Finance and Economics Discussion Series

Federal Reserve Board, Washington, D.C.
ISSN 1936-2854 (Print)
ISSN 2767-3898 (Online)

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2025-062

Please cite this paper as:
Ahn, Hie Joo, and Yunjong Eo (2025). "Hysteresis and the Role of Downward Nominal Wage Rigidity: Evidence from U.S. States," Finance and Economics Discussion Series 2025-062r1. Washington: Board of Governors of the Federal Reserve System, https://doi.org/10.17016/FEDS.2025.062r1.

Hysteresis and the Role of Downward Nominal Wage Rigidity: Evidence from U.S. States[*]

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December 16, 2025

Abstract

This paper empirically investigates the sources of hysteresis, emphasizing the role of downward nominal wage rigidity using U.S. state-level payroll employment growth. U.S. states exhibit heterogeneous recoveries, with L-shaped and U-shaped recessions corresponding to persistent hysteresis and full recovery. L-shaped recessions are importantly driven by demand shocks and reinforced by downward nominal wage rigidity, which prolongs employment losses by raising real wages and deepening downturns. When wage rigidity is strong, expansionary policies are particularly effective in mitigating these effects through labor market adjustment. These mechanisms are validated in a New Keynesian model featuring both hysteresis and downward nominal wage rigidity.

JEL classification: C22; C51; E32; E37.
Keywords: Hysteresis, Regional business cycles; L-shaped recession; U-shaped recession; Downward nominal wage rigidity; Monetary policy; Fiscal policy.

[*]This paper, previously published under the title "Recession Shapes of Regional Evolution: Factors of Hysteresis" has been revised, adding a structural model to supplement the empirical analysis and focusing on the role of downward nominal wage rigidity in hysteresis by dropping discussions related to the gender employment gap. We thank Travis Berge, Joonkyu Choi, Masao Fukui, Francesco Furlanetto, Jordi Galí, Manuel González-Astudillo, Adam Gulan, James Hamilton, Yoon Joo Jo, Mikael Juselius, Antoine Lepetit, Nigel McClung, Michael McCracken, Raven Molloy, Emi Nakamura, Mike Owyang, Damjan Pfajfar, Jeremy Piger, Hannah Rubinton, Jeremy Rudd, Choongryul Yang and seminar participants at the Bank of Finland, the Federal Reserve Board, Hanyang University, the Hitotsubashi Summer Institute, Norges Bank, the Regional Conference of the Federal Reserve System for their helpful comments. Lucas Moyon provided excellent research assistance.

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1 Introduction

Hysteresis refers to a phenomenon where recessionary shocks have permanent or long-lasting effects on the level of economic activity (Cerra et al., 2023). The hysteresis view has gained attention as an explanation for the slow recovery after the Great Recession and, more recently, as a channel through which monetary and fiscal policies have long-lasting effects on the economy (e.g., Jordá et al., 2020; Antolin-Diaz and Surico, 2025). Reflecting this growing interest, recent studies have discussed various drivers of stagnant recoveries or hysteresis, such as wage rigidity (Shimer, 2012; Schmitt-Grohé and Uribe, 2017), gender convergence (Heathcote et al., 2017; Olsson, 2019; Fukui et al., 2023; Albanesi, 2025), structural changes (Jaimovich and Siu, 2020), and secular stagnation (Hall, 2016; Benigno and Fornaro, 2018). Despite its importance, the determinants of hysteresis have been difficult to empirically explore for several reasons. First, economic recessions are relatively rare events, making it difficult to reliably identify and estimate the drivers of hysteresis. Second, and more fundamentally, there is no consensus about how to measure hysteresis, resulting in relatively few empirical analyses of hysteresis based on a formal statistical framework.[1]

This paper empirically investigates factors of hysteresis and explores their policy implications. To overcome the aforementioned empirical challenges, we exploit regional heterogeneity in business-cycle experiences based on state-level payroll employment. We use payroll employment because it is a widely cited monthly cyclical indicator with a long sample period starting in 1960, available for all states, and it contains identifying information about hysteresis (Furlanetto et al., 2025).[2] Adopting the widely accepted notion of hysteresis where recovery from an economic recession does not bring economic activity back to its pre-recession trend, we distinguish between L-shaped and U-shaped recessions: an L-shaped recession reflects hysteresis, whereas a U-shaped recession is characterized by a full return of the economy to its pre-recession growth path.[3] We then estimate the likelihood of the two types of recessions at each point in time for all the U.S. states based on a Markov switching model of Eo and Morley (2022).

[1] Previous studies have primarily relied on structural models to analyze the causes and consequences of hysteresis, due to the empirical challenges of identifying hysteresis episodes and their structural drivers using aggregate data. Recent empirical developments include Furlanetto et al. (2025) based on a structural VAR and Antolin-Diaz and Surico (2025) based on a BVAR model.

[2] Separately, Dupraz et al. (2025) analyze U-shaped recessions through the lens of plucking theory using a structural labor-market model.

[3] Hereafter, the terms hysteresis and L-shaped recession are used interchangeably.

Among the various factors potentially contributing to hysteresis, we focus on downward nominal wage rigidity. In structural models of hysteresis, downward nominal wage rigidity serves as a key mechanism through which recessions inflict lasting damage on the labor market, as downward nominal wage rigidity puts upward pressure on real wages and hence exacerbates job destruction during downturns (e.g., Acharya et al., 2022; Alves and Violante, 2024). In our empirical analyses, we comprehensively control the state-level attributes such as industry composition, minimum wages, and labor market outcomes by gender in a comprehensive manner by building on a coherent empirical framework.[4]

We find that U.S. states exhibit broadly similar but distinct recession dynamics in terms of shape, timing, and magnitude. A state's business cycle does not always align with the national cycle and often displays notable idiosyncrasies; just over half of the states experience recessions that coincide with the national recessions identified by the NBER. Moreover, U-shaped recessions become less frequent after the 1990s. The emergence of jobless recoveries in the aggregate economy during this period is associated with the reduction in the prevalence of U-shaped recessions across states. Nonetheless, our model classifies the sharp and rapid COVID-19 recession as a U-shaped recession.

The two types of recessions are driven by both supply and demand shocks, but they yield markedly different macroeconomic outcomes. The statistically significant link between demand shocks and L-shaped recessions supports the hysteresis hypothesis (e.g., Blanchard and Summers, 1986; Blanchard, 2018), which posits that demand shocks can generate persistent, long-run effects on the economic activity. Moreover, following L-shaped recessions, the negative effects on the employment-to-population ratio and output tend to be more persistent, whereas U-shaped recessions are characterized by a swift rebound in these variables after the recession.

Notably, L-shaped recessions tend to put upside pressure on inflation. This is because they reduce the economy's potential output or raise the natural rate of unemployment. Consequently, the resulting economic slack is much narrower, and the economy's cyclical position improves *ceteris paribus*, exerting upward pressure on price inflation. In contrast, U-shaped recessions exert persistent downward pressure on inflation by worsening the cyclical position of the economy

[4] Recent empirical studies also suggest that higher female employment facilitates faster recoveries (e.g., Cortes et al., 2018; Fukui et al., 2023; Bergholt et al., 2024). This factor and downward nominal wage rigidity are, in fact, correlated: increased female employment raises downward nominal wage rigidity, as women are disproportionately concentrated at the lower end of the wage distribution, where wages are more likely to be downwardly rigid (Jo, 2024). We explicitly consider labor market outcomes by gender in our empirical analyses to control effects of the gender evolution in the labor market.

without permanently altering its potential output.

We also find that downward nominal wage rigidity raises the likelihood of hysteresis. States with a higher share of zero nominal wage changes during economic recessions are more likely to experience L-shaped rather than U-shaped recessions. This pattern holds consistently over the full sample period (1978:Q1–2019:Q4) as well as in the post-2000 period, during which the gender employment gap slowed its narrowing—one factor behind stagnant recession recoveries (Fukui et al., 2023). This finding provides empirical support for structural models where downward nominal wage rigidity serves as a crucial mechanism through which hysteresis arises (Acharya et al., 2022).

Finally, we empirically examine the extent to which demand-side policies—namely, expansionary monetary and tax policies—mitigate hysteresis, and the roles of downward nominal wage rigidity in shaping policy effectiveness. We find that expansionary monetary and tax policies are more effective in mitigating hysteresis when downward nominal wage rigidity is greater. This finding aligns with the conventional view that monetary policy affects the real economy through nominal rigidities. In addition, when labor costs are inflexible due to wage rigidity, tax cuts effectively reduce production costs, helping to sustain labor demand and thereby mitigate hysteresis. This is consistent with Lee (2025), who finds that the multiplier effects of expansionary tax policies are larger when nominal wage rigidity is greater.

To interpret these empirical findings, we turn to a prototype New Keynesian model. The framework builds on Galí (2022), in which an insider–outsider labor market structure gives rise to hysteresis in the form of L-shaped recessions. We extend this model by incorporating downward nominal wage rigidity to emphasize its amplifying role in hysteresis and to evaluate the effectiveness of monetary policy interventions in such an environment. The analysis shows that, in the presence of downward nominal wage rigidity, a negative demand shock results in deeper downturns in employment and output, but a smaller reduction in inflation, compared with the case without such rigidity. It is because downward nominal wage rigidity prevents real wages from falling, which constrains firms' ability to adjust employment flexibly and, consequently, results in a larger employment reduction than would occur under flexible adjustment.

Furthermore, expansionary demand policies effectively mitigate hysteresis, with their impact strengthened by the presence of downward nominal wage rigidity. Therefore, our structural model highlights the dual role of wage rigidity: while it amplifies hysteresis, it also enhances the effectiveness of monetary policy. These results underscore the importance of timely and forceful demand-side interventions to prevent persistent economic scarring.

These findings align with previous research highlighting the long-run effects of monetary and fiscal policies (e.g., Jordá et al., 2020, Antolin-Diaz and Surico, 2025) and the hysteresis-mitigating impact of timely interventions of monetary policy (e.g., Acharya et al., 2022). Distinguishing our analysis from previous studies, we further identify the conditions under which demand-side policies are more effective. Our finding highlights the two faces of nominal wage rigidity. Though downward nominal wage rigidity is the key mechanism through which hysteresis is more likely to be created, demand-boosting policies become more effective in mitigating hysteresis when downward nominal wage rigidity is greater. Previous studies on fiscal multipliers have found that the effects of tax policy or government spending are larger when nominal wages are more downwardly rigid (e.g., de Ridder and Pfajfar, 2017; Shen and Yang, 2018; Jo and Zubairy, 2025; Lee, 2025), although these studies do not account for hysteresis.

This paper lies at the intersection of several strands of the macroeconomic and econometric literature on business cycles.[5] We contribute to the growing body of research on macroeconomic hysteresis by linking it to the long-standing, yet still evolving, econometric literature on recession prediction. To the best of our knowledge, this is the first empirical study to investigate downward nominal wage rigidity as a key source of hysteresis within a formal statistical model of business cycle dynamics using regional data. A further contribution of the paper is to uncover the structural mechanism underlying the empirical findings by developing a New Keynesian model with an insider–outsider labor market structure that incorporates downward nominal wage rigidity.

Unlike the previous studies on hysteresis, we explicitly measure the likelihood of hysteresis using the Markov-switching model developed by Eo and Morley (2022). Previous empirical literature on hysteresis has relied on time series models for aggregate data that distinguish the trend and cycle components of macroeconomic time series and examines effects of structural shocks including supply, demand, and policy shocks on the components of time series (e.g., Furlanetto et al., 2025). Another strand of the empirical literature employs cross-country data to examine country-level differences in long-run growth experiences and the long-run effects of monetary policy (e.g., Cerra and Saxena, 2008; Jordá et al., 2020). In contrast, we first identify the likelihood of hysteresis at the state level based on a Bayesian Markov switching model and then examine the effects of policy shocks and state-specific structural factors on hysteresis.

Our paper also contributes to the literature on the identification of business cycle phases. Pre-

[5] A detailed literature review is provided in Section Appendix A.

vious studies on regional business cycles have typically employed two-regime Markov-switching models that distinguish only between expansions and recessions, focusing primarily on the determinants of business cycle duration (e.g., Owyang et al., 2005; Hamilton and Owyang, 2012; Francis et al., 2018). While Eo and Morley (2022) estimate a three-regime Markov-switching model with two recession types for the aggregate economy based on GDP growth, our study focuses on regional heterogeneity, inferring recession experiences from payroll employment growth for each state. This state-specific approach effectively identifies each state's business cycle phases and recovery shapes, fully capturing regional heterogeneity in business cycles and their determinants in structural factors. To the best of our knowledge, this paper is the first to apply a Markov-switching model with two types of recessions to state-level data and to investigate the sources of recovery shapes.

This paper is organized as follows. Section 2 describes the data used in the empirical analysis. Section 3 explains our strategy for identifying state-level recovery shapes and presents the empirical findings on their heterogeneous patterns. Section 4 documents the key characteristics of the two types of recessions, with a focus on their associations with key macroeconomic variables. Section 5 examines the role of downward nominal wage rigidity in driving hysteresis. Section 6 analyzes how downward nominal wage rigidity influences the effectiveness of policy interventions in mitigating hysteresis. Section 7 validates our empirical findings within a New Keynesian framework that incorporates hysteresis and downward nominal wage rigidity. Section 8 concludes.

2 Data

Section 2.1 discusses state-level payroll employment data. Section 2.2 outlines observed state-level attributes including the measure of nominal wage rigidity.

2.1 State-level Employment Data

As a measure of state-level economic activity, we use nonfarm payroll employment growth by state for several reasons. First, job gains are among the most frequently cited cyclical indicators (e.g., Abraham et al., 2013).[6] Second, monthly payroll employment data are available from 1960.

[6] Owyang et al. (2015) mention that payroll employment is the broadest measure of economic activity.

Identifying hysteresis separately from a bounce-back full recovery and estimating its evolution requires monthly or quarterly data with a sample period long enough to capture the sufficient number of two recession episodes. Given that economic recessions are rare events, losing a few recessions is fairly costly for the identification of two recession types, and hence a longer sample period enhances the precision of estimates. In this sense, the state-level payroll employment data are ideal for our empirical analysis. Alternatively, one might consider state-level output data or unemployment rates from the Local Area Unemployment Statistics. However, quarterly state-level GDP data are only available from 2005, and state-level nominal output data are available from 1963, but only at an annual frequency. In addition, state-level unemployment and labor force participation rates are available monthly, but only from 1976 onward, omitting several important economic recessions that occurred between 1960 and 1975.[7]

For these reasons, previous research on regional business cycles has frequently relied on payroll employment growth to analyze state-level business cycles (e.g., Hamilton and Owyang, 2012). We convert the monthly data to a quarterly frequency to align with the time frame commonly used in the previous literature on business cycles. The sample period spans from 1960:Q1 to 2023:Q4.

Figures B1 - B2 in the appendix display nonfarm payroll employment growth (red lines) and its long-run trend (blue line) by state. Panel (a) also displays the national payroll growth. While states exhibit common procyclical variation in the payroll growth, they differ in both the magnitude and the shape of their recoveries. Additionally, some states experience decelerations in payroll employment not seen elsewhere. For instance, Louisiana (Panel (t) in Figure B1) saw a sharp decline in its payroll employment in 2005 due to Hurricane Katrina. Similarly, due to the boom and subsequent bust of shale oil production, North Dakota (Panel (j) in Figure B2) experienced a rapid payroll growth after the Great Recession, followed by a prolonged decline in its payroll employment.

2.2 Attributes of States: Downward Nominal Wage Rigidity

This section examines state-level covariates that are included in our analyses.

First, as a measure of downward nominal wage rigidity, we employ the fraction of zero

[7] Employment data are relatively free from measurement errors that have cyclical features relative to unemployment rate and labor force participation rate (Ahn and Hamilton, 2022), which is an additional benefit of using the employment data for the identification of recession types.

nominal wage changes out of total wage changes by state constructed by Jo (2024). Based on the Current Population Survey (CPS, 1979–2018), Jo measures the share of workers with no changes in hourly wages, with hourly wage cuts, and with hourly wage increases by state. These three statistics, summarizing the nominal wage change distribution, show asymmetry between wage increases and cuts and a spike at zero, which the author interprets to represent downward nominal wage rigidity.[8] The three shares show sufficient variation across states and over time to permit state-level analyses on the propagation of policy shocks or business cycles (Jo, 2024; Lee, 2025).

