Essays in Labor Economics and Labor Market Policy

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1 Introduction

Since the works of Adam Smith, who is widely considered to be the father of modern economics, labor market outcomes have been at the root of economic reflections. As labor earnings constitute the most important source of income for the majority of households around the world, the importance of labor market outcomes for economic well-being can hardly be overstated.

The last decades have seen rapid technological advancements, growing economic inequality, and severe financial crises. It is a vital task for economic research to expose which developements affected labor market outcomes, to enhance our understanding of how they affected labor market outcomes, and to explore which economic policies can support workers during and after transition periods.

1.1 Overview of the Thesis

This thesis contributes to the existing literature by advancing our understanding of the forces that have shaped labor market outcomes in recent decades. It consists of three independent research papers, which zoom in on the effects of financial frictions, routine-biased technical change, and judicial ideology, respectively. My coauthors and I employ both empirical and analytical methods to analyze the underlying mechanisms and discuss appropriate policy implications.

Collateral Constraints, Wage Rigidity, and Jobless Recoveries The Great Recession has drawn attention to the importance of macro-financial linkages. In Chapter 2, I explore the joint role of imperfections in labor and financial markets for the cyclical adjustment of the labor market. I show that jobless recoveries emerge when, upon exiting a recession, firms are faced with deteriorating credit conditions. On the financial side, collateral requirements affect the cost of borrowing for firms. On the employment side, hiring frictions and wage rigidity increase the need for credit, making the binding collateral constraint more relevant. In a general equilibrium business cycle model with search and matching frictions, I illustrate that tightening credit conditions calibrated from data negatively affect employment adjustments during recovery periods. Wage rigidity substantially amplifies this mechanism, generating empirically plausible fluctuations in employment and output.

A Joint Theory of Polarization and Deunionization Over the past 40 years, the U.S. and several European labor markets have undergone two incisive developments – labor market polarization and deunionization. In Chapter 3, using state-level labor market data, Anna Hartmann and I document a positive relationship between the two phenomena in the U.S.: the decrease in unionization rates has been significantly more pronounced in states with a higher employment share in routine-intensive occupations. Contrary to conventional wisdom, deunionization is mainly driven by large within-industry and within-occupation changes in union membership rates and not only by compositional effects. Building on this observation, we argue that the commonly assumed driver of polarization, routine-biased technical change, is also the main driving force behind the decline in union membership rates. In a model with search and matching frictions where workers choose occupations and endogenously form unions, we illustrate that shifts in the structure of labor demand in favor of low- and high-skill occupations worsen the bargaining position of unions and make participation in collective bargaining less attractive for workers. The ensuing within-industry and within-occupation decline in unionization rates in turn provides incentives for former middle-wage workers to switch to low-wage occupations, which further amplifies job market polarization.

Outlawed: Estimating the Labor Market Effects of Judicial Ideology Most evidence on the economic impact of the judiciary is either case-based or purely anecdotal. In Chapter 4, Christian Bredemeier, Anna Hartmann, and I provide evidence on the systematic labor market effects of judicial ideology, employing broad data on court rulings and labor market outcomes. Our identification strategy uses heterogenous effects of ideological shifts of the U.S. Supreme Court on U.S. district court rulings, which we derive from a theoretical model of judge decision-making and document empirically. Exploiting this heterogeneity, we find that an increase in the share of conservative rulings substantially increases the employment rate and promotes labor market fluidity but also contributes to wage stagnation, employment polarization, deunionization, and rising income inequality. Our main empirical results can be rationalized in a search and matching model with wrongful-termination lawsuits.

1.2 Contribution to Chapters 3 and 4

While Chapter 2 is based on a research paper produced entirely by myself, Chapter 3 is based on joint work with Anna Hartmann and Chapter 4 is based on joint work with Christian Bredemeier and Anna Hartmann.

The research idea for the paper "A Joint Theory of Polarization and Deunionization" was developed in discussions between Anna Hartmann and myself. The empirical analysis was performed by both of us. Anna Hartmann developed the formal representation of the idea and I conducted the quantitative evaluation. We both wrote the first draft of the paper and all subsequent revisions.

The research idea for the paper "Outlawed: Estimating the Labor Market Effects of Judicial Ideology" was developed by Christian Bredemeier. The related literature was evaluated by Christian Bredemeier and myself. Anna Hartmann and Christian Bredemeier conducted most of the empirical analysis. I contributed the assessment of datasets from the law literature and the formal representation and quantitative evaluation of the labor market model. I have written the first draft, which Christian Bredemeier, Anna Hartmann, and I revised.

2 Collateral Constraints, Wage Rigidity, and Jobless Recoveries

Author: Tobias Föll

2.1 Introduction

During the 2008 financial crisis, the unemployment rate in the U.S. doubled from 5.0% to 10.0% within only 18 months. While output fully recovered less than two years after the end of the recession, the unemployment rate took three times as long to reach its pre-crisis level. The marked increase in the unemployment rate was preceded by deteriorating credit conditions and an increase in collateral requirements (cf. Jermann and Quadrini, 2012; Garín, 2015). As collateral constraints directly affect firms' hiring decisions, recessions caused by financial frictions might have particularly large adverse effects for the labor market. Motivated by this, I aim to determine the role of disturbances in the financial sector for jobless recoveries in the U.S. since the 1990s.

In order to to so, I add a financial market friction to the standard DSGE model with search and matching frictions, whereby firms need to provide capital as collateral in order to take on loans. Labor and capital are treated asymmetrical, with capital serving a dual role as production factor and as collateral. These elements generate diverging output and employment dynamics during recovery periods and contribute to the emergence of jobless recoveries.

The key interaction in the model arises from the need for funding: due to a cash flow mismatch, firms are required to finance their working capital requirements, including vacancy posting costs, by taking on loans. The presence of financial frictions makes hiring more costly for constrained firms, as they have to cut investment or dividend payouts to finance their wage bill. This implies that the degree to which firms are affected by wage rigidity also varies with credit tightness.

The model is calibrated to U.S. data and simulated using technology and credit shocks, where credit shocks are meant to capture variation in credit conditions. I compare the model dynamics to business cycle statisitics for the U.S. between 1964 and 2004. The simulated model with only two shocks can account for nearly 50% of the variation in unemployment, roughly 90% of the fluctuations in vacancies, nearly 70% of the variation in labor market tightness, and virtually 100% of the fluctuations in the job-finding rate.

I find that after a negative technology or credit shock, the initial increase in the unemployment rate is stronger and steeper because wage rigidity keeps firms' borrowing needs high. Financial frictions are responsible for flatter decreases in the unemployment rate: following an increase in productivity, firms prioritize investment into the asset used as collateral in order to relax the borrowing constraint. Thus, the initial increase in vacancies and hiring during a recovery period is lower compared to models with perfect credit markets.

The interaction of financial frictions and wage rigidity generates asymmetric unemployment dynamics. I illustrate that the combined effect of these frictions on unemployment dynamics is larger than the sum of the separate effects of financial frictions or wage rigidity. This allows for an amplification of shocks in the model that is close to what is found in the data. In contrast to the results obtained in similar models without financial frictions, even a small and empirically plausible amount of wage rigidity is sufficient to generate highly volatile labor market variables once collateral constraints are taken into account.

Despite the asymmetry in the unemployment rate generated by the combination of a collateral constraint and wage rigidity, recoveries in the model are not jobless unless there is a concomitant erosion of credit conditions. The reason is that credit conditions directly affect the marginal value of an additional worker and thereby the number of hires and the unemployment rate. When credit conditions deteriorate while total factor productivity recovers, unemployment remains above its pre-crisis level. Since capital can be used as production factor and as collateral, the capital stock and output are almost entirely driven by total factor productivity and not by credit conditions. Consequently, recovering total factor productivity, combined with worsening credit conditions, causes jobless recoveries.

This mechanism is consistent with empirical evidence. Analyzing credit conditions during recessions and subsequent recovery periods in the U.S. between 1964 and 2010, I find that prior to 1990 credit conditions started to improve immediately after the end of recessions. During the recent jobless recoveries, credit conditions deteriorated for several quarters after the end of the recessions and the unemployment rate only began to recover once credit conditions had stabilized.

My analysis suggests that low credit availability matters for the occurrence of jobless recoveries after the recent recessions. This has important policy implications. Policies aimed at reducing transitional unemployment through reemployment services, such as the $47 billion dollar spent i.a. on job training in the American Recovery and Reinvestment Act of 2009, might not be as effective as hoped. Alternative policies could be aimed at reducing uncertainty on the credit market in order to make credit more easily available and to facilitate job creation.

The remainder of the paper is organized as follows. Previous research is discussed in the next section. The model outline is presented in Section 2.3. In Section 2.4 the quantitative analysis is described in detail. Jobless recoveries and policy implications are discussed in Section 2.5. To conclude, the results are summarized in Section 2.6.

2.2 Related Literature

This paper adds to the literature trying to understand the role of financial conditions for macroeconomic dynamcis. First steps in this direction have been taken by Kiyotaki and Moore (1997). Wasmer and Weil (2004) introduce search frictions into the credit market and find that the presence of financial frictions increases macroeconomic volatility. Jermann and Quadrini (2012) estimate financial shocks and show that they contribute siginificantly to the dynamics of real and financial variables. While all of these approaches provide interesting insights, I choose to follow Garín (2015) who introduces financial frictions in the style of Kiyotaki and Moore (1997) into a search and matching framework. This approach has the advantage that it provides a direct link between collateral requirements and asset prices and that changes in credit availability directly affect firms' job creation decision.

Closely related to this work are Schoefer (2016) and Moiseeva (2018) who both study the interaction of financial frictions and wage rigidity in a search and matching framework. However, neither of the models presented in these papers is able to generate jobless recoveries. In Moiseeva (2018), since the financial costs of hiring are high in recessions, firms delay hiring until the recession has passed. While this mechanism amplifies fluctuations in labor market variables, the rapid increase in hiring after a recession stands in sharp contrast to the observation of jobless recoveries. Schoefer (2016) explores a channel similar to the one presented here, through which wage rigidity and financial frictions influence a firm's job creation decision. Since labor is the only production factor in Schoefer (2016), any asymmetry in unemployment mechanically spills over to output dynamics as long as technology shocks are symmetric. Thus, after a recession, employment will have fully recovered at the point of output recovery.

Finally, this paper adds to the literature that studies the role of financial conditions for jobless recoveries.[1] Schott (2013) distinguishes between incumbent firms and startups and argues that low credit availibility for young firms is responsible for the lack of job creation. Wesselbaum (2019) emphasizes the role of financial frictions under match efficiency shocks. Calvo et al. (2014) make the case that jobless recoveries are caused by the interaction of financial frictions and wage rigidity. To illustrate their empirical findings, they analyze a stylized competitive model of the labor market with an ad-hoc borrowing constraint. In their model, productivity growth leads to jobless recoveries when the borrowing constraint binds and wages are rigid. I demonstrate that these conditions are not sufficient for the emergence of jobless recoveries in a general equilibrium framework with an endogenous borrowing constraint. Additionally, since the labor market in Calvo et al. (2014) is assumed to be competitive, wage rigidity is essential in generating jobless recoveries. My findings suggest that while wage rigidity amplifies the extent of jobless recoveries, it is not a prerequisite for their occurrence.

[1]The proposed reasons for the joblessness of the most recent recessions in the U.S. are manifold. For example, Meltzer (2003) puts forward a potential downward bias in employment statistics, Groshen and Potter (2003) propose increased speed of structural change, Galí et al. (2012) demand shocks, Shimer (2012) wage rigidity, Schmitt-Grohé and Uribe (2017) liquidity traps, and Jaimovich and Siu (2018) job polarization.

2.3 A Model with Financial Frictions and Wage Rigidity

In this section, I introduce wage rigidity into a simple version of the model presented in Garín (2015).[2] The model economy is populated by two types of agents: workers and capitalists. Capitalists own the firms. Firms produce a homogenous good y_t using labor n_t and physical capital k_t. All dividends d_t are transferred to the capitalists. Workers have access to a one-period riskless bond a_t that is issued by capitalists.

The labor market is subject to search frictions in the sense of Mortensen and Pissarides (1994): hiring workers entails vacancy posting costs that are paid by the firms. Wages are determined by standard Nash bargaining over the entire surplus of a worker-firm match. Households consist of a continuum of workers and are assumed to perfectly share all risks.

2.3.1 The Labor Market

The number of matches on the labor market is determined by $m_t = \nu u_t^{\gamma} v_t^{1-\gamma}$, where ν is the efficiency of the matching technology, u_t is unemployment, and v_t vacancies. The parameter γ governs the elasticity of the matching function with respect to unemployment and vacancies. The job-filling rate, the probability with which a firm fills a vacancy, is given by $m_t/v_t \equiv q(\theta_t) = \nu\theta_t^{-\gamma}$. The job-finding rate, the probability with which an unemployed worker finds a job, is given by $m_t/u_t \equiv f(\theta_t) = \nu\theta_t^{1-\gamma}$. Labor market tightness θ is defined as $\theta_t \equiv v_t/u_t$. When labor market tightness is low, many unemployed workers compete for few vacant jobs. The job-filling rate is high and the job-finding rate is low.

At the beginning of each period, a fraction x of all existing worker-firm matches is exogenously separated. Newly separated workers immediately begin searching for a new job and have the same job-finding rate as all other unemployed workers. Employment evolves according to

$$n_t = (1-x)n_{t-1} + m_t.$$

and at the end of each period

$$u_t = 1 - n_t \tag{2.1}$$

workers remain unemployed. Since search is costless from the household perspective, all non-employed workers search for a job.

Posting a vacancy entails costs of $c(v_t) = \frac{\kappa}{2}v_t^2$ per period, where $\frac{\kappa}{2} \in (0, +\infty)$ represents the resources a firm must spend because of matching frictions.[3] Furthermore, I assume that

[2] The simplifications are made in order to facilitate the understanding of the relevant mechanisms. All of the results are robust to the inclusion of the training costs present in Garín (2015). The model outline is kept brief and derivations are deferred to the appendix.

[3] I follow Merz and Yashiv (2007), Kaas and Kircher (2015), and much of the recent literature in assuming convex vacancy posting costs. While the asymmetry of unemployment dynamics is somewhat dampened, all results are robust to the standard assumption of linear vacancy posting costs. The resource costs are defined as $\kappa/2$ in order to simplify the first order conditions of the firm.

there is no risk on the firm side. Firms can hire h_t workers with certainty by posting $m_t/q(\theta_t)$ vacancies.

2.3.2 Households

The setup allows for the existence of a representative household, consisting of a continuum of workers of measure one. The household aims at maximizing lifetime utility by allocating consumption optimally across all members:

$$\mathbb{E}_t \sum_{n=0}^{\infty} \beta_h^j \left[\ln(c_{t+j}) - \varphi n_{t+j}\right],$$

where c_t is consumption, β_h is the discount factor of the household, φ is the disutility from work, and n_t is the share of workers that is employed at time t.[4] The household's flow of funds constraint is given by

$$c_t + \frac{a_{t+1}}{R_t} + T_t \leq w_t n_t + a_t + (1-n_t)s.$$

Employed workers earn wages w_t and unemployed workers receive benefits s. The benefits are financed through a lump-sum tax $T_t = (1-n_t)s$. The one-period riskless bond a_t pays an interest rate of R_t and is used for consumption smoothing.

The representative household chooses consumption and the number of bonds in order to maximize the expected discounted lifetime utility over consumption and leisure. Since it takes the job-finding rate as given, employment evolution from the household perspective can be described by

$$n_t = (1-x)n_{t-1} + f(\theta_t)u_{t-1}.$$

The complete household maximization problem is given in Appendix A.1. Combining the first order conditions with respect to consumption and bonds results in the standard Euler equation

$$\frac{1}{R_t} = \beta_h \mathbb{E}_t \frac{c_t}{c_{t+1}}. \tag{2.2}$$

Intuitively, the household invests into bonds until the marginal utility of today's consumption is equal to the discounted marginal utility of consuming tomorrow, weighted by the rental rate R_t.

[4] Since the utility function is separable between consumption and leisure and perfect risk-sharing is assumed, all workers have the same level of consumption.

2.3.3 Financial Markets

Due to a cash-flow mismatch firms need to raise funds via intra-period loans l_t in order to finance their working capital requirements.[5] Wage payments, dividend payouts, investments, current debt, and vacancy posting costs all accrue before the realization of revenues. Since contract enforcement is costly, firms are subject to a collateral requirement. Following a default, financial intermediaries cannot seize production. Only the installed capital stock can be recovered and sold at $\eta_t q_{k,t} k_t$, where η_t captures uncertainty regarding the tightness of the credit market and $q_{k,t}$ is the marginal Tobin's Q. As is standard in the literature, financial intermediaries are assumed to have no bargaining power in the debt renegotiation and they do not value the stock of workers in the firm (cf. Garín, 2015; Perri and Quadrini, 2018). η_t is interpreted as an exogenous collateral shock following the stochastic process

$$\ln \eta_t = (1 - \rho_\eta) \ln \bar{\eta} + \rho_\eta \ln \eta_{t-1} + \epsilon_{\eta,t}$$

with $\epsilon_{\eta,t} \sim \mathbb{N}(0, \sigma_\eta)$, where $\bar{\eta}$ is the mean of the stochastic process.

Under these assumptions, a firm's ability to borrow is constrained by

$$l_t + \frac{b_{t+1}}{R_t} \leq \eta_t q_{k,t} k_t. \tag{2.3}$$

The intra-period loan and the newly issued debt must lie below the value of the fraction of the physical capital stock that lenders can recuperate after default. The derivation of the enforcement constraint is provided in Appendix A.2.

2.3.4 Firms

Capitalists are risk-averse and derive utility from the consumption of dividend payouts. They can only access the financial market through the firm and are assumed to be more impatient than households, i.e., $\beta_h > \beta_c$, where β_c is the discount factor of the firm.[6]

Thus, capitalists' expected lifetime utility is a function of dividends,

$$\mathbb{E}_t \sum_{j=0}^{\infty} \beta^{t+j} \frac{d_t^{1-\sigma}}{1-\sigma}.$$

As firms are owned by capitalists, the objective of a firm is to maximize the expected future stream of discounted dividends. Firms own the capital stock k_t and use it together with labor n_t to produce a homogenous good with $y_t = z_t n_t^\alpha k_t^{1-\alpha}$, where $0 < \alpha < 1$ and technology z_t follows the stochastic process $\ln z_t = \rho_z \ln z_{t-1} + \epsilon_{z,t}$ with $\epsilon_{z,t} \sim \mathbb{N}(0, \sigma_z)$. Firms can borrow

[5] Evidence by Buera and Shin (2013) supports the assumption that most of a firm's costs require working capital.

[6] These assumptions are standard in the literature. First, with access to financial markets, capitalists could smooth consumption and reduce the costs associated with changes in dividends. This would dampen any effect of credit frictions. Second, the smaller discount factor compared to households impedes capitalists from saving enough to avoid the borrowing constraint.

using one-period riskless bonds b_{t+1} at the gross interest rate R_t. Since the model does not feature any idiosyncratic shocks, I focus on a symmetric equilibrium and a representative firm. The complete maximization problem is given in Appendix A.3.

The marginal value of an additional worker for the firm $\mathbb{J}_{n,t}$ is obtained by taking the first derivative of the firm's value function $\mathbb{J}_t$ with respect to employment

$$\mathbb{J}_{n,t} = \left[\alpha z_t n_t^{\alpha-1} k_t^{1-\alpha}(1-\mu_{b,t}) - w_t\right] + (1-x)\mathbb{E}_t \Lambda^c_{t|t+1} \mathbb{J}_{n,t+1}, \tag{2.4}$$

where $\Lambda^c_{t|t+j} = \beta^j \left(\frac{d_t}{d_{t+1}}\right)^\sigma$ is the stochastic discount factor of capitalists and $\mu_{b,t}$ is the Lagrange multiplier on the borrowing constraint. The term in square brackets is equal to the net return of an additional worker, while the second term is the present discounted value of the hired worker. Note that without financial frictions $\mu_{b,t}$ is equal to zero. Consider an increase in collateral requirements. The firm is more constrained, which increases the value of relaxing the borrowing constraint, i.e., $\mu_{b,t}$. This reduces the net return of an additional worker and therefore the marginal benefit of hiring. Intuitively, the firm has to finance an additional worker's wage via intra-period loans. When the borrowing constraint is already binding, this can only be done by reducing investment or dividend payouts. This reduces the marginal value of an additional worker.

Proposition 1. *The effect of wage rigidity on the hiring decision is larger for a financially constrained firm.*

Proof. The elasticity of the marginal value of an additional worker with respect to the wage rate is given by

$$\epsilon_{w_t}^{\mathbb{J}_{n,t}} = -\frac{w_t}{\mathbb{J}_{n,t}}.$$

The absolute value of this elasticity increases with $\mu_{b,t}$. As the marginal value of relaxing the borrowing constraint increases proportionally with collateral requirements, the elasticity of the marginal value of an additional worker with respect to changes in the wage increases with collateral requirements, too. □

This means that the marginal benefit of hiring an additional worker reacts more strongly to changes in the wage compared to standard search and matching models. Consequently, even a small amount of wage rigidity has large effects on labor market variables in my model.

2.3.5 Wage Bargaining and Wage Rigidity

As is standard in most of the search and matching literature, wages are determined as the solution of a generalized Nash bargaining problem. The production function exhibits constant returns to scale, which greatly simplifies the bargaining problem.[7] The wage equation is given by[8]

$$\begin{aligned} w_t^* &= \phi \left[\alpha z_t n_t^{\alpha-1} k_t^{1-\alpha}(1-\mu_{b,t}) + (1-x)\mathbb{E}_t \left\{ \Lambda^c_{t+1|t} \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2} \right\} \right] \\ &\quad + (1-\phi)\left[s + \varphi c_t\right] \\ &\quad - \phi(1-x)\mathbb{E}_t \left\{ \Lambda^h_{t+1|t}[1-f(\theta_{t+1})] \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2} \right\}. \end{aligned} \tag{2.5}$$

Since the model economy is subject to two kinds of shocks, wage rigidity in the style of Blanchard and Galí (2010) or Michaillat (2012) is not feasible. Instead, as in Hall (2005) and Krause and Lubik (2007), wage rigidity is introduced through a backward-looking wage norm that limits the adjustment capability of wages

$$w_t = \tau w_t^* + (1-\tau) w_{t-1}, \tag{2.6}$$

where w_t^* is the solution to the generalized Nash bargaining problem given by Equation (2.5) and τ determines the extent of wage rigidity. With this wage schedule, the steady state real wage remains the same regardless of the amount of wage rigidity in the model.

Proposition 2. *Assume that the wage schedule is given by Equation (2.6). Wages are privately efficient if the wage schedule satisfies*

$$\begin{aligned} & s + \varphi c_t - (1-x)\mathbb{E}_t \Lambda^h_{t+1|t}[1-f(\theta_{t+1})]\mathbb{H}_{m,t+1} \\ & \leq w_t \leq \alpha z_t[(1-x)n_{t-1}]^{\alpha-1} k_t^{1-\alpha}(1-\mu_{b,t}) \\ & + (1-x)\mathbb{E}_t \Lambda^h_{t+1|t} \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2}. \end{aligned}$$

Proof. See Appendix A.7. □

This proposition implies that no worker-firm match generating a positive bilateral surplus is separated because of wage rigidity as long as the actual wage remains within the postulated bounds. Thus, the wage schedule in Equation (2.6) is not subject to the Barro (1977) critique that bargaining workers and firms should be able to exploit all possible bilateral gains in long-term worker-firm relationships with reoccuring wage renegotiations. Due to constant returns in production, the model is also not affected by the critique of Brügemann (2017) concerning wage rigidity in search and matching models with diminishing returns.

An alternative way of introducing wage rigidity is provided by Gertler and Trigari (2009). They assume a standard Calvo (1983) wage-setting scenario in the sense that only a fraction τ of

[7]Models with diminishing returns are subject to the critique by Stole and Zwiebel (1996), as each additional worker has a lower marginal product than the last. In addition to constant returns to scale, it is also necessary that firms first hire workers, subsequently bargain about the wages, and only then choose the capital stock (cf. Cahuc and Wasmer, 2001).

[8]The derivation of the wage schedule is provided in Appendix A.4.

all firms is able to renegotiate wages in every period. If a firm is able to adjust wages in a given period, the new wage is determined by generalized Nash bargaining over the total surplus of the match. However, this approach requires a deviation from the standard assumption of fixed vacany posting costs. Since Shimer (2005) and Hall (2005) point out that there is no compelling theory of wage determination in the kind of model presented here, I stick to the analytically very simple form of wage rigidity given by Equation (2.6), which allows me to use the standard labor market setup with constant vacancy posting costs. The derivation of a staggered wage schedule à la Gertler and Trigari (2009) in a model with financial frictions and the corresponding model simulations are provided in Appendix A.8. The results are qualitatively the same and quantitatively very close to the results obtained with the ad-hoc wage norm.

Remarkably, the log-linearized wage index derived in Gertler and Trigari (2009) looks very similar to the wage schedule in Equation (2.6). In particular, the derivations in Appendix A.8 establish that Equation (2.6) is the outcome of their staggered Nash bargaining approach under financial frictions, if neither firms nor workers take into account that they might not be able to renegotiate wages in the subsequent periods. I use this interpretation of the wage norm to calibrate the parameter τ, governing the extent of wage rigidity, in Section 2.4.

2.3.6 Equilibrium

With the model completely described, I define the equilibrium.

Definition 1. *A recursive equilibrium is defined as a set of i) firm's policy functions $d(\omega^c;\Omega)$, $b(\omega^c;\Omega)$, $k(\omega^c;\Omega)$, $i(\omega^c;\Omega)$, and $v(\omega^c;\Omega)$; ii) household's policy functions $c(\omega^h;\Omega)$ and $a(\omega^h;\Omega)$; iii) a lump sum tax $T(\Omega)$, iv) prices $w(\Omega)$ and $R(\Omega)$; and v) a law of motion for the aggregate states, $\Omega' = \Psi(\Omega)$, such that: i) the firm's policies satisfy the firm's first order conditions (Equations (A.7)–(A.11)) and the job creation condition (Equation (2.4)); ii) household's policy function satisfies the household's first order condition (Equation (2.2)), iii) the wage is determined by Equation (2.6); iv) $R(\Omega)$ clears the market for riskless assets such that $a(\Omega) = b(\Omega)$; v) the law of motion $\Psi(\Omega)$ is consistent with individual decisions and with the stochastic processes for z and η, and vi) the government has a balanced budget such that $s(1-n) = T$.*

2.4 Quantitative Analysis

In this section, I calibrate all parameters discussed above to match different aspects of quarterly U.S. data for the time period between the first quarter of 1964 and the fourth quarter of 2004.[9]. I use the calibrated model to simulate time series of all variables. The model performance is evaluated along several dimensions, most importantly with respect to unemployment dynamics.

[9]The 2008 financial crisis is deliberately left out of the sample to guarantee that the results are not driven by this particular recession.

2.4.1 Calibration

Table 2.1 lists the exact parameter values as well as the source that encourages the specific choice. The discount factors are set to $\beta_h = 0.996$ and $\beta_c = 0.983$, to match an annual interest rate of 1.6% and an annual return on equity of 7%.

Next, I calibrate the labor market variables. For the separation rate, I choose a conventional value of 0.1 (cf. Shimer, 2005). Regarding vacancy posting costs, there is a relatively wide range of admissable values in the literature. Silva and Toledo (2009) estimate recruitment costs equal to 3.6% of a worker's monthly wage. Using microdata by Barron et al. (1997), Michaillat (2012) estimates the costs of posting a vacancy at 9.8% of a worker's steady state wage. Vacancy costs calibrated to match the latter value imply steady state vacancy posting costs of 0.28% of the total wage bill and 0.17% of GDP in Michaillat (2012). I calibrate $\kappa/2$ to 0.18, which is slightly more than 9% of a worker's steady state wage. Given this value, steady state vacancy posting costs account for 0.31% of the total wage bill and 0.2 % of GDP.

