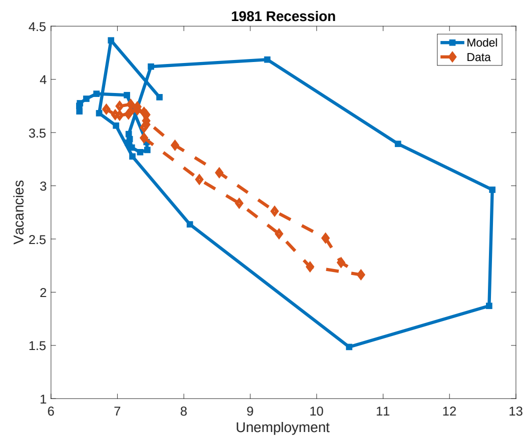
(a) 1969 Recession



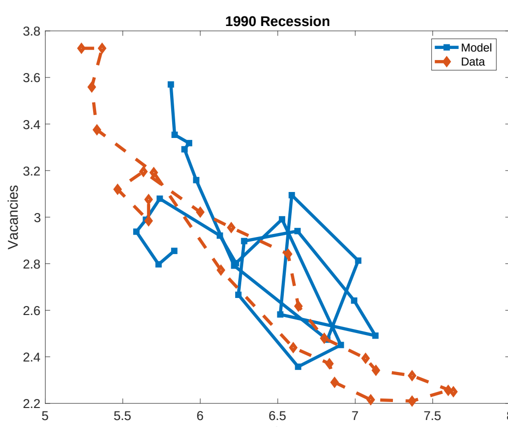
(b) 1973 Recession



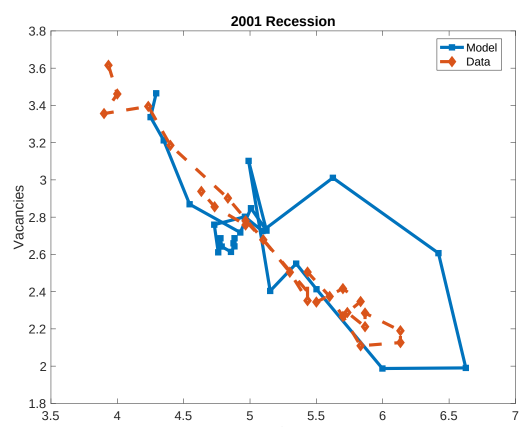
(c) 1980 Recession



(d) 1981 Recession



(e) 1990 Recession



(f) 2001 Recession

Figure 12: Beveridge Curve in Recessions (levels)

D Computation of the Stochastic Model

The challenge in computing the model is that the state space, Ω, is large. It contains exogenous productivity z, whether an emergency extension is currently in place, I^{EUC}, eligible and ineligible employment l^E, l^I, and the full distribution of unemployed across eligibility, u^I, u^E, and u_m. Since the maximum extension in our simulation is for 23 months, we have $m = 0, 1, \cdots, 23$. To make progress, we exploit the economics of the problem and note that the firm vacancy posting decision within the period depends only on the total unemployment rate and the share of unemployed that are eligible (and the expectation of those going forward). Thus, we approximate the full Ω, with $\hat{\Omega} = \{z, D, I^{EUC}, \varpi, u\}$. We approximate $H(\Omega)$ by conjecturing log-linear laws of motion for u and ϖ, with intercept and slope coefficients that can depend explicitly on the remaining agggregate states:

$$
\begin{aligned}
\log(u_{t+1}) &= a_0(z_t, D_t(I_t^{EUC}), D_{t+1}(I_{t+1}^{EUC})) \\
&+ a_1(z_t, D_t(I_t^{EUC}), D_{t+1}(I_{t+1}^{EUC})) \log(u_t) \\
&+ a_2(z_t, D_t(I_t^{EUC}), D_{t+1}(I_{t+1}^{EUC})) \log(\varpi_t) \\
\log(\varpi_{t+1}) &= b_0(z_t, D_t(I_t^{EUC}), D_{t+1}(I_{t+1}^{EUC})) \\
&+ b_1(z_t, D_t(I_t^{EUC}), D_{t+1}(I_{t+1}^{EUC})) \log(u_t) \\
&+ b_2(z_t, D_t(I_t^{EUC}), D_{t+1}(I_{t+1}^{EUC})) \log(\varpi_t)
\end{aligned}
$$

We conjecture coefficients a_0, a_1, and a_2, and solve b_0, b_1, and b_2, solve the worker and firm value functions problem, and then simulate the economy. When we simulate the economy we keep track of the full state Ω_t and use it to compute the approximated aggregate state $\hat{\Omega}_t$. Then, using the realized sequence of $\hat{\Omega}$, we perform the previous regression and check whether the implied coefficients are the same as the conjectured ones. If they are, we have found the law of motion; if not, we update our guess and repeat. We discretize the productivity process into 21 states. The benefit schedule consists typically of three levels of benefit entitlement, depending on the unemployment rate. Thus, we solve for $21 \times 3 \times 3 = 189$ laws of motion for both u and ϖ. Thus, while we assume that u and ϖ follow log-linear laws of motion conditional on productivity and benefits, we effectively allow the laws of motion to vary non-parametrically in the productivity and benefits states.

We repeat these 17 times, for each of the different configurations of emergency benefit extensions outlined in the previous section. Thus, in total we estimate 6,426 laws of motion. We find that our approximation procedure is highly accurate, with all of the R^2 in excess of 0.9999.