As additional state-level attributes, we consider factors reflecting industry structure such as oil production and the employment shares of manufacturing, professional and business services, and finance. Specifically, we construct an indicator variable for oil-producing states, which takes the value of one if a state has positive oil production and zero otherwise. For the other industry measures, we use the industry's employment share out of the state's nonfarm payroll employment. For union membership, we use the fraction of workers who are union members out of the state's nonfarm payroll employment.[9] As a proxy for market competition or monopsony power, we use the fraction of workers employed by firms with 500 or more employees. In addition, we include indicators for Census regions to account for heterogeneity across broader geographic areas. We also consider the gender gap in labor market outcomes, measured by the difference in the employment-to-population (EPOP) ratio between men and women.

In addition, we take state-specific policy variables into consideration, including the minimum wage and the tax-to-income ratio (for a state's minimum wage, we use the higher of the state-level minimum wage and the federal minimum wage). Detailed information on the sources of the state-level data is provided in Appendix B.

For state-level inflation in nontradables, tradables, and all categories, we use the estimates from Hazell et al. (2022). We also employ the state-level macroeconomic dataset compiled by Jo and Zubairy (2025), which includes nominal gross state product (GSP), employment levels,

[8] The estimates are found in the author's website (`https://sites.google.com/view/yoonjoojo/rsearch`).

[9] For the employment shares of the three industries, we use BEA's annual nonfarm wage-and-salary employment and its industry breakdown based on the SIC classification (1969–2001). These data exclude proprietors' employment (self-employment) and therefore conceptually align with CES (establishment) employment measures, which does not include the self-employed. For the more recent period, we use nonfarm payroll employment (CES concept) from the BLS' *State and Metro Area Employment, Hours, & Earnings*, which is part of the CES program. Although these data are available beginning in 1990, we use this source to compute the industry employment shares only from 2002 onward, corresponding to the years for which the BEA sectoral data are unavailable. We address potential structural breaks arising from the change in data source by including time fixed effects, as discussed further in Section 5.

and the consumer price index at an annual frequency to construct state-level real GSP and labor productivity.

For externally identified policy shocks, we use monetary policy shocks from Romer and Romer (2004), as extended by Wieland and Yang (2020), and tax shocks constructed by Romer and Romer (2010).[10]

3 The Shapes of Recoveries at the State Level

Section 3.1 presents the model and the identification of the two recession types. Section 3.2 describes our estimation procedure. Section 3.3 reports the national-level results, and Section 3.4 reports the state-level results.

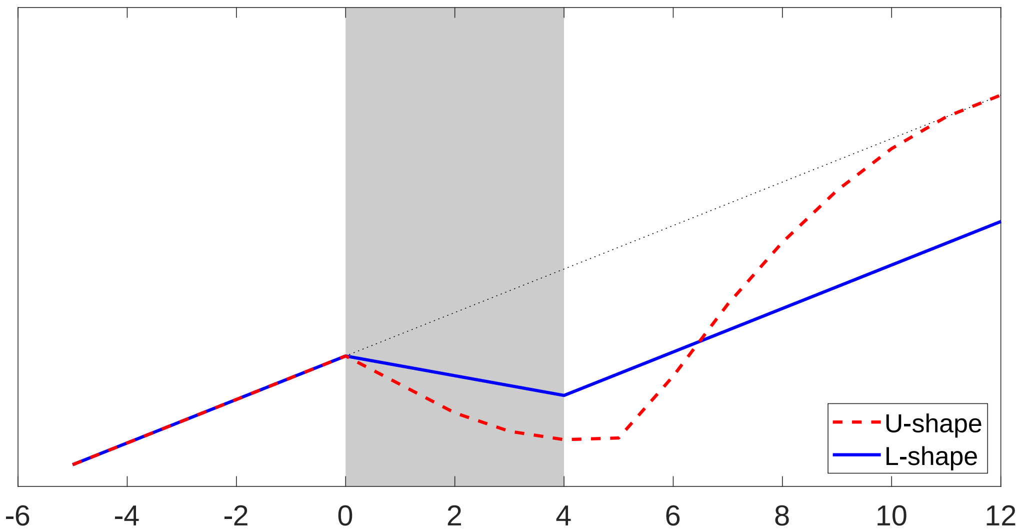
3.1 Model and Identification

To identify the shapes of recession recoveries across U.S. states, we employ the Markov-switching model of business cycles from Eo and Morley (2022). The model allows a given recession to either permanently alter the level of employment (i.e., an L-shaped recession) or only have a temporary effect (i.e., a U-shaped recession). An L-shaped recovery is characterized by an initial decline in economic activity, followed by an expansion that does not restore employment to its pre-recession level. In contrast, a U-shaped recovery also begins with a sharp decline but is followed by a strong rebound, during which growth temporarily exceeds its pre-recession pace, allowing the economy to return to its trend growth path. Illustrations of L-shaped and U-shaped recoveries are provided in Figure 1.

The change in payroll employment at the state level is assumed to follow a first-order Markov-switching process with three regimes: expansion, U-shaped recession, and L-shaped recession.

[10] Since monetary policy shocks identified from high-frequency data are only available from the late 1980s, we use Romer and Romer (2004)'s estimates, which cover a longer sample period starting in 1969. For tax shocks, Lee (2025) construct state-level average marginal tax shock estimates and corresponding instrumental variables for the period 1980–2000. However, these series are annual and thus not directly applicable to our empirical analysis without strong assumptions regarding the frequency or timing of the shocks. Therefore, for the sake of transparency, we directly employ aggregate tax shock estimates from Romer and Romer (2010).

Figure 1: ILLUSTRATION OF L-SHAPED AND U-SHAPED RECOVERIES



Notes to figure: The X-axis denotes periods after recession. The black dotted line indicates the hypothetical employment level if a recession had not occurred. The shaded area represents periods of recession.
Source: Authors' calculation.

For the geographic state i,

$$\begin{aligned}\Delta y_{i,t} - \tilde{\mu}_{i,t} = \mu_{i,E} \quad &+ \quad \mu_{i,L} \cdot \chi_{i,t} \cdot \mathbf{1}(S_{i,t} = L) \\ &+ \quad \mu_{i,U} \cdot \chi_{i,t} \cdot \mathbf{1}(S_{i,t} = U) + \lambda_i \cdot \sum_{k=1}^{m} \chi_{i,t-k} \cdot \mathbf{1}(S_{i,t-k} = U) \\ &+ \quad \chi_{i,t} \cdot e_{i,t}, \quad e_{i,t} \sim i.i.d.N\left(0, \sigma_i^2\right) \end{aligned} \tag{3.1}$$

where $\Delta y_{i,t}$ is the quarterly growth rate of payroll employment for state i, $S_{i,t} = E, L, U$ is the regime indicator, with $S_{i,t} = E$ corresponding to the expansion regime, $S_{i,t} = L$ to the L-shaped recovery regime, and $S_{i,t} = U$ to the U-shaped recovery regime, $\mathbf{1}(\cdot)$ is an indicator function, and $\mu_{i,E}$, $\mu_{i,L}$, and $\mu_{i,U}$ are the conditional means for the respective regimes. The bounceback effect λ_i represents the strong recovery over m quarters for $t+1, \ldots, t+m$ following the U-shaped recession shock, $\mu_{i,U}$, at time t, such that $\lambda_i \cdot m + \mu_{i,U} = 0$. Therefore, λ_i is calibrated as a form of the restriction without estimation. To account for state-specific changes in long-run trend employment growth, we use "dynamic demeaning" for employment growth.[11] Dynamic demeaning is essentially the 10-year moving average of employment growth: $\tilde{\mu}_{i,t} \equiv \frac{1}{40} \sum_{j=0}^{39} \Delta y_{i,t-j}$.[12]

[11] Eo and Morley (2022) consider this treatment in their robustness analysis. Relatedly, see Eo and Kim (2016) for a discussion of the importance of allowing for time variation in long-run trend growth when estimating Markov-switching models of the business cycle.

[12] Kamber et al. (2018) provides in-depth discussion on the usefulness of dynamic demeaning and the choice of 10

Moreover, to address extreme outliers arising from the COVID-19 pandemic, as shown in Figure 2, we adopt a decay function following Lenza and Primiceri (2022) and Eo and Morley (2023). The scaling parameter $\chi_{i,t}$ is specified as follows. For the period prior to the onset of the COVID-19 pandemic ($t^* = 2020Q2$), we set $\chi_{i,t} = 1$ for all $t < 2020Q2$. After this period, we employ a scaling factor defined as $\chi_{i,t^*+j} = c_i + (1 - c_i)\rho_i^j$, where j denotes the number of periods since the pandemic began, reflecting the gradual decline in the size of COVID-19 shocks. We estimate these parameters without imposing any restrictions on their values and do not preassign them to any particular type of recession *ex ante*. The decay parameter ρ_i is constrained to lie between 0 and 1 during estimation.

For employment growth during contractionary regimes, we set the conditional means to $\mu_{i,L}$ and $\mu_{i,U}$ for L-shaped and U-shaped recessions, respectively, when $t <$ 2020Q2, and to $(\chi_{i,t}\mu_{i,L})$ and $(\chi_{i,t}\mu_{i,U})$ otherwise. Accordingly, the bounceback effect for the U-shaped recovery is defined as $\lambda_{i,t-k} = -\mu_{i,U}/m$ for $t <$ 2020Q2, and $\lambda_{i,t-k} = -\chi_{i,t}\mu_{i,U}/m$ otherwise.

The indicator $S_{i,t}$ is a latent Markov-switching reggime variable governed by transition probabilities where the probability of moving from regime k at time $t-1$ to regime j at time t is $Pr[S_{i,t} = j \mid S_{i,t-1} = k] = p_{i,kj}$ for $k, j \in \{E, L, U\}$. In addition, we rule out direct transitions between L-shaped and U-shaped recession regimes, requiring that any movement between them passes through an expansionary regime. This restriction is imposed as

$$p_{i,LU} = Pr[S_{i,t} = U \mid S_{i,t-1} = L] = 0, \qquad p_{i,UL} = Pr[S_{i,t} = L \mid S_{i,t-1} = U] = 0.$$

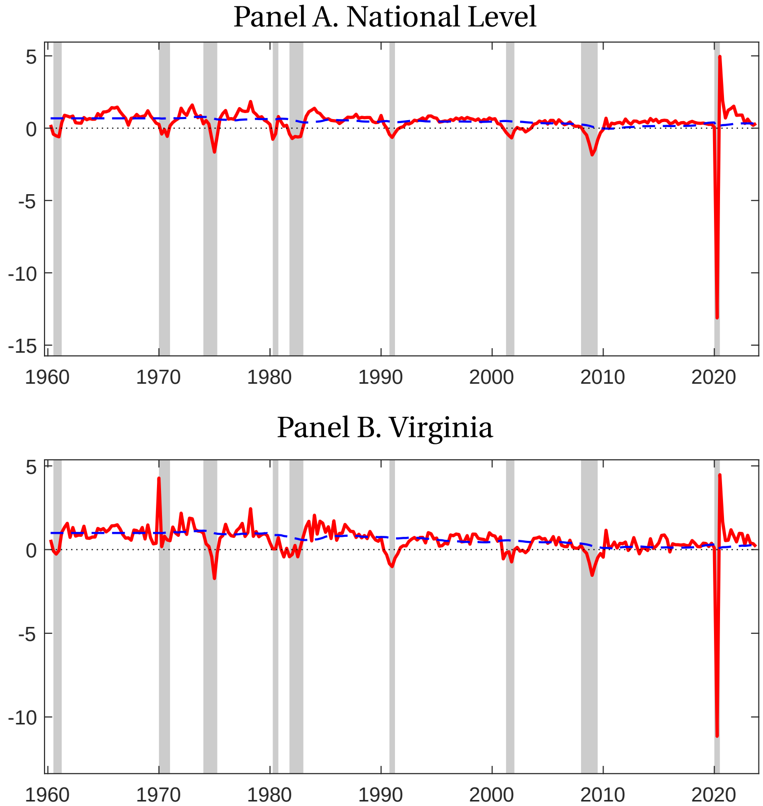
Therefore, the regime transition matrix for state i is given by

$$\Pi_i = \begin{bmatrix} 1 - p_{i,EL} - p_{i,EU} & p_{i,EL} & p_{i,EU} \\ 1 - p_{i,LL} & p_{i,LL} & 0 \\ 1 - p_{i,UU} & 0 & p_{i,UU} \end{bmatrix},$$

where the sum of elements in the kth row satisfies $\sum_{j \in \{E,L,U\}} p_{i,kj} = 1$. For example, the $(1,1)$ element of the matrix is $p_{i,EE} = 1 - p_{i,EL} - p_{i,EU}$.

years window.

Figure 2: NONFARM PAYROLL EMPLOYMENT GROWTH AT THE NATIONAL LEVEL AND IN VIRGINIA



Notes to figure: The figures plot nonfarm payroll employment growth at the national level (Panel A) and in Virginia (Panel B). The red line represents the payroll employment growth, while the blue dashed line shows the long-run growth rate, calculated using a 40-quarter rolling average.

Source: BLS, Haver, and authors' calculation

3.2 Bayesian Estimation

The length of the post-recession bounce-back period was set to $m = 5$ quarters following Eo and Morley (2022).[13] The model parameters are estimated using a Bayesian approach, which accommodates the irregular likelihood function inherent in regime-switching models. This estimation strategy is selected for its robustness in capturing parameter uncertainty and its capacity to incorporate prior information (Owyang et al., 2005).

3.2.1 Priors

The prior distribution for the transition probabilities from the expansion regime $(p_{E,E}, p_{E,L}, p_{E,U})$ follows a Dirichlet distribution with parameters $Dirichlet(36, 1, 1)$, while those for the L-shaped

[13] As a robustness check, we consider alternative values for m, and the results remain robust to these alternative values.

and U-shaped recession regimes $\left(p_{L,L}, p_{U,U}\right)$ follow a Beta distribution with parameters *Beta*$(5,1)$.[14] The prior distributions for the conditional means $\left(\mu_E, \mu_L, \mu_U\right)$ are given by *Normal*$(1,1)$, *Normal*$(-2,1)$, and *Normal*$(-2,1)$, respectively. The prior distributions for the scaling parameters (c, ρ) are *Normal*$(5,1)$ and *Beta*$(8,2)$, respectively. Finally, the prior distribution for the error variance σ^2 is given by *Inverse Gamma*$(10,5)$. These priors are relatively diffuse to allow for flexibility in estimation.

3.2.2 MCMC Procedure

For Bayesian estimation, we employ Markov Chain Monte Carlo (MCMC) sampling techniques to estimate the model parameters in each state. For notational convenience, we suppress the state indicator i from the parameters. Specifically, Metropolis-Hastings sampling with a random walk proposal is utilized for the COVID-19 scaling parameters, denoted as c and ρ. For the remaining parameters, Gibbs sampling is implemented. The priors for all parameters are set according to established and standard values found in the literature. The MCMC procedure involves generating 10,000 draws, with the initial 5,000 draws discarded as burn-ins to ensure the convergence of the sampling process. This sampling approach allows for efficient exploration of the posterior distributions of the parameters, particularly given the high dimensionality of the model and the complexity introduced by the COVID-19 scaling parameters. Let $\mathbf{Y} = \left\{\Delta y_t\right\}_{t=1}^{T}$, $\Theta \equiv \left(\mu_E, \mu_L, \mu_U, \sigma^2\right)$, $\mathbf{P} \equiv \left(p_{E,L}, p_{E,U}, p_{L,L}, p_{U,U}\right)$, $\mathbf{S} = \{S_t\}_{t=1}^{T}$, and $\Gamma = (c, \rho)$. The following summarizes the posterior sampling algorithm.