The efficiency of the matching function is chosen to match a quarterly job-finding rate of 0.8 and the elasticity of the matching function with respect to unemployment to match empirical evidence from Petrongolo and Pissarides (2001). Unemployment benefits are set to 0.4. This value implies a steady state replacement rate of about 0.2, which is at the lower end of the values found in the literature. The parameter φ, governing the disutility of labor, is set to match a steady state unemployment rate of 11%.[10]

Next, I calibrate the parameter governing wage rigidity based on the interpretation of the wage schedule arising from a staggered Nash bargaining setting. With this calibration strategy, τ can be interpreted as representing an upper bound on wage rigidity. Taylor (1999) argues that medium sized and large firms typically readjust wages anually. Additional evidence is provided by Gottschalk (2005), who finds that wage adjustments are most common one year after the last change. Thus, I set τ to 0.25, implying an average renegotiation frequency of once per year.

Since investment adjustment costs can potentially generate asymmetric unemployment dynamics, they are cautiosly set to $\xi = 0.05$, a value at the very low end of the values found in the literature.[11] The mean of the credit shock process, $\bar{\eta}$, is set to match the empirical ratio of outstanding debt in the non-financial corporate business sector to output of 1.75. The parameters for the persistence and standard deviation of the technology and credit shock sequence are estimated using the dataset constructed by Jermann and Quadrini (2012).[12]

[10] I choose a value twice the size of the actual unemployment rate over the considered time period as the model does not accout for workers that are out of the labor force. For similar reasons Barnichon (2010), Chugh (2013), Garín (2015), and Petrosky-Nadeau (2014) choose a steady state unemployment rate of 10%.

[11] A more detailed discussion of the role of investment adjustment costs is provided in Appendix A.5.

[12] For the estimation I use the code provided by Pfeifer (2016).

Table 2.1. Calibration of the Model Parameters

Symbol	Interpretation	Value	Source/Target
β_h	Household's discount factor	0.996	Annual interest rate of 1.6%
β_c	Firms' discount factor	0.983	Annual return on equity of 7%
x	Separation rate	0.1	Shimer (2005)
$\frac{\kappa}{2}$	Recruiting costs	0.18	Michaillat (2012)
ν	Matching efficiency	0.651	Job-finding rate of 0.8
γ	Unemployment-elasticity of matching	0.5	Petrongolo and Pissarides (2001)
s	Unemployment benefits	0.4	Replacement rate of 0.2
φ	Disutility of labor	0.85	Unemployment rate of 11%
τ	Renegotiation probability	0.25	Taylor (1999); Gottschalk (2005)
ξ	Investment adjustment cost	0.050	Lower end of literature values
$\bar{\eta}$	Steady state credit market tightness	0.3086	Debt-to-output ratio of 1.75
σ	Agents relative risk aversion	2	Standard in the literature
ϕ	Worker's bargaining power	0.4	Midpoint of literature values
α	Marginal returns to labor	0.66	Labor share of 0.66
ρ_z	Autocorrelation of technology shocks	0.9508	Jermann and Quadrini (2012)
σ_z	Standard deviation of technology shocks	0.0083	Jermann and Quadrini (2012)
ρ_η	Autocorrelation of credit shocks	0.9788	Jermann and Quadrini (2012)
σ_η	Standard deviation of credit shocks	0.0126	Jermann and Quadrini (2012)
δ	Capital depreciation rate	0.025	Jermann and Quadrini (2012)

2.4.2 Simulated Moments

I compare the simulated moments of the model to business cycle statistics for U.S. data. For the vacancy series I take data from Michaillat (2012), who merged the Job Openings and Labor Turnover Survey (JOLTS) for 2001–2004 with the Conference board help-wanted advertising index for 1964–2001. Unemployment data is taken from the Bureau of Labor Statistics (BLS) and labor market tightness is calculated as the ratio of vacancies to unemployment. For each of these series I take the quarterly average. The real wage estimates are average hourly earnings in the nonfarm business sector constructed by the BLS Current Employment Statistics. Output is quarterly real output from the BLS Major Sector Productivity and Costs program. In order to isolate business cylce fluctuations, I use a Hodrick-Prescott filter with smoothing parameter

Table 2.2. Summary Statistics – Quarterly US Data, 1964–2004

	u	v	θ	w	y	z
Standard deviation	0.166	0.186	0.339	0.021	0.030	0.020
Autocorrelation	0.918	0.946	0.934	0.949	0.902	0.890
Correlation	1	-0.888	-0.968	-0.114	-0.820	-0.514
	—	1	0.975	0.162	0.762	0.488
	—	—	1	0.140	0.810	0.514
	—	—	—	1	0.499	0.639
	—	—	—	—	1	0.883
	—	—	—	—	—	1

Note: All data are seasonally adjusted. The sample period is 1964:I - 2004:IV. The unemployment rate u is the quarterly average of the monthly series constructed by the BLS from the Current Population Survey (CPS). Vacancies are taken from Michaillat (2012) and constructed as detailed in the text. Labor market tightness θ is the ratio of vacancies to unemployment. The real wage is quarterly average hourly earnings in the nonfarm business sector, constructed by the BLS Current Employment Statistics program, and deflated by the quarterly average of the monthly Consumer Price Index for all urban households, constructed by the BLS; y is the quarterly real output in the nonfarm business sector constructed by the BLS Major Sector Productivity and Costs dataset; $ln(z)$ is constructed as a residual. Following Haefke et al. (2013), fluctuations in the capital stock are ignored. All variables are log deviations from an HP trend with smoothing parameter 10^5.

100.000 as recommended in Shimer (2005).[13] Table 2.2 displays the second order moments for key labor market variables that will be used to evaluate the performance of the model.

I simulate 264 quarters of data corresponding to the empirical sample size of 1964:I to 2004:IV.[14] The data is detrended using the same HP filter. The simulation is repeated 500 times and each repetition provides an estimate of the means of the simulated data. Standard deviations are calculated to judge the precision of the estimates. While the technology and credit shock processes are calibrated to match the empirical data, all other simulated moments are outcomes of the mechanics of the model. All simulations are performed using a second-order perturbation method. Since I am interested in asymmetric unemployment dynamics, a first order approximation is obviously not feasible. As the results remain virtually unchanged when using third- or fourth-order approximations, a second-order approximation seems to capture most of the relevant dynamics.

The model performs well along most dimensions that a model without financial frictions and without wage rigidity fails to capture.[15] While the standard deviation of unemployment

[13]The results remain virtually unchanged when using a smoothing parameter of 1600.
[14]The first 100 quarters are discarded as a burn-in period.
[15]The simulation results for a benchmark model without financial frictions and without wage rigidity are given in Appendix A.6.

Table 2.3. Simulated Moments – Financial Frictions and Wage Rigidity τ= 0.25

	u	v	θ	w	y	z
Standard deviation	0.080	0.164	0.227	0.011	0.021	0.015
	(0.018)	(0.017)	(0.029)	(0.003)	(0.004)	(0.002)
Autocorrelation	0.810	0.388	0.533	0.960	0.849	0.829
	(0.061)	(0.102)	(0.100)	(0.022)	(0.061)	(0.070)
Correlation	1	-0.782	-0.881	-0.494	-0.839	-0.747
		(0.021)	(0.016)	(0.090)	(0.054)	(0.094)
	—	1	0.975	0.223	0.624	0.576
			(0.004)	(0.073)	(0.051)	(0.072)
	—	—	1	0.330	0.726	0.665
				(0.081)	(0.052)	(0.082)
	—	—	—	1	0.820	0.788
					(0.040)	(0.037)
	—	—	—	—	1	0.973
						(0.015)
	—	—	—	—	—	1

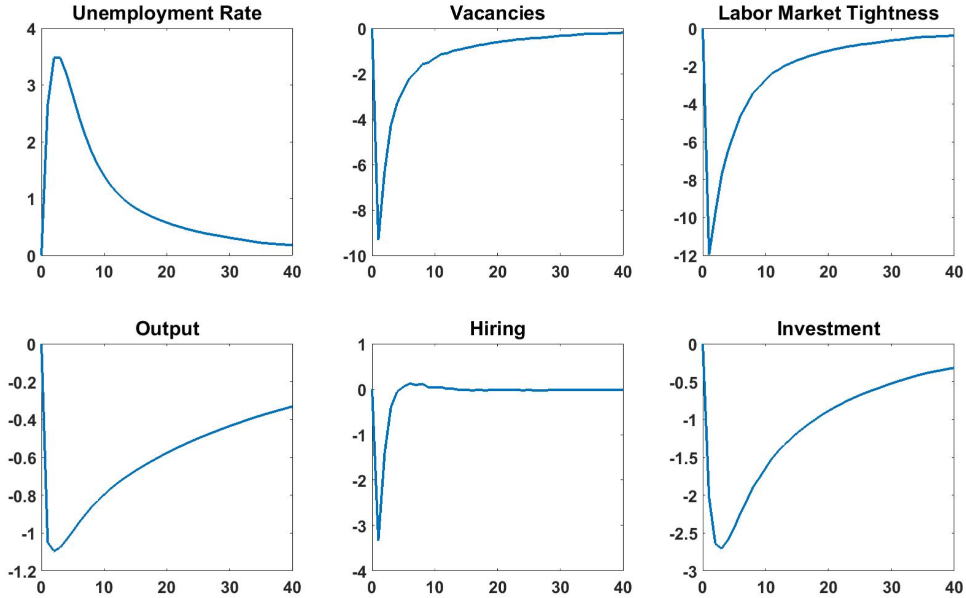
Note: Results from simulating the model with stochastic technology with a second-order perturbation method. All variables are log deviations from an HP trend with smoothing parameter 10^5. Simulated standard errors (standard deviations across 500 simulations) are reported in parentheses.

is still too low compared to U.S. data, it is about four times the standard deviation of output. In addition, the model accounts for roughly 70% of the volatility of labor market tightness, 90% of the volatility in vacancies, and nearly 100% of the fluctuations in the job-finding rate. Robustness exercises in the form of business cycle statistics for a model with financial frictions but without wage rigidity and for a model with wage rigidity but without financial frictions are provided in Appendix A.9. These simulations confirm that the interaction between wage rigidity and financial frictions, and not only the sum of the separate effects, plays an important role in matching business cycle statistics and in explaining unemployment dynamics.

Shocks are amplified considerably in the model: a 1% decrease in productivity increases unemployment by 3.9%, decreases vacancies by 6.2%, and decreases labor market tightness by 9.9%.[16] In the data, a 1% decrease in productivity increases unemployment by 4.2%, decreases

[16]The elasticity of unemployment with respect to technology ϵ_a^u is the coefficient obtained in an ordinary least squares regression of log unemployment on log technology. This coefficient can be calculated as $\epsilon_a^u = \rho(u, a) \times \sigma(u)/\sigma(a) = -0.514 \times 0.166/0.020$. All other elasticities can be calculated accordingly.

Figure 2.1. Impulse Response Functions: Negative Technology Shock



Note: The scale represents percentage deviations from the steady state. The size of the technology shock is one standard deviation.

vacancies by 4.5%, and decreases labor market tightness by 8.6%. The response of vacancies and labor market tightness is a bit higher in the model than in the data, which might be due to a lower elasticity of wages with respect to changes in technology. Haefke et al. (2013) find an elasticity of about 0.7, while the presented business cycle statistics for the U.S. suggest a value of 0.65. The simulated elasticity is a bit lower with a value of 0.58.

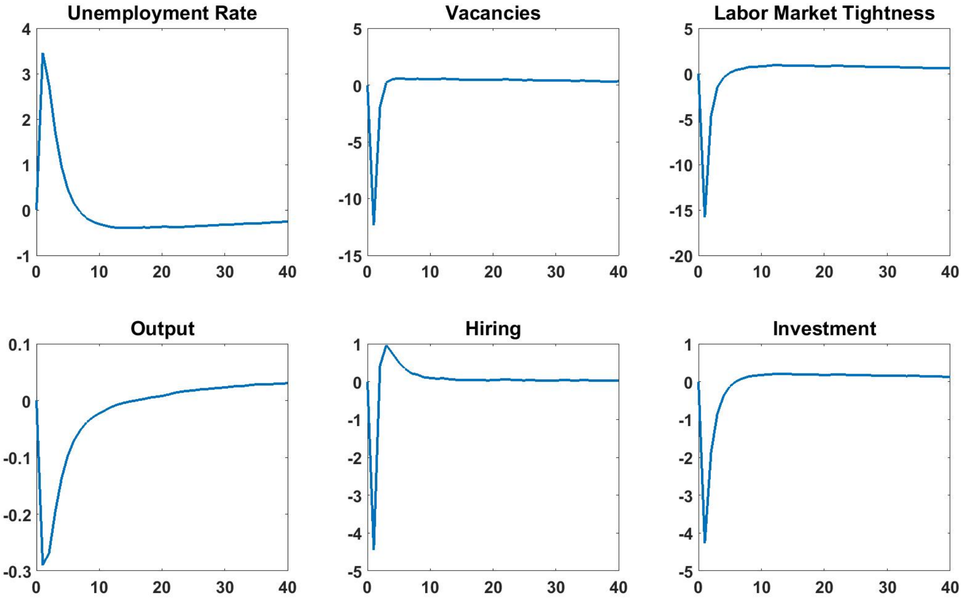
Comparing the elasticity of unemployment to technology in this model with the elasticity in a model with perfect credit markets, I find that the effect of wage rigidity is six times larger when firms are constrained in their ability to borrow. Additionally, the effect of financial frictions on the elasticity of unemployment to technology is more than twice as large under wage rigidity compared to a model with flexible wages. As they reinforce each other, the combined effect of wage rigidity and financial frictions is two times larger than the sum of the two separate effects.

2.4.3 Impulse Response Functions

In this section, I present the impulse response functions of several variables to a negative one standard deviation shock to total factor productivity and a negative one standard deviation shock to credit tightness. The scale represents log deviations from steady state. The impulse response functions for a positive and a negative shock are not symmetric for the unemployment rate. This asymmetry is discussed in detail in Subsection 2.4.4.

The impulse response functions for a negative shock to technology comply with the literature. Following a negative shock, firms decrease their hiring with vacancies dropping by nearly

Figure 2.2. Impulse Response Functions: Negative Credit Shock



Note: The scale represents percentage deviations from the steady state. The size of the credit shock is one standard deviation.

9% on impact. The unemployment rate increases, leading to an even larger decrease in labor market tightness. The marginal value and the collateral value of the capital stock drop, which triggers the decrease in investment. Note that the model captures unemployment dynamics quite well: after a negative technology shock unemployment peaks around four months after the initial shock. This is in line with the empirical findings in Stock and Watson (1999).

Figure 2.2 depicts the response of the model to a negative one standard deviation shock to credit market tightness. Firms are able to borrow less against their collateral and respond by cutting hiring and investment. This lowers the future capital stock and further tightens the credit constraint. The drop in vacancies is not persistent, but still large enough to generate a persistent increase of the unemployment rate. After dropping on impact, hiring increases above its steady state value long before output has recovered. This is in line with Blanchard and Diamond (1990), Fujita and Ramey (2006), and Elsby et al. (2009), who all document an increase in the number of hires in recessions. These dynamics are not present in standard search and matching models as the number of hires tends to follow production closely.

Note that neither technology nor credit shocks generate dynamics in the unemployment rate that are more persistent than output dynamics. Therefore, neither a simple shock to credit tightness nor a simple shock to total factor productivity is able to induce a jobless recovery.

2.4.4 Unemployment Dynamics

In this section, I turn to the asymmetric behavior of the cyclical component of the unemployment rate documented and analyzed in, for example, McKay and Reis (2008), Barnichon (2010), and

Table 2.4. Skewness of the Simulated Unemployment Rate

		$\tau = 1$	$\tau = 0.5$	$\tau = 0.25$
With Financial Frictions	Levels	0.42	0.44	0.56
	Changes	0.23	0.25	0.30
Without Financial Frictions	Levels	0.12	0.13	0.18
	Changes	0.02	0.03	0.05

Note: The amount of wage rigidity τ implies an average renegotiation frequency of three months, six months, and twelve months, respectively.

Atolia et al. (2018). Following Sichel (1993), I measure asymmetry in unemployment dynamics with the skewness coefficient.[17] For U.S. data in the time period between 1964 and 2004, the skewness of the unemployment rate is 0.72 in levels and 1.30 in changes.[18] These values suggest that the unemployment rate is characterized by short periods of sharp increases and long periods of flat decreases.

Table 2.4 displays the skewness of the simulated unemployment rate in levels and in changes for different amounts of wage rigidity.[19] A standard search and matching model with symmetric shocks is unable to match these observations despite the asymmetry resulting from costly vacancy posting. The simulated unemployment series in a benchmark model without financial frictions and without wage rigidity displays a skewness of 0.12 in levels and 0.02 in changes, explaining only about 17% and 2% of the respective skewness in the data.[20] For the model with financial frictions and with wage rigidity, the skewness of the simulated unemployment series is 0.56 in levels and 0.30 in changes. Over 75% of the skewness in levels and nearly 25% of the skewness in changes in the data can be explained by combining both frictions in a search and matching framework.[21]

As for the elasticity of unemployment to technology shocks, financial frictions and wage rigidity reinforce their respective effects. The combined effect of wage rigidity and financial frictions on the skewness in levels is 22% larger than the sum of the two separate effects. For the skewness in changes the combined effect is 17% larger.[22]

The mechanism generating asymmetry is intuitively simple and depends on both financial

[17] Positive skewness in levels implies that the unemployment rate is more often above than below its trend. Positive skewness in changes implies that the unemployment rate is more likely to decrease than to increase.

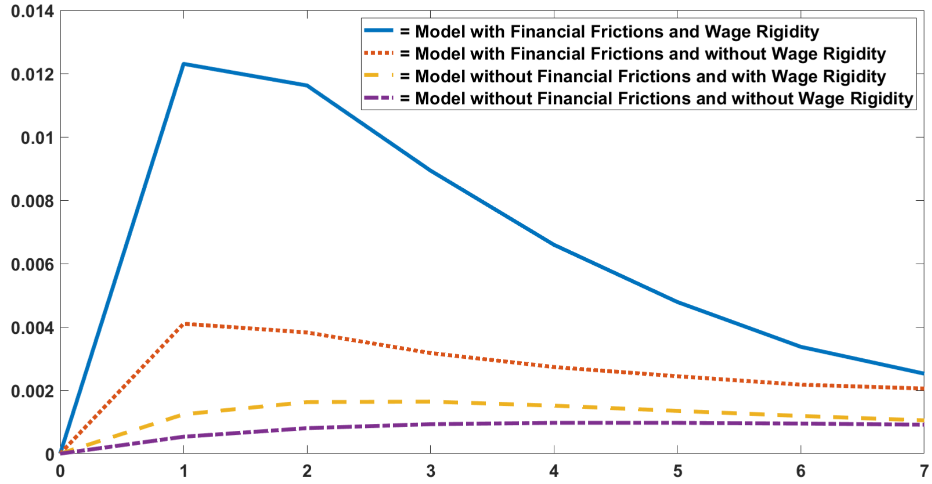
[18] The unemployment rate is the quarterly average of the monthly unemployment series constructed by the BLS from the CPS. The series is detrended using a HP filter with smoothing parameter 100.000.

[19] For this exercise τ is adjusted, leaving all other parameter values unchanged. Independent of the size of τ, the simulated volatility of output is always below the actual volatility of output. Thus, higher wage rigidity does not come at the cost of counterfactually large fluctuations in output.

[20] The results remain virtually unchanged when the model without financial frictions and without wage rigidity is calibrated to match the volatility of the unemployment rate in the model with financial frictions and wage rigidity instead of the same steady state targets.

[21] Other important aspects for explaining asymmetric unemployment dynamics might be demand shocks (cf. Barnichon, 2010), goods market frictions (cf. Petrosky-Nadeau and Wasmer, 2015), or creative destruction (cf.

Figure 2.3. Asymmetric Impulse Response Functions: Unemployment



Note: This figure plots the difference between the deviation of the unemployment rate from its steady state following a negative one standard deviation shock to technology and the absolute value of the deviation of the unemployment rate from its steady state following a positive one standard deviation shock to technology.

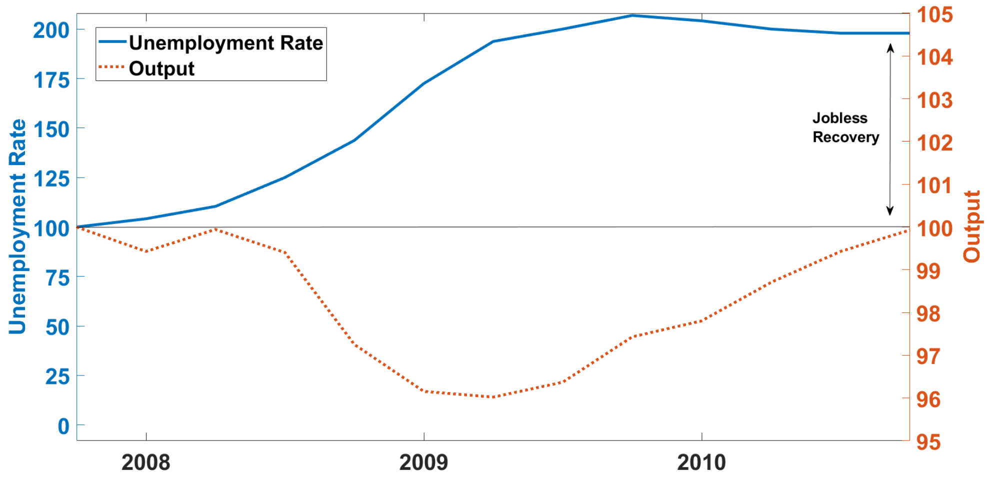
frictions and wage rigidity. With rigid wages, the response of the unemployment rate to negative shocks increases as the firm's surplus reacts more strongly to changes in technology and credit tightness. This implies steeper increases in the unemployment rate on impact compared to models with flexible wages. With financial frictions, a positive technology shock tightens the credit constraint as it increases working capital requirements. Firms invest in the asset used as collateral in order to loosen the borrowing constraint. The increase in employment is delayed and the decrease in unemployment is flatter compared to standard search and matching models.

Figure 2.3 illustrates the asymmetry of the unemployment rate based on the impulse response functions. The absolute value of the deviation of the unemployment rate from its steady state following a positive one standard deviation shock to technology is substracted from the deviation of the unemployment rate from its steady state following a negative shock to technology of the same size. A value larger than zero implies that the deviation of the unemployment rate is larger after a negative shock. The asymmetry is most pronounced for the model with financial frictions and wage rigidity. The stronger reaction of the unemployment rate visualizes the argument that firms proiritize investment into collateral following a positive technology shock. In contrast, the impulse response functions for output are almost completely symmetric.

McKay and Reis, 2008).

[22]In line with the data, simulated output displays no skewness in levels or in changes (cf. Barnichon, 2010).

Figure 2.4. Jobless Recovery During the Great Recession



Note: The unemployment rate is the quarterly average of the seasonally adjusted monthly unemployment series constructed by the BLS from the CPS. Output is the seasonally adjusted quarterly real GDP from the Bureau of Economic Analysis. For both series the value in 2007Q4 is normalized to 100.

2.5 Jobless Recoveries

The recessions and subsequent recoveries in 1990–1991, 2001, and more recently the recovery after the Great Recession in 2007–2009, have sparked a debate about so called jobless recoveries in the U.S. (cf. Galí et al., 2012; Shimer, 2012; Calvo et al., 2014; Jaimovich and Siu, 2018). Following the definition used in, among others, Calvo et al. (2014), I classify recoveries as jobless if the unemployment rate is above its pre-crisis level by the time output has fully recovered.[23]

Figure 2.4 depicts the joblessness of the recovery following the Great Recession. After the end of the Great Recession in June 2009 it took slightly less than two years for output to fully recover. At the point of output recovery the unemployment rate had only recovered by about 15% and was still about four percentage points above its pre-crisis level.

2.5.1 Mechanism

In my model, regardless of whether the recession is caused by a technology or by a credit shock, output and unemployment behave very similarly as long as only one shock drives the economy. Jobless recoveries emerge when credit conditions continue to erode while total factor productivity recovers. They are jobless because worsening credit conditions are an important driver of unemployment dynamics and keep unemployment high, but play only a minor role for fluctuations in output. In the variance decomposition exercise in Table 2.5, over 25% of the fluctuations in unemployment are caused by credit shocks, while virtually all of the fluctuations

[23]Note that in the model this statement is equivalent to employment being below its pre-crisis level at the point of output recovery.

Table 2.5. Variance Decomposition

	y	u	v	θ	w
TFP Shocks	98.61	73.87	56.08	63.77	99.35
Credit Shocks	1.39	26.13	43.92	36.23	0.65

Note: The variance decomposition is used to assess the relative importance of technology and credit shocks for generating volatility in the simulated model.

in output are due to productivity shocks.[24]

The mechanism behind this result is also evident in the first order conditions of the firm. Credit conditions directly affect the marginal value of an additional worker. A tightening of future credit conditions increases $\mu_{b,t+1}$, the Lagrange multiplier on the borrowing constraint, and thus the marginal value of relaxing this constraint. This increase in credit tightness reduces the marginal value of an additional worker by

$$L_{n,\mu_{b,t+1}} = -\alpha z_{t+1} n_{t+1}^{\alpha-1} k_{t+1}^{1-\alpha}.$$

Now consider the effect of the same increase in credit tightness on the marginal value of an additional unit of capital

$$L_{k,\mu_{b,t+1}} = -(1-\alpha) z_{t+1} n_{t+1}^{\alpha} k_{t+1}^{-\alpha} + \mathbb{E}_t \Lambda_{t|t+1}^{c} \eta_{t+1} q_{k,t+1}.$$

As for the marginal value of an additional worker, the increase in credit tightness reduces the marginal value of the capital stock in the production process. However, capital can also be used as collateral, the value of which increases with credit tightness. Thus, the effect of credit shocks on unemployment is larger than the effect of credit shocks on capital and output.[25] When total factor productivity recovers after a recession, firms increase their capital stock which in turn increases production. If this recovery is accompanied by tightening credit conditions, the marginal value of an additional worker stays low and at the point of output recovery the unemployment rate will be above its pre-crisis level.

[24]The findings in Garín (2015) suggest that the role played by credit shocks is even larger. The presence of wage rigidity in my model reduces the importance of credit shocks for explaining variation in the key labor market variables. Since wages are already relatively more rigid with respect to credit conditions, wage rigidity increases the relative importance of productivity shocks for fluctuations in the considered variables. Note, however, that the validity of results from a variance decomposition with only two shocks is limited.

[25]In this sense joblessness arises because labor cannot be used as collateral. This is in line with empirical evidence. Calvo (2015) shows that collateral requirements are lower for firms possesing easily recognizeable collateral and that the majority of this easily recognizeable collateral is physical capital.

Table 2.6. Credit Shocks During Recoveries

	Average Credit Shock
Recessions Prior to 1990	-0.00248
1990-1991 Recession	-1.3968
2001 Recession	-0.8237

Note: The average credit shock is the average size of the credit shocks in the four quarters following the end of a recession divided by the standard deviation of the credit shock series. The credit shock series is estimated from the dataset provided by Jermann and Quadrini (2012) for the time period between 1964 and 2010. The recession in 1980 is left out of the sample as the recovery period is overlaid by the start of the recession in 1981. The Great Recession is left out as the dataset only covers the time period up to the first quarter of 2010.

2.5.2 Empirical Evidence

Evidence for deteriorating credit conditions during jobless recoveries can be found in the dataset provided by Jermann and Quadrini (2012) as well as in the Senior Loan Officer Opinion Survey on Bank Lending Practices from the Federal Reserve Board.