MCMC Sampling Procedure

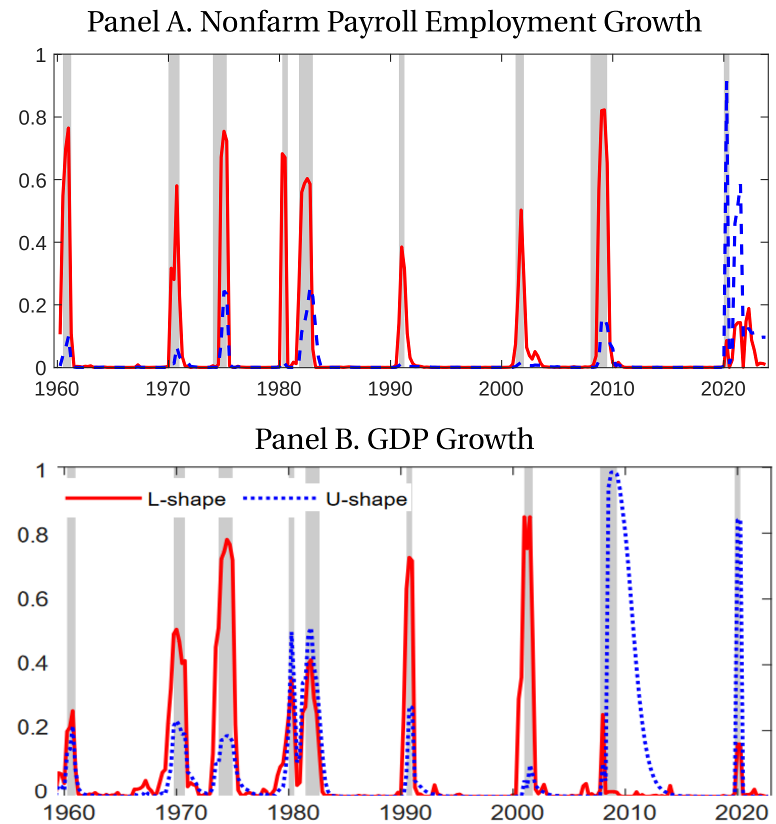
- Step 1: Gibbs Sampling $\Theta|\mathbf{Y}, \mathbf{S}, \Gamma$
- Step 2: Gibbs Sampling $\mathbf{S}|\mathbf{Y}, \Theta, \Gamma, \mathbf{P}$
- Step 3: Gibbs Sampling $\mathbf{P}|\mathbf{S}$
- Step 4: Metropolis-Hastings Sampling $\Gamma|\mathbf{Y}, \mathbf{S}, \Theta$

In the Metropolis–Hastings algorithm, we aim for an acceptance rate for the scaling parameters (c, ρ) between 0.15 and 0.40.[15]

[14] These priors imply that the expected duration of expansions is 19 quarters, while the expected durations of L-shaped and U-shaped recessions are both 6 quarters.

[15] Gelman et al. (1997) suggest an optimal acceptance rate of 0.234 with a random walk proposal.

Figure 3: National Level Recession Probabilities



Notes to figure: The figures display the probabilities of L-shaped and U-shaped recessions at the national level, estimated using nonfarm payroll employment growth (Panel A) and real GDP growth (Panel B). The blue lines represent the probability of a U-shaped recession, while the red lines represent the probability of an L-shaped recession. The Y-axis indicates the probability, and the X-axis represents calendar time in quarter.

Source: Authors' calculation

3.3 National Level

Figure 3 presents the national-level recession probabilities estimated based on the payroll employment growth (Panel A) and the real GDP growth (Panel B). We compare the two national-level estimates to investigate the extent to which payroll employment and real GDP provide consistent signals about business-cycle phases. A few important points are worth noting.

First, the employment-based estimates closely align with the NBER recession dates, highlighting their timeliness and effectiveness in identifying phases of the business cycle. One exception is the second wave of COVID-19. Although both estimates and the NBER recession chronology agree on the pandemic recession itself, the employment-based estimate interprets the second wave as another U-shaped recession that is less severe than the pandemic recession. In contrast, the GDP-based estimate and the NBER recession chronology both classify the second wave as part of expansion. Indeed, consistent with the employment-based estimate, initial unemployment claims increased between November 2020 and January 2021. This discrepancy underscores the usefulness of payroll employment in assessing the cyclical position of the economy.

Second, overall, L-shaped recessions are more pronounced in employment growth than in real GDP growth, possibly reflecting the phenomenon of jobless recoveries, a pattern that becomes increasingly evident since the 1990s.[16] Notably, the employment-based estimate classifies the Great Recession as an L-shaped recession, while the GDP-based estimate categorizes it as a prolonged U-shaped recession. Following the Great Recession, both overall and long-term unemployment recovered slowly, taking nearly a decade to return to pre-recession levels. Although the GDP-based estimate labels the recession as U-shaped, it still indicates a recovery duration of approximately five years, the longest on record.[17] In contrast, the employment-based estimate classifies it as L-shaped, capturing near-permanent employment losses that align with the prolonged weakness in both labor market and output recovery in the years that followed.

3.4 State Level

Figure 4 presents heatmaps illustrating the evolution of recession probabilities across all states.[18] The figure highlights variation in the types of recessions across states and over time, revealing several notable patterns.

First, roughly half of the states exhibit recession patterns that are broadly consistent with the national recessions identified by the NBER, while others show deviations, either missing some national recessions or experiencing additional episodes not recognized at the national level. States also differ in both the shape and timing of their economic downturns. As an example, Figure 5 shows the estimated recession probabilities for two states: New York and Wisconsin. New York did not experience recessions during the 1980s, whereas Wisconsin, a state with a large manufacturing sector, underwent pronounced L-shaped recessions. In contrast, New York faced an L-shaped recession in the early 1990s, while Wisconsin did not experience a downturn during that period.[19] Moreover, some states undergo independent economic downturns not shared by

[16] Jaimovich and Siu (2020) show that the declining employment share of routine occupations has driven jobless recoveries.

[17] For estimates of potential output and their implications for assessing the impact of the Great Recession, see Coibion and Ulate (2018).

[18] The full set of state-level recession probability estimates is provided in Figures C3 and C4 in the appendix.

[19] Both New York and Wisconsin experienced U-shaped recessions during the COVID-19 pandemic. In 2021, New York underwent a pronounced L-shaped recession, reflecting severe damage in sectors such as hospitality,

Figure 4: PROBABILITIES OF L- AND U- SHAPE RECESSIONS ACROSS STATES OVER TIME

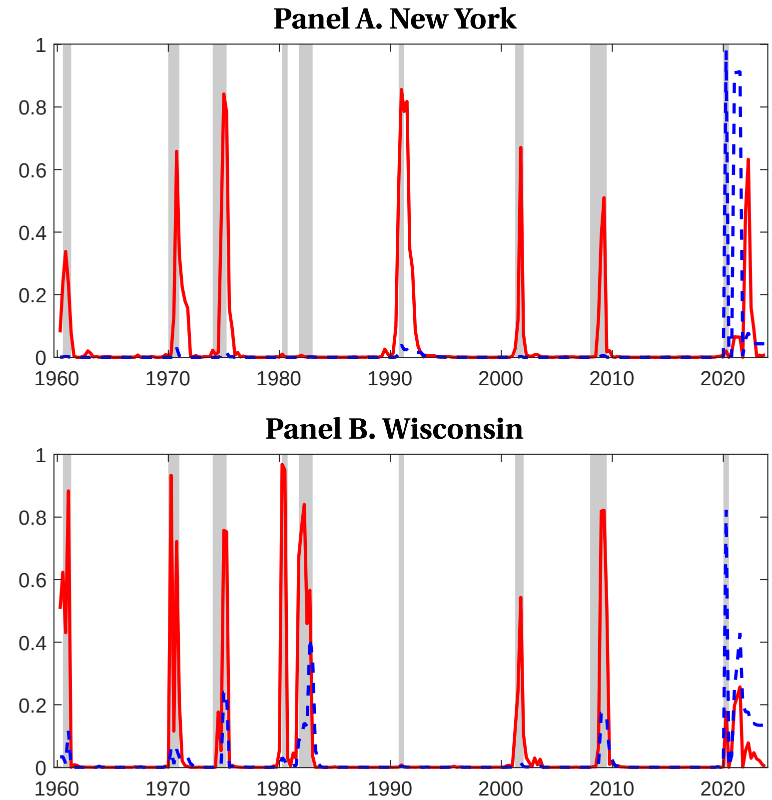
Panel A. L-shape

Panel B. U-shape

Notes to figure: The figures show the estimated probabilities of states experiencing L-shaped recessions (Panel A) and U-shaped recessions (Panel B). The Y-axis represents the states, while the X-axis gives calendar time. Darker colors in each heat map indicate higher recession probabilities, as shown in the color legends.

Source: Authors' calculation

Figure 5: L-SHAPED AND U-SHAPED RECESSION PROBABILITIES: NEW YORK AND WISCONSIN



Notes to figure: The figures display the estimated probabilities of L-shaped and U-shaped recessions for New York state (Panel A) and Wisconsin (Panel B). The blue lines represent the probability of a U-shaped recession, while the red lines represent the probability of an L-shaped recession. The Y-axis indicates the probability, and the X-axis represents calendar time in quarter. The shaded areas denote the NBER recessions.
Source: Authors' calculation

others, as shown by isolated dark dots in both heatmaps (Figure 4). For example, North Dakota experienced an L-shaped recession in 2015 due to a sharp decline in oil prices, while Louisiana faced a U-shaped recession in 2005 due to Hurricane Katrina. Overall, these patterns highlight substantial heterogeneity in the incidence and nature of recessions across states, indicating that regional business cycles do not always coincide with the national ones.

Second, U-shaped recessions became less frequent after the 1990s, a pattern that is widespread across states (Panel B of Figure 4). This pattern coincides with the increased prevalence of jobless recoveries in the labor market. Nevertheless, it is important to note that our model identifies the COVID-19 recession as a U-shaped recession rather than an L-shaped one for most states, effectively capturing the sharp, short-lived, and rapidly evolving nature of the pandemic-induced

tourism, retail, and commercial real estate. Meanwhile, Wisconsin experienced a milder U-shaped recession after the pandemic recession.

downturn.

In sum, the state-level recession probabilities highlight significant heterogeneity in business-cycle experiences and hysteresis across states. Building on this variation, we explore the drivers of hysteresis using observable state-level characteristics in the subsequent sections

4 Features of U- and L- Shaped Recessions

This section discusses the characteristics of the two types of recessions. Section 4.1 explores the roles of supply and demand factors in shaping U- and L-shaped recessions, while Section 4.2 analyzes their effects on labor market outcomes and broader macroeconomic variables.

4.1 Supply and Demand Shocks

To examine the extent to which supply and demand factors drive each type of recession, we consider the following model:

$$p_{it}^{j} = \beta_s^j I_{it}^s + \beta_d^j I_{it}^d + \gamma^j g_{it} + \alpha_i^j + \epsilon_{it}^j \quad \text{for} \quad j \in \{l, u\}, \tag{4.1}$$

where p_{it}^{l} and p_{it}^{u} are the probabilities that state i experiences L-shaped and U-shaped recessions in quarter t, respectively, obtained from the estimates described in Section 3. The supply factor indicator, I_{it}^s, equals one if both the unemployment rate and price inflation in state i increase in quarter t, and zero otherwise. The demand factor indicator, I_{it}^d, equals one if the unemployment rate increases while price inflation decreases in state i during quarter t, and zero otherwise. We use state-level total inflation, including prices of both tradables and nontradables, from Hazell et al. (2022). The parameters β_s^j and β_d^j denote the coefficients of I_{it}^s and I_{it}^d, respectively. We further examine the magnitude of the recessionary shock, g_{it}, to account for the possibility that L-shaped recessions are more likely to result from larger shocks. Specifically, we measure the magnitude of the shock using the absolute value of changes in the unemployment rate, interacted with the probability of being in a recession (i.e., the combined probability of U-shaped and L-shaped regimes). The parameter γ^j is the coefficient of g_{it}. The sample period spans 1978:Q1–2017:Q4, consistent with the availability of state-level CPI data from Hazell et al. (2022).

Table 1 reports the estimated effects of supply and demand factors on recession shapes. Overall, both supply and demand factors contribute significantly to the likelihood of L-shaped

Table 1: EFFECTS OF SUPPLY AND DEMAND FACTORS

	(1) p^l_{it}	(2) p^u_{it}	(3) p^l_{it}	(4) p^u_{it}
[1] β^j_s (I^s_{it})	0.042***	0.010***	0.020***	0.002**
	(0.002)	(0.001)	(0.002)	(0.001)
[2] β^j_d (I^d_{it})	0.088***	0.014***	0.024***	0.003***
	(0.004)	(0.002)	(0.002)	(0.001)
[3] γ^j (g_{it})			0.739***	0.127***
			(0.006)	(0.003)
State FE	✓	✓	✓	✓
No. of obs.	12,189	12,189	8,925	8,925
R^2	0.064	0.029	0.701	0.252

Notes to table: This table presents the coefficient estimates from Equation (4.1). The variables in parentheses denote the regressors corresponding to each coefficient. State FE denotes state fixed effects. The notation ***, **, and * indicates statistical significance at the 1%, 5%, and 10% levels, respectively. Numbers in parentheses are standard errors.
Source: Authors' calculation.

as well as U-shaped recessions. In addition, L-shaped recessions tend to be associated with larger shocks: the coefficient on g_{it} for the L-shaped recession probability is substantially larger than that for the U-shaped recession probability (columns 3–4). It is also noteworthy that the responsiveness of the L-shaped recession probability to the demand factor (β^l_d) is the largest among the four coefficients ($\beta^l_d, \beta^l_s, \beta^u_d, \beta^u_s$), regardless of whether g_{it} is included or not. With g_{it}, β^l_d is marginally larger than β^l_s, the supply factor's effect on the L-shaped recession probability; without g_{it}, β^l_d is roughly twice as large as β^l_s. This pattern is consistent with the hysteresis theory of Blanchard (2018), which posits that negative demand shocks can have persistent adverse effects on economic activity.

4.2 Macroeconomic Outcomes of U- and L-shaped Recessions

This section examines the post-recession macroeconomic outcomes of U-shaped and L-shaped recessions, over the four quarters (one year) following each recession episode.[20]

Consider the following model:

$$y_{i,t+4} - y_{it} = \beta_u p^u_{it} + \beta_l p^l_{it} + \alpha_i + e_{i,t+4}, \tag{4.2}$$

where $y_{i,t+4} - y_{i,t}$ captures the change in state i's macro variable of interest between quarters t

[20] We consider the one-year horizon, because the bounce-back phase of a U-shaped recession is set to be 5 quarters.

Table 2: POST-RECESSION DYNAMICS OF MACROECONOMIC VARIABLES

y_{it}	(1) Log Real GSP	(2) EPOP Ratio	(3) Total Price Inflation	(4) Tradables Inflation	(5) Nontradables Inflation
β_l (p^u_{it})	-0.046***	-1.627***	1.321***	1.037***	1.734***
	(0.003)	(0.098)	(0.319)	(0.368)	(0.386)
β_u (p^l_{it})	0.112***	3.009**	-3.696***	-5.093***	-2.391
	(0.008)	(0.341)	(1.253)	(1.448)	(1.517)
State FE	✓	✓	✓	✓	✓
No. of Obs.	11,220	8,772	4,479	4,479	4,479
R^2	0.057	0.036	0.043	0.022	0.047

Notes to table: This table reports the coefficient estimates from Equation (4.2). The dependent variables are the annual change in log output (yearly data) and the four-quarter change in the EPOP ratio and four-quarter changes in the inflation rates of total, tradable, and nontradable goods (quarterly data) for columns (1)–(5), respectively. The notations ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. Numbers in parentheses are standard errors.
Source: Authors' calculations.

and $t+4$; β_u and β_l are the coefficients on p^u_{it} and p^l_{it}, respectively; α_i is a state fixed effect; and $e_{i,t+4}$ is the prediction error.

Our analysis focuses on the post-recession dynamics of log real gross state product (GSP), EPOP ratio, and inflation rates of total, tradable, and nontradable goods. The real GSP is measured annually, so the dependent variable is the annual output growth (yearly data). The other variables are available quarterly, and hence we consider four-quarter changes in the EPOP ratio and four-quarter changes in the inflation rates of total, tradable, and nontradable goods. The state-level inflation, measured in four-quarter percent changes, are taken from Hazell et al. (2022).[21]

Table 2 summarizes the estimated macroeconomic effects of L-shaped and U-shaped recessions. It is notable that the two types of recessions exhibit opposite effects on the output growth: an L-shaped recession lowers output one year ahead, whereas a U-shaped recession raises it (column 1). A similar pattern emerges for the EPOP ratio (column 2), consistent with the well-documented features of macroeconomic hysteresis (e.g., Furlanetto et al., 2025).[22] In addition, the two types of recessions also display contrasting effects on price inflation. While L-shaped recessions increase total price inflation one year ahead, U-shaped recessions lower it (column 3). This pattern is primarily driven by tradables (column 4) and, to a lesser extent, by

[21] In this exercise, we consider inflation rates rather than price levels because prior studies on the Phillips curve (e.g., Hazell et al., 2022) relate the inflation rate to the level of the unemployment rate (the level of economic activity).
[22] Comparable results are found for the labor force participation rate (LFPR), and both EPOP ratio and LFPR by gender (see Appendix D).

nontradables (column 5).

These findings suggest that L-shaped recessions are likely to lower potential output or raise structural unemployment and, *ceteris paribus*, widen the output gap or narrow the unemployment rate gap, thereby putting upward pressure on inflation. In contrast, U-shaped recessions—being less likely to induce structural changes—tend to reduce the output gap or widen the unemployment rate gap, consequently exerting downward pressure on inflation.

5 Downward Nominal Wage Rigidity and Hysteresis

This section examines the role of downward nominal wage rigidity (DNWR) in hysteresis (L-shaped recession). Section 5.1 outlines the econometric methodology for this analysis; Section 5.2 presents and discusses the estimation results.

5.1 Linear Competing Risks Model

In this section, we examine the extent to which an observable state-level factor, such as DNWR, influence the likelihood of L-shaped relative to U-shaped recessions. In general, when the dependent variable is an unobserved probability across three or more categorical outcomes, a competing risks framework is used, typically implemented as a multinomial logit model. In such models, outcome probabilities are inferred from observed counts of outcomes. In our case, however, the probabilities of the three business cycle phases are directly observed. Accordingly, we estimate a linear competing risks model using OLS, where the dependent variable is defined as the difference in probabilities of the two types of recessions.[23]

Consider the following model for the relative risk of an L-shaped recession compared to a U-shaped recession:

$$\left(p^l_{it} - p^u_{it}\right) = \beta_e p^e_{it} + \beta_z Z^r_{it} + \mathbf{\Gamma}_x X_{it} + \alpha_i + D_t + \epsilon_{it}. \tag{5.1}$$

[23] While the multinomial logit model can be numerically challenging to estimate, particularly when it includes a large number of parameters, the linear competing risks model avoids such difficulties. For example, individual fixed effects can be readily incorporated into the linear model using standard panel regression techniques. Such a treatment is often infeasible in a multinomial logit framework, as numerical optimization frequently fails to converge in the presence of a high-dimensional parameter space. In this regard, the linear competing risks model enables the comprehensive inclusion of state-level covariates within a coherent statistical framework, allowing for the analysis of the determinants of hysteresis without imposing significant computational burdens.

In the above equation, we include p^e_{it}, the estimated probability of being in an expansion, to control for business cycle phases, as the dependent variable tends to approach zero both during expansions and during economic downturns when the probabilities of the two recession types can be similar.[24] The coefficient on p^e_{it} is denoted by β_e.

Our main focus is on DNWR. To evaluate the effect of DNWR on the relative risk, we construct an indicator of greater DNWR, denoted Z^r_{it}, as follows. We first construct Z_{it}, an indicator of greater nominal wage rigidity, which equals one if the change in the share of zero nominal wage inflation at time t exceeds the cross-state average at t, and zero otherwise. This state-level indicator captures larger nominal wage rigidity relative to the cross-state average. Since DNWR is primarily relevant during economic downturns, we interact Z_{it} with the state's recession probability $(1-p^e_{it})$ to produce Z^r_{it}, an indicator of greater *downward* nominal wage rigidity.[25]

The vector X_{it} contains control variables capturing state-specific attributes, and $\Gamma_{\mathbf{x}}$ denotes the vector of corresponding coefficients. Additional control variables are listed in the panel labeled "Controls" in Table 3, with details on each measure provided in Section 2.2.[26] The term α_i represents the state fixed effect, D_t the time fixed effect, and ϵ_{it} the residual.

The sample period spans 1978:Q1 – 2019:Q4, based on the availability of large-firm share data beginning in 1978:Q1. We exclude the pandemic period from the analysis, as the correlations between the regressors and recession probabilities during that time can substantially differ from pre-pandemic patterns.

5.2 Estimation Results

Table 3 summarizes the estimation results for the model's ability to distinguish between L-shaped and U-shaped recessions. To begin with, the model effectively captures the relative probability of an L-shaped recession compared to an U-shaped recession, with an R^2 around 0.8.

It is notable that greater DNWR (Z^r_{it}) significantly increases the likelihood of an L-shaped recession relative to a U-shaped one. This relationship is robust across both the full sample and the post-2000 subsample. We consider the post-2000 period to examine the role of DNWR

[24]One might consider constructing a variable by dividing the dependent variable by p^e_{it}. However, this measure is undefined when p^e_{it} is zero or near zero. To address this issue and account for business-cycle effects, we instead include p^e_{it} as a regressor in the model.

[25]We do not include the indicator of nominal wage rigidity during economic expansions, as an increased share of zero wage changes in expansions may also reflect upward nominal wage rigidity.

[26]We include the state-level gender gap in the labor market as it is considered as an important factor for the macroeconomic hysteresis (e.g., Fukui et al., 2023).

Table 3: Effects of DNWR on the relative risks of recessions

	(1) All	(2) 2000-2019
Indicator of DNWR	0.062***	0.052***
(Z^r_{it})	(0.009)	(0.009)
(Controls)		
p^e_{it}	✓	✓
gender gap	✓	✓
oil-producing	✓	✓
minimum wage	✓	✓
union	✓	✓
manufacturing	✓	✓
prof. services	✓	✓
finance	✓	✓
large-firm share	✓	✓
tax-income share	✓	✓
State fixed effect	✓	✓
Time fixed effect	✓	✓
No. of Obs.	8,256	3,856
R^2	0.796	0.867

Notes to table: This table presents the coefficient estimates from Equation (5.1). The notations ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. Numbers in parentheses are standard errors. The panel labeled '(Controls)' lists the control variables included in the regression model.

Source: Authors' calculation.

during the period of slowed gender convergence, given that gender convergence is significantly associated with the stagnant recession recoveries in the labor market (e.g., Fukui et al., 2023).[27] The finding supports the central premise of previous research that DNWR constitutes a key mechanism through which hysteresis effects emerge.

Taken together, the state-level evidence suggests that DNWR is an important source of hysteresis in economic activity.[28]

6 Effectiveness of Policy Interventions in Mitigating Hysteresis

This section evaluates the extent to which DNWR affects the effectiveness of monetary and fiscal policies in mitigating hysteresis. Section 6.1 provides an empirical model for this analysis and Section 6.2 discusses the estimation results.

6.1 Model

We examine the effects of monetary and tax shocks on the relative likelihood of an L-shaped versus a U-shaped recession, depending on the degree of DNWR. To do so, we estimate the following nonlinear local projection:

$$\begin{aligned} y_{i,t+h} - y_{i,t-1} &= \alpha_i^h + \beta_1^h (s_{i,t-1}) z_t + \beta_2^h (1 - s_{i,t-1}) z_t \\ &+ (\boldsymbol{\gamma}_1^h)'(s_{i,t-1}) \boldsymbol{x}_{it} + (\boldsymbol{\gamma}_2^h)'(1 - s_{i,t-1}) \boldsymbol{x}_{it} + \epsilon_{i,t+h} \qquad \text{for} \quad h = 1, 2, 3, \cdots H, \end{aligned} \tag{6.1}$$

where $y_{i,t} = (p^l_{it} - p^u_{it})$ and hence $(y_{i,t+h} - y_{i,t-1})$ captures changes between $t-1$ and $t+h$ in the relative probability of an L-shaped recession over a U-shaped recession. The notation α_i^h captures the state fixed effect at horizon h, $s_{i,t-1}$ captures the probability that state i is in regime 1 at time $t-1$, and $(1 - s_{i,t-1})$ corresponds to the probability of being in regime 2 at time $t-1$, which we will define below. The externally identified policy shock is denoted by z_t. We use monetary monetary policy shocks constructed by Romer and Romer (2004) and extended by

[27] We find that a larger male-female employment gap is positively associated with the risk of hysteresis with statistically significant. This result is reported in Appendix D.

[28] As a robustness check, we re-estimate the model excluding all controls except for state and time fixed effects, using the same sample period. Overall, the results remain consistent, with the exception of the gender gap coefficient, which becomes statistically insignificant for the 1978–2019 period. However, this coefficient regains statistical significance when we include the employment shares of the manufacturing and professional services industries, suggesting that gender differences are correlated with the industry composition of employment.

Wieland and Yang (2020), along with tax shocks from Romer and Romer (2010). Since the focus is on the effects of expansionary demand policy shocks, we retain only negative realizations of monetary and tax shocks. The coefficients β_1^h and β_2^h capture effects of an unexpected policy shock on the dependent variable h-quarter ahead after the shock's impact in regimes 1 and 2, respectively. Negative values of β_1^h and β_2^h indicate that a demand policy helps to mitigate hysteresis, while positive values indicate that it does not. We set $H = 12$ quarters.

We construct the regime indicator for DNWR (s_{it}) in a manner analogous to the indicator used in equation (5.1). Specifically, we interact Z_{it}—the indicator of greater nominal wage rigidity—with state i's binary recession indicator at t. The recession indicator equals one if $(1 - p_{i,t}^e) > 0.2$, and zero otherwise, for the monetary policy experiment. For the experiment with tax shocks, we adopt a threshold of $(1 - p_{i,t}^e) > 0.15$ for the recession indicator to equal one, and zero otherwise.[29] Accordingly, the DNWR indicator is binary: it equals 1 in regime 1 ($s_{it} = 1$), when state i exhibits nominal wage rigidity above the cross-state average and is in recession in quarter t, and 0 otherwise, denoting regime 2 ($s_{it} = 0$).[30]

For the vector of controls $\boldsymbol{x}_{it}$, we include eight quarterly lags of variables capturing state-level characteristics. Specifically, these controls comprise the employment shares in manufacturing, finance, and professional and business services; the tax-to-income ratio; the employment share of unionized workers; and the indicator of oil production. In addition, we include state i's probability of being in an expansion at time t to account for the scaling effect of the current business-cycle phase on the relative likelihood of recessions, in line with our treatment in equation (5.1). As our primary interest lies in the effects of monetary policy shocks on the relative likelihood of an L-shaped versus a U-shaped recession outside of expansions, we control for each state's concurrent probability of expansion.[31]

Note that we also interact $\boldsymbol{x}_{it}$, the vector of state-level controls at t, with the lagged regime

[29] The purpose of interacting the two indicators is to characterize DNWR, which is prevalent during a recession, while excluding upward nominal wage rigidity—a less likely occurrence during an economic downturn. Because the state-level wage data are annual, we use four-quarter moving averages of the expansion probability to construct the DNWR indicator for this experiment. Although these thresholds yield the largest differences between the two regimes, imposing the same threshold (0.85) to both experiments.

[30] For this empirical analysis, we consider the DNWR indicator to be binary for the clean identification of monetary policy pass-through.

[31] We further consider the magnitude of shocks measured with the negative changes in the EPOP ratio to control for effects of large shocks in determining the interaction between policy effectiveness and downward nominal wage rigidity (as considered by Jo and Zubairy, 2025), for robustness analyses. We consider both the contemporaneous value alone and the combination of contemporaneous and lagged values of these shock measures. The estimated impulse responses remain robust to these specifications.

indicators $s_{i,t-1}$ and $(1-s_{i,t-1})$ to comprehensively account for the state-dependent underlying dynamics of each state's economy, in line with Auerbach and Gorodnichenko (2012). The coefficients $\boldsymbol{\gamma}_1^h$ and $\boldsymbol{\gamma}_2^h$, are the vector of coefficients on the controls in regimes 1 and 2, respectively.

The sample period for this experiment spans from 1980:Q1 to 2007:Q4. The start date reflects the availability of state-level large-firm share data beginning in 1978:Q1, combined with the use of eight lags of the control variables. The end date corresponds to the availability of monetary policy shocks from Wieland and Yang (2020), which extends through 2007. For consistency and comparability, we use the same sample period in the analysis of tax shocks.

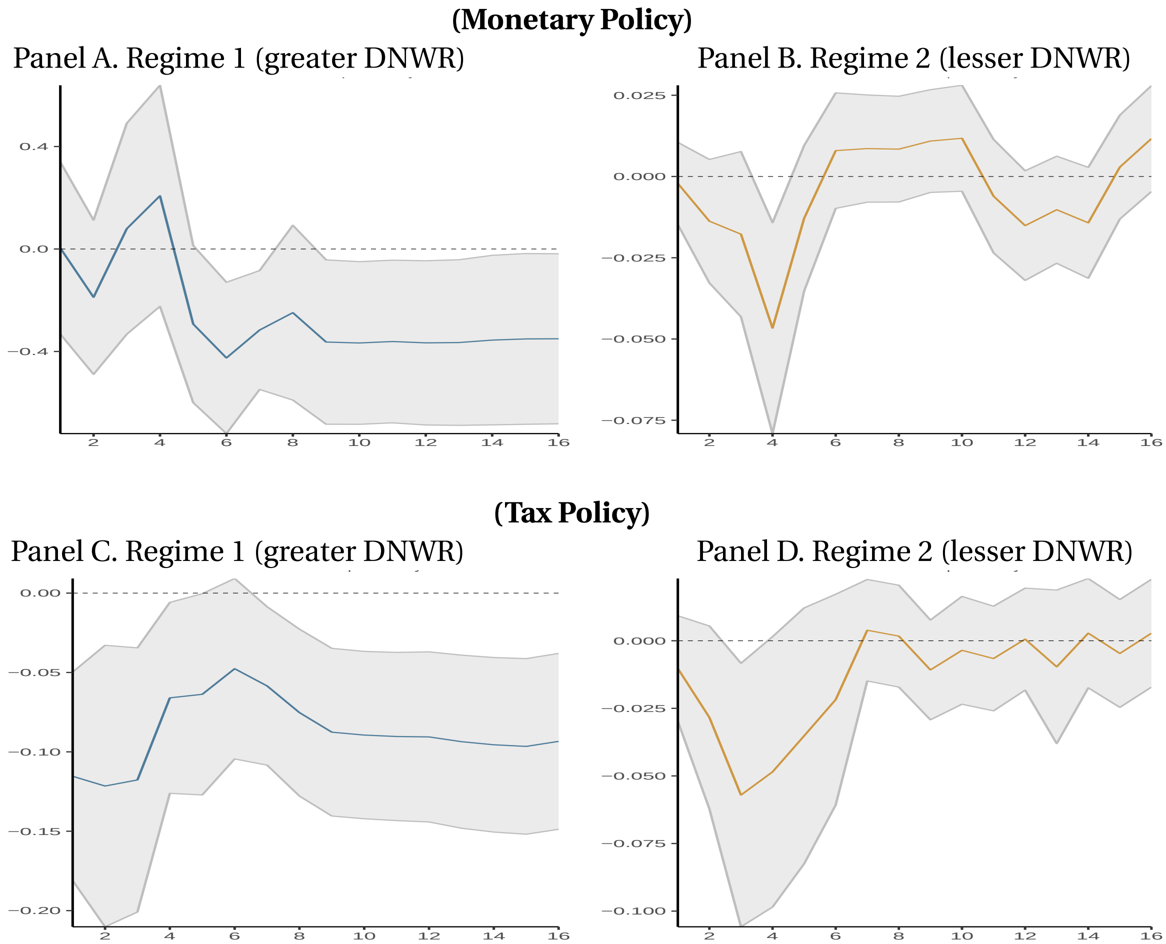
6.2 Estimation Results

This section discusses the estimation results. Upper panels in Figure 6 show the response of the relative probability of an L-shaped recession compared with a U-shaped recession following a one-unit expansionary monetary policy shock: Panel A presents estimates under greater DNWR (regime 1), and Panel B under lower DNWR (regime 2). Negative values indicate that such a policy shock significantly reduces the likelihood of an L-shaped recession relative to a U-shaped one, thereby mitigating hysteresis, while positive values indicate the opposite. In the regime of greater DNWR, an expansionary monetary policy reduces the relative probability of an L-shaped recession with statistical significance approximately four quarters after the shock, with effects that persist over time. In contrast, these hysteresis-mitigating effects are weaker and shorter-lived in the regime of lesser DNWR. This finding underscores the long-run effectiveness of monetary policy in mitigating hysteresis and highlights that its impact is stronger when nominal wages are more downwardly rigid.

Next, we examine the effects of expansionary tax shocks in mitigating hysteresis (lower panels of Figure 6). Tax cuts exhibit substantial hysteresis-reducing effects, when nominal wages are more downwardly rigid (Panel C), as reflected by their statistically significant negative effects on the relative probability of an L-shaped recession over a U-shaped recession. In contrast, the expansionary tax shocks are not statistically significant and shorter-lived under lesser DNWR (Panel D). These findings indicate that tax cuts are more effective at mitigating hysteresis in environments with greater DNWR.