Table 2.6 displays the average size of credit shocks following the end of a recession estimated from the dataset provided by Jermann and Quadrini (2012). A negative value implies a tightening of credit conditions. Prior to the two jobless recoveries in 1990–1991 and in 2001, credit conditions remained virtually constant immediately after the end of a recession.[26] Following the end of the recessions in 1990–1991 and in 2001, credit conditions continued to worsen with the average credit shocks amounting to -1.4 and -0.82 times the standard deviation.

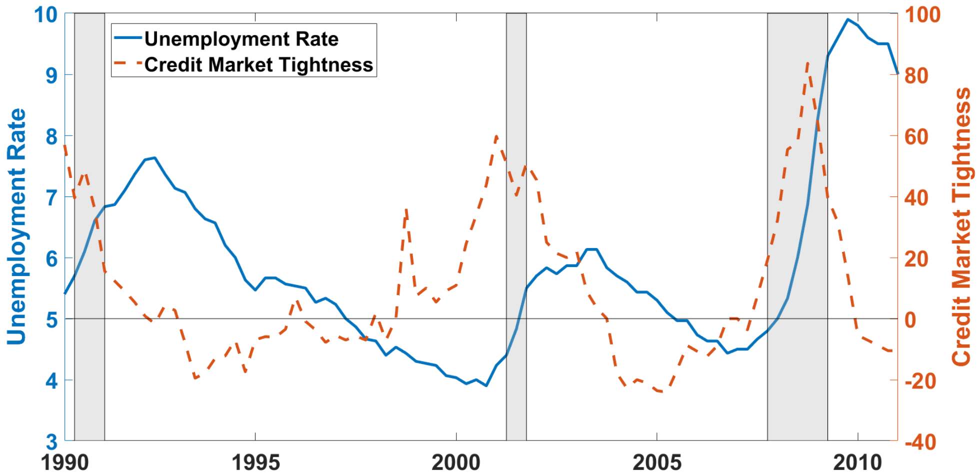
A tightening of credit conditions during the recent recoveries, including the recovery after the Great Recession, is also reported in the Senior Loan Officer Opinion Survey on Bank Lending Practices from the Federal Reserve Board depicted in Figure 2.5.[27] Credit market tightness is calculated as the fraction of surveyed banks reporting to tighten credit standards minus the fraction of banks reporting to lower their standards. A positive value therefore implies a tightening of credit conditions.

Two observations are striking. First, following the end of all three recessions, credit conditions continued to deteriorate – for several months after the recessions in 1990–1991 and in 2007–2009, and for nearly two years after the recession in 2001. Second, following the end of the recessions, the unemployment rate only began to decrease after credit conditions had stabilized.

[26] Prior to 1990 the largest negative average credit shock following a recession was one fourth of a standard deviation.

[27] The Senior Loan Officer Opinion Survey on Bank Lending Practices is available from 1990 onwards.

Figure 2.5. Unemployment Rate and Credit Market Tightness, 1990–2004



Note: The unemployment rate is the quarterly average of the seasonally adjusted monthly unemployment series constructed by the BLS from the CPS. Credit market tightness is measured as the Net Percentage of Domestic Respondents Tightening Standards for Commercial and Industrial Loans for Medium and Large Firms obtained from the Senior Loan Officer Opinion Survey on Bank Lending Practices from the Federal Reserve Board. Periods classified as recessions by the National Bureau of Economic Research (NBER) are highlighted in gray.

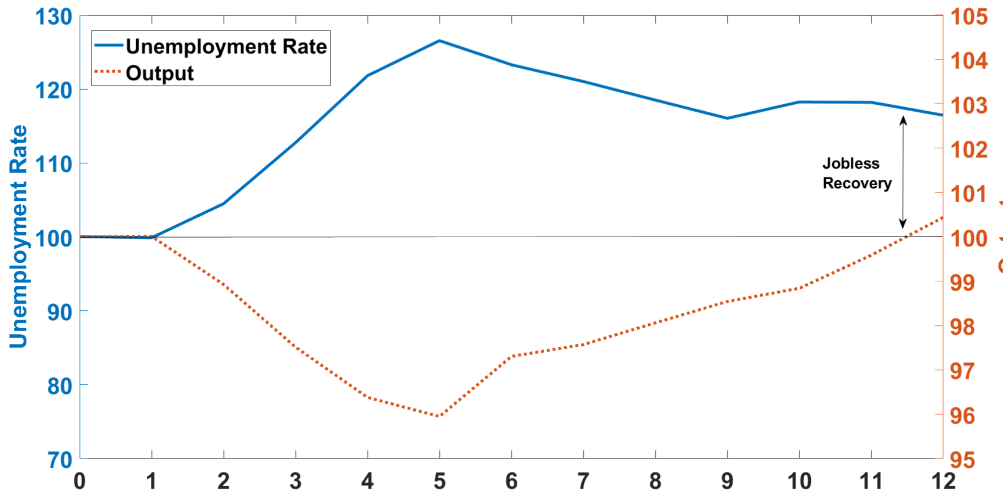
2.5.3 Simulated Recoveries

Figure 2.6 plots a jobless recovery in the model. Technology and credit shock series are calibrated to match the behaviour of output during the Great Recession. Negative shocks to credit tightness and to technology cause a recession in which output decreases by 4% during the first four quarters. Technology recovers but credit conditions continue to deteriorate. Firms invest into capital to offset the increasing credit tightness and output fully recovers after six to seven quarters. At the point of output recovery the unemployment rate has only recovered by around 31% relative to its peak (compared to a recovery by 15% after the Great Recession) and is two percentage points above its pre-crisis level (compared to four percentage points after the Great Recession).

In contrast to the results obtained by Calvo et al. (2014) in a competitive model of the labor market, wage rigidity is not necessary to generate joblessness in a DSGE model with financial frictions. Nonetheless, the joblessness of the recovery period is more pronounced under wage rigidity. Without wage rigidity, the unemployment rate would have recovered by about 45% at the point of output recovery.

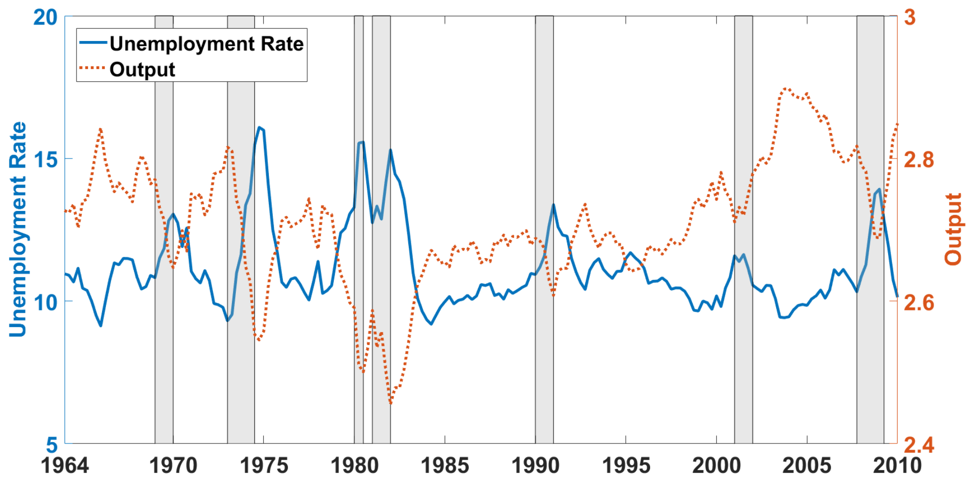
Next, I consider the ability of the model to account for jobless recoveries given technology and credit shocks estimated from the data. To that end I again use the dataset provided by Jermann and Quadrini (2012) to estimate credit and technology shock series for the time period between 1964 and 2010. Figure 2.7 plots the simulated series for unemployment and output. In line with the data, the model displays no sign of joblessness for the four recovery periods after recessions prior to 1990. At the point of output recovery after the recession in 1990–1991, unemployment is slightly above its pre-crisis level. The model predicts a more pronounced

Figure 2.6. Simulated Jobless Recovery



Note: The jobless recovery is generated using series of technology and credit shocks in the simulated model with financial frictions and wage rigidity. The shock series are calibrated to match the behaviour of output during the Great Recession. The pre-recession levels of unemployment and output are normalized to 100.

Figure 2.7. Simulated Unemployment and Output, 1964–2010



Note: Output and unemployment are simulated using technology shock and credit shock series estimated from the dataset to Jermann and Quadrini (2012). Periods classified as recessions by the NBER are highlighted in gray.

Table 2.7. Changes in Employment and Output: Model versus Data

		Output		Employment	
		Data	Model	Data	Model
1900–1991 Recession	1 year	-0.7%	-1.6%	-1.1%	-2.8%
	2 years	2.3%	0.5%	-0.5%	-0.6%
2001 Recession	1 year	1.5%	0.1%	-1.2%	-1.8%
	2 years	3.3%	1.8%	-0.3%	-0.7%
2007–2009 Recession	1 year	-3.3%	-2.5%	-1.7%	-2.1%
	2 years	-3.8%	-1.2%	-5.5%	-1.4%

Note: The growth rates are calculated by comparing peak output and employment at the start of the recession (or within one quarter) with output and employment one and two years later. Employment is total nonfarm payroll employment from the BLS Current Employment Statistics. Output is the seasonally adjusted quarterly real GDP from the Bureau of Economic Analysis.

joblessness for the recovery following the recession in 2001: at the point of output recovery the unemployment rate has only recovered by about 50%. Similarly, by the time simulated output recovered after the Great Recession, the model predicts a recovery of unemployment by 56% compared to its peak.

Finally, Table 2.7 displays the development of output and employment in the model and in the data for the three U.S. recessions after 1990. In line with the data, the model predicts employment to be below its peak value two years after the peak for all three recessions. Comparing the recent recessions with recessions prior to 1990, lower employment after two years is a unique feature of jobless recoveries.[28]

2.6 Conclusion

Incorporating financial frictions and wage rigidity considerably improves the performance of the standard search and matching model. Besides increasing the volatility of key labor market variables, the interaction of the two frictions facilitates the replication of important aspects of unemployment dynamics.

The simulated model with only technology and credit shocks can account for nearly 50% of the variation in unemployment, roughly 90% of the fluctuations in vacancies, nearly 70% of the variation in labor market tightness, and virtually 100% of the fluctuations observed in the job-finding rate. I find that wage rigidity is responsible for the steeper increase in the

[28] The only exception is the recession in 1980. However, the recovery period after this recession is overlaid by the beginning of the recession in 1981.

unemployment rate after negative shocks, whereas credit constraints ensure that decreases after positive shocks are flatter. Jobless recoveries are induced by eroding credit conditions during recovery periods.

While the explored mechanism provides an easy way to add important dynamics to search and matching models, it might also be able to reconcile models with endogenous separations with the highly negative correlation between unemployment and vacancies, i.e., the Beveridge curve. In models with endogenous seperations, the unemployment pool increases disproportionately after a negative technology shock due to the large inflow of separated workers. This decreases the labor market tightness and makes hiring in recessions cheap. Most models with endogenous separations therefore entail on-the-job search in order to reconcile the model with the data. In a model with endogenous separations and financial frictions, unemployment will also increase disproportionately after a negative credit shock. However, the incentive to post vacancies is reduced by a tightening of the borrowing constraint. It is an interesting task for further research to explore whether this mechanism is strong enough to generate a highly negative vacancy-unemployment correlation.

3 A Joint Theory of Polarization and Deunionization

Authors: Tobias Föll and Anna Hartmann

3.1 Introduction

Job market polarization and deunionization have radically changed the labor market over the last decades. Job market polarization refers to the falling employment shares in middle-skill occupations and increasing employment shares in low-skill and high-skill occupations. The share of employment in the middle range of skills has been continuously decreasing in the U.S. and is now more than 10 percentage points below its value in the 1980s (cf. Autor and Dorn, 2013). Deunionization describes the ongoing decline in union membership rates. According to the Union Membership and Coverage Database constructed by Hirsch and Macpherson (2003), U.S. union membership rates declined from 24.0% of all employed workers in 1973 to 10.5% in 2018.

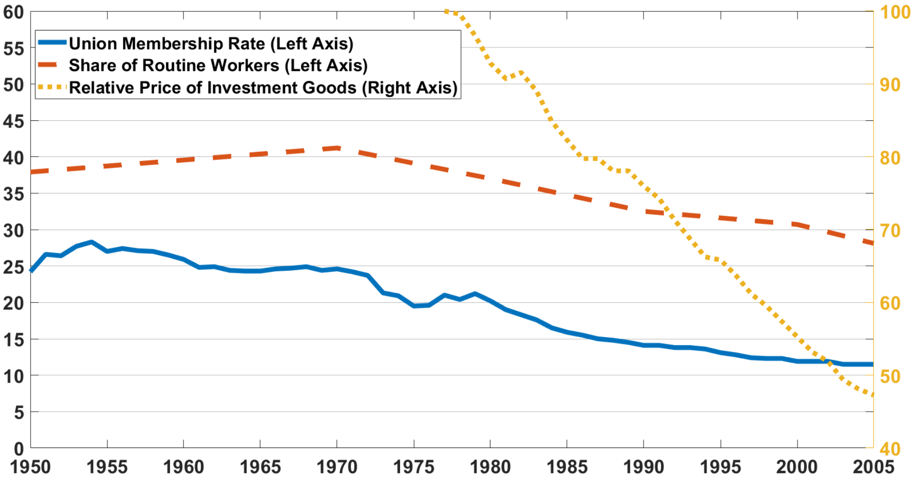
Polarization and deunionization have both proven to be especially harmful for low-wage and middle-wage workers: job market polarization because the relative shifts in labor demand away from routine occupations have suppressed wage growth in that area and deunionization because unionization rates are typically highest among lower middle-skill workers. American middle class workers have been in the focus for U.S. politicians not just since President Barack Obama declared himself "a warrior for the middle class".[1] Even though the share of U.S. households classified as middle class by the American Institute for Economic Research (AIER) has declined steadily since the 1980s, roughly 50% of households still counted as middle class in 2013.[2] Thus, identifying and implementing suitable policies to support the middle class has become an ever more pressing issue for today's policymakers, especially considering the recent trends of political radicalization among this group (cf. Post, 2017).

The prevalent explanation for polarization is the routinization hypothesis, which states that machines or computers replace middle-wage workers in occupations performing routine tasks (cf. Autor et al., 2003, 2006b; Autor and Dorn, 2013; Autor et al., 2015; Michaels et al.,

[1] Remarks by the president on the economy, Knox College, Galesburg, IL, 24.06.2013.

[2] The AIER defines households with a disposable income of two thirds to twice the median income for their household size as middle class.

Figure 3.1. Relative Price for Investment Goods, Share of Routine Workers, and U.S. Union Membership Rate



Note: The share of workers in routine occupations is constructed using the dataset and the occupational classification by Autor and Dorn (2013). Data for the union membership rates are taken from Mayer (2004), who merges data from the Current Population Survey, the Union Membership and Coverage Database constructed by Hirsch and Macpherson (2003), and from the Bureau of Labor Statistics Handbook and Employment and Earnings Survey. The membership rate includes all wage and salary workers. Public sector and agricultural workers are included in order for the data to be comparable to the data used in Autor and Dorn (2013). Missing data points are extrapolated from adjoining data points. The FRED series for the relative price of investment goods is measured as the investment deflator divided by the consumption deflator and displayed as an index with 1980 = 100. We display the relative price for investment goods rather than the price for computer capital since data on the former is more reliable and available for a longer time period.

2014; Feng and Graetz, 2015). The non-routine nature of tasks performed by low-wage and high-wage workers means that their jobs are more difficult to automate. In contrast to job polarization, no consensus has yet emerged regarding the source of deunionization (cf. Dinlersoz and Greenwood, 2016; Ortigueira, 2013; Aghion et al., 2011; Lee and Roemer, 2005). In this paper, we argue that routine-biased technical change is also the main driving force behind falling unionization rates. Figure 3.1 depicts the falling relative price for investment goods (proxying routine-biased technical change), the employment share of workers in routine occupations, and the union membership rate for the U.S. between 1950 and 2005. The union membership rate and the share of routine workers display a very similar negative trend over the last decades (with a correlation of 0.92).

To estimate the causal effect of routine-biased technical change on unionization, we borrow methodology from the trade (cf. Autor et al., 2013) and migration literatures (cf. Dustmann et al., 2017). Specifically, we use an interaction term between time-varying relative prices for investment goods and time-invariant state-specific routine employment shares in regressions of unionization rates that include both time and state fixed effects. Using state-level labor market data, we document that the effect of falling prices for investment goods on unionization rates is more pronounced in U.S. states with a larger share of workers employed in routine-intensive occupations. Thus, states that are more strongly affected by routine-biased technical change also experience larger declines in unionization rates. Additionally, and in contrast to conventional wisdom, we illustrate that the decrease in union membership is not mainly driven by changes in the industry or occupational composition.

Motivated by this, we develop a joint theory of polarization and deunionization. We endogenize both the occupational choice of workers, who differ with respect to their ability, and the union status of a firm in a search and matching model of the labor market. The occupational choice is modeled by giving previous routine workers the option to switch to low-skill manual occupations upon becoming unemployed. The union status of a firm is determined through an election, in which the employees decide whether they want to form a union, and consequently a collective bargaining unit, or whether they want to bargain individually over their wages with their employer.[3] If a union is established, it enters into a negotiation with the firm and distributes its share of the negotiated surplus equally across its members.

The main mechanism behind our results is quite simple. Relative prices for computer capital, which is able to replace routine tasks, fall (proxying for routine-biased technical change). This reduces the demand for routine workers, whereas manual and abstract workers, who are complementary to routine tasks, are in greater demand. The change in the labor demand structure implies that wages in manual occupations increase by more than wages in routine occupations. Manual workers, who benefit from the changing demand structure, are discouraged from voting in favor of a collective bargaining agreement because the lower demand for routine workers dampens the growth of union wages. The lowest-skilled previously unionized routine workers, when faced with lower wages compared to manual workers, decide to switch occupations. This amplifies the initial polarization caused by routine-biased technical change. Notably, our mechanism is in line with the empirical literature on union membership decisions. DiNardo et al. (1996) and Rueda et al. (2002), among others, document a decreasing effectiveness of unions in redistributing earnings over the last decades. Building on this argument, Baccaro and Locke (1998) and Checchi et al. (2010) both emphasize disillusion about potential wage growth as the main reason for declining union membership rates among the least skilled workers.

The increasing relative skill of union members (cf. Farber et al., 2018), the constant union wage premia over time (cf. Bryson, 2002; Hirsch and Schumacher, 2004; Breda, 2015; Farber et al., 2018), and the large contribution of within-industry and within-occupation changes to deunionization constitute three importamnt empirical observations that existing models of technical change and deunionization are unable to explain jointly (cf. Acemoglu et al., 2001; Açıkgöz and Kaymak, 2014; Dinlersoz and Greenwood, 2016).[4] We illustrate that all three of these observations can be rationalized in our model, where routine-biased technical change is driving deunionization. Low-skilled workers endogenously decide to vote against union coverage because of low wage growth in unionized firms. This leads to large within-industry

[3] A bargaining unit is commonly defined as a group of employees that shares a set of interests and may reasonably be represented by a collective bargaining agreement.

[4] Related literature is discussed in detail in Section 3.3.

and within-occupation changes in unionization rates and to a larger share of high-skilled union members. As only the weakest unions, i.e., those providing the lowest wage growth for their members, are terminated, average union wage premia are only mildly affected by declining membership rates.

We assess quantitatively the effect of routine-biased technical change on occupational decisions and on union formation. The model is calibrated to match U.S. data for the time period between 1983 and 2005. Predicted changes in the employment and wage distribution are close to the data. Routine-biased technical change, through changes in the labor demand structure, leads to a drop of 9.3 percentage points in overall union density in the model compared to a drop of 6.6 percentage points in the data. The falling union density amplifies polarization. As previously unionized routine workers are more likely to switch occupations when they are unable to find a routine job that is covered by a collective bargaining agreement, about 15% of the simulated changes in low- and middle-skilled employment are driven by deunionization.

In line with the empirical literature, the predicted overall effect of deunionization on inequality, measured by the Gini index, is small (cf. Frandsen, 2012; Checchi et al., 2010; DiNardo and Lee, 2004). However, deunionization has substantial effects for the lowest-skilled previously unionized routine workers. For this group of workers, wage growth in the model would be 60% larger if they were covered by one of the remaining collective bargaining contracts.

In our model, unions could dampen polarization and deunionization if they were able and willing to adjust the wage distribution, allowing for less equality inside the collective bargaining agreement. However, empirical evidence suggests that unions are characterized by rigid structures that partly prevent them from adjusting to recent developments on the labor market (cf. Waddington, 2005; Bryson et al., 2016). Bryson et al. (2016) argue that the decline in union membership rates across countries is strongly related to the degree of progressiveness of the unions. Thus, it seems that unions are lacking the modern and progressive structures necessary to attract more and especially younger members.

The remainder of the paper is organized as follows. Empirical evidence on job market polarization and deunionization is presented in Section 3.2 and previous research is discussed in Section 3.3. We give an overview of the union framework in the U.S. in Section 3.4. The model and analytical results are presented in Section 3.5. In Section 3.6 we provide a quantitative evaluation of the model. Policy implications are discussed in Section 3.7. To conclude, the results of this paper are summarized in Section 3.8.

3.2 Empirical Evidence

In this section, we present empirical evidence on the within-industry and within-occupation contribution to deunionization and on the relationship between polarization and deunionization.

3.2.1 A Decomposition Analysis

Conventional wisdom holds that the decline in unionization rates since the 1980s is mainly driven by a composition effect: routine-biased technical change reduces employment in the heavily unionized routine-manufacturing occupations while increasing employment in the less-unionized service and information technology occupations. We illustrate that changing employment shares between industries and between occupations contributed only little to declining union membership rates between 1983 and 2005, which are mainly driven by strong within-industry and within-occupation declines in unionization.

Borrowing the methodology used in, among others, Farber and Krueger (1992) and Baldwin (2003), we conduct a decomposition exercise to assess the relative importance of within- and between-industry and within- and between-occupation changes for deunionization. The within-industry (within-occupation) component measures the effect of a change in the union membership rate for a specific industry (occupational group), keeping the employment share in that industry (occupational group) constant. The between-industry (between-occupation) component measures the effect of a change in the employment share of a specific industry (occupational group), keeping the union membership rate in that industry (occupational group) constant. Summing up both components over all industries (occupational groups) yields the estimated overall change in the union membership rate.

For the analysis, we use data on industry-specific and data on occupation-specific union membership rates for several industries and occupational groups provided in the Union Membership and Coverage Database described in Hirsch and Macpherson (2003). The results are summarized in Table 3.1. Nearly 95% of the decline in unionization rates is accounted for by the within-industry component, with changing industry employment shares only contributing about 5%. These results are in line with previous empirical findings (cf. Baldwin, 2003). A similar picture emerges for the within- and between-occupation contribution to deunionization. Over 80% of the overall decline in unionization rates is driven by within-occupation declines in membership rates, with between-occupation changes accounting for less than 20%. When the occupational groups are reduced to abstract, routine, and manual, using the classification by Autor and Dorn (2013), the contribution of the between-occupation component drops to below 5%. Thus, contrary to conventional wisdom, deunionization is mainly driven by within-industry and within-occupation changes in union membership rates and not by simple composition effects.

3.2.2 Linking Polarization and Deunionization

A first look at the detailed statistics on union creation and union termination in the 20th century in Troy and Sheflin (1985) reveals that 1970 has been the year with the highest number of newly founded unions, while the most union terminations are observed in 1980. The accelerated decline in union membership rates in the late 1970s to early 1980s fits well with the documented

Table 3.1. Changing Unionization Rates – Decomposition, 1983–2005

Industry	Percentage Point	Share
Total Change	-9.18	100%
Within-industry	-8.70	94.87%
Between-industry	-0.47	5.13%

Occupation	Percentage Point	Share
Total Change	-11.01	100%
Within-occupation	-8.93	81.07%
Between-occupation	-2.08	18.93%

Note: Data for industry employment shares, occupational employment shares and union membership rates are taken from the Union Membership and Coverage Database constructed by Hirsch and Macpherson (2003). Industries include mining, construction, manufacturing, transportation, trade, services, finance, insurance, real estate, and public administration. Occupational groups include executive, managerial, professional, sales, machine operating, construction, transportation, and service.

starting point of job polarization.[5] Job polarization, and to a lesser extent also wage polarization, can be observed in the U.S. and several European countries at least since the 1980s (cf. Autor and Dorn, 2013; Goos et al., 2009). Additionally, and supporting our argument, Dinlersoz and Greenwood (2016) document that the steep decline in union membership rates in the 1980s followed the emergence and diffusion of early advanced technologies.

Cross-Country Evaluation

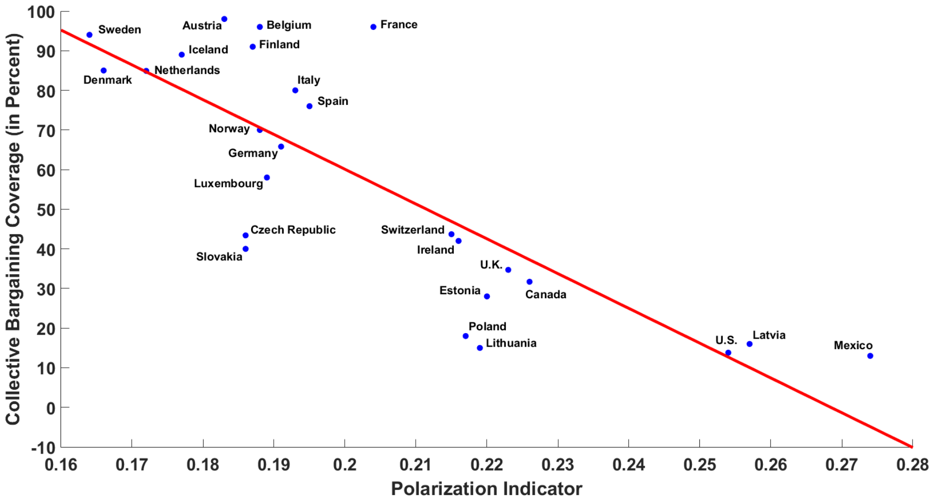
Looking at cross-sectional evidence, the degree of unionization is more pronounced in countries with larger degrees of job and wage polarization (cf. Meyer, 2019; von Brasch et al., 2018). This is visible when comparing the U.S. to Europe, but also within the group of European countries. The Nordic countries, which experienced upgrading rather than polarization, exhibit constant or even increasing union membership rates.[6]

Figure 3.2 plots the polarization indicator developed in Duclos et al. (2004), which evaluates

[5]The decline in union membership rates in the 1950s is usually explained by political resistance and the sharp increase in labor force participation of women, who tend to be less unionized (cf. Oh, 1989; Troy and Sheflin, 1985).

[6]The term upgrading refers to a specific pattern of changes in the employment structure, where employment growth is positively correlated with the required skill level.

Figure 3.2. Polarization and Collective Bargaining Coverage across Countries, 2004



Note: Figure 3.2 plots the polarization indicator developed in Duclos et al. (2004) against the collective bargaining coverage for the U.S., Canada, Mexico, and several European countries. Country selection is based on data availability. For all countries the polarization indicator is calculated for the year 2004. The collective bargaining coverage is the share of employed workers covered by a collective bargaining agreement in 2004 from the OECD data. The red line is the result of an OLS regression of the polarization indicator on the collective bargaining coverage. The regression coefficient is $\beta = -8.78$ and R^2 is 0.66.

the distance between and the distinction of income groups, against the collective bargaining coverage for the U.S., Canada, Mexico, and several European countries.[7] Despite the small sample size, the negative coefficient in an OLS regression of the collective bargaining coverage on the polarization indicator is statistically significant at the 0.1%-level.

Polarization and Deunionization Across U.S. States

Due to vast differences in the institutional frameworks of the considered countries, and due to the small number of countries for which reliable estimates can be obtained for the entire sample period, the previous results are only suggestive of a relationship between polarization and deunionization. We establish a causal relationship between routine-biased technical change and deunionization, using broad state-level labor market data for the U.S.