In summary, both expansionary monetary policy and tax shocks are more effective at alleviating hysteresis when DNWR is greater. This finding is consistent with the conventional view that real effects of monetary policy stem from nominal rigidities. In addition, tax cuts can also

Figure 6: EFFECTS OF EXPANSIONARY DEMAND POLICIES ON THE RELATIVE RISK OF L-SHAPED RECESSION OVER U-SHAPED RECESSION



Notes to figure: This figure displays the estimated responses of relative risk of an L-shaped recession over a U-shaped recession to a one-unit decrease in monetary policy shock (Panels A and B) and in tax shock (Panels C and D). Panels A and C show the responses in the regime of greater DNWR, and Panels B and D show those in the regime of lesser DNWR. Monetary policy shock estimates are from Romer and Romer (2004) extended by Wieland and Yang (2020). Tax shocks are from Romer and Romer (2010). The X-axis shows quarters after the shock, and the Y-axis indicates the policy shock's pass-through to relative risk. The shaded area captures 95 percent confidence intervals based on Newey-West standard errors.

Source: Authors' calculation

mitigate hysteresis by reducing production costs when labor costs are inflexible due to DNWR or help maintain labor demand.[32]

[32]This result aligns with Lee (2025), who finds that macroeconomic variables respond nearly twice as strongly and more persistently in states with higher nominal wage rigidity compared to those with more flexible wages. Our analysis differs from Lee's in that we focus specifically on the response of hysteresis, whereas Lee examines the effects on general macroeconomic variables generally following a marginal tax shock.

7 A New Keynesian Model of Hysteresis with DNWR

In this section, we use a prototype New Keynesian model to validate our empirical findings. Building on Galí (2022), who develops an insider–outsider labor market framework that generates hysteresis in the form of L-shaped recessions, we extend the model by incorporating DNWR.

In this model, wage bargaining is conducted by unions that represent only last period's employed workers (the insiders). Because these unions set wages to protect insiders' jobs while disregarding the unemployed (the outsiders), their employment targets become anchored to past employment levels, endogenously generating persistence in both unemployment and output. As a result, employment losses caused by temporary shocks become persistent, giving rise to hysteresis. Downward nominal wage rigidity further amplifies these hysteresis effects by hindering wage adjustment in the labor market.

Our quantitative analysis is motivated by the empirical results in Sections 4–6. In this analysis, we focus solely on the interaction between monetary policy shocks and DNWR. Fiscal policy shocks can independently generate permanent distortions and long-run effects on the economy even in the absence of nominal wage rigidity (Burnside et al., 2004; Fatás and Summers, 2018). In contrast, monetary policy shocks are typically transitory, but our empirical results show that their effects become more effective at mitigating hysteresis when they interact with DNWR. To ensure a clean identification of this mechanism, we therefore focus on monetary policy.

We model hysteresis as driven by the contractionary demand shock documented in Section 4, calibrating the size of the shock to match the magnitude of the employment hysteresis observed at the national level in Section 3. The quantitative experiment aims to: (1) show that DNWR amplifies hysteresis, producing larger negative effects on employment and output and smaller declines in inflation; (2) evaluate the effectiveness of monetary policy in mitigating these effects; and (3) assess how this effectiveness strengthens with greater DNWR. These mechanisms, documented in Sections 5 and 6, highlight the amplifying role of DNWR and the conditional effectiveness of monetary policy interventions.

7.1 Model

7.1.1 Households

We consider a large number of identical households. Each household contains a continuum of members, uniformly distributed over the unit square. Each member is identified by a pair $(j, s) \in [0,1] \times [0,1]$. The first coordinate, j, denotes the type of labor service (or "occupation") in which the member is specialized, while the second coordinate, s, captures her disutility from work. This disutility is given by $\frac{\Theta_t}{\overline{C}_t}\chi s^{\varphi}$ if she is employed and zero otherwise with preference parameters $\chi > 0$ and $\varphi > 0$. Each employed worker supplies a fixed number of hours, normalized to one unit of time. For each occupation j, employment $N_t(j) \in [0,1]$ is determined by labor demand and is taken as given by the household. Employment slots are assigned to the members with the lowest disutility from work, that is, those with $s \in [0, N_t(j)]$. We assume separable preferences with logarithmic utility in consumption and complete risk sharing within the household (i.e., $C_t(i, j) = C_t$ for all $(j, s) \in [0,1] \times [0,1]$). As a result, all members consume the same amount, regardless of their occupation or employment status.

This preference specification implies that each household maximizes its lifetime utility:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left(\log C_t - \frac{\Theta_t}{\overline{C}_t} \chi \int_0^1 \frac{N_t(j)^{1+\varphi}}{1+\varphi} dj \right) Z_t,$$

subject to the budget constraint

$$\int_0^1 P_t(i) C_t(i) di + Q_t B_t \le B_{t-1} + \int_0^1 W_t(j) N_t(j) dj + D_t,$$

where $C_t = \left(\int_0^1 C_t(i)^{1-1/\epsilon_p} di \right)^{\epsilon_p/(\epsilon_p - 1)}$ is the Dixit–Stiglitz consumption index, ϵ_p is the elasticity of substitution across differentiated goods, $\beta \in (0,1)$ is the household's discount factor, $P_t(i)$ denotes the price of variety i, $W_t(j)$ the nominal wage for occupation j, B_t the household's holdings of a nominal riskless one-period discount bond, Q_t its price, and D_t dividends from firm ownership.

Following Galí et al. (2012), the preference shifter Θ_t is defined as

$$\Theta_t = \Theta_{t-1}^{1-\nu} \overline{C}_t^{\nu},$$

where $\overline{C}_t$ denotes aggregate consumption, treated as exogenous from the household's perspective. We interpret Θ_t as a smoothed consumption measure that affects labor supply: when actual consumption exceeds Θ_t, the marginal disutility of labor falls. Consequently, the ratio $\frac{\Theta_t}{C_t}$ determines the strength of the household's wealth effect. A lower ν implies a weaker short-run wealth effect, allowing the model to better capture the observed comovement of labor and real wages. For separable utility, this modification follows Galí et al. (2012), building on Jaimovich and Rebelo (2009). Further implications are discussed in Section 7.2 in the context of the labor supply equation.

When lowercase letters denote logs, the demand shock $z_t \equiv \log Z_t$ follows an AR(1) process:

$$z_t = \rho_z z_{t-1} + \varepsilon_t^z, \tag{7.1}$$

where $\rho_z \in [0,1)$ and ε_t^z is white noise.

The household's optimization problem yields the intertemporal Euler equation, which in logs is written as:

$$c_t = E_t\{c_{t+1}\} - (i_t - E_t\{i_{t+1}\} - \rho) + (1-\rho_z)z_t. \tag{7.2}$$

Let $L_t(j)$ denote the marginal participant in occupation j. The labor supply optimality condition is

$$\frac{1}{C_t}\frac{W_t(j)}{P_t} = \frac{\Theta_t}{C_t}\chi L_t(j)^{\varphi}.$$

Taking logs and aggregating across occupations gives the average real wage and preference shifter equations:

$$\omega_t = \theta_t + \varphi l_t + \log\chi, \tag{7.3}$$

$$\theta_t = (1-\nu)\theta_{t-1} + \nu c_t, \tag{7.4}$$

where $\omega_t = w_t - p_t$ is the average log real wage, $w_t = \int_0^1 w_t(j)dj$ is the average log nominal wage, and $l_t = \int_0^1 l_t(j)dj$ is the log labor force.

Then, unemployment is defined as

$$u_t = l_t - n_t, \tag{7.5}$$

where $n_t = \int_0^1 n_t(j)dj$ is the log of aggregate employment, determined by firms' labor demand.

7.1.2 Firms

A monopolistically competitive intermediate goods-producing firm $i \in [0,1]$ produces output according to

$$Y_t(i) = N_t(i)^{1-\alpha},$$

where $0 < \alpha < 1$.

Prices are set following Calvo stickiness: a fraction θ_p of firms cannot adjust their prices each period. At time t, a firm that can reoptimize chooses the reset price $P_t^*(i)$ to maximize expected discounted profits:

$$E_t \sum_{k=0}^{\infty} (\beta\theta_p)^k \Lambda_{t,t+k} \left[\left(P_t^*(i) - MC_{t+k} \right) Y_{t+k|t}(i) \right],$$

subject to the demand function

$$Y_{t+k|t}(i) = \left(\frac{P_t^*(i)}{P_{t+k}} \right)^{-\epsilon_p} Y_{t+k},$$

where $\Lambda_{t,t+k}$ is the stochastic discount factor, ϵ_p is the elasticity of substitution between intermediate goods, and P_{t+k} is the aggregate price level at time $t+k$.

Firms' profit-maximization problem yields the New Keynesian Phillips curve in logs:

$$\pi_t^p = \beta E_t \pi_{t+1}^p + \chi_p \tilde{y}_t + \lambda_p \tilde{\omega}_t, \tag{7.6}$$

where the output gap $\tilde{y}_t = y_t - y_t^n$ and the real wage gap $\tilde{\omega}_t = \omega_t - \omega_t^n$ are log deviations from their natural levels under flexible prices and wages, which are defined in the wage-setting section. The parameters are defined as $\chi_p = \frac{\alpha\lambda_p}{1-\alpha}$ and $\lambda_p = \frac{(1-\theta_p)(1-\beta\theta_p)}{\theta_p} \frac{1-\alpha}{1-\alpha+\alpha\epsilon_p}$.

Finally, in equilibrium, the log of employment satisfies

$$(1-\alpha) n_t = y_t, \tag{7.7}$$

and goods market clearing requires

$$y_t = c_t. \tag{7.8}$$

7.1.3 Wage Setting

Whereas in a standard New Keynesian model unions set nominal wages to be equal to a weighted average of expected markups, the insider–outsider model assumes that they instead set wages to ensure that the weighted average of expected employment equals the measure of insiders. A constant fraction $1-\theta_w$ of occupations (or their unions) are allowed to reset their wage in any period. When setting $W_t^*(j)$, a union representing occupation j considers the labor demand for its members: for $k = 1, 2, 3, \ldots$,

$$N_{t+k|t}(j) = \left(\frac{W_t^*(j)}{W_{t+k}}\right)^{-\epsilon_w} N_{t+k}.$$

The evolution of the average log nominal wage is

$$w_t = \theta_w w_{t-1} + (1-\theta_w) w_t^*,$$

where w_t^* denotes the log of the average newly set wage in period t.

Following Galí (2022), unions set wage optimally so that a weighted average of expected (log) employment equals the target insider employment level:

$$(1-\beta\theta_w)\sum_{k=0}^{\infty}(\beta\theta_w)^k E_t\{n_{t+k|t}(j)\} = n_t^*(j)$$

where the target insider stock evolves as

$$n_t^*(j) = \gamma n_{t-1}(j) + (1-\gamma) n^*.$$

Here, n^* denotes the union's long-run employment target, common across occupations, and $\gamma \in [0,1]$ captures the persistence of the target insiders, that is, the extent to which current employment shapes future employment targets and generate hysteresis. The target $n_t^*(j)$ represents the measure of insiders in occupation j with higher γ implying stronger hysteresis effects. Under fully flexible wages, the equilibrium condition reduces to $n_t^n = n_t^*$ for all t. If initial employment equals its target, $n_0 = n^*$, then $n_t^n = n^*$. Consequently, the natural level of employment equals the target n^*, and the corresponding natural level of output and wage are

$$y_t^n = (1-\alpha) n^*, \tag{7.9}$$

$$\omega_t^n = log(1-\alpha) - \alpha n^* - \mu_p - log(1-\tau), \tag{7.10}$$

where $\mu_p = log\frac{\epsilon_p}{\epsilon_p - 1}$ is the natural wage markup, and τ is a constant wage subsidy.

With stickiness, wage setting rule is given by

$$w_t^*(j) = -\frac{1}{\epsilon_w} n_t^*(j) + (1-\beta\theta_w)\sum_{k=0}^{\infty}(\beta\theta_w)^k E_t\left[w_{t+k} + \frac{1}{\epsilon_w} n_{t+k}\right].$$

Averaging over unions resetting in t gives

$$\begin{aligned} w_t^* &= -\frac{1}{\epsilon_w} n_t^* + (1-\beta\theta_w)\sum_{k=0}^{\infty}(\beta\theta_w)^k E_t\left[w_{t+k} + \frac{1}{\epsilon_w} n_{t+k}\right] \\ n_t^* &= \gamma n_{t-1} + (1-\gamma) n^*. \end{aligned}$$

Combining the equations above with wage evolution equation, we obtain the wage Phillips curve:

$$\pi_t^w = \beta E_t\{\pi_{t+1}^w\} + (1-\gamma)\lambda_n(1-\beta\theta_w)\hat{n}_t + \gamma\lambda_n \Delta n_t \tag{7.11}$$

where $\hat{n}_t \equiv n_t - n^*$ denotes the deviation of employment from the union's long-run target, and $\lambda_n \equiv \frac{1-\theta_w}{\theta_w \epsilon_w}$ measures the sensitivity of wage inflation to employment deviation and its gaps.

The wage Phillips curve above implies that when γ is high (i.e., hysteresis is strong), wage inflation responds less to to employment gaps, so persistent deviations from the target are not self-correcting. Wage inflation, price inflation, and the real wage satisfy the identity

$$\omega_t = \omega_{t-1} + \pi_t^w - \pi_t^p. \tag{7.12}$$

In addition, following Schmitt-Grohé and Uribe (2022), we incorporate DNWR by imposing the constraint, for all firms $j \in [0,1]$,

$$W_t(j) \geq W_{t-1}(j),$$

which is equivalent to

$$\pi_t^w \geq 0. \tag{7.13}$$

Table 4: CALIBRATED PARAMETERS

Parameter	Description	Value
φ	Curvature of labor disutility	3.4
β	Discount factor	0.99
α	Decreasing returns to labor	0.25
ϵ_p	Elasticity of substitution (goods)	3.8
ϵ_w	Elasticity of substitution (labor)	4.3
θ_p	Price stickiness	0.75
θ_w	Wage stickiness	0.10
ϕ_i	Interest rate rule lag	0.9
ϕ_π	Interest response to inflation	1.5
ϕ_y	Interest response to output growth	0.1
ρ_z	Demand shock AR(1)	0.90
γ	Persistence of the target insiders	0.99
ν	Wealth effect parameter	0.02

7.1.4 Monetary Policy

Monetary policy is modeled using a Taylor-type interest rate rule, in which the central bank adjusts the nominal interest rate in response to deviations of inflation and output growth from their respective targets. The rule includes interest rate smoothing to reflect the gradual adjustment behavior commonly observed in practice:

$$i_t = \phi_i i_{t-1} + (1-\phi_i)\left[\rho + \phi_\pi \pi_t^p + \phi_y \Delta y_t\right] + \varepsilon_t^m \tag{7.14}$$

where i_t is the nominal interest rate, ρ is the steady-state real interest rate, π_t^p is the inflation rate of prices, and Δy_t denotes output growth. The parameters ϕ_π and ϕ_y govern the responsiveness of policy to inflation and output growth, respectively, while $\phi_i \in [0,1]$ captures the degree of interest rate smoothing. The term ε_t^m represents an exogenous monetary policy shock.

7.2 Calibration

Simulations use the log-linearized structural equations (7.2)–(7.12) and (7.14), incorporating the demand shock process (7.1) and monetary policy shock (7.14) under downward nominal wage rigidity constraint (7.13). We solve the model under this constraint using the OccBin methodology of Guerrieri and Iacoviello (2015).

Table 4 presents the calibrated parameter values used in the simulations. The model is calibrated at a quarterly frequency. Most parameter values are taken from Galí (2022), except for the wage stickiness parameter θ_w and the preference shifter smoothness parameter ν. We briefly discuss the implications of these parameter choices, along with other parameters, below. For further details on the calibration, see Galí (2022).

In our framework, we incorporate downward nominal wage rigidity in addition to standard nominal wage rigidity. Specifically, when a shock puts downward pressure on nominal wages, all workers receive unchanged wages. With upward pressure, only a fraction θ_w receive unchanged wages. We set $\theta_w = 0.10$ so that this lower wage rigidity parameter aligns with the model featuring only downward nominal wage rigidity, as in Daly and Hobijn (2014) and Jo and Zubairy (2025). The price rigidity parameter is set to $\theta_p = 0.75$, so the average duration of prices is four quarters.