Data Sources We use labor market data from the Current Population Survey (CPS). Data on union membership and union coverage is taken from the CPS and from the the Union Membership and Coverage Database constructed by Hirsch and Macpherson (2003) using CPS data. For capital prices we use the Relative Price of Investment Goods, which is the investment deflator divided by the consumption deflator. For minimum wage laws we use the minimum

[7] In contrast to the U.S., the differences between union membership rates and the percentage of workers covered by a collective bargaining agreement are large for most of the European countries. Thus, when looking at union influence, the share of workers covered by a collective bargaining agreement seems to be more appropriate here. The results also hold when exchanging the collective bargaining coverage for union density. The results are very similar when using changes in collective bargaining coverage instead of collective bargaining coverage.

wage rates by state. Both series are taken from Federal Reserve Economic Data (FRED). Data on the federal intergovernmental revenue is taken from the State and Local Government Finance Dataset constructed by the Census Bureau. The tax burden is constructed by the Tax Foundation and calculated as the total amount of paid taxes divided by the state's total income. Data on state legislatures is obtained through the State Partisan Composition collected by the National Conference of State Legislatures.

Sample Selection We choose 1983 as the starting date for our analysis, as union membership estimates by detailed occupation are provided in the Union Membership and Coverage Database from this date onwards. 2005 is chosen as the endpoint because Beaudry et al. (2016) document a reversal in the demand for cognitive skills since the early 2000s and accounting for this reversal goes beyond the scope of our analysis.

An observation is a state-year combination, as union membership rates and detailed labor market data can only be constructed at the state level from the CPS. In principle, our sample thus contains 23 years × 50 states = 1150 state-year observations.[8] After excluding observations for which we lack information on certain control variables, we are left with a consistent sample of 1116 observations.

Methodology We estimate

$$u_{s,t} = \gamma \cdot p_{K,t} \cdot rsh_{s,83} + \beta \cdot X_{s,t} + \delta_s + \eta_t + \varepsilon_{s,t}, \tag{3.1}$$

where $u_{s,t}$ is the union membership rate or union coverage rate in state s in year t, $p_{K,t}$ is the relative price of investment goods in year t, and $rsh_{s,83}$ is the employment share in routine-intensive occupations in state s in year 1983.[9] $X_{s,t}$ is a vector of control variables, including controls for state policy (minimum wage laws, tax burden), state government and state legislation (party of governor, majority party in state senate and state house), state demographics (age, education, gender, ethnic composition), industry composition, and occupational composition.[10] The complete list of control variables is provided in Appendix B.7. δ_s and η_t are state and time fixed effects and $\varepsilon_{s,t}$ is the residual. Observations are weighted by the average state population over our sample period.

We explicitly address two potential concerns about our methodology. First, the effect of routine-biased technical change might work through changes in employment composition. To adress this concern, we run seperate regressions with and without controls for the industy and occupational composition in a state. Comparing these regressions allows us to gauge the relative importance of composition effects. Second, standard errors are not clustered as our

[8] The District of Columbia is excluded because of its specific labor market structure.

[9] Occupations are classified using the categorization in Autor and Dorn (2013).

[10] The state legislature in Nebraska is unicameral and officially non-partisan. However, since there has been a de facto Republican majority from 1983 to 2005, effects of this specific state legislature will be absorbed by the state fixed effect.

Table 3.2. Regression Results for Changes in the Routine Employment Share

	(1)	(2)	(3)	(4)
Initial routine employment share	-0.8430***	-0.8174***	-0.6639***	-0.7881***
	(0.1029)	(0.0845)	(0.2267)	(0.0896)
Observations	50	50	50	50
R^2	0.8960	0.8648	0.8306	0.6174
Industry and occupation controls	yes	yes	no	no
State policy controls	yes	no	yes	no
State legislation controls	yes	no	yes	no
State demographic controls	yes	no	yes	no

Note: Observations are weighted by the average state population over our sample period. The standard errors are reported in parentheses. *** indicates significance at the 1 percent level, ** indicates significance at the 5 percent level, and * indicates significance at the 10 percent level.

regressions include state fixed effects and there is no reason to expect heterogeneity in the sampling or in the treatment effects (cf. Abadie et al., 2017).

Results In a first step we confirm that the negative relationship between the initial employment share in routine-intensive occupations and the subsequent change in the share of routine-intensive occupations documented for U.S. commuting zones by Autor and Dorn (2013) holds on the state-level as well. Column (1) in Table 3.2 reports the results for our most preferred specification, including the entire set of controls. The other three columns illustrate that the results do not depend on the specific set of controls. In all four columns, the initial routine employment share in 1983 is highly predictive of the change in the routine employment share between 1983 and 2005. States with a higher initial routine employment share are the ones that experience a more pronounced employment polarization.

In a second step we use the interaction term between the time-invariant initial routine employment share and the time-variant relative price of investment goods in regressions with unionization rates as the dependent variable.[11] States with a larger initial employment share in routine-intensive occupations are more strongly affected by routine-biased technical change. Thus, a positive coefficient on the interaction term implies that routine-biased technical change (measured as the relative price of investment goods) causes deunionization.

[11] Several robustness checks are discussed in Appendix B.6.

Table 3.3. Regression Results for Unionization Rates

	(1)	(2)	(3)	(4)
Capital prices	0.3104***	0.4267***	0.2914***	0.3588***
× routine employment share	(0.0509)	(0.0516)	(0.0465)	(0.0459)
Observations	1116	1116	1116	1116
R^2	0.9870	0.9833	0.9864	0.9819
Industry and occupation controls	yes	yes	no	no
State policy controls	yes	no	yes	no
State legislation controls	yes	no	yes	no
State demographic controls	yes	no	yes	no
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: Observations are weighted by the average state population over our sample period. The standard errors are reported in parentheses. *** indicates significance at the 1 percent level, ** indicates significance at the 5 percent level, and * indicates significance at the 10 percent level.

The results are reported in Table 3.3. Column (1) constitutes our most preferred specification, featuring the full set of control variables. Column (2) excludes all control variables except the industry and occupation controls, Column (3) excludes industry and occupation controls, and Column (4) excludes all control variables. The coefficient on the interaction term is positive and highly statistically significant in all four specifications. This means that an increase in capital prices has a larger positive effect on unionization rates in states with a larger routine employment share. Consequently, following a decrease in capital prices, the fall in unionization rates is more pronounced in states with a high share of routine employment.

Quantitatively, the relative price of investment goods has dropped by 48% between 1983 and 2005. Consider two states that differ by ten percentage points in their routine employment share in 1983. When capital prices fall by 48%, our analysis suggests that the drop in the unionization rate will be about 1.6 percentage points larger in the state with the higher share of routine workers in 1983, controlling for both industry and occupational composition.

Columns (2) to (4), which leave out control variables, illustrate that our results do not depend on the set of controls. Specifically, the industry and occupation controls do not substantially change the size of the coefficient. Thus, the effect of routine-biased technical change on unionization rates across U.S. states is also not mailny driven by composition effects.

3.3 Related Literature

The evidence presented in the previous section motivates us to develop a joint theory of polarization and deunionization. We add to the literature by providing the first model of technical change and deunionization that is in line with the empirical literature on declining union membership rates (cf. Farber et al., 2018). Furthermore, our paper is the first to theoretically evaluate how routine-biased – as opposed to skill-biased – technical change affects union membership decisions.

Until now, technical change as a cause for deunionization has received only limited attention in the theoretical literature. Acemoglu et al. (2001) show that skill-biased technical change can trigger deunionization by increasing the outside option of skilled workers. In their model, deunionization is counterfactually entirely driven by quitting high-skilled workers: skill-biased technical change weakens the incentives for skilled workers to join the unionized sector, which they interpret as the manufacturing industry. Furthermore, the lower share of high-skilled workers in the unionized sector in Acemoglu et al. (2001) counterfactually implies declining union wage premia and less skilled union members over time.

Açıkgöz and Kaymak (2014) study deunionization in a search and matching framework with endogenous union membership. In their model, an exogenous increase in the skill premium encourages the most skilled workers to leave the union, while firms avoid to hire the least skilled union workers. This contrasts with evidence in Baccaro and Locke (1998) and Checchi et al. (2010), who argue that disillusion about potential wage growth is the main driving force behind declining union membership rates among the least-skilled workers. Additionally, and counterfactually, the decline in the union membership rate in Açıkgöz and Kaymak (2014) is stronger for high-skilled than for low-skilled workers, implying a decrease in the relative skill level of union members.

Dinlersoz and Greenwood (2016) focus on the connection between technology, unionization, and inequality. In a general equilibrium model of unionization with heterogeneous firms, skilled, and unskilled labor, they show that when the productivity of unskilled labor is high, unions decide to organize a lot of firms and demand generous wages for their members. Thus, skill-biased technical change leads to counterfactually declining union membership premia. While union members are exclusively drawn from low-skilled workers in Dinlersoz and Greenwood (2016), the inclusion of union members of other skill types would, as in Acemoglu et al. (2001), Açıkgöz and Kaymak (2014), and in basically any model of skill-biased technical change, lead to union members becoming less skilled over time.

3.4 Unions in the U.S.

In this section, we provide a brief overview of how labor unions work in the U.S. These institutional features will be used when setting up the model.

In the U.S., unions base their right to represent workers through collective bargaining on the voting decision of a so called *bargaining unit*. The National Labor Relations Act (NLRA) specifies the structure through which union organization and legal recognition takes place. This structure focuses on a system of petitions and elections to determine whether a majority of employees in the workplace wants to be represented by a union. The union then becomes the exclusive representative of all employees in the bargaining unit, whether they are union members or not. If a majority of the employees votes against union representation, the unit is not represented by the union, no matter if workers individually choose to be union members or not. In the event of a lawfully-called strike, unions are allowed under the NLRA to fine members that still decide to work.

An appropriate bargaining unit, according to the NLRA, is a group of employees in a workplace that meets the legal test of sufficient community of interest to be represented by the union, whereby managers and supervisors are excluded from any bargaining unit. According to the National Labor Relations Board (NLRB), professional employees who engage in predominantly intellectual and not in routine mental, manual, or mechanical work are excluded from bargaining units with manual and routine workers, since they do not share a community of interests.

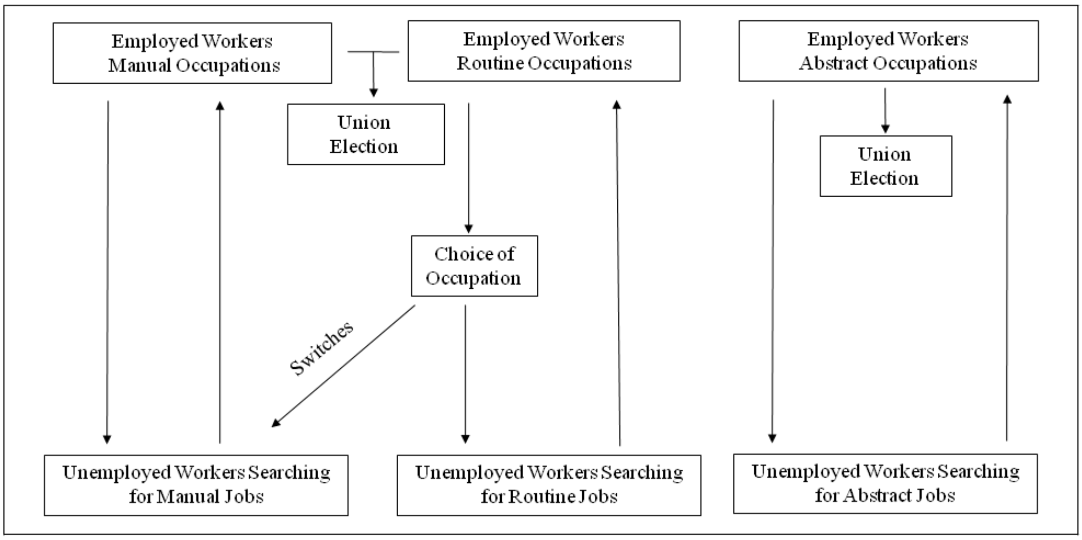
The structure of bargaining in the U.S. is highly decentralized with the estimated number of separate collective bargaining agreements ranging between 170,000 and 190,000 according to the Bureau of Labor Statistics. Most collective bargaining in the private sector takes place at the level of the individual firm (cf. Traxler, 1994; Nickell and Layard, 1999).

3.5 A Model of Occupational Decisions and Union Formation

In this section, we introduce labor unions into the multi-sectoral search and matching model developed by Albertini et al. (2017). There are two types of workers, abstract and non-abstract. Non-abstract workers are heterogeneous and differ with respect to their ability η, which is uniformly distributed. For each ability level, there is a continuum of workers. Abstract workers are assumed to be homogenous. As depicted in Figure 3.3, workers can be specialized in manual, routine, or abstract tasks. Upon becoming unemployed, workers previously employed in routine tasks can choose to switch occupations and join the unemployment pool of manual workers.[12]

[12] To ease notation, and in line with the empirical evidence in Smith (2013), we abstract from other switches. Thus, in our model there will be 'overqualified' routine workers in manual occupations, but we rule out the case of 'underqualified' manual workers in routine occupations and 'underqualified' routine workers in abstract occupations. Neither the results on deunionization nor the results on polarization depend on the assumption that manual workers are unable to switch to routine occupations. Note that because of falling prices for computer capital, the relative demand for manual workers increases. Thus, switches from manual to routine occupations only occur whenever the job-finding rate for routine workers is larger than the job-finding rate for manual workers in a unionized environment. These inefficient switches would only increase the speed

Figure 3.3. Graphical Representation of the Model



In our model, unions arise endogenously through elections within firms.[13] When a simple majority of the respective bargaining unit votes in favor of a union, wages are bargained collectively between the respective firm and the union. The collective bargaining agreement covers all workers in the bargaining unit regardless of the individual voting decision.

3.5.1 Labor Market Frictions

The labor market is characterized by search and matching frictions à la Mortensen and Pissarides (1994). Search is directed, as there are labor sub-markets for each of the three occupations $i = a, r, m$, where a, r, and m refer to abstract, routine, and manual occupations, respectively. Within each pool, vacancies and unemployed workers are matched randomly in any period and firms learn about the ability level of a worker upon matching. Given the number of vacancies v_i posted and the share of unemployed workers u_i for every occupation i, the number of matches is determined by a Cobb-Douglas matching technology with matching efficiency Ψ_i

$$m_i = \Psi_i v_i^{\psi} u_i^{1-\psi} \text{ where } 0 < \psi < 1 \text{ and } i = a, r, m.$$

Following Petrongolo (2001), constant returns to scale are assumed. A vacancy is filled with probability $q_i = \frac{m_i}{v_i}$ and the job finding probability is $f_i = \frac{m_i}{u_i}$. The labor market tightness is defined as the ratio $\theta_i \equiv \frac{v_i}{u_i}$. When the labor market is tight, many firms compete for few unemployed workers. The job finding probability is high, but the job filling rate is low.

with which deunionization occurs. Additionally, Smith (2013) shows that the increase in abstract employment is mainly driven by increased educational attainment and not by occupational switches. Thus, we let the labor supply of abstract workers increase exogenously in our model.

[13] Our production function features constant returns to scale. In contrast to Taschereau-Dumouchel (2017), firms have no incentive to overhire high-wage and low-wage workers and to underhire middle-wage workers in our model.

3.5.2 Occupational Choice

Workers can be employed in an abstract, a routine, or a manual occupation. Existing jobs are destroyed at the exogenous rates s_i, with $i = a, r, m$. The value function for unionized (superscript u) manual workers is given by

$$W_m^u(\eta) = w_m^u(\eta) + \beta[(1 - s_m)\big(\mathbb{1}_{u,+1} W_{m,+1}^u(\eta) + (1 - \mathbb{1}_{u,+1}) W_{m,+1}^n(\eta)\big) + s_m U_{m,+1}(\eta)],$$

where β is the discount factor and $w_m^u(\eta)$ denotes the wage received by a manual union worker with ability η. $\mathbb{1}_u$ is an indicator function with $\mathbb{1}_u = 1$ if and only if the worker is a union member. Thus, the term $\mathbb{1}_{u,+1}$ indicates whether a worker in the firm is covered by a collective bargaining regime in the next period.

In turn, the non-union (superscript n) manual workers' value function is given by

$$W_m^n(\eta) = w_m^n(\eta) + \beta[(1 - s_m)\big(\big(\mathbb{1}_{u,+1} W_{m,+1}^u(\eta) + (1 - \mathbb{1}_{u,+1}) W_{m,+1}^n(\eta)\big) + s_m U_{m,+1}(\eta)],$$

where $w_m^n(\eta)$ is the wage received by a manual non-union worker with ability η.

When unemployed, workers lose their union membership.[14] Therefore, the union and non-union value functions for an unemployed manual worker are identical and given by

$$U_m(\eta) = z_m(\eta) + \beta[(1 - f_m) U_{m,+1} + f_m\big(\mathbb{1}_{u,+1} W_{m,+1}^u(\eta) + (1 - \mathbb{1}_{u,+1}) W_{m,+1}^n(\eta)\big)],$$

where $z_m(\eta)$ denotes the unemployment benefits received from the government by a manual worker with ability η.

Analogously, the value functions for abstract workers and routine workers are

$$\begin{aligned}
W_a^u &= w_a^u + \beta[(1 - s_a)\big(\mathbb{1}_{u,+1} W_{a,+1}^u + (1 - \mathbb{1}_{u,+1}) W_{a,+1}^n\big) + s_a U_{a,+1}], \\
W_a^n &= w_a^n + \beta[(1 - s_a)\big(\mathbb{1}_{u,+1} W_{a,+1}^u + (1 - \mathbb{1}_{u,+1}) W_{a,+1}^n\big) + s_a U_{a,+1}], \\
U_a &= z_a + \beta[(1 - f_a) U_{a,+1} + f_a\big(\mathbb{1}_{u,+1} W_{a,+1}^u + (1 - \mathbb{1}_{u,+1}) W_{a,+1}^n\big)]
\end{aligned}$$

and

$$\begin{aligned}
W_r^u(\eta) &= w_r^u(\eta) + \beta\left[(1 - s_r)\big(\mathbb{1}_{u,+1} W_{r,+1}^u(\eta) + (1 - \mathbb{1}_{u,+1}) W_{r,+1}^n(\eta)\big)\right] \\
&\quad + \beta s_r \max\{U_{m,+1}(\eta), U_{r,+1}(\eta)\}, \\
W_r^n(\eta) &= w_r^n(\eta) + \beta\left[(1 - s_r)\big(\mathbb{1}_{u,+1} W_{r,+1}^u(\eta) + (1 - \mathbb{1}_{u,+1}) W_{r,+1}^n(\eta)\big)\right] \\
&\quad + \beta s_r \max\{U_{r,+1}(\eta), U_{m,+1}(\eta)\}, \\
U_r(\eta) &= z_r(\eta) + \beta[(1 - f_r) \max\{U_{m,+1}(\eta), U_{r,+1}(\eta)\} + f_r\big(\mathbb{1}_{u,+1} W_{r,+1}^u(\eta) \\
&\quad + (1 - \mathbb{1}_{u,+1}) W_{r,+1}^n(\eta)\big)].
\end{aligned}$$

[14] This is in line with Lewis (1989), who finds that unions are not perceived to represent the interests of the unemployed.

The term max $\{U_{m,+1}(\eta), U_{r,+1}(\eta)\}$ governs the occupational choice of routine workers. Whenever the value of being an unemployed manual worker is larger than the value of being an unemployed routine worker, the worker switches occupations. Thus, the equation defining the endogenous occupational threshold between manual and routine occupations, η_m, is given by

$$U_r(\eta_m) = U_m(\eta_m). \tag{3.2}$$

3.5.3 Firms

The model features a continuum of final good firms and intermediate firms. As the setup admits the presence of a representative firm on each level, firm indices are dropped. To further ease notation, we only use indices related to the union status of a firm when they are necessary to understand the model mechanics.

The final good-producing firm uses three homogeneous intermediate goods, Z_a, Z_r, and Z_m, as input factors to produce the final product Y. Intermediate goods are acquired at their competitive factor prices.[15] Z_a is produced with abstract jobs L^a, Z_r with computer technology K and routine workers $L^r(\eta)$, and Z_m with manual jobs $L^m(\eta)$. Routine workers and computer technology K are close substitutes, whereas abstract workers are complementary to the intermediate good Z_r. The maximization problem of the goods-producing firm is given by[16]

$$\begin{aligned} \Pi = \max_{Z_a, Z_r, Z_m} & \{Y - p_{Z_a} Z_a - p_{Z_r} Z_r - p_{Z_m} Z_m\} \\ \text{s.t. } & Y \leq [(A Z_a^{\alpha} Z_r^{1-\alpha})^{\rho} + (A_m Z_m)^{\rho}]^{1/\rho}, \end{aligned}$$

where $0 < \alpha < 1$, $-\infty < \rho < 1$, A, and A_m are parameters of the production function.

Intermediate firms maximize profits by choosing employment next period and the number of vacancies to be posted subject to the firm-level employment constraint. Job creation comes at a flow cost of c_a, c_r, or c_m. The behavior of the intermediate firm in producing the intermediate good Z_a, which is paid at price p_{Z_a}, is described by

$$\Pi^{Z_a} = \max \left\{ p_{Z_a} Z_a - \mathbb{1}_u w_a^u L_a - (1 - \mathbb{1}_u)\, w_a^n L_a - c_a v_a + \beta \Pi_{+1}^{Z_a} \right\}$$

$$\begin{aligned} \text{s.t. } & Z_a \leq L_a \\ & L_{a,+1} = (1 - s_a) L_a + q_a v_a, \end{aligned}$$

[15] This production structure is chosen in order to facilitate representation, as it allows for solving the maximization problems of the good-producing firm and the intermediate firms consecutively. The job creation conditions are identical if we instead assume that the good-producing firm directly uses manual, routine, and abstract workers as input factors.

[16] A nested production function is chosen in order to allow for larger complementarity in production between abstract and routine than between routine and manual tasks.

where $L_{a,+1}$ denotes the total abstract workforce next period. $\mathbb{1}_u$ is again an indicator function with $\mathbb{1}_u = 1$ indicating if the workforce in the firm is covered by a collective bargaining regime.

The behavior of the firm producing the intermediate good Z_r, which is paid at price p_{Z_r}, is described by

$$\Pi^{Z_r} = \max \left\{ p_{Z_r} Z_r - p_K K - \mathbb{1}_u \int_{\eta_m}^{\bar{\eta}} w_r^u(\eta) L_r(\eta) - (1 - \mathbb{1}_u) \int_{\eta_m}^{\bar{\eta}} w_r^n(\eta) L_r(\eta) - c_r v_r + \beta \Pi_{+1}^{Z_r} \right\}$$

$$\begin{aligned} \text{s.t. } Z_r &\leq \left[\left((1-\mu) \int_{\eta_m}^{\bar{\eta}} \eta L_r(\eta) \,\mathrm{d}\,\eta \right)^{\sigma} + (\mu K)^{\sigma} \right]^{\frac{1}{\sigma}} \\ L_{r,+1} &= (1 - s_r) L_r + q_r v_r, \end{aligned}$$

where $0 < \mu < 1$ and $-\infty < \sigma < 1$ are production parameters, $\bar{\eta}$ denotes the upper bound on the ability distribution for non-abstract workers, and η_m the endogenous ability threshold between manual and routine workers. Following Albertini et al. (2017), firms can freely choose their desired level of computer capital K at the price p_K.

The behavior of the intermediate firm in producing the intermediate good Z_m, which is paid at price p_{Z_m}, is described by

$$\Pi^{Z_m} = \max \left\{ p_{Z_m} Z_m - \mathbb{1}_u w_m^u L_m - (1 - \mathbb{1}_u)\, w_m^n L_m - c_m v_m + \beta \Pi_{+1}^{Z_m} \right\}$$

$$\begin{aligned} \text{s.t. } Z_m &\leq L_m \\ L_{m,+1} &= (1 - s_m) L_m + q_m v_m. \end{aligned}$$

As in Autor and Dorn (2013), workers in manual occupations are homogenous with respect to their productivity in performing manual tasks. This implies that wages for manual workers are constant, while wages for routine workers are increasing in the skill level η. Combining this with the definition of η_m in equation (3.2), it is straightforward to see that workers with an ability level lower than η_m work in manual occupations. The first order conditions and the job creation conditions are derived in Appendix B.1 and Appendix B.2.

3.5.4 Wage Bargaining Regimes

Since we focus on the U.S., we want our union framework to be as close as possible to the institutional framework presented in Section 3.4. Workers can decide to form a union on the level of the good-producing firm, which bargains with the firm and distributes the surplus according to a union wage schedule. Once new workers are hired, all workers vote to decide whether to form a union or not. Abstract workers are excluded from the collective bargaining unit with manual and routine workers. Thus, our model features two types of unions: one industrial union - aiming to cover workers of two different skill groups - and one craft union, covering only abstract workers. If a union is established, the collective bargaining agreement covers all workers in the bargaining unit, regardless of whether or not the individual worker voted in favor of the union. The voting decision of an individual worker is endogenously determined and depends directly upon the potential union wage premium. Workers vote in favor of a union if the value of being a worker in a unionized firm is higher than the value of being a worker in a non-unionized firm

$$W_i^u(\eta) > W_i^n(\eta), \text{ with } i = a, r, m.$$

In the model, the number of voting thresholds above or below which workers in a bargaining unit vote against the union depend on the choice of the union wage schedule. The thresholds are denoted by η_l^u and $\eta_l^{u,a}$ with $l \in [1, 2, ...]$, where the superscript a denotes the union for abstract workers.

If a majority of the bargaining unit votes against a collective bargaining agreement, workers in this bargaining unit are not represented by the union and wages are negotiated individually. Union and non-union wages are both determined by generalized Nash bargaining over the match surplus. However, the surplus that is bargained over differs between the two bargaining regimes: non-union workers bargain individually over their marginal product, whereas the union bargains over the total match surplus of all workers in the bargaining unit.

Individual Bargaining

If a majority of the manual and routine workers votes against a union, each worker bargains individually with the firm. Denoting the worker's weight in the bargaining process by $\gamma^i \in [0, 1]$, this implies the following sharing rule for individual bargaining

$$\begin{aligned} W_i^n(\eta) - U_i(\eta) &= \frac{\gamma^i}{1-\gamma^i} J_i^n(\eta), \\ \text{with } i &= a, r, m, \end{aligned}$$

where $W_i^n(\eta)$ is the asset value of employment for non-union members, $U_i(\eta)$ is the value of being unemployed, and $J_i^n(\eta)$ is the value of the marginal non-union worker of type i and ability η to the firm. The resulting wage schedules are

$$w_a^n = \gamma^a p_{Z_a} + \gamma^a c_a \theta_a + (1-\gamma^a)\, z_a \tag{3.3}$$

for workers in abstract jobs,

$$w_r^n(\eta) = \gamma^r p_{Z_r} y_r(\eta) + \gamma^r c_r \theta_r + (1 - \gamma^r)\, z_r(\eta) \tag{3.4}$$

for workers in routine jobs, and

$$w_m^n = \gamma^m p_{Z_m} + \gamma^m c_m \theta_m + (1 - \gamma^m)\, z_m(\eta) \tag{3.5}$$

for workers in manual jobs.[17]

It follows that the wages resulting from individual bargaining are given by the sum of the marginal productivity of the workers in each occupation, the search returns, and the outside option.

Collective Bargaining

We consider unions which negotiate wages on behalf of all covered workers within a firm and thus bargain over the total surplus of all union members. We make the following assumptions based on the union framework in the U.S. outlined in Section 3.4:

Assumption 1. *All workers that are covered by a collective bargaining agreement are union members.*

Assumption 2. *The union can force all of its members to strike.*

Under these assumptions, if no agreement on wages can be reached, all members of the respective bargaining unit in the unionized firm go on a strike and the firm can only produce using the remaining workers and computer capital.