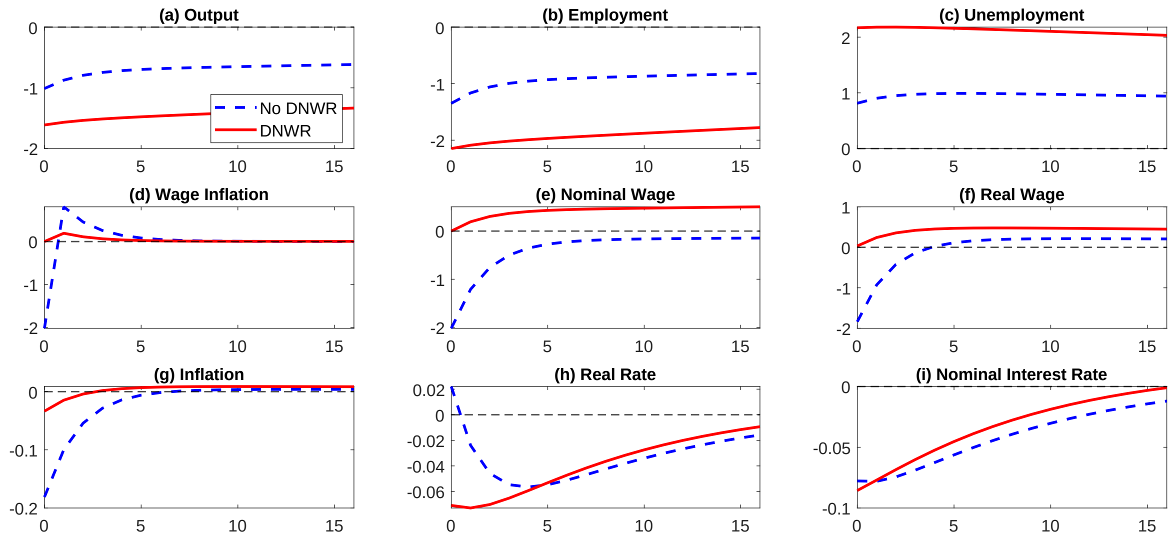
We also consider a small wealth effect in labor supply to match the empirical evidence on the effects of monetary policy shocks reported by Christiano et al. (2021). Their VAR results show that labor supply increases in response to expansionary monetary policy shocks. However, the conventional King-Plosser-Rebelo (1988) (KPR, henceforth) preferences, which imply a strong wealth effect through $w_t - p_t - c_t = \varphi l_t$, could generate a decline in labor supply following an expansionary monetary policy shock, which is counterfactual. To address this, following Galí et al. (2012), we set $\nu = 0.02$, a value close to Greenwood-Hercowitz-Huffman (1988) (GHH, henceforth) preferences (see Jo and Zubairy 2025).[33] A smaller wealth effect (i.e., lower ν) strengthens the procyclical response of labor supply to monetary policy shocks through the labor supply condition

$$w_t - p_t - \theta_t = \varphi l_t, \quad \theta_t = (1-\nu)\theta_{t-1} + \nu c_t.$$

The elasticities of substitution for goods and labor are set to $\epsilon_p = 3.8$ and $\epsilon_w = 4.3$, corresponding to steady-state markups of 35% and 30%, respectively. The curvature parameter of labor disutility is set to 3.4, which implies a steady-state unemployment rate of approximately 7.8%. The discount factor is fixed at 0.99, following standard practice in the literature. The labor income share is calibrated to $\alpha = 0.25$. The monetary policy rule parameters are set to $\phi_\pi = 1.5$, $\phi_y = 0.5$, and $\phi_i = 0.9$. The persistence of the demand shock is given by $\rho_z = 0.9$. The sizes of the demand and monetary policy shocks will be discussed in the exercise based on impulse response functions.

[33] The parameter value $\nu = 1$ corresponds to KPR preferences.

Figure 7: THE ROLE OF DNWR IN AMPLIFYING THE HYSTERESIS EFFECT



Notes to figure: This figure shows the impulse response functions to a contractionary demand shock under two scenarios: with the DNWR constraint (red solid line) and without it (blue dashed line).
Source: Authors' calculation

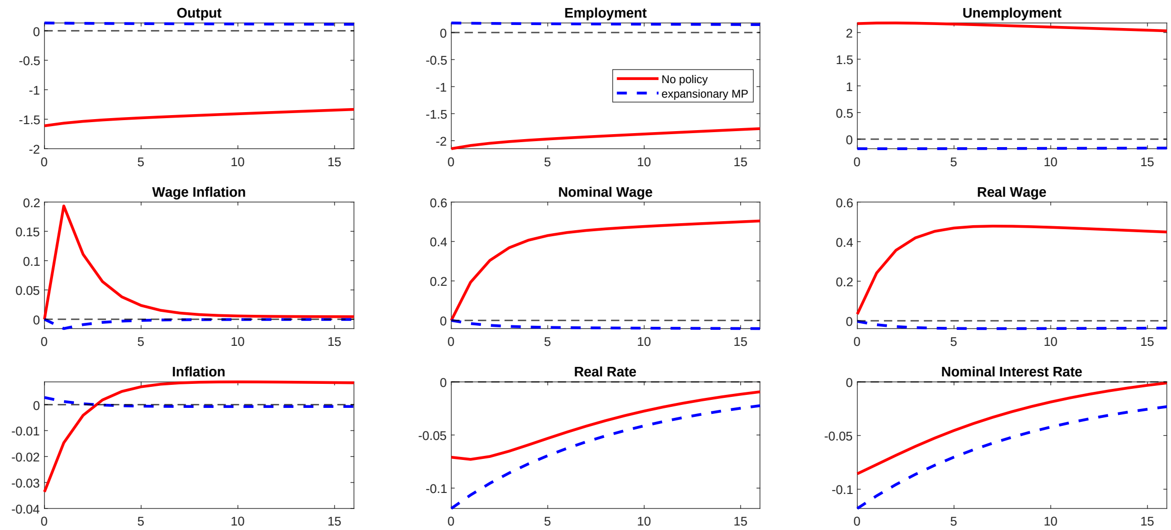
7.3 Quantitative Analysis

We examine how DNWR amplifies hysteresis in output, employment, and unemployment during recessions triggered by contractionary demand shocks, confirming the empirical patterns documented earlier.

Figure 7 presents the effects of DNWR by comparing the impulse responses to a contractionary demand shock under scenarios with and without the rigidity constraint. The size of the demand shock is calibrated to match the estimated national-level employment growth rate for an L-shaped recession (i.e., a recession accompanying hysteresis), μ_L. Employment on impact declines by 1.20 percent without DNWR and by 1.92 percent with DNWR, yielding an average of $\hat{\mu}_L = 1.56$. See the estimated value of μ_L reported in Table C1 of Section Appendix C. The impulse responses of employment and output show that hysteresis effects are highly persistent and that DNWR amplifies the impact of the shock, leading to declines in employment and output that are approximately 60 percent larger, consistent with stronger hysteresis effects.

Unemployment rises under DNWR because higher real wages increase labor supply, but also suppress labor demand. In the New Keynesian Phillips curve, inflation is driven by output and real wage gaps. The upward pressure on real wages outweighs the recessionary effect on output,

Figure 8: HYSTERESIS AND EXPANSIONARY MONETARY POLICY UNDER DNWR



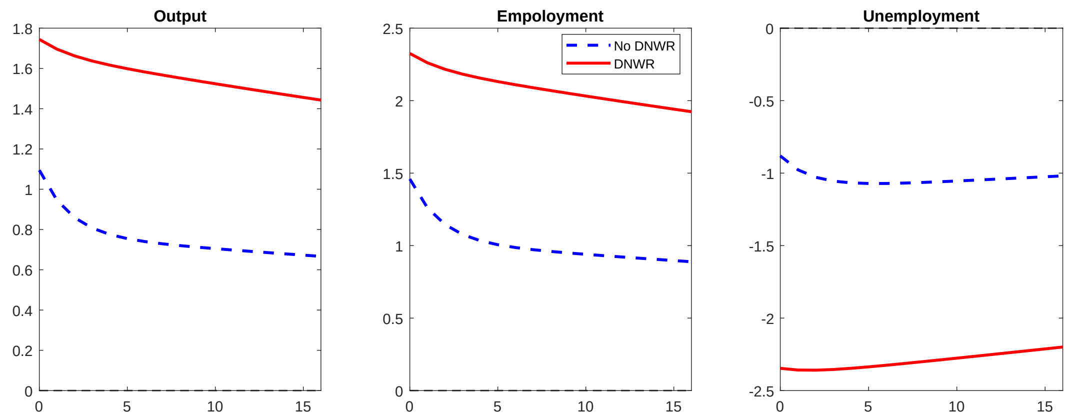
Notes to figure: The figure displays the impulse response functions to a contractionary demand shock in the presence of DNWR, comparing the case with expansionary monetary accommodation (blue dashed line) to the case without it (red solid line).
Source: Authors' calculation

producing a smaller decline in inflation. Consequently, the larger output contraction and smaller inflation decline largely offset each other in the monetary policy rule, so the nominal interest rate behaves similarly with or without DNWR. This leads to a substantially larger decline in the real interest rate under this rigidity.

We now conduct a counterfactual analysis of an expansionary monetary policy implemented alongside hysteresis driven by a contractionary demand shock under DNWR. This expansionary policy is modeled as a 50 basis points per annum ($\varepsilon^m = -0.125\%$ at a quarterly rate), occurring simultaneously with the onset of the contractionary demand shock. Figure 8 shows that this intervention substantially mitigates the shock's adverse effects. In particular, it nearly offsets the rise in real wages induced by the price decline, preventing sharp contractions in employment and output that would otherwise generate persistent hysteresis. Consequently, hysteresis in employment and output is largely eliminated, underscoring the importance of timely and forceful monetary accommodation when DNWR limits labor market adjustment.

Figure 9 plots the difference in impulse responses to a contractionary demand shock with and without an accompanying expansionary monetary policy, comparing outcomes under DNWR and without the constraint. The shocks are calibrated as in Figures 7 and 8. When

Figure 9: The Amplification of Expansionary Monetary Policy Effects by DNWR



Notes to figure: The figure plots the effects of an expansionary monetary policy implemented during hysteresis, obtained by taking the difference between the impulse responses to a contractionary demand shock with and without the monetary policy shock, under two scenarios: with the DNWR constraint (red solid line) and without it (blue dashed line).
Source: Authors' calculation

DNWR binds, nominal wages cannot fall, causing real wages to rise, depressing labor demand, and amplifying the recession. Expansionary monetary policy mitigates this effect by raising the price level, lowering real wages, and restoring labor market equilibrium. Consequently, as shown in Figure 9, the same monetary expansion under DNWR substantially cushions the downturn in employment and output and lowers unemployment relative to the case without it. Overall, DNWR makes monetary policy roughly 1.6 times more effective in stabilizing output, employment, and unemployment.

8 Conclusion

This paper examines the role of DNWR in macroeconomic hysteresis, exploiting state-level heterogeneity in recession experiences. To this end, we first estimate a Bayesian Markov-switching model that distinguishes between U-shaped recessions—characterized by full recovery—and L-shaped recessions, which exhibit hysteresis. Our empirical findings can be summarized as follows. First, state-level business-cycle experiences exhibit substantial heterogeneity, and this rich variation allows us to identify the key contributors to hysteresis. Second, DNWR emerges as an important and statistically significant factor that magnifies hysteresis. Third, expansionary monetary and fiscal policies can mitigate hysteresis when implemented in a timely manner.

Finally, we develop a calibrated New Keynesian model in which hysteresis arises from a contractionary demand shock. The quantitative results replicate the empirical patterns documented in state-level data, underscoring the central role of DNWR in amplifying hysteresis and validating the effectiveness of policy interventions in such environments.

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Online Supplement to “Hysteresis and the Role of Downward Nominal Wage Rigidity: Evidence from U.S. States"

Appendix A Additional Literature Review

Our paper is related to several strands of the literature. The first is the literature on business cycles and hysteresis. A dominant view in macroeconomics has been “the independence assumption” where shocks to trends or structural aspects of the economy and cyclical fluctuations are independent from each other (e.g., Lucas, 1977; Blanchard and Quah, 1989).[34] In this traditional framework, innovations to trends are treated as supply shocks and those to cyclical components are viewed as demand shocks.

Meanwhile, the hysteresis view allows for permanent effects of cyclical shocks. Blanchard and Summers (1986) propose a structural mechanism of hysteresis building on the insider–outsider model to characterize labor market hysteresis and account for sclerosis in European labor markets. Galí (2022) incorporates the insider–outsider framework into a New Keynesian model to show that the source of hysteresis is an inefficiently high equilibrium wage. Abbritti et al. (2021) construct a New Keynesian model with downward nominal wage rigidity and endogenous growth, demonstrating how cyclical shocks can permanently affect trend growth. Similarly, Acharya et al. (2022) build a structural model featuring a search-and-matching labor market block and show that downward nominal wage rigidity, combined with skill depreciation among displaced workers, can drive the economy toward a steady-state unemployment trap. Alves and Violante (2024, 2025) develop a heterogeneous-agent New Keynesian model with three labor market states and nominal wage rigidity, illustrating how a recessionary shock persistently reduces both employment and labor force participation—particularly among low-skilled workers—along with income.

Empirically, recent studies analyze macro hysteresis by focusing on the long-lasting effects of

[34] Cerra et al. (2023) present an extensive review of the literature on hysteresis.

demand shocks or monetary policy shocks. Examples are Cerra and Saxena (2008), Blanchard et al. (2015) and Jordá et al. (2020), who base their analyses on multi-country data. Ma and Zimmermann (2023) estimate effects of monetary policy on R&D investment, interpreting this effect as the source of hysteresis. Relatedly, Furlanetto et al. (2025) identify a permanent demand shock based on a structural VAR model; this shock has long lasting effects on long-term unemployment and also employment. Cajner et al. (2021) empirically show persistent effects of cyclical shocks on the labor force participation rate based on state-level data. Antolin-Diaz and Surico (2025) uncover long-run positive effects of government spending using a BVAR framework and a newly constructed series of military spending disaggregated by category. Bhattarai et al. (2021) examine local hysteresis effects with a specific focus on the housing crisis in the U.S.

Our paper also intersects with the literature on nominal wage rigidity. Tobin (1972), Akerlof et al. (1996), and Benigno and Ricci (2011) demonstrate that nominal wage rigidity is an important source of nonlinearity in the business cycle. Dupraz et al. (2025) highlight downward nominal wage rigidity as a key channel through which asymmetries in unemployment dynamics arise, echoing the predictions of the plucking theory. Other studies—such as Abbritti et al. (2021) and Acharya et al. (2022)—also emphasize that downward nominal wage rigidity is a critical mechanism behind the lasting economic damage caused by recessions, as it raises real wages during downturns and amplifies the effects of recessionary shocks. Empirically and theoretically, Daly and Hobijn (2014) show that downward nominal wage rigidity is essential for understanding nonlinearities in the Phillips curve and the effectiveness of monetary policy. Jo (2024) develops unique state-level measures of downward nominal wage rigidity and investigates its determinants, and Jo and Zubairy (2025) analyze the role of downward nominal wage rigidity in the transmission of government spending shocks. Our empirical findings suggest that downward nominal wage rigidity plays a critical role in understanding hysteresis and the effectiveness of demand-side policies in mitigating it.

Our paper also relates to the literature on regional business cycles and recession prediction. In this literature, the Markov-switching model has been widely used to detect business cycle phases (e.g., Hamilton, 1989). However, as noted by Francis et al. (2018) and Hamilton and Owyang (2012), this standard approach typically assumes that the depth of recessions, the trajectories of recoveries, and the lengths of recovery periods are the same across all recessions. To provide a more nuanced characterization of recovery patterns—particularly to distinguish between bounce-back recoveries and hysteresis— Francis et al. (2018) model the duration of recessions using an accelerated failure time framework, while Eo and Morley (2022) develop a Markov-switching model that allows for two distinct types of recessions.

The literature on recession prediction has increasingly relied on regional data to investigate the effectiveness of policies in stabilizing business cycles. Francis et al. (2018) and Berge et al. (2021) use state-level data to examine the transmission of monetary and fiscal policy, respectively. This paper is closely related to that work in its use of state-level data to analyze business cycles and policy effectiveness; it differs by focusing on hysteresis and its attempt to distinguish between U-shaped and L-shaped recessions using a Bayesian Markov-switching model.

Appendix B Data

B.1 Data sources

Data were amalgamated from a variety of sources including HAVER, FRED, BLS, and Census. For several variables, data were collected from one source for the earlier part of the sample (1960s, 70s, 80s) and from another for more recent years.

1. State-level payroll employment (1960:M1 - 2023:M12): retrieved from HAVER. See Figures B1 and B2.

2. Employment share of manufacturing, finance, and professional services (1969-2023, yearly): Haver (BEAEMPL) 1969-2001; BLS State employment and unemployment (retrieved from FRED, ALFRED) 2002-2023. Employment shares for the three industries are available in FRED starting in 1990. However, to ensure consistency in industrial classification, we use

data from HAVER through 2001 and switch to FRED/ALFRED data from 2002 onward. The employment share for the finance industry is unavailable in ALFRED for New Mexico and South Dakota, so we exclude these two states from the finance series starting in 2002.