Our approach only pins down the share of the surplus going to the workers, not how it is distributed among them. It is well-established in the literature that unions induce wage compression, that individual union wage premia decrease in the skill level of the worker, and that craft unions tend to negotiate higher union wage premia compared to industrial unions (cf. Card et al., 2004; Streeck, 2005). To keep the degrees of freedom in choosing the wage schedule small, we assume the simplest wage schedule that is in line with these observations: unions set a constant wage for all workers in the bargaining unit.[18] This accords with evidence in Fitzenberger et al. (2006), who show that unions tend to prefer wage equality over higher average wages. It follows that union wages are given by

$$w^u = S^u / (L_m + L_r) \tag{3.6}$$

[17] See Appendix B.3 for a detailed derivation of the wage schedules.

[18] The evaluation in Appendix B.5 establishes that the main mechanism behind falling union membership rates in our model is robust to alternative union wage schedules.

and

$$w_a^u = S_a^u / L_a. \tag{3.7}$$

Industrial Union

Under collective bargaining the outside option of a union member is not the value of being unemployed but the value of being a union member during a strike. Therefore, denoting the industrial union's weight in the bargaining process by $\gamma^u \in [0,1]$, the following surplus sharing rule holds in the case of collective bargaining

$$\max_{w^u} \left(\sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) \left[W_i^u(\eta) - W_i^{u,s}(\eta) \right] \mathrm{d}\,\eta \right)^{\gamma^u} \left(\sum_i \left\{ p_{Z_i} Z_i - p'_{Z_i} Z'_i - \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^u(\eta) \,\mathrm{d}\,\eta \right\} \right)^{1-\gamma^u}$$
$$\text{with } i = r, m,$$

where $W_i^u(\eta)$ is the asset value of employment for manual and routine union members with productivity η and $W_i^{u,s}(\eta)$ is the value of being a union member during a strike. Z_i is the production of the manual or routine intermediate good and Z'_i is the production in the manual or routine sector when workers are on a strike, which is compensated at price p'_{Z_i}.

It follows that the total surplus received by the industrial union S^u is given by [19]

$$S^u = \gamma^u \sum_i (p_{Z_i} Z_i - p'_{Z_i} Z'_i) + (1-\gamma^u) \sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w^{u,s} \,\mathrm{d}\,\eta \tag{3.8}$$
$$\text{with } i = r, m,$$

where $w^{u,s}$ denotes the wage received by a worker during a strike, regardless of occupation and ability. Note that the total surplus of the industrial union is a function of the productivity of all manual and routine workers, while the non-union wage is a function of the individual productivity of the respective worker.

Craft Union

Analogously, denoting the craft union's weight in the bargaining process by $\gamma_a^u \in [0,1]$, the following surplus sharing rule holds in the case of collective bargaining

$$\max_{w_a^u} \left(L_a \left[W_a^u - W_a^{u,s} \right] \right)^{\gamma^u} \left(p_{Z_a} Z_a - p'_{Z_a} Z'_a - L_a w_a^u \right)^{1-\gamma^u},$$

where W_a^u is the asset value of employment for craft union members and $W_a^{u,s}$ is the value of being a union member during a strike. Z_a is the production of the abstract intermediate

[19] See Appendix B.4 for a detailed derivation.

good and Z'_a is the production in the abstract sector when workers are on a strike, which is compensated at price p'_{Z_a}.

Thus, the total surplus received by the craft union S^u_a is given by

$$S^u_a = \gamma^u_a (p_{Z_a} Z_a - p'_{Z_a} Z'_a) + (1 - \gamma^u) L_a w^{u,s}_a. \tag{3.9}$$

3.5.5 Households, Government Expenditures, and Transfers

In the model, there is one household for each occupation and for each employment status, i.e., employed and unemployed. Households own the firm and consume the final good Y. There are no savings. For each worker the budget constraint is given by

$$C = I$$

with $I \in \{w^n_a, w^n_r(\eta), w^u_r, w^n_m, w^u_m, z_a, z_r(\eta), z_m(\eta)\}$.

Since the government pays out unemployment benefits, government expenditures are

$$G = z_a u_a + \int_{\underline{\eta}}^{\bar{\eta}} (z_r(\eta) u_r + z_m(\eta) u_m) \,\mathrm{d}\,\eta.$$

Firms can generate profits, which are given by

$$\Omega = \Pi^{Z_a} + \Pi^{Z_r} + \Pi^{Z_m}.$$

Transfers received by households are therefore

$$\Gamma = -G + \Omega.$$

Total consumption in the economy is then given by the sum of individual wages, individual benefits, and the transfers.[20]

3.5.6 Equilibrium

With the model completely described, we define the equilibrium.

Definition 2. *An equilibrium is defined as a set of i) firms' policy functions; ii) households' policy functions; iii) a union wage schedule; iv) prices; and v) a law of motion for the aggregate states, such that: i) for each firm the firm's policies satisfy the firms' first order conditions and the job creation conditions; ii) for each household the households' policy functions satisfy the households' first order conditions; iii) the wage is determined through individual or collective bargaining; iv) the choices given the aggregate states clear the markets; and v) the law of motion for the exogenous*

[20] This allows us to abstract from the distribution of transfers to households. The results remain unchanged when lump-sum transfers are assumed instead.

aggregate states is consistent with individual decisions and with the process for computer capital prices.

3.5.7 Effects of Routine-Biased Technical Change

It is well-established in the literature, that routine-biased technical change generates polarization in models of the labor market (cf. Autor and Dorn, 2013; Albertini et al., 2017). In our model, polarization is driven by occupational switches from previous routine workers to manual occupations. This result is formalized in Proposition 3.

Proposition 3. *Routine-biased technical change increases the incentives for previous routine workers to switch to manual occupations if $\sigma > 0$ and $\sigma > (1-\alpha)\rho$.*

Proof. See Appendix B.5 for a proof of Proposition 3. □

Thus, our model features polarization, as long as σ, the elasticity of substitution between computer capital and routine labor, is large enough. Intuitively, in order for routine-biased technical change to increase the incentives for occupational switches, capital and routine tasks need to be substitutes and they need to be better substitutes than routine and manual tasks in the production of the final good.

Routine-biased technical change, by increasing the capital stock, raises the productivity of manual workers by more compared to the productivity of routine workers. This leads to higher relative wages and job-finding rates for manual workers. Thus, the incentives for previous routine workers to switch to manual occupations increase. We add to this well-known result by demonstrating that routine-biased technical change additionally leads to deunionization in our model. Proposition 4 summarizes the main mechanism.

Proposition 4. *Routine-biased technical change reduces the incentives for manual workers to vote in favor of union coverage if the intermediate good produced by abstract labor, routine labor, and computer capital is a substitute to the intermediate good produced by manual labor, i.e. $\rho > 0$.*

Proof. See Appendix B.5 for a proof of Proposition 4. □

Intuitively, falling computer capital prices imply lower marginal costs of production. This increases the demand for workers in all three occupations. However, because of the complementarity of computer capital and routine workers in production, there is a negative substitution effect that reduces the demand for routine workers. Their marginal productivity increases by less than the marginal productivity of manual workers. Thus, the non-union wages of manual workers increase by more than the non-union wages of routine workers. The increasing relative demand for manual workers in response to the drop in the price of computer capital increases the size of the surplus the union can extract, while the negative substitution effect on the relative demand for routine workers tends to work in the opposite direction. Since unions set identical wages for manual and routine workers, routine workers benefit from the higher relative demand for manual workers, while manual workers suffer from the lower relative demand for routine workers, i.e. in the union the positive demand effect for manual workers is partially absorbed by routine workers. This directly implies that non-union wages for manual workers grow by more than union wages. Furthermore, the increase in the amount of capital used in production worsens the bargaining position of unions, as a potential strike becomes less harmful for the firm. This additionally dampens union wage growth compared to non-union wage growth. Thus, the incentives to unionize decrease unambiguously for manual workers.

Note that the mechanism we emphasize here is in line with the literature on union membership decisions. Empirical studies, including DiNardo et al. (1996) and Rueda et al. (2002), document that unions have become less effective in redistributing earnings over the last decades. This argument is taken up and extended in Baccaro and Locke (1998) and Checchi et al. (2010), who both highlight disillusion about potential wage growth as the driving force behind declining union membership rates among the least skilled workers.

The effect of routine-biased technical change on the voting incentives for routine workers is ambiguous and depends on the larger union wage growth due to the relatively larger productivity growth of manual workers and the lower union wage growth due to the larger amount of capital. In the quantitative evaluation, the incentives for routine workers to vote in favor of a collective bargaining agreement monotonically decrease with falling computer capital prices. However, even if the incentives were to increase for the lower-skilled routine workers, manual workers would still drive deunionization, as they make up between 46% and 53% of the bargaining unit inside firms.

3.6 Quantitative Analysis

In this section all the parameters discussed above are calibrated to match different aspects of U.S. data for 1983. In line with empirical data, we let computer capital prices fall by 48% until 2005. We use the calibrated model to quantify the effect on the occupational choice of workers and on union elections. For the simulation we choose a setting with heterogeneous unions that differ with respect to their bargaining powers γ^u and γ^u_a. We consider an economy that consists of a number N of independent islands, where each island represents a set of firms in an industry. This is in line with the empirical literature, as evidence in Tüzemen and Willis (2013) suggests that polarization is also mainly driven by within-industry changes. All islands are identical except for the bargaining power of the potential union. The performance of the model is evaluated along several dimensions, especially with regard to the empirical evidence on deunionization in the U.S. We focus on steady states as we are mainly interested in the long-run effect of routine-biased technical change on the economy.

3.6.1 Calibration

The model is calibrated to quarterly frequency. Target values pertain to economy-wide averages. Table 3.4 lists the exact parameter values as well as the source that encourages the specific choice. We first calibrate the discount factor β to a conventional value of 0.99, which implies an annual interest rate of 4%. Next, we calibrate the labor market variables. The separation rates of manual and routine workers are set to the standard value of $s_m = s_r = 0.1$ (Shimer, 2005). Following Albertini et al. (2017), we set the separation rate of abstract workers to the lower value of $s_a = 0.05$.

The matching efficiencies are calibrated in order to match the average job-finding rate between 1983 and 2005 reported in Shimer (2005). Under this calibration the job-finding rate increases with the skill level of workers. A large literature documents no or only small effects of unionization on employment: Frandsen (2012) and Montgomery (1989) on the aggregate level, Boal and Pencavel (1994) on the industry level, and DiNardo and Lee (2004) on the firm level. Furthermore, using linked employer-employee data, Brändle and Goerke (2018) argue that negative employment effects might be caused by selection in cross-sectional studies. We take this evidence into account by calibrating the matching efficiency on unionized islands to match the same job-finding rates as on non-unionized islands.

Vacancy posting costs are chosen to correspond on average to 35% of a worker's quarterly steady state wage, which lies well in the range of values found in the literature (cf. Garín, 2015; Michaillat, 2012). For simplicity, unemployment benefits and strike pay are both set to zero.[21]

All production and skill specific parameters are set in order to match data on employment shares in 1983 (30.7% manual, 35.7% routine, and 33.6% abstract workers), as well as the abstract employment share of 40.9% in 2005. This leaves manual and routine employment shares in 2005 as untargeted moments to gauge the performance of the model. The growth rates of computer capital prices g_{p_K} and abstract labor supply $g_{L_a^S}$ are calibrated to match a drop in computer capital prices by 48% and an increase in the abstract employment share of 7.3 percentage points.

Depending on birth cohort, age group, and survey data (Census/ACS, CPS, NLSY, PSID, and SIPP), the difference in wages between high school graduates and college graduates amounts to 10%-29%. The average Mincer college wage premium – over age groups, birth cohorts, and survey data – amounted to roughly 15% to 20% in the U.S. in 1983 (cf. Ashworth and Ransom, 2019).[22] Setting the bargaining power of abstract workers to $\gamma_a^n = 0.8$ and the bargaining power of manual and routine workers to $\gamma_m^n = \gamma_r^n = 0.5$ yields a college wage premium of 17% in the model in 1983 while leaving the average worker bargaining power in the standard range between 0.4 and 0.6.[23]

[21] The results are robust to alternative parameter choices.

[22] Mincer college wage premium refers to a wage premium that is adjusted for observable skills using the model proposed by Mincer (1974). Typically, the Mincer wage premium is roughly half the size of the raw wage premium.

[23] The college wage premium can be calculated when assuming that the individual skill η refers to the educational

Table 3.4. Calibrated Parameters

Symbol	Interpretation	Value	Source
β	Discount factor	0.99	Annual interest rate of 4%
s_m	Manual eparation rate	0.1	Garín (2015)
s_r	Routine separation rate	0.1	Garín (2015)
s_a	Abstract separation rate	0.05	Albertini et al. (2017)
Ψ_m	Manual matching efficiency	0.25	Job-finding rate of 0.56
Ψ_r	Routine matching efficiency	0.33	Job-finding rate of 0.56
Ψ_a	Abstract matching efficiency	0.8	Job-finding rate of 0.56
ψ	Unemployment-elasticity of matching	0.5	Petrongolo and Pissarides (2001)
c_m	Manual recruiting costs	0.3	35% of wages
c_r	Routine recruiting costs	0.3	35% of wages
c_a	Abstract recruiting costs	0.5	35% of wages
A	Productivity routine and abstract input	3.4	Occupational shares in 1983
A_m	Productivity of manual input	0.77	Occupational shares in 1983
α	Marginal return to abstract labor	0.45	Occupational shares in 1983
ρ	Production parameter	0.65	Occupational shares in 1983
σ	Production parameter	0.74	Albertini et al. (2017)
μ	Production parameter	0.5	Albertini et al. (2017)
$\underline{\eta}$	Lower bound on skill	0.48	Occupational shares in 1983
$\bar{\eta}$	Upper bound on routine skill	1.44	Occupational shares in 1983
$g_{L_a^S}$	Growth rate of abstract labor supply	0.015	Abstract employment in 2005
g_{p_K}	Growth rate of computer capital prices	-0.029	Investment prices in 2005
γ^m	Manual Worker's bargaining power	0.5	Midpoint of literature values
γ^r	Routine Worker's bargaining power	0.5	Midpoint of literature values
γ_a	Abstract Worker's bargaining power	0.8	College Wage premium 1983
γ^u	Union bargaining power	0.51 - 1	Non-Abstract Union Membership
γ_a^u	Craft Union bargaining power	0.88 - 1	Abstract Union Membership

The union bargaining power of the potential unions is assumed to be equally distributed – on the interval between 0.51 and 1 for the potential industrial unions and on the interval between 0.88 and 1 for the potential craft unions.[24] With the bargaining power of the most powerful unions set to one, a lower bound of 0.88 on the bargaining power of the unions for abstract workers matches the union membership rate of 16.6% in 1983 reported in the Union Membership and Coverage Database constructed by Hirsch and Macpherson (2003) for workers in abstract occupations. Given this calibration, a lower bound of 0.51 for industrial

attainment of otherwise identical workers. If we further assume that on average manual workers have high school education, abstract workers a college degree, and routine workers some college or an associates degree, than the college wage premium is given by the ratio of abstract to manual wages in the model.

[24]The large differences between the union bargaining powers and the individual bargaining power of a worker are necessary because under collective bargaining workers are not lost to the firm when bargaining breaks down. If we instead assume that the firm loses its workforce when no agreement is reached, the calibration targets for the union bargaining power would be substantially lower than under individual bargaining. The reason behind this is that the union bargains over the average product of all workers in the bargaining unit, while each individual worker only bargains over his or her marginal product. The results are robust to alternative intervals of the union bargaining power.

Table 3.5. Unionization Rates: Model versus Data

	1983		2005	
	Data	Model	Data	Model
Overall	19.5%	19.5%	12.9%	10.2%
Manual Workers	24.8%	21.0%	14.5%	6.3%
Routine Workers	17.7%	21.0%	10.2%	6.3%
Abstract Workers	16.6%	16.6%	13.4%	15.9%

Note: Data for union membership rates by occupations are calculated using the Union Membership and Coverage Database constructed by Hirsch and Macpherson (2003) and include all wage and salary workers. We use the occupational classification by Autor and Dorn (2013). The overall union membership rate is calculated using the employment shares reported in Autor and Dorn (2013) and the union membership rates by occupation.

unions matches the overall union membership rate of 19.5% in 1983, calculated using the Union Membership and Coverage Database and the employment shares from Autor and Dorn (2013).

3.6.2 Deunionization

As capital prices fall, the unions with the lowest bargaining power fail to gain majority support in the subsequent elections and are terminated.[25] Our model performs well in generating declining union membership rates between 1983 and 2005. The predicted and actual changes are given in Table 3.5, with the only targeted values being the overall and the abstract union membership rate in 1983.

The union membership rate falls by 9.3 percentage points from 19.5% to 10.2% in the model, compared to a drop by 6.6 percentage points from 19.5% to 12.9% in the data. The union membership rate for manual workers drops by 14.7 percentage points (10.3 in the data), the membership rate for routine workers by 14.7 percentage points (7.5 in the data), and the membership rate for abstract workers by 0.7 percentage points (2.5 in the data).[26]

As abstract workers are unionized in a homogenous group, the higher marginal productivity due to technical change affects union and non-union wages for these workers similarly. However, under individual bargaining the higher demand for abstract workers increases the cost

[25] This model prediction is supported by evidence in the 2004 NLRB Performance and Accountability Report. Going from 1994 to 2004, the number of filed representation petitions has dropped by 25% and the share of won elections has increased by over five percentage points.

[26] The model slightly overpredicts the decline in the membership rates for manual and routine workers and underpredicts the decline in membership rates for abstract workers. Possible explanations for the former are workers that remain union members despite declining monetary incentives out of habit, due to peer pressure, or because of other non-monetary membership advantages. The latter might arise because we ignore heterogeneity among abstract workers.

Table 3.6. Simulated Changes in Unionization Rates – Decomposition, 1983–2005

	Percentage Point	Share
Total Change	-10.27	100%
Within-occupation	-9.97	97.08%
Between-occupation	-0.30	2.92%

Note: The relative contribution of the within-occupation and between-occupation component is calculated using the methodology described in Section 3.2.

of hiring a worker in the next period. The outside option under collective bargaining, i.e., a strike of abstract workers, is associated with the same costs as before. Thus, the incentives to unionize decrease slightly for abstract workers but by less compared to manual and routine workers.

Result 1. *The drop in the overall union membership rate between 1983 and 2005 is mainly driven by decreasing membership rates within occupations and not by changing employment shares.*

Using the same methodology as in Section 3.2, we calculate the within-occupation and between-occupation component for the three occupations in our model. The results are summarized in Table 3.6. Deunionization does not only work entirely through changes within industries (by construction) but also mainly through changes within occupations rather than through changing employment shares: over 95% of the changes in union membership rates between 1983 and 2005 are driven by the within-occupation component in our model.

Result 2. *Despite falling union membership rates, the average union wage premium stays roughly constant between 1983 and 2005.*

Estimates of the average union - non-union wage differential across workers range from close to zero (cf. Bryson, 2002; Booth and Bryan, 2004) to 25% (cf. Hirsch and Schumacher, 2004). Recent studies by DiNardo and Lee (2004) and Frandsen (2012), who focus on employer and union election data, find only very small or even negative union wage premia on average. Additionally, Streeck (2005) argues that because of its structure, industrial unions tend to exhibit even lower wage premia on average compared to craft unions.

Farber et al. (2018) emphasize that existing models of union formation have trouble explaining the observation of a relatively constant union wage premium in times of rapidly declining union membership rates, as the increasing use of capital and high-skilled workers reduces the value of low-skilled workers for the firm and thus worsens the bargaining position of unions. A similar effect is at work in our model. However, as our model predicts that unions with the lowest bargaining power will be the ones that are terminated, union termination in the model is associated with an increasing average union bargaining power. These countervailing effects

Table 3.7. Simulated Union Wage Premium and Skill Ratio

	1983	2005
Union Wage Premium	0.6%	-0.6%
Skill Ratio Non-unionized Workers	0.53	0.62
Skill Ratio Unionized Workers	0.4	1.75

Note: The skill ratio in the model is defined as the ratio of abstract to non-abstract workers.

imply relatively constant union wage premia despite a sharp decline in union membership rates. The evolution of the average union wage premium in the model is given in Table 3.7. Despite a drop of close to 10 percentage points in the union membership rate, the average union wage premium decreases by only 1.2 percentage points in the model.

Result 3. *The relative skill level of union members increases between 1983 and 2005.*

Existing models of union formation mostly rely upon declining membership rates among the highest-skilled workers in order to explain deunionization. This stands in sharp contrast to the union membership data in Hirsch and Macpherson (2003). In our model the union membership rate of abstract workers decreases only slightly. Consider an increase in the skill level of a worker and how this affects his or her probability of being a union member. Given the predicted changes in unionization rates between 1983 and 2005, an increase in the skill level of a worker decreases the probability of being a union member in 1983 but increases the probability of being a union member in 2005. This coincides with evidence on the effect of educational attainment on the union status of workers in Farber et al. (2018). The reason is that the union membership rate of abstract workers decreases by less compared to the union membership rates of the less-skilled manual and routine workers, both in the data and in our model. The ratio of abstract to non-abstract workers inside and outside of unions in our model is reported in Table 3.7.

3.6.3 Polarization

As shown in Section 3.5.7, falling computer capital prices lead to employment adjustments, with the lowest-skilled routine workers deciding to switch to manual occupations upon becoming unemployed.

The employment shares of the three occupational groups in the model and in the data are given in Table 3.8. The share of manual workers increases from 30.7% to 31.1% in the data and to 31.0% in the model between 1983 and 2005, while the employment share of routine workers decreases from 35.7% to 28.0% in the data and to 27.9% in the model. Figure 3.4 displays the respective percentage point changes in the employment share for each occupation.

Table 3.8. Employment Shares in 1983 and 2005: Model versus Data

	1983		2005	
	Data	Model	Data	Model
Manual	30.7%	30.7%	31.1%	31.0%
Routine	35.7%	35.7%	28.0%	27.9%
Abstract	33.6%	33.6%	40.9%	40.9%

The share of workers in each occupation is constructed using the dataset and the occupational classification by Autor and Dorn (2013).

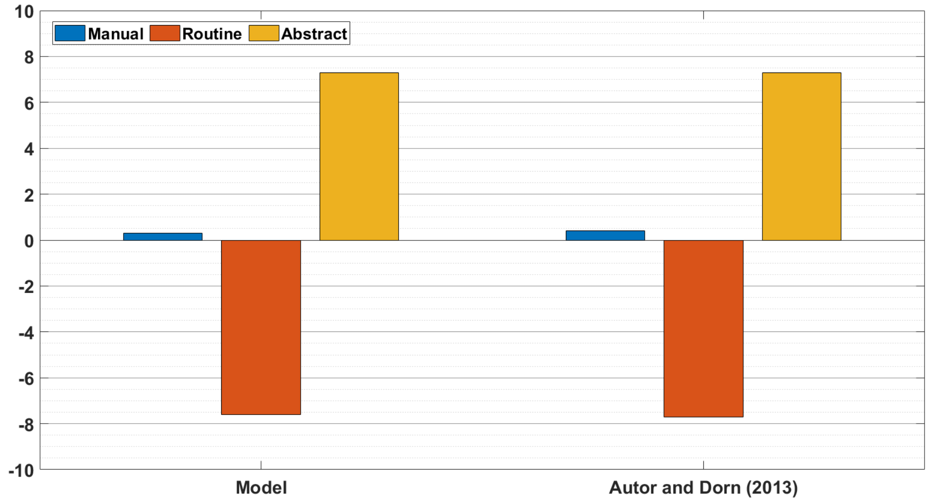
Employment changes are less pronounced in unionized firms: as wages for manual and routine workers grow equally, the lowest-skilled unionized routine workers have no incentive to switch to manual jobs.[27] While there is no direct evidence on the polarization of the employment structure in unionized versus non-unionized firms, our model prediction is supported by two strands of the literature. First, Calmfors et al. (2001) and Rogers and Streeck (1995) argue that in many countries the management is under the obligation to at least consult with the relevant unions over restructuring and layoff plans. In these cases union officials tend to prefer policies that favor those workers who are most likely to be union members, in order to improve their chances in future elections. Thus, unions will likely oppose plans that reinforce polarization. Second, Connolly et al. (1986), Hirsch and Link (1987), and more recently Bradley et al. (2017) argue that unions have detrimental effects on innovation and technology adaption. As technical change is the most important driving force behind polarization, less innovation is likely to be accompanied by less polarization. This implies, as our model predicts, that deunionization amplifies polarization.

Even though the manual employment share remains roughly unchanged, there has been substantial employment reallocation with more than 10% of all routine workers in 1983 deciding to switch to manual occupations. About 15% of the occupational switches in our model are triggered by the termination of unions. When low-skilled routine workers are unable to find unionized jobs, which would pay them a substantial union wage premium, their incentives to switch occupations increase. While the model predicts routine-biased technical change to be the main explanation for job market polarization, deunionization substantially amplifies employment changes.

The changes in employment are accompanied by wage changes. The model predicts wages for abstract, routine, and manual workers to grow by 10%, 8%, and 8.5%, respectively. Although a bit smaller, these changes accord with the pattern of wage changes by skill levels reported in Autor and Dorn (2013) for the time period between 1980 and 2005.

[27] This result is independent of the specific choice of the union wage schedule and holds as long as union wages for routine workers are higher compared to union wages of manual workers.

Figure 3.4. Percentage Point Changes in Employment Shares, 1983–2005: Model versus Data



Note: The share of workers in each occupation is constructed using the dataset and the occupational classification by Autor and Dorn (2013).

3.6.4 Inequality

In contrast to the large effect on employment changes, deunionization has only modest effects on wage changes. Going from 1983 to 2005, the Gini index in our model increases by 18% compared to an increase of 12% for U.S. data.[28] However, since union wage premia are small on average and the unions with the lowest bargaining power are terminated, this increase in inequality is almost entirely driven by the increasing employment share of abstract workers and by their increasing relative wages. The small overall effects of deunionization on wage inequality in our model accord with the empirical findings in DiNardo et al. (1996), Frandsen (2012), and Farber et al. (2018).

The effects of deunionization for those groups that traditionally receive a high union wage premium, the lower middle-skilled workers, are substantial. For the lowest-skilled previously unionized routine workers, i.e., those workers that lose their union wage premium going from 1983 to 2005 and subsequently switch occupations, the wage growth would be 60% larger if they were covered by one of the remaining unions.

3.7 Discussion and Policy Implications

While routine-biased technical change hurts middle-wage workers, job market polarization per se, in the sense of changing employment shares, does not. In the model, the possibility to switch occupations allows labor supply to adjust to the changes in labor demand and thereby to partly offset the wage effects of routine-biased technical change. However, Kambourov

[28] The Gini index in our model is computed using wage ventiles.

and Manovskii (2009), Gathmann and Schönberg (2010), and Cortes and Gallipoli (2017) show that occupational switching costs are large. Therefore, as proposed for example in Autor et al. (2003), policies that simplify job switches or that aim at making them less costly for workers could serve to dampen income inequality caused by routine-biased technical change.

Additionally, our analysis has shown that, even though the overall effect of deunionization on income inequality is small, there are large effects for lower middle-skilled workers. Taking into account evidence from Frandsen (2012), who reports that most union elections are very closely contested, even small policy changes could lead to large effects on income inequality for these workers.

We briefly consider the effects of two policies that aim at supporting lower middle-skilled workers. The first policy simply abolishes union elections after the first election in 1983 and maintains the established unions regardless of worker preferences. While this approach prevents deunionization, it also prevents efficient deunionization in the sense that even unions generating a highly negative average wage premium would be maintained. The second policy lowers the necessary voting threshold for unions. For specific voting thresholds, this policy achieves the same results as abolishing elections, with identical downsides. However, such an intervention is not well suited to stop the overall trend of declining union membership rates and the threshold would have to be regularly adjusted to changes in the economy. Furthermore, low threshold values, apart from being difficult to justify, could in principle lead to the founding of further inefficient unions.