3. Unemployment rate and labor force participation rate by gender and by state : BLS. The employment-to-population ratio is calculated based on the unemployment rate and the labor force participation rate. (https://www.bls.gov/lau/ex14tables.htm; https://www.bls.gov/opub/geographic-profile/.)

4. Employment share by firm size (1978-2021, yearly): Census Bureau's Business Dynamics Statistics (BDS)

5. Total tax and total income (yearly, 1960-2022), Annual Survey of State Govt Tax Collections (Census) and Personal Income By State (BEA) data, respectively.

Beyond publicly available data provided by the statistical agencies, some series are sourced from websites maintained by individual researchers or non-profit institutions.

1. Oil production by state (1960-2022 yearly): EIA and Hamilton (2011)[35]

2. Union membership (1964-2021, yearly): Barry Hirsch (http://www.unionstats.com/)

3. Fraction of wage cuts, no wage changes, and wage increases: Yoon Joo Jo (https://sites.google.com/view/yoonjoojo/rsearch).

4. Minimum wage data (by state, and of the federal level): Vaghul and Zipperer (2016)(https://equitablegrowth.org/working-papers/historical-state-and-sub-state-minimum-wage-da

5. Monetary policy shocks of Romer and Romer (2004) extended by Wieland and Yang (2020) (https://www.openicpsr.org/openicpsr/project/135741/version/V1/view?path=/openicpsr/135741/fcr:versions/V1/Monetary_shocks.zip&type=file)

[35] We thank professor James Hamilton for sharing the dataset used in Hamilton (2011).

Table C1: POSTERIOR ESTIMATES FOR THE MARKOV-SWITCHING MODEL OF EMPLOYMENT AT THE NATIONAL LEVEL

Parameter	Posterior Mean	90% Credible Interval
p_{00}	0.93	[0.85, 0.95]
p_{01}	0.04	[0.00, 0.06]
p_{02}	0.03	[0.00, 0.04]
p_{11}	0.73	[0.28, 0.86]
p_{22}	0.77	[0.24, 0.92]
μ_0	0.05	[–0.17, 0.12]
μ_1	–1.30	[–4.34, –1.06]
μ_2	–1.56	[–5.27, –1.18]
σ^2	0.18	[0.11, 0.23]
c_0	7.22	[4.27, 8.50]
ρ	0.61	[0.37, 0.72]
Acceptance Rate	0.35	

Note to table: The posterior estimates are based on the regime-switching model in (3.1) for employment growth at the national level. **Source**: Authors' calculation.

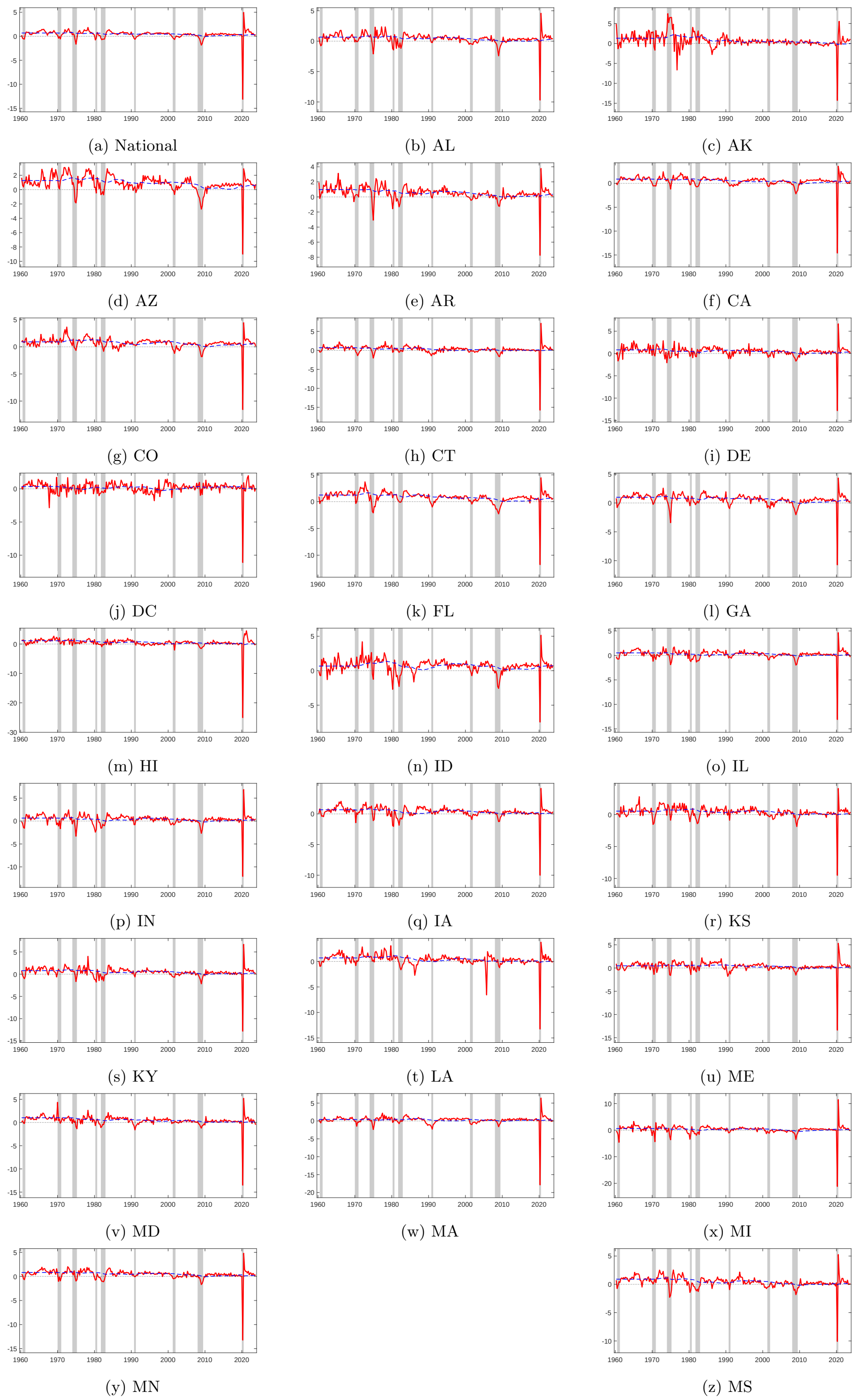
6. Military spending news shock of Ramey and Zubairy (2018) (https://www.openicpsr.org/openicpsr/project/135741/version/V1/view)

Appendix C More Estimation Results

C.1 National Level

Table C1 provides the posterior estimates for the regime-switching model at the national level. The transition probabilities from an expansion regime to L-shaped and U-shaped recession regimes, p_{01} and p_{02}, are 0.04 and 0.03, respectively. This indicates that L-shaped and U-shaped recessions are almost equally likely to occur in employment growth at the national level. The posterior means for the recession shocks, μ_1 and μ_2, are estimated at -1.30 and -1.56 for the L-shaped and U-shaped regimes, respectively, suggesting that the magnitude of the recession shocks is quite similar. The COVID-19 scaling parameter, c_0, is 7.22, indicating that COVID-19 had approximately seven times the recessionary impact on U.S. employment compared to conventional recessions. The decay parameter, ρ, is estimated at 0.61, suggesting that the unusually

Figure B1: NONFARM PAYROLL EMPLOYMENT GROWTH BY STATE (1)

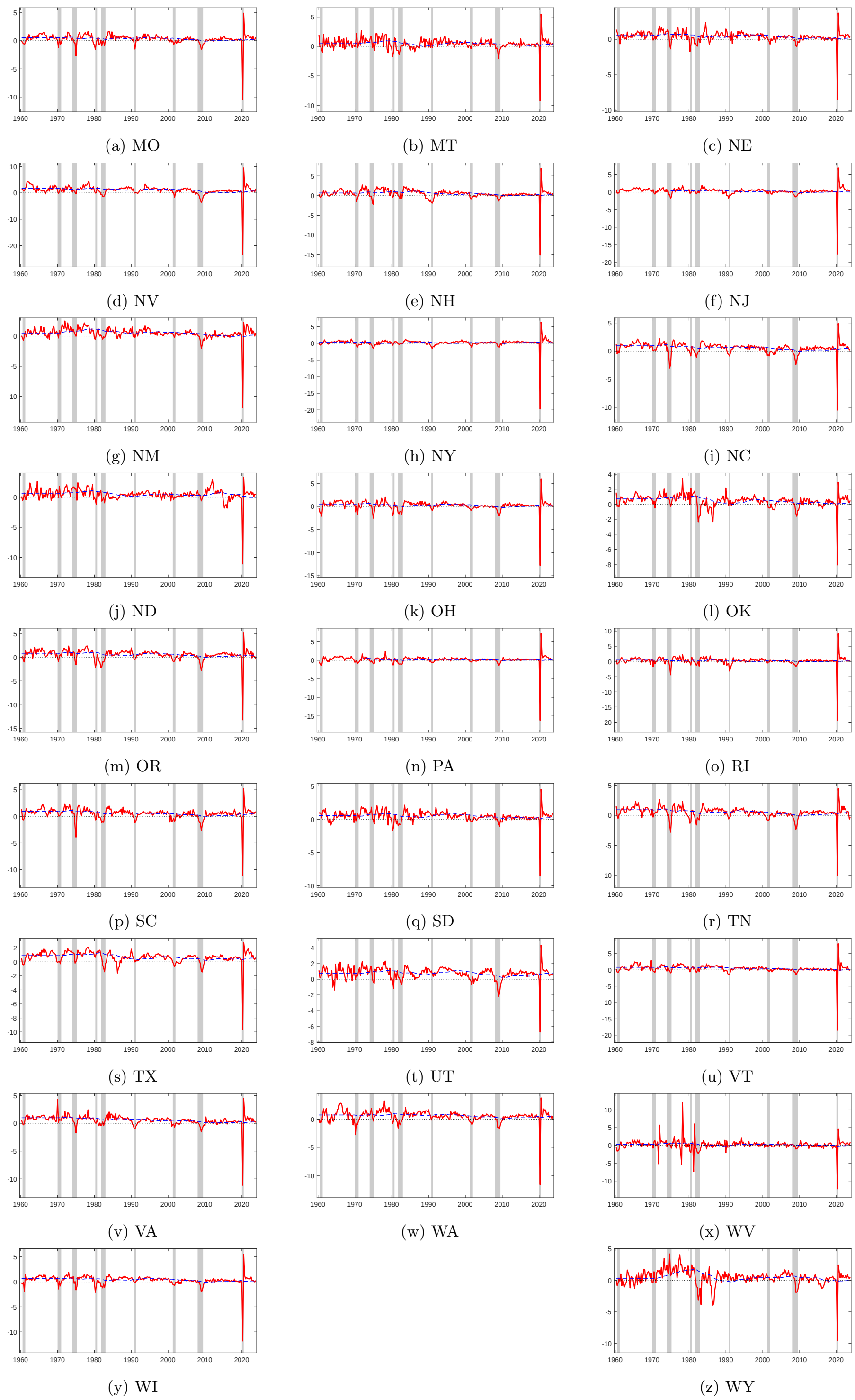


(a) National (b) AL (c) AK (d) AZ (e) AR (f) CA (g) CO (h) CT (i) DE (j) DC (k) FL (l) GA (m) HI (n) ID (o) IL (p) IN (q) IA (r) KS (s) KY (t) LA (u) ME (v) MD (w) MA (x) MI (y) MN (z) MS

Notes to figure: The figures show nonfarm payroll employment growth by state, with the blue dashed line representing the 40-quarter moving average of employment growth for each state. The shaded areas denote the NBER recessions.

Sources: BLS, Haver, and authors' calculation

Figure B2: NONFARM PAYROLL EMPLOYMENT GROWTH BY STATE (2)



(a) MO (b) MT (c) NE

(d) NV (e) NH (f) NJ

(g) NM (h) NY (i) NC

(j) ND (k) OH (l) OK

(m) OR (n) PA (o) RI

(p) SC (q) SD (r) TN

(s) TX (t) UT (u) VT

(v) VA (w) WA (x) WV

(y) WI (z) WY

Notes to figure: The figures show nonfarm payroll employment growth by state, with the blue dashed line representing the 40-quarter moving average of employment growth for each state. The shaded areas denote the NBER recessions.

Source: BLS, Haver, and authors' calculation large impact of COVID-19 diminished rapidly. The acceptance rate of the Metropolis–Hastings sampler is approximately 0.35, which lies within the target range of 0.15 to 0.40 suggested by Gelman et al. (1997).

C.2 State-level Results

Figures C3 and C4 present the estimated probabilities of L-shaped and U-shaped recessions for each state. The blue lines represent the probability of a U-shaped recession, and the red lines represent that of an L-shaped recession. The Y-axis indicates the probability, while the X-axis shows calendar time in quarters. The shaded areas denote NBER-dated recessions.

Table C2 reports the estimated transition probabilities and the expected duration of each regime for all states, along with the national-level estimates for comparison. The expected duration of regime i is computed as $1/(1-p_{ii})$. Across all states, we find that (i) each regime is more likely to persist than to transition to another, indicating strong regime persistence, and (ii) the expected duration of expansions is much longer than that of either L-shaped or U-shaped recessions, suggesting that expansions dominate most sample periods. The expected duration of recessions is about four quarters (one year) on average, and in many states, U-shaped recessions tend to last longer than L-shaped ones.

Appendix D Gender and Hysteresis

D.1 Outcome by Gender

In this section, we examine effects of U-shaped and L-shaped recessions on the labor force participation rate (LFPR) and EPOP ratio in total and by gender based on equation (4.2). Table D3 presents the estimation results. Only L-shaped recessions are associated with statistically significant declines in the LFPR four quarters ahead; in contrast, U-shaped recessions show no statistically significant correlation with changes in the labor force participation rate four quarter

Table C2: PERSISTENCE OF STATE-LEVEL EXPANSIONS AND RECESSIONS

Panel A: States A–M

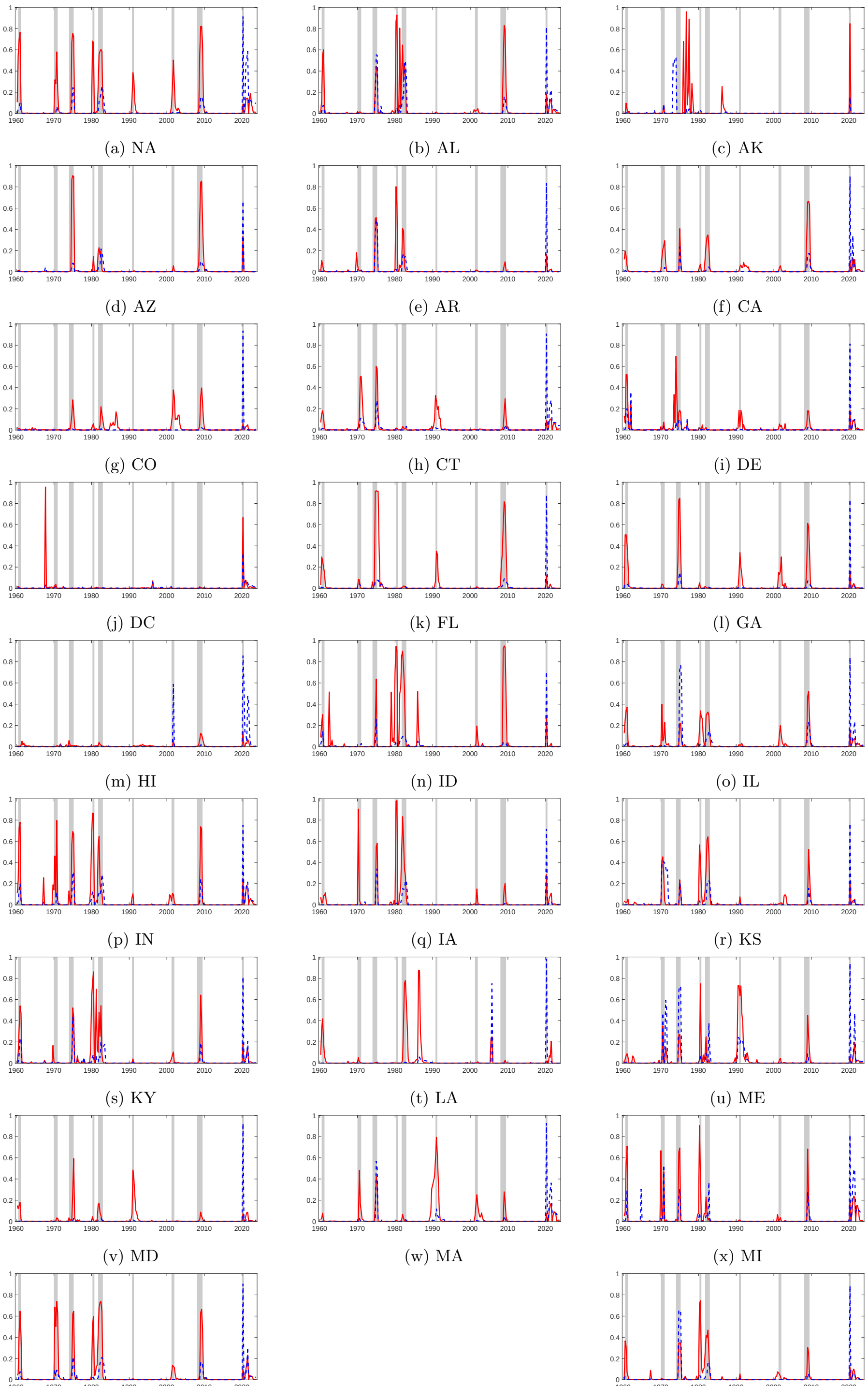
State	**Transition Probabilities**					**Expected Durations**		
	p_{EE}	p_{EL}	p_{EU}	p_{LL}	p_{UU}	Expans.	L-shape	U-shape
National	0.93	0.04	0.03	0.73	0.77	14.69	3.74	4.43
Alabama	0.94	0.04	0.03	0.70	0.76	15.70	3.28	4.10
Alaska	0.95	0.03	0.02	0.64	0.82	18.67	2.74	5.63
Arizona	0.95	0.03	0.02	0.77	0.81	20.27	4.38	5.22
Arkansas	0.95	0.03	0.02	0.76	0.78	19.98	4.09	4.63
California	0.95	0.03	0.02	0.77	0.76	19.50	4.41	4.25
Colorado	0.95	0.03	0.02	0.79	0.80	21.71	4.79	5.07
Connecticut	0.95	0.03	0.02	0.77	0.78	19.35	4.39	4.60
Delaware	0.94	0.03	0.02	0.71	0.75	18.01	3.45	4.05
Florida	0.95	0.03	0.02	0.79	0.80	19.95	4.84	5.04
Georgia	0.95	0.03	0.02	0.76	0.80	19.71	4.17	4.89
Hawaii	0.96	0.02	0.02	0.83	0.76	22.37	5.81	4.18
Idaho	0.94	0.04	0.02	0.70	0.78	15.72	3.31	4.61
Illinois	0.94	0.03	0.03	0.76	0.76	17.77	4.14	4.25
Indiana	0.93	0.04	0.03	0.68	0.77	14.67	3.10	4.32
Iowa	0.94	0.03	0.02	0.69	0.79	17.68	3.25	4.73
Kansas	0.95	0.03	0.03	0.75	0.79	18.19	3.98	4.77
Kentucky	0.94	0.04	0.03	0.69	0.75	15.97	3.22	3.92
Louisiana	0.95	0.03	0.02	0.77	0.75	19.41	4.28	3.97
Maine	0.93	0.03	0.03	0.72	0.70	15.05	3.61	3.35
Maryland	0.95	0.02	0.02	0.78	0.80	21.71	4.47	4.91
Massachusetts	0.95	0.03	0.03	0.78	0.76	18.44	4.46	4.19
Michigan	0.93	0.04	0.03	0.65	0.70	14.72	2.84	3.37
Minnesota	0.94	0.04	0.03	0.74	0.77	15.92	3.83	4.31
Mississippi	0.95	0.03	0.02	0.76	0.77	18.66	4.20	4.40
Missouri	0.94	0.04	0.03	0.68	0.72	16.05	3.12	3.61
Montana	0.95	0.03	0.02	0.74	0.79	20.32	3.81	4.79

Panel B: States N–Z

State	**Transition Probabilities**					**Expected Durations**		
	p_{EE}	p_{EL}	p_{EU}	p_{LL}	p_{UU}	Expans.	L-shape	U-shape
Nebraska	0.95	0.03	0.02	0.72	0.80	19.77	3.53	5.05
Nevada	0.95	0.03	0.02	0.79	0.79	21.08	4.73	4.87
New Hampshire	0.95	0.03	0.03	0.81	0.73	18.19	5.23	3.64
New Jersey	0.95	0.03	0.03	0.78	0.75	18.61	4.64	3.98
New Mexico	0.95	0.02	0.02	0.78	0.79	21.45	4.49	4.86
New York	0.94	0.04	0.03	0.72	0.77	16.07	3.61	4.42
North Carolina	0.95	0.03	0.02	0.75	0.78	19.38	4.06	4.54
North Dakota	0.95	0.02	0.02	0.79	0.80	22.18	4.82	5.02
Ohio	0.93	0.04	0.03	0.69	0.79	14.08	3.18	4.74
Oklahoma	0.95	0.03	0.02	0.73	0.82	19.87	3.70	5.55
Oregon	0.93	0.04	0.02	0.71	0.80	15.23	3.46	4.88
Pennsylvania	0.95	0.02	0.03	0.79	0.76	19.63	4.70	4.13
Rhode Island	0.94	0.03	0.03	0.74	0.75	16.88	3.87	3.96
South Carolina	0.95	0.03	0.02	0.72	0.80	18.67	3.59	5.00
South Dakota	0.95	0.03	0.02	0.77	0.80	21.44	4.27	4.92
Tennessee	0.94	0.03	0.03	0.73	0.77	17.56	3.64	4.43
Texas	0.94	0.04	0.02	0.75	0.80	16.64	4.06	5.03
Utah	0.95	0.03	0.02	0.71	0.78	18.54	3.47	4.59
Vermont	0.95	0.02	0.02	0.79	0.77	20.74	4.79	4.41
Virginia	0.95	0.02	0.02	0.78	0.78	21.71	4.64	4.55
Washington	0.95	0.03	0.02	0.76	0.80	19.89	4.23	4.93
West Virginia	0.95	0.02	0.03	0.83	0.62	19.23	6.04	2.65
Wisconsin	0.93	0.04	0.03	0.66	0.78	13.88	2.91	4.47
Wyoming	0.95	0.03	0.02	0.79	0.84	20.33	4.68	6.22
Washington, DC	0.96	0.02	0.02	0.73	0.84	22.82	3.77	6.16

Notes to table: Estimates of transition probabilities for each state capture the persistence and transitions across regimes: p_{EE} (expansion→expansion), p_{EL} (expansion→L-shape), p_{EU} (expansion→U-shape), p_{LL} (L-shape→L-shape), and p_{UU} (U-shape→U-shape). Expected durations are calculated as $(1-p_{ii})^{-1}$ for $i \in \{E, L, U\}$ and are measured in quarters. **Source**: Authors' calculation.

Figure C3: ESTIMATED RECESSION PROBABILITIES BY STATE (1)



(a) NA (b) AL (c) AK

(d) AZ (e) AR (f) CA

(g) CO (h) CT (i) DE

(j) DC (k) FL (l) GA

(m) HI (n) ID (o) IL

(p) IN (q) IA (r) KS

(s) KY (t) LA (u) ME

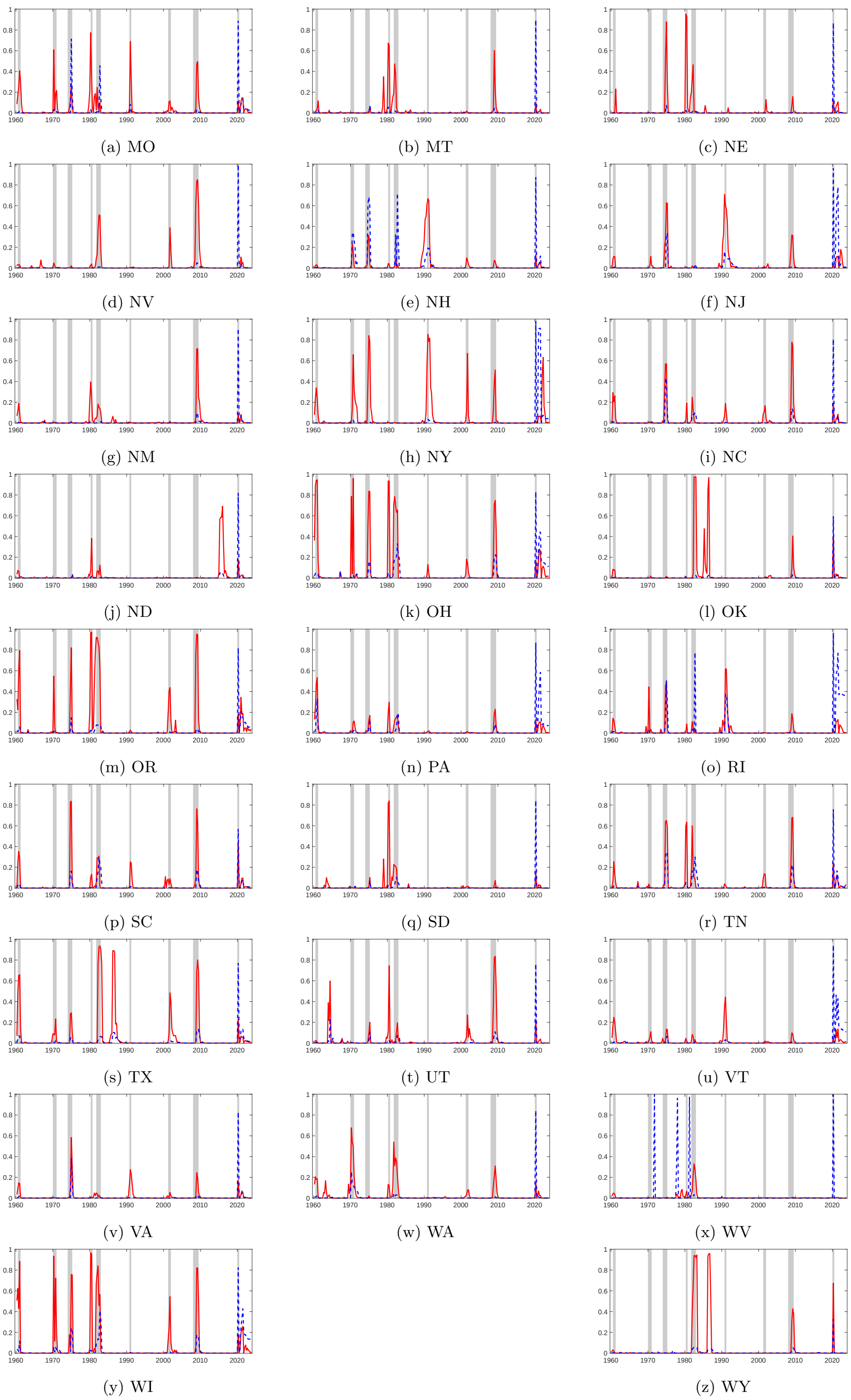
(v) MD (w) MA (x) MI

(y) MN (z) MS

Notes to figure: The figures display state-level recession probabilities. Red solid lines indicate the probability of an L-shaped recession, blue dashed lines indicate a U-shaped recession, and shaded areas denote NBER recessions.

Source: Authors' calculation

Figure C4: ESTIMATED RECESSION PROBABILITIES BY STATE (2)



(a) MO (b) MT (c) NE
(d) NV (e) NH (f) NJ
(g) NM (h) NY (i) NC
(j) ND (k) OH (l) OK
(m) OR (n) PA (o) RI
(p) SC (q) SD (r) TN
(s) TX (t) UT (u) VT
(v) VA (w) WA (x) WV
(y) WI (z) WY

Notes to figure: The figures display state-level recession probabilities. Red solid lines indicate the probability of an L-shaped recession, blue dashed lines indicate a U-shaped recession, and shaded areas denote NBER recessions.

Sources: Authors' calculation

Table D3: PREDICTABILITY OF LFPR AND EPOP RATIO

LFPR	Men	Women	Total
[1] β_l (p^l_{it})	-0.675***	-0.251**	-0.476***
	(0.106)	(0.123)	(0.070)
[2] β_u (p^u_{it})	0.288	-0.080	-0.223
	(0.370)	(0.429)	(0.244)
State fixed effects	✓	✓	✓
R^2	0.010	0.007	0.019
No. of obs.	8,772	8,772	8,772
EPOP	Men	Women	Total
[1] β_l (p^l_{it})	-2.170***	-1.120***	-1.627***
	(0.146)	(0.132)	(0.098)
[2] β_u (p^u_{it})	3.438***	1.325***	3.009**
	(0.509)	(0.459)	(0.341)
State fixed effects	✓	✓	✓
R^2	0.027	0.013	0.036
No. of obs.	8,772	8,772	8,772

Notes to table: This table presents the coefficient estimates from Equation (5.1), excluding observations with zero recession probabilities. The notations ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. Numbers in parentheses are standard errors.

Source: Authors' calculation.

ahead.[36]

We also see quite different effects on the EPOP ratio. L-shaped recessions predict declines in the EPOP ratio four quarters ahead, mirroring their effect on labor force participation, while U-shaped recessions predict a rise in the EPOP ratio, primarily driven by the swift recovery of the unemployment rate. The significant positive association between U-shaped recessions and the four-quarter-ahead EPOP ratio — contrasted with the significant negative association for L-shaped recessions — provides further evidence that our methodology effectively distinguishes between the two distinct recession experiences.

We further examine heterogeneous responses of the EPOP ratio and LFPR by gender. Since the labor data by gender are available only at an annual frequency, we analyze one-year changes in these measures to align with the forecasting horizons used for the aggregate data. Overall, the aggregate patterns hold for both gender, though the magnitudes differ. Specifically, L-

[36] The result remains robust for $h = 8$.

shaped recessions have more negative effects on men's LFPR than on women's, suggesting that shocks that trigger L-shaped recessions have persistent negative effects on men's labor force participation. Meanwhile, U-shaped recessions lead to stronger positive effects on the EPOP ratio for men than for women, suggesting that the rebound of employment is more concentrated among men. Overall, men's employment and labor force participation are more cyclically sensitive and more heavily affected by the nature of the recovery in the labor market. This result further suggests that the disappearance of U-shaped recessions and the increased prevalence of L-shaped recessions since the 1990s are likely to be associated with the stagnation of male employment (Cortes et al., 2018).

All told, the persistent negative effects of L-shaped recessions on the LFPR and the EPOP ratio are consistent with previous empirical findings on macroeconomic hysteresis discussed by Furlanetto et al. (2025) and Alves and Violante (2025). The results validate our empirical methodology's ability to distinguish effectively between U-shaped and L-shaped recessions and confirm that L-shaped recessions effectively capture the phenomenon of hysteresis.

D.2 Effects of Gender Gap on Hysteresis

Table D4 reports the coefficients on the gender gap in the labor market. We consider two measures of the gender gap: the difference in the EPOP ratio between men and women (Gap_{it}), and an indicator of the recession gender gap (E^r_{it}), which equals one when Gap_{it} exceeds the cross-state average and is interacted with $(1-p^e_{it})$. The statistically significant and positive coefficients indicate that a larger gender gap increases the relative risk of hysteresis over a full recession recovery.

Table D4: EFFECTS OF GENDER EMPLOYMENT GAP ON THE RELATIVE RISKS OF RECESSIONS

	(1) All	(2) 2000-2019	(3) All	(4) 2000-2019
Indicator of DNWR	0.062***	0.052***	0.050***	0.039***
(Z_{it}^r)	(0.009)	(0.009)	(0.009)	(0.009)
Gender gap	0.082**	0.178***		
(Gap_{it})	(0.037)	(0.043)		
Indicator of recession gender gap			0.116***	0.081***
(E_{it}^r)			(0.010)	(0.015)
(Controls)				
p_{it}^e	✓	✓	✓	✓
oil-producing	✓	✓	✓	✓
minimum wage	✓	✓	✓	✓
union	✓	✓	✓	✓
manufacturing	✓	✓	✓	✓
prof. services	✓	✓	✓	✓
finance	✓	✓	✓	✓
large-firm share	✓	✓	✓	✓
tax-income share	✓	✓	✓	✓
State fixed effect	✓	✓	✓	✓
Time fixed effect	✓	✓	✓	✓
No. of Obs.	8,256	3,856	8,256	3,856
R^2	0.796	0.867	0.799	0.868

Notes to table: This table presents the coefficient estimates from Equation (5.1) with the zero recession probabilities replaced with the minimum of the corresponding estimates and with the share of wage cuts. The notations ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels, respectively. Numbers in parentheses are standard errors. Panels labeled "(Controls)" indicate the inclusion of control variables in the regression model.

Source: Authors' calculation.