In our simulation, deunionization can always be prevented by adjusting the union wage schedule towards less equality inside the unionized firms. However, empirical evidence suggests that rigid organizational structures partly prevent unions from meeting today's challenges. Waddington (2005) contends that trade union practices are perceived as formal and old-fashioned and that the representative structures inside unions are often inappropriate for the participation of all members. Bryson et al. (2016) argue that union representatives have very long tenure and tend to become less representative of the membership over their term of office.

While unionization rates decline across all age groups, according to data from the Bureau of Labor Statistics, membership rates for workers aged between 16 and 24 declined at twice the rate of overall membership between 2002 and 2012. Data on the evolution of the median age of union members points in the same direction: Dunn and Walker (2016) stress that over half of all U.S. union members are between 45 and 64 years of age. Thus, it seems that unions are mostly controlled and influenced by older members that might display a tendency to stick to established practices. Bryson et al. (2016) argue that the decline in union membership rates across countries is negatively related to the degree of progressiveness of the unions. Thus, one straightforward policy suggestion is to restrict the tenure of union representatives to ensure that union officials are drawn from the current membership.

3.8 Conclusion

This paper explores the effect of routine-biased technical change on both the occupational and the union-membership choice of workers. We use broad state-level labor market data to illustrate that the decline in unionization rates is more pronounced in U.S. states with a larger decline in the employment share of routine-intensive occupations. We additionally show that this decline is not driven by a simple composition effect but mainly by within-industry and within-occupation changes. Building on this observation, we develop a model that endogenizes both the occupational and the union membership decision in a search and matching framework.

We provide analytical results and use the calibrated model to show that routine-biased technical change, represented by a sharp drop in computer capital prices, not only generates employment and wage polarization but also deunionization. The drop in computer capital prices reduces the demand for routine workers, while the demand for abstract and manual workers increases. The changing demand structure influences the surplus unions can extract and thereby also the individual union wage premium of workers. Manual workers, who benefit from the changing demand structure, are discouraged from voting in favor of a collective bargaining agreement. As wage gains for manual workers would be distributed more equally between manual and routine workers by the union, manual workers are better off bargaining individually with the firm. Former routine workers, when faced with lower wages compared to manual workers, decide to switch occupations.

We demonstrate that this effect can lead to a change in the voting outcome, with the majority of the workforce of previously unionized firms now voting against unionization and in favor of individual bargaining. In an economy in which unions differ with respect to their bargaining power, routine-biased technical change leads to a large decrease in union membership rates, because those unions with the lowest bargaining power are terminated. As about 15% of all job switches are triggered by deunionization, this contributes substantially to employment polarization. While overall effects on income inequality are small, low- to middle-skilled previously unionized workers are severely affected.

4 Outlawed: Estimating the Labor Market Effects of Judicial Ideology

Authors: Christian Bredemeier, Tobias Föll, and Anna Hartmann

4.1 Introduction

Do ideological tendencies influence court rulings? An exhaustive literature suggests that the answer to this question is: Yes! (cf. Cohen and Yang, 2019; Taha, 2004; Songer et al., 1994). However, do general ideological tendencies of the judiciary also have direct economic effects? And if yes, how large are these? In this paper, we aim to fill a gap in the literature by providing answers to these important questions.

Ideological tendencies of the judiciary are generally considered to be of paramount importance in the United States. Among others, U.S. Circuit Judge Michael McConnell, a potential nominee to the Supreme Court during the presidency of George W. Bush, is frequently cited arguing that Supreme Court nominations are among the most important decisions of a U.S. president. The confirmation battles regarding President Trump's Supreme Court nominees Neil Gorsuch and Brett Kavanaugh in the U.S. Senate corroborate this view. Unsurprisingly, the appointment of conservative federal judges has been one of most prominent topics in both of Donald Trump's presidential campaigns and one of the major appeals to moderate Republicans.

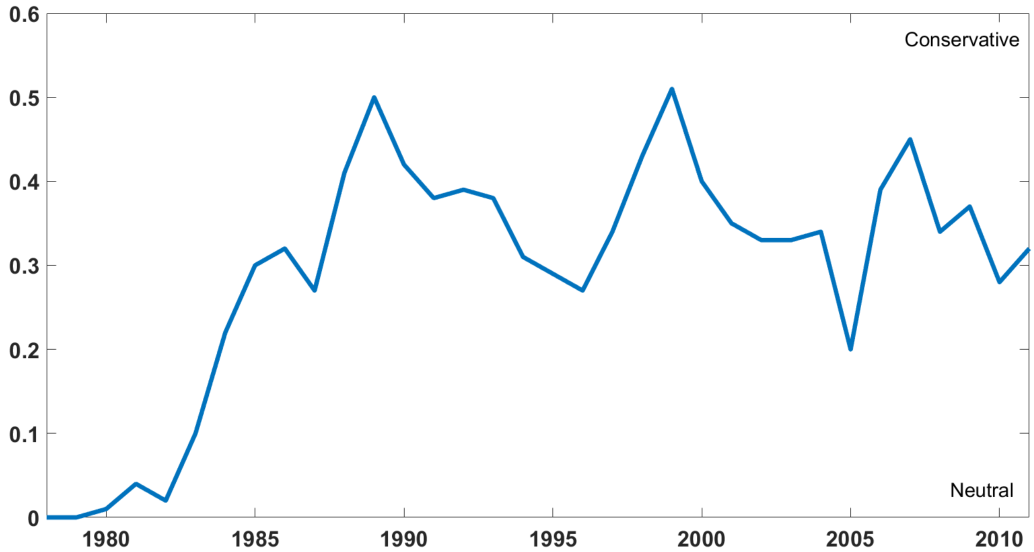
In light of this perceived importance, it is surprising that all existing evidence on the economic impact of the Supreme Court is either case-based or purely anecdotal (cf. Epstein et al., 2013; Gilman, 2014). In this paper, we document heterogenous effects of Supreme Court ideology on district court rulings across U.S. states and exploit these differences in order to identify the economic impact of jurisdiction. Our contribution to the literature is twofold. First, we move the focus of the literature away from the individual judge and towards the judiciary in general. Second, using an extensive dataset, we provide estimates of the degree to which economic conditions in the U.S. are influenced by ideological tendencies of the judiciary.

While the importance and influence of the Supreme Court is undisputed, it can only hear about 150 cases every year. The decisions made by federal courts thus constitute the last word in thousands of cases every year. A large literature (cf. Boyd, 2015a; Benesh and Reddick, 2002; Cannon and Johnson, 1984; Wasby, 1970; Songer et al., 1994) establishes that lower courts tend to follow the path set by the Supreme Court when the Supreme Court's orientation is clear and unambiguous, wheras an ideologically rather neutral or ambiguous approach of the Supreme Court gives judges some leeway which they can use to follow their own ideology. This behavior of judges is usually attributed to reversal aversion (cf. Miceli and Coşgel, 1994; Posner, 2005; Gennaioli and Shleifer, 2008; Randazzo, 2008). We conclude that one can expect regional heterogeneity in the effect of changes in the ideology of the Supreme Court on court decisions, which can be used to identify the effects of judicial ideology. Building on Miceli and Coşgel (1994), we develop a model of judge decision-making with ideological preferences and reversal aversion that makes this argument explicit. The model predicts that a state is more strongly affected by the changes in Supreme Court ideology since the late 1970s (from center to clearly conservative, see Figure 4.1) the more liberal its district court judges are. The intuition is as follows. In the late 1970s, with the Supreme Court rather balanced ideologically, both more conservative and more liberal district court judges were, at least partly, able to influence court rulings according to their own ideology. With the Supreme Court shifting towards being more conservative, all district courts issue rather conservative rulings. While rulings in conservative districts remain rather conservative, liberal judges shy away from the risk of reputational damage due to overturned rulings by also issuing more conservative rulings.

We confirm the predicted regional heterogeneity in the effects of Supreme Court ideology on decisions by lower courts using an econometric procedure derived from the model. To this end, we use data on rulings of federal district courts in close to 24,000 economic or labor-related cases from the Carp-Manning U.S. District Court Database compiled by Carp and Manning (2016). District court rulings are chosen because of three reasons. First, the federal court system hears cases involving the laws and treaties of the U.S. Hence, a large share of lawsuits related to economic issues are filed in federal courts, while the state courts are mostly concerned with traffic, criminal, and civil cases. Second, rulings issued by the district courts are much more likely to create a precedent than rulings at state courts and are thus relevant to a large number of additional cases. Third, district courts have the last word in about 99% of the filed federal court cases, as only about 1% of all district court cases are reversed by higher courts (cf. Cohen and Yang, 2019; Edwards, 2019; Eisenberg, 2004). We find that an increase in conservatism at the Supreme Court, in line with our model, strongly and significantly increases the share of conservative rulings in states with rather liberal district courts relative to the rulings in states with rather conservative district courts.

Having established that the interaction between Supreme Court ideology and district court ideology can be used as an instrument for exogenous variations in district court rulings along the ideological spectrum, we use it to analyze the effect of court rulings on the labor market. This borrows methodology from the trade and migration literatures, where researchers exploit regional variation in the exposure to import competition (cf. Autor et al., 2013) or migrant inflows (cf. Dustmann et al., 2017) to identify causal effects of these phenomena. Specifically, we use an interaction term between time-varying Supreme Court ideology and a time-invariant state-specific measure of the ideology of district court judges in regressions of labor market

Figure 4.1. Ideological Leanings of the Supreme Court



Note: This graph depicts the ideal point estimates provided in the dataset to Bailey (2013) for the ideological leanings of the median Supreme Court justice between 1978 and 2011. The estimates from Bailey (2013) are chosen over the estimates from Martin and Quinn (2002), as the former estimates explicitly take into account the problem of agenda changes over time by using bridging information. This allows for the use of the scores in a cardinal sense, whereas the estimates in Martin and Quinn (2002) can only be used as ordinal measures. However, both estimates clearly show the shifts in ideological leanings of the Supreme Court towards the conservative end of the ideological spectrum since the 1970s.

outcomes that include both time and state fixed effects.[1] This exploits that court rulings in more liberal states are more strongly affected by the Supreme Court's rising conservatism, such that the coefficient on the interaction term is to be interpreted as a causal effect of ideological tendencies of the judiciary. Put differently, the econometric procedure isolates the part of the change in regional district court rulings that is driven by developments at the U.S. Supreme Court in Washington D.C. and therefore arguably exogenous to regional labor market conditions.

Our empirical analysis suggests that an increase in the share of pro-business rulings at district courts increases labor market fluidity. Unemployment and unemployment duration fall, while the job-finding rate and employment increase. However, on the downside, we find that more pro-business rulings tend to reduce wages and other measures of job quality while accelerating the hollowing-out of the middle class, as union coverage and employment shares in routine-intensive occupations and industries fall. Moreover, we also find that conservative

[1] We focus on labor market outcomes as labor earnings are the major source of income for most households and thus a primary determinant of life satisfaction. With more conservative judges and justices tending to be rather pro-business and more liberal judges rather pro-worker, ideological shifts in Supreme Court composition affect decisions in cases regarding affirmative action, union rights, worker compensation upon firings, layoffs, and the like. The Business Litigant Dataset for the terms between 1946 and 2011 and the fraction of votes in favor of business in Epstein et al. (2013) reveal large effects of changes in Supreme Court composition on rulings, especially for cases concerning economic issues. Seven of the ten Supreme Court justices least favorable to businesses served between 1960 to 1970. In contrast, in 2011 five of the nine serving Supreme Court justices counted among the ten justices most favorable to businesses.

court rulings contribute to increasing income inequality. Quantitatively, a ten percentage point increase in the share of pro-business rulings in a state is associated with a reduction in the state's unemployment rate by about 0.7 percentage points relative to other states. Average hourly wages fall by 1.7%, union coverage by 1.3 percentage points, and the employment share in routine-intensive occupations by 0.6 percentage points. Income inequality, measured as the 90/10 ratio in family income, increases by 3.7%.

Over the 34 years in our sample, the trend towards pro-business rulings increased their share by about 6 percentage points. A back-of-the-envelope calculation, assuming that general-equilibrium effects of court rulings in other states are small, indicates that the rise in the share of conservative court decisions from 1978 to 2011 can be expected to be responsible for a decrease of 0.4 percentage points in the unemployment rate, a 1.1% reduction in wages, a fall in union coverage of 0.8 percentage points, a 0.4 percentage point reduction in the routine employment share, and an increase in the 90/10 income ratio of about 2.2%. In this light, increasing judicial conservatism seems to have contributed, although in a moderate way, to important long-run economic developments such as wage stagnation, deunionization, job market polarization, and rising inequality.

Our main empirical results can be rationalized in a simple search and matching framework which we extend by wrongful-termination lawsuits upon separation. In the model, a larger share of pro-business rulings induces falling wages by eroding the bargaining power of workers. Lower labor costs result in a larger number of posted vacancies and consequently in a higher job-finding rate and in a lower unemployment rate.

Our results have important implications regarding the appointment and retirement of federal judges. Due to lifetime appointments and increasingly strategic retirements on federal courts, changing an established majority in the judiciary has become ever more difficult over the last decades. This means that today's decisions regarding the composition of the judiciary influence peoples' lives for decades to come, even though future generations might have very different preferences regarding societal trade-offs, especially when taking into consideration the rapidly changing composition of the U.S. population. Given that our results reveal quite strong effects of judicial ideology, they lend support to term limits for federal judges, as they are proposed by politicians from both sides of the aisle.[2]

The remainder of the paper is organized as follows. In Section 4.2 we give an overview of the related literature. The effect of Supreme Court ideology on district court rulings is discussed and estimated in Section 4.3. The findings from this section are used to estimate the effect of ideological tendencies of the judiciary on the labor market in Section 4.4. The results are summarized in Section 4.5.

[2]Prominent advocates include Senators Sanders, Warren, Bennet, Rubio, and Cruz.

4.2 Related Literature

This paper is related to different strands of the literature, in particular to those analyzing the determinants of labor market outcomes and of court rulings, respectively.

A number of important determinants of labor market fluidity have been identified by the literature. For example, firing costs have been shown to reduce job-finding rates both theoretically (cf. Wasmer, 2006) and in the data (cf. Kugler and Saint-Paul, 2004). Kugler and Saint-Paul (2004) and Autor et al. (2006a) document that exceptions to the employment at-will doctrine (wrongful-discharge laws) reduce job-creation and lead to lower employment rates. Acemoglu et al. (2001), among others, illustrate that employment protection laws reduce the job-finding probability for affected groups. Cahuc et al. (2019) show that pro-worker rulings in wrongful-termination cases reduce job-creation rates in affected firms. We contribute to this field by emphasizing that increasingly conservative court rulings in economic cases increase both the employment rate and the job-finding rate.

We also contribute to the debate about the causes of incisive developments witnessed over the last decades. Computerization, skill-biased technical change, and routine-biased technical change are put forward as explanations for rising inequality (cf. Autor et al., 2006c), structural change away from manufacturing industries (cf. Autor et al., 2003), polarizing changes in the occupational employment structure at the expense of routine-intensive jobs (cf. Autor and Dorn, 2013), and deunionization (cf. Dinlersoz and Greenwood, 2016). Our results complement these explanations by showing that increasing conservatism of the judiciary accelerates all of these developments.

Literature on the economic effects of court rulings is rare. Analyzing case composition, rulings, and votes of Supreme Court justices over time, Epstein et al. (2013) conclude that the Supreme Court has indeed become more favorable to businesses over the last decades. The analysis does however not extend to the effect of the larger share of pro-business rulings on actual economic conditions. Gilman (2014) argues that the Supreme Court reinforces economic inequality by verbally analyzing selected Supreme Court rulings. Neither Epstein et al. (2013) nor Gilman (2014) provide a systematic statistical evaluation of the economic impact of the Supreme Court.

Due to our identification of exogenous variation in court rulings, our paper is also related to the literature that discusses determinants of court rulings which are not directly related to the case at hand. This literature has established that court rulings, conditional on case characteristics, depend on aggregate conditions such as outside temperatures (cf. Heyes and Saberian, 2019), media coverage on crime (cf. Philippe and Ouss, 2018), the success of local sports teams (cf. Eren and Mocan, 2018), and the aggregate business cycle (cf. Ichino et al., 2003; Marinescu, 2011). Furthermore, there is ample evidence that, conditional on case characteristics, individual characteristics of judges at various levels of the judiciary have substantial effects on court rulings. Kling (2006), Dahl et al. (2014), French and Song (2014), Aizer and Doyle Jr. (2015), Dobbie et al. (2018), Bernstein et al. (2019b), Bernstein et al. (2019a), and Cahuc et al. (2019) all construct measures of judge leniency which differ substantially across judges.[3] Knepper (2017) and Boyd et al. (2010) focus on judge gender, Welch et al. (1988), Chew and Robert E (2009), Yang (2015), and Kastellec (2013) on judge race, and Glynn and Sen (2015) on the effect of parenthood. The ideology or political affiliation of judges is an exceptionally important determinant of rulings. While this is undisputed for U.S. Supreme Court justices, empirical evidence also emphasizes an important role of ideology in the lower courts, including the federal district courts on which our analysis focuses. For example, Cohen and Yang (2019) exploit random case assignment of judges within district courts and document substantial effects of party affiliation on criminal-sentencing decisions. Similar ideology effects in district court rulings are reported by Tiede et al. (2010), Schanzenbach and Tiller (2007), Schanzenbach and Tiller (2008), and Taha (2004).

A number of studies have addressed the interplay between a judge's own ideological preferences and the preferences of the judge's superiors at higher courts, which is at the core of our identification strategy. In particular, judges are generally considered to be reversal-averse which lets them put their own ideological preferences last when these stand in sufficiently strong conflict with the ideologies of their superiors at higher courts. Our theoretical model builds on Miceli and Coşgel (1994), who construct a model of judge decision-making under reversal aversion. Posner (2005) and Gennaioli and Shleifer (2008) also consider theoretical models where judges' decisions are determined in a conflict between their ideological preferences and their reversal aversion. Empirical evidence for reversal aversion by district court judges is provided by, among others, Randazzo (2008). Boyd (2015b) documents that ideological differences between a district court judge and direct superiors at the circuit court indeed raises the probability of reversal. Songer et al. (1994) understand the relation between the Supreme Court and the courts of appeals as a principle-agent relation where appellate judges try to pursue their own ideological interests, but face incentives to follow the Supreme Court's lead. They document that appellate judges are highly responsive to the Supreme Court's monitoring practices, but they also find a strong independent effect of appellate judges' own ideologies which indicates that they satisfy their own policy interests when weaker monitoring offers them leeway. Similarly, Cohen and Yang (2019) document that the influence of district court judges' ideology on rulings increases with the amount of posessed discretion. Zorn and Bowie (2010) illustrate that the importance of judge ideology for rulings decreases as one moves down the federal judicial hierarchy, which is in line with reversal aversion, as district courts are more strongly monitored than courts of appeals and the Supreme Court is not monitored at all. Choi et al. (2012) document that district court judges in circuits with ideologically uniform circuit court judges follow the orientation of their superiors, suppressing their own ideology, while

[3] These studies exploit the random case assignment of heterogeneous judges to identify the effects of criminal sentencing (cf. Dobbie et al., 2018; Aizer and Doyle Jr., 2015; Kling, 2006), bankruptcies (cf. Bernstein et al., 2019b,a), disability payments (cf. Dahl et al., 2014; French and Song, 2014), and firing costs (cf. Cahuc et al., 2019).

district court judges in circuits with ideologically diverse circuit judges find it more difficult to minimize their risk of reversal. All of these observations support the mechanism underlying our identification approach.

Our results additionally contribute to this literature by exposing that the ideology of the Supreme Court influences decisions of district court judges in a way consistent with reversal aversion.[4]

4.3 The Effect of Supreme Court Ideology on District Court Rulings

Analyzing district court rulings across U.S. states, we establish that a conservative ideological shift of the Supreme Court induces an increase in the share of pro-business rulings in states with liberal district courts relative to states with conservative district courts.

4.3.1 Theory

To fix ideas, we present a simple model which guides our identification of the effects of Supreme Court ideology on district court rulings. We build upon the model of judge decision-making developed by Miceli and Coşgel (1994). In this model, we focus on a specific factor that potentially determines case outcomes: judge ideology. While existing laws and precedent are undoubtedly the most important predictors of case outcomes, there is a large literature that exposes substantial effects of seemingly unrelated factors, see Section 4.2.

Judges have two sources of utility from a particular ruling r. The first source of utility originates from private preferences over the case at hand, $V(r)$. This source of utility reflects what we summarize as ideological leanings and includes, for example, the political views and the theory of the law of the judge. This utility component is larger, the closer the actual decision r resembles the private preferences. The second source of utility originates from the judge's reputation. Reputational utility is given by $R(r)$ and is meant to capture increased promotion chances of the judge due to a better reputation. While Miceli and Coşgel (1994) focus on future citations, our focus is on the probability of a decisions being reversed by higher courts, i.e., by the circuit courts or in the last instance by the Supreme Court.[5] See Section 4.2 for an overview

[4]While Songer et al. (1994) document that rulings at courts of appeals strongly respond to Supreme Court orientation, Choi et al. (2012) argue that the Supreme Court only weakly affects courts of appeals, stating the low rate at which decisions of courts of appeals are reversed at the Supreme Court. However, what they interpret as a low risk of reversal on the side of appellate judges might simply be a sign of compliance. If appellate judges are reversal-averse, they can be expected to issue decisions in a way that reduces the risk of reversals, such that reversals will be rare in equilibrium. In this case, the threat of reversal is still an important determinant of the decisions of appellate judges. Our empirical results clearly indicate that ideological leanings of Supreme Court justices affect district court rulings – arguably passing through the courts of appeals – in a way that is consistent with reversal aversion of both district and appellate judges.

[5]For simplicity, the model only includes district court judges and the Supreme Court. However, the results from a nested model version including circuit courts are qualitatively the same. The circuit courts can be thought of

of the literature on reversal aversion.

We consider the representative (average) district court judge in state s, called judge s. The utility of judge s at time t is given by

$$U_s\left(r_{s,t}\right) = V_s\left(r_{s,t}\right) + R\left(r_{s,t}\right).$$

Overall utility is given by the private utility of judge s

$$V_s\left(r_{s,t}\right) = -\frac{\kappa}{4} \cdot \left(r_{s,t} - dci_s\right)^2,$$

where dci_s (for district court ideology) summarizes the ideological leaning and $\kappa > 0$ determines the preference weight on ideology, and by the reputational utility of judge s

$$R\left(r_{s,t}\right) = -q\left(r_{s,t}, sci_t\right),$$

where q is the probability of reversal and sci_t is the ideology of the Supreme Court.

Based on Songer et al. (1994), we postulate

$$q\left(r_{s,t}, sci_t\right) = sci_t^2 \cdot \left(x - sci_t\right)^2/4.$$

This implies that a neutral Supreme Court ($sci = 0$) overturns neither clearly liberal nor clearly conservative decisions and that an ideologically clear Supreme Court ($sci = -1$ or $sci = 1$) overturns every decision that is fully at odds with its own ideology.

Under this assumption for the behavior of the Supreme Court, the optimal behavior of a district court judge can be expressed as the following maximization problem

$$\max_{r_{s,t}} -\kappa/4 \cdot \left(r_{s,t} - dci_s\right)^2 - sci_t^2 \cdot \left(r_{s,t} - sci_t\right)^2/4.$$

Maximization with respect to the decision $r_{s,t}$ results in the first order condition

$$-2\kappa/4\left(r_{s,t} - dci_s\right) - 2 \cdot sci_t^2 \cdot \left(r_{s,t} - sci_t\right)/4 = 0,$$

which can be solved for the optimal decision

$$r_{s,t}^* = \frac{\kappa}{\kappa + sci_t^2} \cdot dci_s + \frac{sci_t^2}{\kappa + sci_t^2} \cdot sci_t.$$

It follows that the optimal decision of a district court judge is a weighted average of the judge's own ideology dci_s and Supreme Court ideology sci_t. The respective weights depend on the preference parameter κ and on the unambiguity of the ideological orientation of the Supreme

as passing through the guidelines set by the Supreme Court to the district courts – potentially imperfectly so because appellate judges may be able to incorporate their own ideological orientation.

Court. Specifically, when the Supreme Court is rather balanced ideologically (i.e., sci_t takes values close to zero), the weight on dci_s is close to one and rulings are mainly based on district court judges own preferences. By contrast, when the Supreme Court has a clear ideological leaning (i.e., sci takes values close to -1 or close to 1), rulings mainly depend on Supreme Court guidance.

Next, assume that the ideological leaning of the Supreme Court changes by $\Delta sci = sci_{t+\tau} - sci_t$. Taking the ideology dci_s of a district court judge as given, the effect of this change in Supreme Court ideology on the optimal decision of judge s can be calculated as

$$\begin{aligned}\Delta r_s =& r_{s,t+\tau} - r_{s,t} = \frac{\kappa}{\kappa + sci_{t+\tau}^2} \cdot dci_s + \frac{\kappa}{\kappa + sci_{t+\tau}^2} \cdot sci_{t+\tau} - \frac{\kappa}{\kappa + sci_t^2} \cdot dci_s - \frac{\kappa}{\kappa + sci_t^2} \cdot sci_t \\ =& \kappa \cdot \left(\frac{1}{\kappa + sci_{t+\tau}^2} - \frac{1}{\kappa + sci_t^2} \right) \cdot dci_s + \kappa \cdot \left(\frac{1}{\kappa + sci_{t+\tau}^2} \cdot sci_{t+\tau} - \frac{1}{\kappa + sci_t^2} \cdot sci_t \right).\end{aligned} \tag{4.1}$$

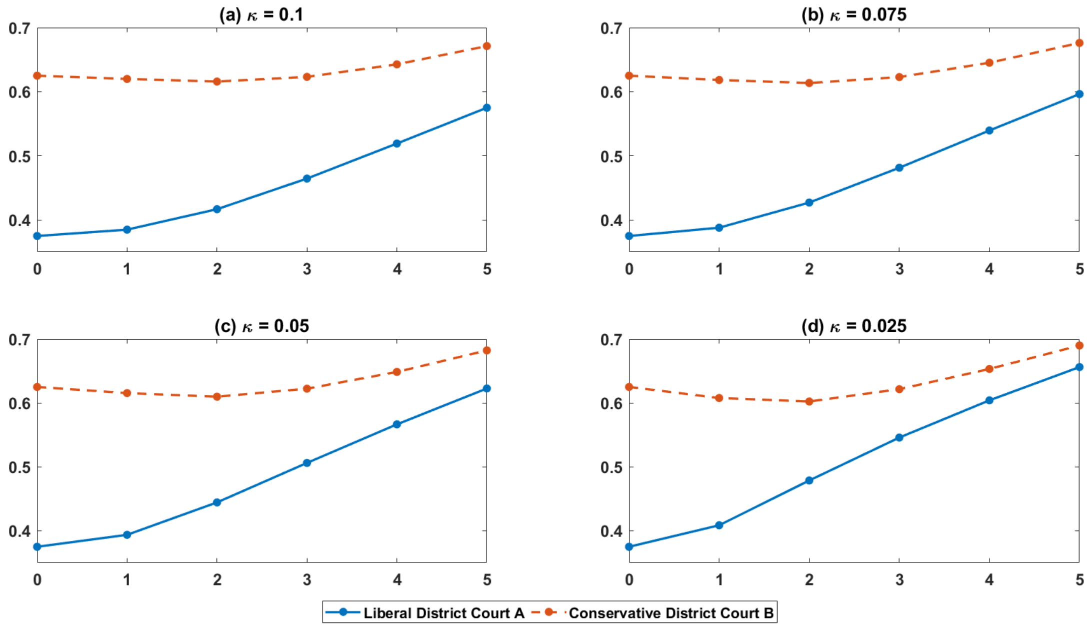
Suppose that Supreme Court ideology is positive and increases, i.e., $sci_{t+\tau} > sci_t > 0$ as in our empirical sample, see Figure 4.1. Then, as Equation (4.1) illustrates, this change in Supreme Court ideology induces an increase in the conservatism of district court rulings that is more pronounced the more liberal the considered district court judge (i.e., the lower dci_s) is. To see this, note that the second summand of the final expression in Equation (4.1) is independent of dci_s and that the first bracket in this expression is negative if $sci_{t+\tau} > sci_t > 0$, such that Δr_s decreases in dci_s.

Figure 4.2 illustrates this point in an example where Supreme Court ideology increases linearly from zero to 0.4 (a stylized description of the empirical development illustrated in Figure 4.1). We compare the share of conservative rulings of a rather liberal district court judge A with $dci_A = -0.25$ with the share of conservative rulings of a rather conservative judge B with $dci_B = 0.25$ (in our empirical sample this is roughly a comparison of New York and Wyoming). The share of pro-business rulings in a district court $\rho_{s,t}$ is linked to rulings $r_{s,t}$ through the definition $\rho_{s,t} = (1+r_{s,t})/2$. Accounting for the large reversal aversion documented in the literature, see Section 4.2, we use four relatively small values of the preference parameter κ. While rulings turn more conservative in both courts, the increase in conservatism of the Supreme Court induces rulings of the liberal district court judge A to become substantially more conservative relative to rulings of the conservative district court judge B.

In our econometric analysis, we make use of this differential impact of Supreme Court ideology across district courts. We estimate a regression with average district court rulings (where s now represents a state instead of a judge) $r_{s,t}$ as the dependent variable, year fixed effects η_t, state fixed effects δ_s, and the interaction between Supreme Court ideology and district court ideology, $sci_t \cdot dci_s$, (and control variables $X_{s,t}$ in the empirical analysis) as independent variables

$$r_{s,t} = \gamma \cdot sci_t \cdot dci_s + \beta \cdot X_{s,t} + \delta_s + \eta_t + \varepsilon_{s,t}. \tag{4.2}$$

Figure 4.2. Model-Predicted Share of Conservative District Court Rulings



Note: These graphs depict the simulated share of conservative rulings in two different district courts: a court with a liberal district court judge A ($dci_A = -0.25$) and a court with a conservative district court judge B ($dci_B = 0.25$). Supreme Court ideology increases linearly from $sci_0 = 0$ to $sci_5 = 0.4$.

To understand the role of the interaction effect in Equation (4.2), suppose we observe two states, A and B, in two years, t and $t + \tau$. In such a setting, the estimated coefficient on the interaction term $\widehat{\gamma}$ is given by

$$\widehat{\gamma} = \frac{\Delta\Delta r}{\Delta sci \Delta dci} = \frac{r_{A,t+\tau} - r_{A,t+\tau} - (r_{B,t+\tau} - r_{B,t})}{(sci_{t+\tau} - sci_t) \cdot (dci_A - dci_B)}, \tag{4.3}$$

where $\Delta\Delta r$ is the difference of the change in average rulings in the two states, Δsci is the change in Supreme Court ideology, and Δdci is the difference in ideological leanings of the two states' district courts.

In our model, $\Delta\Delta r$ is given by

$$\Delta\Delta r = \kappa \cdot \left(\frac{1}{\kappa + sci^2_{t+\tau}} - \frac{1}{\kappa + sci^2_t} \right) \cdot \Delta dci. \tag{4.4}$$

Hence, Equation (4.3) evaluates as

$$\widehat{\gamma} = -\frac{\kappa}{(\kappa + sci^2_t)(\kappa + sci^2_{t+\tau})} \cdot (sci_t + sci_{t+\tau}), \tag{4.5}$$

which is derived by substituting Equation (4.4) into Equation (4.3) and rearranging terms. Consequently, when the Supreme Court is rather conservative, the model predicts the coefficient on the interaction term to be negative.

Table 4.1 illustrates this estimation approach using the example from Figure 4.2. In particular,

Table 4.1. Illustration of the Econometric Procedure

(a) Average Rulings for Two States and Two Years

Year	State	sci	dci	r			
				$\kappa = 0.1$	$\kappa = 0.075$	$\kappa = 0.05$	$\kappa = 0.025$
0	A	0	-0.25	-0.25	-0.25	-0.25	-0.25
5	A	0.4	-0.25	0.15	0.1926	0.2452	0.3122
0	B	0	0.25	0.25	0.25	0.25	0.25
5	B	0.4	0.25	0.3423	0.3521	0.3643	0.3797

(b) Calculation of the Interaction Effect

	$\kappa = 0.1$	$\kappa = 0.075$	$\kappa = 0.05$	$\kappa = 0.025$
$\Delta r_A = r_{A,5} - r_{A,0}$	0.4	0.4426	0.4952	0.5622
$\Delta r_B = r_{B,5} - r_{B,0}$	0.0923	0.1021	0.1143	0.1297
$\Delta\Delta r = \Delta r_A - \Delta r_B$	0.3077	0.3404	0.3810	0.4324
$\Delta sci = sci_5 - sci_0$	0.4	0.4	0.4	0.4
$\Delta dci = dci_A - dci_B$	-0.5	-0.5	-0.5	-0.5
$\widehat{\gamma} = \Delta\Delta r/(\Delta sci \cdot \Delta dci)$	-1.5385	-1.7021	-1.9048	-2.1622

the upper part of the table shows the average rulings in the two states in years 0 and 5 of the example in panel form. Rulings in both states become more conservative, but the increase in conservatism is more pronounced in the state with the liberal district court judges, $\Delta r_A > \Delta r_B > 0$, $\Delta\Delta r > 0$. This implies that the interaction term is assigned a negative coefficient, $\widehat{\gamma} = \Delta\Delta r/(\Delta sci \cdot \Delta dci) < 0$. Quantitatively, for the considered values of the preference weight κ, resulting coefficients lie between -1.5 and -2.2.[6]

4.3.2 Evidence

In this Section, we empirically assess the model prediction that increasing conservatism of the Supreme Court renders district court rulings relatively more conservative in states with rather liberal district court judges by estimating Regression (4.2). In Section 4.4, we will use the thus identified ideological variation in state-specific court rulings (caused by changes at the

[6] Note from Equation (4.5) that the difference between the district court ideologies in the two states is irrelevant for the value of the estimated coefficient $\widehat{\gamma}$, as $\Delta\Delta r$ is proportional to Δdci (see Equation (4.4)). Hence, although the two district court ideologies are chosen arbitrarily for an illustrative example, the values for $\widehat{\gamma}$ in Table 4.1 are informative about what to expect for a sample where the ideology of the Supreme Court develops in a way as displayed in Figure 4.1.

Supreme Court and thus arguably exogenous to state-specific developments) as independent variable in regressions seeking to explain labor market outcomes. This makes the state the relevant level of our analysis, as detailed labor market data from the Current Population Survey is not available on a less aggregate level.

Variables, Data Sources, and Sample Selection

In the following, we describe our sample, give an overview of the variables used in our regressions, and state the sources from which these variables are obtained.

District Court Rulings For district court rulings $r_{s,t}$, we use the Carp-Manning U.S. District Court Database compiled by Carp and Manning (2016). All cases in this dataset are taken from the Federal Supplement, which is the primary source of published U.S. district court decisions. In practice, even though the publisher has no legal monopoly over the court opinions, any decision that a sitting federal district judge submits has been published in the Federal Supplement. Decisions to publish are mainly determined by the official publication guidelines and not by the judges' ideological tendencies (cf. Swenson, 2004).[7] These official guidelines generally encourage publication if the opinion lays down a new rule of law, alters an existing rule, critizises an existing rule, or changes the way in which an existing rule has been applied (cf. West Publishing Company, 1994). Hence, rulings in our dataset are rulings on cases with a high precedential value and are thus bound to be influential for a large number of other (unpublished) cases.

The database contains a total of 23,135 rulings of district courts in the 50 states from 1978 to 2011 that can be clearly labeled as either conservative (+1) or liberal (-1) and that can be categorized as Economic Regulation and/or Labor Cases. The majority of cases falling into this category are employee versus employer cases, which make up over one third of all included rulings. Cases of company versus either a union or the NLRB make up close to 15%. In general, pro-business decisions are considered to be conservative rulings. In a dispute between workers and their employer decisions in favor of the workers are regarded as liberal, whereas decisions in favor of the employer are regarded as conservative. In regulation cases, decisions for the government are considered to be liberal. Our dependent variable $r_{s,t}$ is the average ideological leaning of rulings in state s and year t. This variable would take the value 1 (-1) if all cases were decided in a liberal (conservative) way.

Supreme Court Ideology For the ideology of Supreme Court justices sci_t, we use the ideal point estimates calculated by Bailey (2013). The ideology scores from Bailey (2013) are chosen over the more common Martin-Quinn scores, since the former are able to distinguish between shifts in ideologies and shifts in case composition by using bridging information such as positions of justices on previous cases. With changing ideological leanings of Supreme Court

[7] About 20% of all cases decided in district courts are eventually published in the Federal Supplement.

justices, the case composition is bound to change as well.[8] If ideological leanings and case composition change simultaneously, the effect on liberal voting percentages of Supreme Court justices, on which the Martin-Quinn scores are based, is unclear. The use of briding information allows Bailey (2013) to disentangle the two effects.[9] In the regressions, we define sci_t as the median Bailey score of Supreme Court justices.

District Court Ideology For the ideology of district court judges dci_s, we use information on ideologies provided by Boyd (2015a).[10] The lack of data on district court judges does not allow for the use of the methodology developed in Bailey (2013) here. As the rulings of district court judges will arguably be influenced by Supreme Court ideology, we refrain from using ideology scores that are based on rulings and use scores that are calculated based on the appointment process for federal judges instead. The Boyd scores exploit the norm of senatorial courtesy: if a judge is appointed from a state where the president and a senator (both senators) share a political party, the judge is assigned the ideology score of the senator (the average of the senators), else the judge is assigned the ideology score of the president. We link the Boyd data to information on confirmation, reassignment, and retirement dates of district court judges from the Biographical Directory of Article III Federal Judges provided by the Federal Judicial Center.

While the ideology of judges at a district court is a constant in our model, in reality it can change over time due to changes in judge preferences, the confirmation of new judges, and the retirement of old ones. In order to avoid endogeneity in the ideology measure for district courts, we use the average Boyd ideology score of district court judges in state s that have been serving between 1936 (the first year for which there are ideology scores available from the Boyd dataset) and 1977 (the year before our regression sample begins). Formally, we define

$$dci_s = \frac{1}{\sum_{j \in J_s} y_j} \sum_{j \in J_s} y_j B_j,$$

where dci_s is the average ideology score for state s, J_s is the set of district judges serving in state s between 1936 and 1977, y_j is the term length for judge j in this time frame, and B_j is judge j's ideology score.

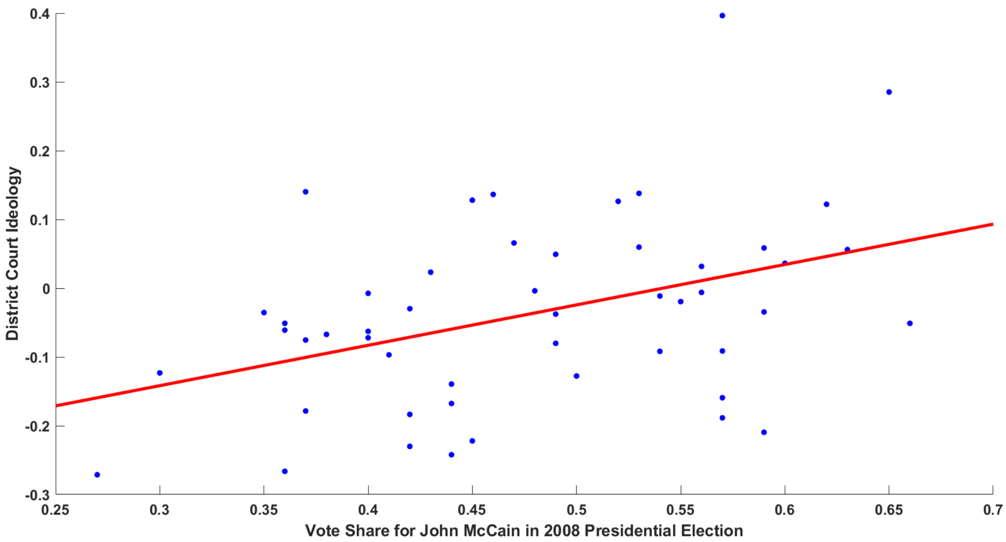
This pre-sample ideology measure is informative about the ideological leanings of a state's

[8] Cases are heard by the Supreme Court if they are supported by at least four Supreme Court justices. Thus, with more conservative justices, one would expect some cases to be chosen that would not be heard by a more liberal Supreme Court. This pertains, for example, to cases with liberal rulings of the lower courts, that a liberal Supreme Court would be very unlikely to overturn.

[9] The Bailey scores are bounded between -2 and 2, with a clearly liberal and a clearly conservative justice fixed at -1.5 and 1.5 for reference. In the data, median Baily score of the Supreme Court between 1950 and 2011 has never been below -1.1 and has never exceeded a value of 0.6. Thus, a value of -1 already constitutes an exceedingly liberal Supreme Court that can be expected to overturn overly conservative rulings at lower courts. The reverse argument holds for the value 1.

[10] The calculation of the ideology scores from Boyd (2015a) follows the methodology developed in Giles et al. (2001) and extended in Epstein et al. (2007). Ideology scores are bounded between -1 and 1.

Figure 4.3. District Court Ideology and 2008 Voting Shares for John McCain



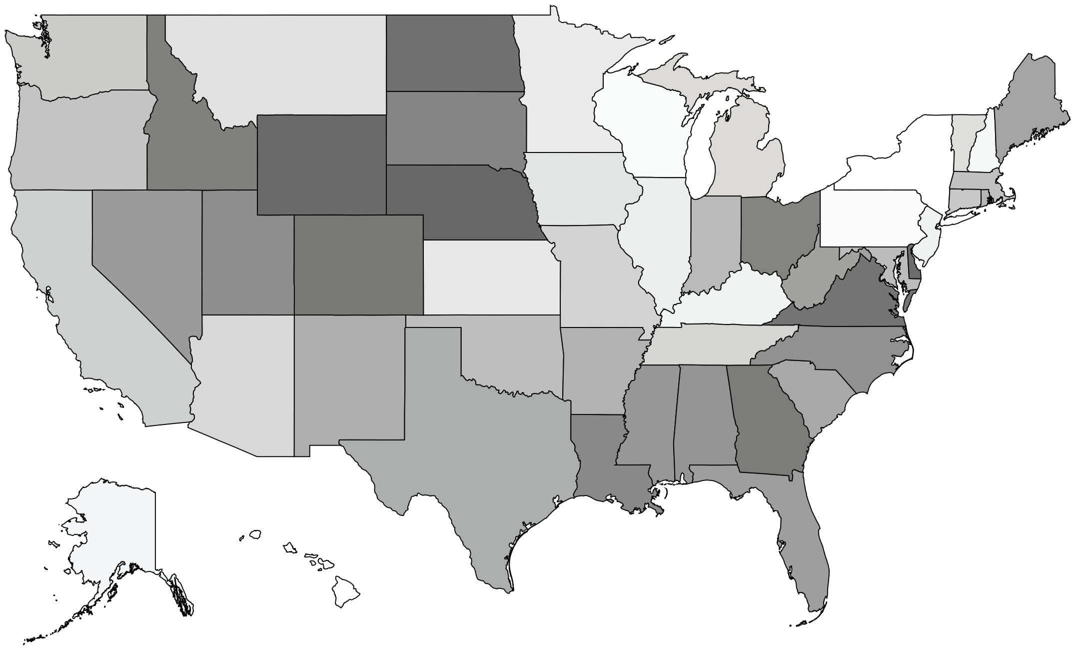
Note: This graphs plots the average Boyd ideology score of district court judges by state for the time period between 1936 and 1977 against the voting share for John McCain in the 2008 presidential election, obtained from the Federal Election Commission.

district courts judges in the regression period, as ideological leaning of judges dispay substantial persistence. Federal judges are appointed for life and hence serve (on average) long terms until they retire voluntarily. Reappointments to other courts are rare. Further, strategic retirement plays an important role at district courts, perpetuating ideological leanings beyond the current judges' retirements. Specifically, district court judges tend to retire when the current President is ideologically similar to themselves. As an extreme case, the district court for the district of North Dakota has never had a judge who was appointed by a Democratic president since 1954. In the Boyd database, the average ideology of judges at a district court is highly autocorrelated, with most district courts displaying a yearly autocorrelation of about 0.8 and some an autocorrelation of over 0.95.[11]. Thus, while the pre-sample ideology measure we use is indeed informative about judge ideology within our regression sample, it is unrelated to potentially endogenous ideological changes occurring within our sample.

In most cases, a state's district court ideology coincides with the perceived political ideology in that state, see Figure 4.3. However, there are a few exceptions like Kentucky (which has rather liberal district courts) or Delaware (which has a rather conservative district court). The correlation between the 2008 general-election voting for John McCain from the Federal Election Commission (as an indicator for a states general conservatism) and our district court ideology measure is 0.4. For our analysis, it is advantageous that this correlation is not too high, such that we can actually disentangle a state's district court ideology from the general political leaning of the state. A map depicting the liberalism/conservatism of states according to their district courts is provided in Figure 4.4. The map shows some concentration of rather liberal

[11]The evolution of the average ideology score by district court is shown in Appendix C.1

Figure 4.4. Average Ideology Score of District Court Judges by State, 1936–1977



Note: This graphs depicts the average Boyd ideology score of district court judges by state for the time period between 1936 and 1977. Darker colors indicate conservatism and lighter colors indicate liberalism.

district courts in the northeast, with New York, Pennsylvania, and New Hampshire belonging to the five states with the most liberal district court judges according to our measure. We have tested for regional variation in treatment effects, which would indicate a need to cluster standard errors despite the fact that we include state fixed effects in the regressions (cf. Abadie et al., 2017), but could not find any systematic pattern.

Control Variables We control for variables that can be expected to affect court rulings beyond the interplay between Supreme Court ideology and district court ideology. We account for the case composition, using information from the Carp-Manning database, and for judge composition along characteristics such as age, race, gender, and experience (which have been identified as determinants of rulings by the literature, see Section 4.2), using information from the Biographical Directory of Article III Federal Judges and the Carp-Manning database. We take the role of circuit courts into account by controlling for the average Boyd score of the responsible appellate judges. To ensure that our results are not driven by compositional changes at the district courts, we include the share of district court judges appointed by a Republican president.[12]. To capture non-judicial ideological forces potentially affecting district court rulings, we also control for well-known determinants of ideological leanings of the state's population, such as population size, urban density, age, and racial composition, from the CPS. Furhter determinants, like the political party of the governor and the majority parties in the state's legislative chambers, are obtained from the State Partisan Composition collected by the

[12] A detailed description of the evolution of this share by district court is provided in Appendix C.2

National Conference of State Legislatures and additionally included as controls.

Finally, since the autocorrelation of the average judge ideology score at a district court is below one, average rulings by district courts might show some tendency to converge towards the middle, i.e., rulings at initially rather liberal district courts might tend to become more conservative over time independent of developments at the Supreme Court. To pick up such mean-reverting tendencies, we also include the lagged dependent variable in the set of control variables.

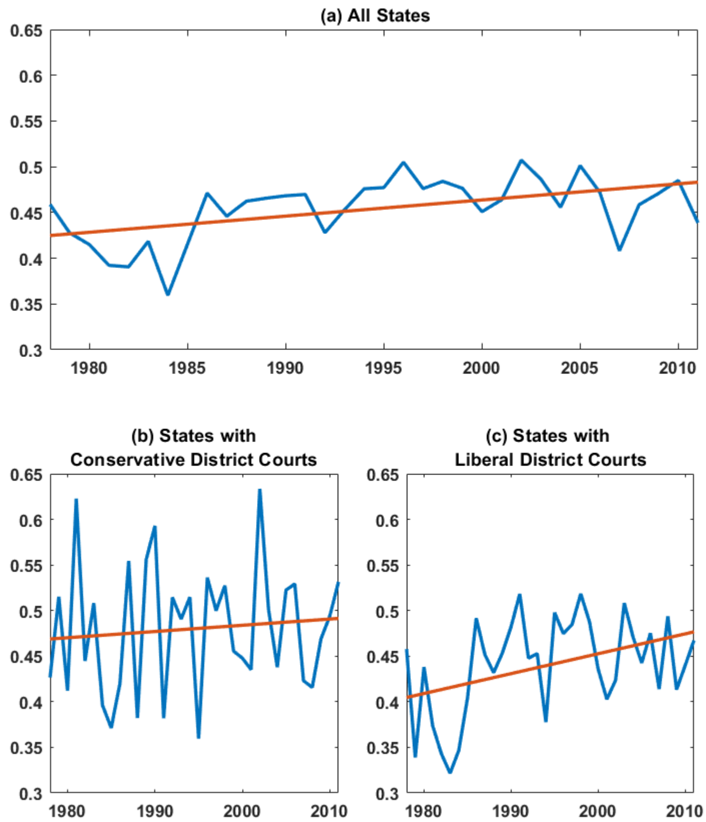
The literature, see Section 4.2, has also documented that court rulings can be affected by economic conditions. In our preferred specification we leave out economic indicators as control variables for two reasons. First, there is no obvious correlation between changes in state-specific economic outcomes and changes in Surpreme court ideology, such that the omission of economic variables is unlikely to bias the coefficient on the interaction between Supreme Court ideology and our constant measure of district court ideology. Second, we argue that economic outcomes are themselves affected by court rulings, such that including economic variables as controls would erroneously take out the correlation between economic outcomes and court rulings that is driven by causal effects *from* court rulings *to* economic outcomes. To corroborate our findings, we consider additional specifications where we include state-specific labor market outcomes and state GDP growth as controls in Appendix C.3.

Sample Selection Our sample runs from 1978 to 2011. We choose 1978 as the starting date because of two reasons. First, our measure of Supreme Court ideology reaches a value of zero in 1978 and stays above this value for the entire sample period. Thus, liberal district courts will be unambigously more affected by the shifts in Supreme Court Ideology over our entire sample period. Second, state-level labor market data is only available from the late 1970s onwards in the (CPS). The end date is chosen because ideology scores for the Supreme Court by Bailey (2013) are only available until 2011. We concentrate on the 50 states and exclude the District of Columbia because many cases heard at the district court for D.C. do not specifically relate to the D.C. labor market but concern the federal government.

In principle, our sample contains 34 years $\times$ 50 states = 1700 state-year observations. However, there are 79 state-year combinations with no rulings falling into the Economic Regulation and/or Labor Cases category. Since we also use lagged rulings as a control variable in our regressions, we lose another 62 observations due to years without rulings in certain states.[13]

[13]The Carp-Manning U.S. District Court Database does not include rulings in the economic category for Alaska in 1980, 1987, 1990, 1991, 1997, 2001, 2004, and 2009, for Arizona in 1981, for Arkansas in 2000, for Delaware in 2011, for Hawaii in 1985, for Idaho in 1977, 1979, 1991, 1994, 1997, 1999, 2000, and 2006, for Iowa in 1978, for Kentucky in 1980 and 2000, for Maine in 1977, 1978 and 1981, for Montana in 1984, 1990, 1991 and 1993, for Nebraska in 1983, 1988, 1991, 2006 and 2007, for Nevada in 1994, 2007, and 2008, for New Hampshire in 1979, 2001, 2003, and 2011, for New Mexico in 1979, 1981, 1982, 1988, 1998, and 2006, for North Dakota in 1977, 1979, 1988, 1990, 1992, 1993, 1994, 1997, 1998, 1999, and 2001, for Rhode Island in 1981, for South Dakota in 1987, 1988, 1991, 1998, and 1999, for Utah in 1978, for Vermont in 1981, 1982, 1984, 1986, 1988, 2010, and 2011, for Washington State in 1978 and 1979, and for Wyoming in 1977, 1981, 1984, 1998, 2003, 2006, 2007, and 2011.

Figure 4.5. Share of Conservative District Court Rulings in Economic and/or Labor Cases



Note: This graphs depicts the share of conservative rulings for cases in the Economic and/or Labor Cases category in the Carp-Manning U.S. District Court Database compiled by Carp and Manning (2016) for all states, for states with conservative district courts ($dci_s > 0$), and for states with liberal district courts ($dci_s < 0$). The red lines are linear trends.

Missing values for other control variables induce the loss of another 60 observations.[14] This leaves us with a consistent sample of 1499 state-year observations for which we observe all our variables.

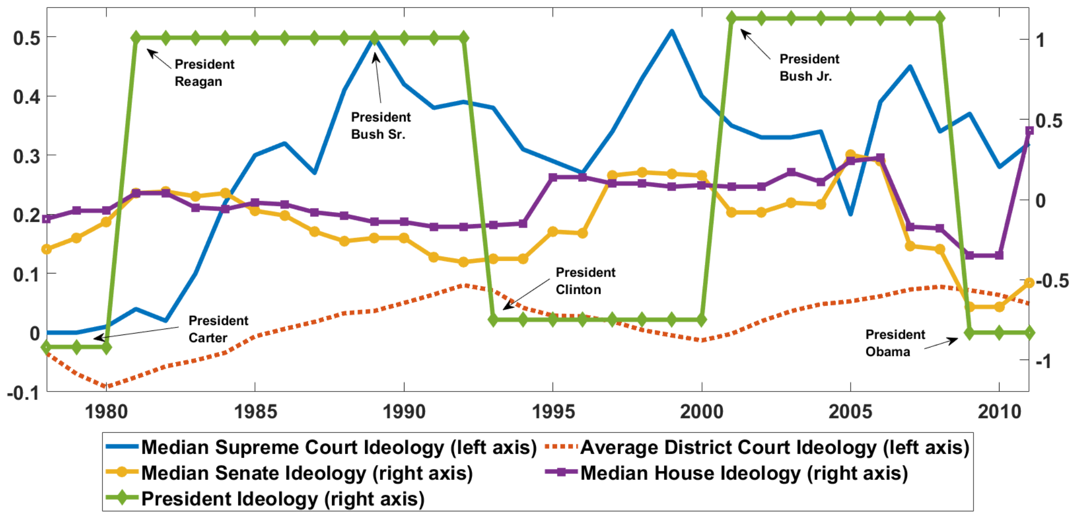
Descriptive Developments

We begin our analysis by looking descriptively at the evolution of the share of conservative rulings in the district courts. Figure 4.5 is a clear first indication that, as predicted by our model, the share of conservative rulings has increased in states with liberal district courts relative to states with conservative district courts between 1978 and 2011.

Figure 4.6 compares the evolution of the ideological leanings of Supreme Court justices to the evolution of the ideological leanings of district court judges.[15] As both Supreme Court justices and district court judges are appointed by the president, the two series naturally display a high positive correlation. Still, the ideology scores depicted in Figure 4.6 suggest that the

[14]Our urban-density variables are not reported in the CPS before 1986 for Delaware, Idaho, Maine, Montana, Nevada, New Hampshire, North Dakota, and South Dakota, as well as between 1986 and 1995 for Wyoming.
[15]Keep in mind that the ideology scores of district court judges are based on their appointment process and are thus unaffected by changes in rulings or case composition.

Figure 4.6. Ideological Leanings of Supreme Court Justices, District Court Judges, the President, the Senate, and the House of Representatives



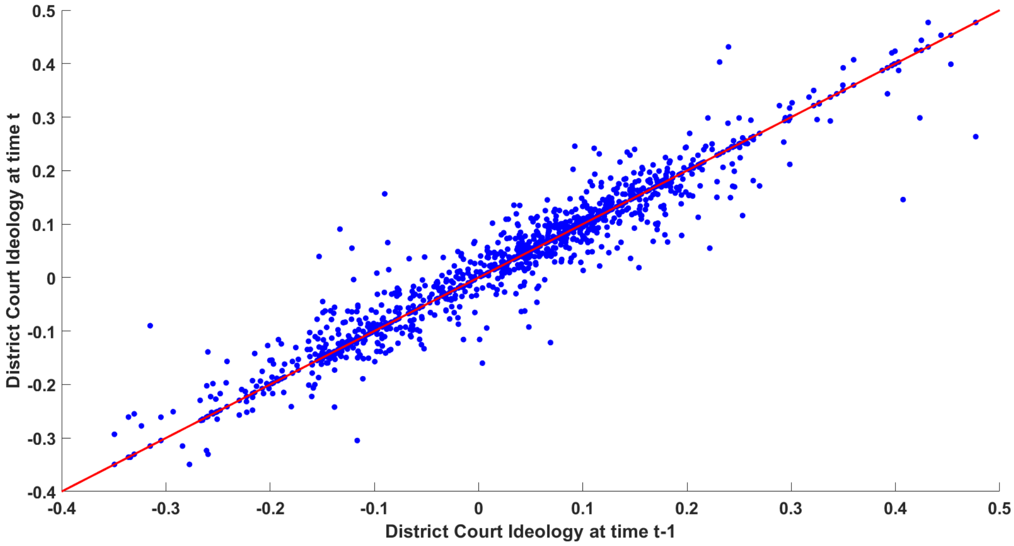
Note: This graph depicts the ideal point estimates provided in the dataset to Bailey (2013) for the ideological leanings of the median Supreme Court justice, the median senator on the U.S. Senate, the median representative in the House of Representatives, and the president between 1978 and 2011, as well as the average Boyd ideology scores for district court judges between 1978 and 2011. Again, positive values are tantamount to conservative ideological leanings, while negative values imply liberal ideological leanings.

conservative shift of district court judges is much more modest over the entire sample period. This ameliorates potential concerns that the relative increase in conservative rulings in liberal states might not be driven by ideologically unchanged district court judges following the increasingly conservative guidelines set by the Supreme Court but by a concomitant shift of district court ideology towards the conservative end of the ideological spectrum. To further address this concern, we include the share of district court judges appointed by a Republican president as a control variable in our regressions as explained above. Median ideology scores for the Senate and the House of Representatives experience only very modest changes over our sample period (which will arguably be completely captured by our time fixed effects).

Figure 4.7 plots the average ideology score of district court judges by state and year against last year's value. Most observations concentrate around the 45-degree line, indicating a high persistence in district court ideology by state. This persistence is key to our identification, which relies on long-run ideological differences between persistently rather liberal courts and persistently rather conservative courts.

A simple regression of the average ideology score of district court judges by state and year on its own lag and state fixed effects gives a coefficient on the lag of 0.92. Thus, while district court ideology is highly persistent, it displays some tendency to revert to the middle of the ideological spectrum over time, reflecting that some rather liberal (conservative) judges retire during the presidency of a Republican (Democratic) president in each year. One may argue that this induces rulings in initially rather liberal district courts to become more conservative

Figure 4.7. Correlation of Ideology Scores of District Court Judges by State and Year, 1978–2011



Note: This graph plots the average Boyd ideology score by state and year against last year's value. The 45-degree line is indicated in red.

over time, independent of ideological developments at the Supreme Court. For this reason, we include the lagged dependent variable as a control in our regressions to capture mean-reverting tendencies in rulings by state, as described above. Additionally, we also directly control for the share of judges appointed by a Republican president in our regressions.

Econometric Results

The regression results for district court rulings are reported in Table 4.2. Column (1) constitutes our most preferred specification, featuring the full set of control variables. Column (2) excludes the lagged dependent variable, Column (3) excludes all control variables except the lagged dependent variable, and Column (4) excludes all control variables.

As predicted, the coefficients on the interaction term between Supreme Court ideology and district court ideology are negative in all four specifications. This means that the shift in Supreme Court ideology did indeed induce rulings to become more conservative in states with rather liberal district courts relative to states with rather conservative district courts. Quantitatively, estimates are about -1.8 to -2 and hence fall in the range suggested by our model (see Table 4.1). Column (2) shows a somewhat larger coefficient in absolute value than Column (1), indicating that taking into account the tendency of rulings in a state to converge to the center over time is indeed important. However, as the coefficients are fairly similar, this tendency does not seem to matter too much. Columns (3) and (4), which leave out certain control variables, illustrate that our results do not depend on the specific set of included controls.

We perform several checks in order to assess the robustness of our findings. Specifically, we

Table 4.2. Regression Results for District Court Rulings

	(1)	(2)	(3)	(4)
Supreme Court ideology	-1.9673	-2.0102	-1.7978	-1.8778
$\times$ district court ideology	(0.7104)	(0.7095)	(0.6904)	(0.6905)
	p=0.0057	*p=0.0047*	*p=0.0093*	*p=0.0066*
Observations	1499	1499	1499	1499
R^2	0.2619	0.2612	0.0918	0.0885
Lagged dependent variable	yes	no	yes	no
State demographics	yes	yes	no	no
Court, judge, and case characteristics	yes	yes	no	no
State gov. and leg. controls	yes	yes	no	no
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

include controls for local labor market conditions, include higher lags of the dependent variable, weigh observations by the number of rulings per state population, and use a moving average of our measure of Supreme Court ideology sci_t. In all of these specifications the coefficient on the interaction term between Supreme Court ideology and district court ideology remains distinctly negative and highly statistically significant. See Appendix C.3 for the detailed results.

4.4 Labor Market Effects of Judicial Ideology

In this section, we exploit that our instrument $sci_t \cdot dci_s$ induces an increase in the share of pro-business rulings in states with liberal district courts relative to states with conservative district courts to estimate the effect that ideological tendencies in court rulings exert on the labor market. After presenting the empirical results, we rationalize them in a simple search and matching model with wrongful termination lawsuits.

4.4.1 Evidence

In order to identify the economic effects of jurisdiction on the labor market, we now use our instrument $sci_t \cdot dci_s$ in regressions with labor market outcomes as the dependent variable. Specifically, we estimate

$$z_{s,t} = \gamma^z \cdot sci_t \cdot dci_s + \beta^z \cdot \tilde{X}_{s,t} + \delta^z_s + \eta^z_t + \varepsilon^z_{s,t}. \tag{4.6}$$

where $z_{s,t}$ is a specific labor market outcome of interest in state s and year t. $\tilde{X}_{s,t}$ is the set of time-varying state-specific variables that can be expected to affect labor market outcomes directly. State and time fixed effects are captured by δ^z_s and η^z_t. $\varepsilon^z_{s,t}$ is the residual.

Variables and Data Sources

The interaction term $sci_t \cdot dci_s$ remains the regressor of interest.[16] Using this interaction term instead of a direct measure of court rulings in a state isolates the change in state-specific court rulings which is driven by a nation-wide development, i.e., the changing Supreme Court ideology. This strongly ameliorates any concerns about reverse causality. Judge decisions have been shown to be affected by economic conditions (cf. Ichino et al., 2003; Marinescu, 2011), but our interaction term is arguably unaffected by changing labor-market conditions in the specific state. Since the measure of district court ideology is time-invariant and determined from pre-sample data, it does by construction not react to changes in the state's economy. Furthermore, economic conditions may also affect the ideology of the Supreme Court. For example, the Great Recession with its high levels of unemployment is believed to have contributed to the election of Barack Obama in the 2008 presidential elections and thus also to the appointments of the rather liberal justices Sonia Sotomayor and Elena Kagan. Nevertheless, it is likely national economic conditions which affect the Supreme Court and, for our results to be affected by reverse causality, Supreme Court ideology would have to be affected by changes in (the distribution of) state-level labor market conditions.

Instruments such as ours have recently been criticized for causing biases, as they might be correlated to previous shocks (cf. Jaeger et al., 2018; Goldsmith-Pinkham et al., 2018). For such a correlation between instrument and responses to past shocks to drive our results, one would have to argue that unfavorable past shocks to a state's economy have led to the appointment of more liberal district court judges and are still driving economic performance in our sample, such that the recovery from those shocks drives the positive correlation between the increase in Supreme Court ideology and economic performance in states with rather liberal district court judges. We are confident that the long time period we can use for the calculation of the pre-sample measure of district court ideology makes this a minor issue for our analysis. The average judge (weighted by years in office) who influences our pre-sample measure of district

[16] See Section 4.3 for definitions and sources of sci_t and dci_s.

court ideology was appointed in 1956, more than 20 years before the start of our regression sample. Business cycle shocks are usually considered to fade a lot quicker and permanent shocks to a state's economy are taken into account by using state fixed effects.

Labor Market Outcomes For labor market outcomes, which are the dependent variables of our regressions, we draw on the Current Population Survey (CPS). The CPS is a monthly survey of about 60,000 U.S. households conducted by the United States Census Bureau. The sample is representative of the civilian noninstitutional population. We construct yearly data on state-specific unemployment rates, job-finding rates, employment rates, hourly wage rates, other job attributes, employment shares by industry and occupational group, and inequality measures using weights from the Integrated Public Use Microdata Series (IPUMS). More information on the dependent variables can be found in Appendix C.5. Due to the small sample size of the CPS in some smaller states, variables for these states are measured rather noisily. We address this issue by weighing observations by state population.

Control Variables We include the following time-varying state-specific variables that can be expected to affect labor market outcomes directly. Note that all variables that are either state-specific but constant or time-varying but determined at the national level are captured by the respective fixed effects. For example, the party holding the Presidency, which is correlated with Supreme Court ideology, see Figure 4.6, does not vary by state and its effects are hence captured by the year fixed effect.

A first set of control variables, taken from the CPS, describes the state's industry and occupational composition. It includes the employment shares in the construction, manufacturing, transportation, trade, financial, and services industries as well as employment share in abstract, routine, and manual occupations, following the categorization by Autor and Dorn (2013).

We further control for a set of state-specific policy measures. This set includes a measure of the tax burden, the state minimum wage, the state's federal intergovernmental revenue and a measure of employment protection laws in the state. The tax burden is the total amount paid in taxes by a state's residents divided by the state's total income computed by the Tax Foundation. Minimum wages are the minimum wage rates by state from Federal Reserve Economic Data. Data on the federal intergovernmental revenue of a state is taken from the State and Local Government Finance Dataset constructed by the Census Bureau through the Annual Survey of State and Local Government Finances. These revenues consist of all monies a state obtains from the federal government. Regarding employment protection, dummies for exceptions from the doctrine of at-will employment are constructed using the data provided in Autor et al. (2006a).

We include controls for state government and legislative majorities. Specifically, we add dummies indicating the party of the governor, the majority party in the state senate, and the majority party in the state house.

In robustness checks, we also include control for state demographics (like in the regressions

Table 4.3. Regression Results for Measures of Labor Market Fluidity

	(1)	(2)	(3)
Dependent variable	Unemployment rate	Job-finding rate	Employment rate
Supreme Court ideology sci_t	0.0705	-0.0565	-0.0831
$\times$ district court ideology dci_s	(0.0208)	(0.0247)	(0.0297)
	p= 0.0007	*p= 0.0223*	*p= 0.0052*
Effect of conservatism (γ^z/γ)	-0.0358	0.0287	0.0422
Observations	1499	1499	1499
R^2	0.7561	0.6414	0.8623
Industry and occupation controls	yes	yes	yes
State policy controls	yes	yes	yes
State gov. and leg. controls	yes	yes	yes
Year fixed effects	yes	yes	yes
State fixed effects	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

for court rulings) and take into account that a Republican president might affect labor market outcomes in liberal states through other channels than the judiciary. Since the party of the President is correlated with Supreme Court ideology, see Figure 4.6, and states with liberal district courts tend to be also otherwise liberal states, see Figure 4.3, such channels would bias our results. To address this possibility, we control for the interaction between a dummy for a Republican President and the 2008 general-election voting share for John McCain (as a continuous "blue-state/red-state" measure) taken from the Federal Election Commission. More information on the control variables is provided in Appendix C.5.

Table 4.4. Regression Results for Job Attributes

	(1)	(2)	(3)	(4)
Dependent variable	Avg. hourly wage rate	Vol. PT share	PT/FT wage rate	Union coverage
Supreme Court ideology sci_t	0.1739	0.0375	0.4265	0.1305
$\times$ district court ideology dci_s	(0.0707)	(0.0186)	(0.2065)	(0.0267)
	p= 0.0140	*p= 0.0447*	*p= 0.0391*	*p= 0.0000*
Effect of conservatism (γ^z/γ)	-0.0884	-0.0191	-0.2168	-0.0663
Observations	1499	1499	1499	1499
R^2	0.9933	0.8071	0.4180	0.9666
Industry and occupation controls	yes	yes	yes	yes
State policy controls	yes	yes	yes	yes
State gov. and leg. controls	yes	yes	yes	yes
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Results

Our main results are summarized in Tables 4.3–4.7. In order to discuss the direction of the effects of court rulings on the labor market, it is most intuitive to consider the effect of conservatism, calculated as the ratio of the regression coefficients γ^z from the labor market regression in Equation (4.6) and γ from the court-ruling regression in Equation (4.2). This ratio gives the change in the dependent variable per increase in conservatism of district court rulings. While the sign of this statistic facilitates understanding the effects of judical ideology, its size relates to an extreme thought experiment, as a one-unit increase in the conservatism of district court rulings corresponds to, e.g., a change in the share of pro-business decisions from 50% to 100%. We will therefore discuss the implied effects of more moderate changes in judicial ideology below.

Table 4.5. Regression Results for Occupational Employment Shares

	(1)	(2)	(3)
Dependent variable	Abstract emp. share	Routine emp. share	Manual emp. share
Supreme Court ideology sci_t	-0.0645	0.0633	0.0220
$\times$ district court ideology dci_s	(0.0248)	(0.0250)	(0.0209)
	p= 0.0093	*p= 0.0116*	*p= 0.2932*
Effect of conservatism (γ^z/γ)	0.0328	-0.0322	-0.0112
Observations	1499	1499	1499
R^2	0.9055	0.8267	0.7951
Industry and occupation controls	no	no	no
State policy controls	yes	yes	yes
State gov. and leg. controls	yes	yes	yes
Year fixed effects	yes	yes	yes
State fixed effects	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Table 4.3 shows that judicial conservatism tends to promote labor market fluidity. We find that more conservative, i.e. more pro-business, rulings increase the employment rate, reduce the unemployment rate, and increase the probability for unemployed people to find a new job. Quantitatively, a change from neutral decision making ($r = 0$), where half the cases are decided pro-business, to a situation of completely conservative decision making ($r = 1$), where all cases are decided pro-business, reduces the unemployment rate by about 3.5 percentage points, raises the job-finding rate by close to 3 percentage points, and raises the employment rate by about 4 percentage points. While these numbers appear stark at first glance, bear in mind the strong change in court decision-making considered here. A more modest increase in the share of conservative district court decisions, e.g., by ten percentage points, reduces the unemployment rate by about 0.7 percentage points.

Table 4.6. Regression Results for Industry Employment Shares

	(1)	(2)	(3)
Dependent variable	Construction emp. share	Manufacturing emp. share	Service emp. share
Supreme Court ideology sci_t	-0.0456	0.1428	-0.0930
$\times$ district court ideology dci_s	(0.0160)	(0.0297)	(0.0260)
	p= 0.0044	*p= 0.0000*	*p= 0.0004*
Effect of conservatism (γ^z/γ)	0.0232	-0.0726	0.0473
Observations	1499	1499	1499
R^2	0.6692	0.9287	0.9013
Industry and occupation controls	no	no	no
State policy controls	yes	yes	yes
State gov. and leg. controls	yes	yes	yes
Year fixed effects	yes	yes	yes
State fixed effects	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Over the 34 years in our sample, the share of pro-business rulings increased by 6 percentage points. Our findings indicate that such a development can reduce the unemployment rate by about 0.4 percentage points. Of course, such a transfer of the results from our state-level analysis to the aggregate level neglects potential general equilibrium effects of court rulings in other states. Nevertheless, we find this transfer informative and helpful to put our results into perspective.

Turning to job attributes (Table 4.4), we find that a larger share of pro-business rulings reduces average hourly wages, the employment share of voluntary part-time workers, and union coverage, while increasing the part-time hourly wage penalty. Hence, as employment increases, labor earnings, workplace flexibility (voluntary part-time employment), and job security (union coverage) all decrease. The rise in the part-time penalty can be seen as an

Table 4.7. Regression Results for Inequality

	(1)	(2)	(3)
Dependent variable	90/10 percentiles	90/50 percentiles	50/10 percentiles
Supreme Court ideology sci_t	-0.3674	-0.1973	-0.1701
$\times$ district court ideology dci_s	(0.1592)	(0.0753)	(0.1331)
	p= 0.0211	*p= 0.0089*	*p= 0.2013*
Effect of conservatism (γ^z/γ)	0.1868	0.1003	0.0865
Observations	1499	1499	1499
R^2	0.8228	0.8281	0.7006
Industry and occupation controls	yes	yes	yes
State policy controls	yes	yes	yes
State gov. and leg. controls	yes	yes	yes
Year fixed effects	yes	yes	yes
State fixed effects	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

increase in firms' ability to discriminate between different groups of workers in terms of pay, which is brought about by a lower risk of losing lawsuits.

The results concerning occupational employment shares (Table 4.5) and industry composition (Table 4.6) indicate that conservative court rulings also lead to a decline in the routine-manufacturing employment share while increasing the employment share of abstract workers and of employees in the construction and in the service sector.[17] In this sense they accelerate the hollowing out of the middle-class, as workers in routine-manufacturing jobs typically rank in the middle of the income distribution.

Finally, more pro-business rulings also contribute to rising income inequality. Table 4.7 shows results for the 90/10, 90/50, and 50/10 percentile ratios of the distribution of family

[17]Obviously, we cannot control for the state's industry-occupation composition in these regressions. For completeness, Table C.2 in Appendix C.4 shows results for further industry groups not included in Table 4.6.

income.[18] The coefficient on the interaction term is only significant for the 90/10 and 90/50 family income ratios, indicating that the increase in income inequality due to increasingly conservative district court rulings is mainly driven by increasing inequality at the top half of the income distribution.

Quantitatively, a ten percentage point increase in the share of conservative rulings decreases hourly wages by 1.7%, union coverage by 1.3 percentage points, the routine employment share by 0.6 percentage points, and income inequality, measured as the 90/10 family income ratio, by 3.7%. A back-of-the-envelope calculation hence indicates that the rise in the share of conservative court decisions from 1978 to 2011 (about 6 percentage point) can be expected to be responsible for a decrease of about 1.1% in wages, a fall in union coverage of about 0.8 percentage points, a 0.4 percentage point reduction in the routine employment share, and a 2.2% increase in the 90/10 income ratio. In this light, increasing judicial conservatism seems to have contributed, although in a moderate way, to important long-run economic development such as wage stagnation, deunionization, job market polarization, and rising inequality.

Again, we perform several checks to test the robustness of our findings.[19] Specifically, we vary the set of included controls in order to illustrate that our results are not driven by the inclusion or exclusion of certain control variables, include controls for state demographics, and include controls for the effect of a Republican president. All of these checks support our results.

Intuitively, more pro-business decisions lower costs for firms while also improving their bargaining position. This reduces unemployment at the cost of lower wages and higher inequality. In the subsequent Section 4.4.2 we develop a theoretical model of the labor market that makes this argumentation explicit. The results regarding the other considered variables can be understood in a similar way. As the union bargaining power is depressed by higher chances of pro-business rulings, incentives to join a union fall, which further contributes to lower wages and larger income inequality. Furthermore, as the adoption of new technologies proceeds slower in unionized firms due to employment protection (cf. Connolly et al., 1986; Bradley et al., 2017), lower unionization rates might also explain (at least part of) the documented changes in industry employment shares and in the occupational composition.

4.4.2 Explanation

In this section, we extend the canonical search and matching model presented in Michaillat (2012) by including the possibility of wrongful-termination lawsuits. To keep the model simple, lawsuits are introduced in a way that proceeds analogous to standard firing costs. The purpose of this exercise is to theoretically evaluate the economic effects of a conservative shift of the ideological leanings of the judiciary. We confirm that our empirical findings can be rationalized in this simple model.

[18] We consider the 80/20, 80/50, and 50/20 income ratios in Table C.3 in Appendix C.4 and find similar effects of judicial ideology.

[19] See Appendix C.4 for a detailed description.

Labor Market

The model is populated by a unit mass of risk-neutral workers that can either be employed or unemployed and searching for a job.[20] On the labor market, a continuum of firms $i \in [0, 1]$ hire workers by posting vacancies. Existing worker-firm matches are destroyed at the exogenous rate s. Newly separated workers begin searching for a job in the next period. The number of unemployed workers is given by

$$u_t = 1 - (1 - s)n_t$$

and the number of employed workers evolves according to

$$n_t = (1 - s)n_{t-1} + h_t.$$

h_t is the number of new matches, which is given by the constant-returns Cobb-Douglas matching function

$$h_t = \mu u_t^{\eta} v_t^{1-\eta},$$

where μ is the matching efficiency and η is the elasticity of the matching function with respect to the number of unemployed workers. The labor market tightness is defined as $\theta \equiv v_t/u_t$, such that the job-finding probability of a worker is given by $f(\theta_t) = h_t/u_t$ and the job-filling probability for a firm is $q(\theta_t) = h_t/v_t$. The cost of opening a vacancy is c and there is no randomness on the firm side. It follows that a firm can hire a new worker with certainty by opening $1/q(\theta_t)$ vacancies.

Firms

The setting allows for the existence of a representative firm. The real profit of this firm is given by

$$\pi_t = g(n_t) - w_t n_t - \frac{c}{q(\theta_t)} h_t,$$

where $g(n_t) = n_t$ is the production function and w_t are wages. As the production function implies that the Nash-bargained wages will not depend on the number of employed workers, the first order condition for employment is

$$1 = w_t + \frac{c}{q(\theta_t)} - \delta(1 - s)\mathbb{E}_t \left[\frac{c}{q(\theta_{t+1})} \right],$$

where δ is the discount factor. The firm hires new workers until the marginal product of labor and the discounted costs of hiring next period are equal to the marginal cost of labor, i.e., the wage and the hiring cost.

[20]There is no saving technology, which implies that workers consume their entire income in each period.

Wage Bargaining

As is standard in most of the search and matching literature, wages are determined as the solution of a generalized Nash bargaining problem. Without loss of generality, we assume that new workers are paid the same wage as incumbent workers and only enter wage negotiations in the next period. It follows that for a worker the value of being employed is

$$W_t = w_t + \delta \mathbb{E}_t \left[(1-s)W_{t+1} + sU_{t+1}\right]$$

and the value of being unemployed is

$$U_t = \delta \mathbb{E}_t \left[(1 - f(\theta_{t+1})U_{t+1} + f(\theta_{t+1})W_{t+1}\right].$$

The difference between these two value functions then gives the worker's surplus from a successful renegotiation. We explicitly assume that the costs of lawsuits are lost to the worker-firm pair (think of a firing cost as opposed to a severance payment from the firm to the worker). This choice is based on two observations. First, compared to the legal fees and court fees on both sides, actual payments from firms to workers make up a relatively small part of the costs of employee lawsuits according to the 2017 Hiscox Guide to Employee Lawsuits. Second, as the settlement payment is meant to compensate the employee for forgone earnings and emotional damage due to illegal employer behavior, it would be misleading to include these payments in the worker's value function.

On the firm side, an unsuccessful wage renegotiation and a subsequent termination of the match entails the risk of a wrongful-termination lawsuit. Thus, the firms' surplus from a successful renegotiation is the hiring cost per worker $c/q(\theta_t)$, plus the expected costs of a wrongful-termination lawsuit $\mathbb{L}$, times the probability of losing the lawsuit (1-ρ), where ρ is the probability of a pro-business ruling. Because the possibility of losing a lawsuit enters a firm's value function as a cost, it facilitates exposition to summarize judicial ideology by the loss probability from the perspective of firms in this model. The probability of a pro-business ruling ρ is linked to the average ruling r from the model presented in Section 4.3 through the definition $\rho = (1+r)/2$. Denoting the bargaining power of the worker with β, Nash bargaining solves

$$L_t - U_t = \frac{\beta}{1-\beta}\left[\frac{c}{q(\theta_t)} + (1-\rho)\cdot \mathbb{L}\right].$$

The resulting steady state wage schedule is

$$w = \frac{\beta(1-s-f(\theta))}{1-\beta}\left[\frac{c}{q(\theta)} + (1-\rho)\cdot \mathbb{L}\right].$$

Table 4.8. Parameter Calibration

Symbol	Interpretation	Value	Source/Target
δ	Discount factor	0.999	5% Annual discount rate
s	Separation rate	0.0095	Michaillat (2012) using JOLTS
μ	Matching efficiency	0.233	Michaillat (2012) using JOLTS
η	Unemployment-elasticity of matching	0.5	Petrongolo and Pissarides (2001)
c	Vacancy posting costs	0.32	0.32 x steady state wage
β	Bargaining power of workers	0.5	Shimer (2005)
ρ	Probability of pro-business ruling	0.55	Carp and Manning (2016)
$\mathbb{L}$	Cost of lost lawsuit	0.78	6.4% unemployment rate

Theoretical Effects of More Pro-Business Rulings

Now that we have derived both the first order condition for employment and the wage schedule, we consider the effects of an increase in the probability of pro-business decisions on labor market outcomes in our simple model.

As the lower expected average cost of a lawsuit reduces the employer's surplus from wage negotiations, the employer is able to enforce lower wages. Intuitively, the employer can credibly claim that the continuation of the worker-firm relationship is less valuable, as a lawsuit upon termination hurts the firm less. The effect of an increase in the share of pro-business rulings ρ on the steady state wage is given by

$$\begin{aligned}\frac{\partial w}{\partial \rho} &= -\frac{\beta}{1-\beta}\frac{\partial f(\theta)}{\partial \theta}\frac{\partial \theta}{\partial w}\frac{\partial w}{\partial \rho}\left[\frac{c}{q(\theta)} + (1-\rho)\cdot \mathbb{L}\right] \\ &\quad - \frac{\beta(1-s-f(\theta))}{1-\beta}\left[\frac{c}{q(\theta)^2}\frac{\partial q(\theta)}{\partial \theta}\frac{\partial \theta}{\partial w}\frac{\partial w}{\partial \rho} + \mathbb{L}\right].\end{aligned}$$

As the increase in ρ reduces labor costs, firms will post more vacancies which attenuates the negative effect of ρ on the wage rate. This leads to an increase in the steady state labor market tightness θ by

$$\frac{\partial \theta}{\partial w} = \frac{1}{\eta}\left[\frac{1-w}{[1-\delta(1-s)]\,c}\right]^{\frac{1-\eta}{\eta}}\frac{1}{[1-\delta(1-s)]\,c}$$

and consequently to an increase in the job-finding rate. Using the Beveridge curve, steady state employment increases by

$$\frac{\partial n}{\partial \theta} = \frac{1}{((1-s)+s/f(\theta))^2}\frac{s}{f(\theta)^2}\frac{\partial f(\theta)}{\partial \theta}.$$

Table 4.9. Theoretical Effects of Pro-Business Rulings

	Probability of a pro-business ruling increases by...		
	5 ppt.	10 ppt.	15 ppt.
Unemployment	-0.35 ppt.	-0.66 ppt.	-0.95 ppt.
Employment	+0.38 ppt.	+0.67 ppt.	+0.95 ppt.
Job-finding rate	+3.71 ppt.	7.43 ppt.	11.17 ppt.
Wage	-0.06 ppt.	-0.12 ppt.	-0.17 ppt.

Note: The entries in this table represent percentage point changes in a specific labor market outcome following an increase of 5, 10, and 15 percentage points in the probability of a pro-business ruling.

Quantitative Evaluation

In this section, we calibrate the model to match quarterly U.S. data for the time period between 1978 and 2011. The calibrated model is used to quantitatively evaluate the effect of an increase in the share of pro-business rulings in the model.

Calibration In calibrating the model we follow the calibration strategy used in Michaillat (2012). Table 4.8 lists the parameter values and the source that encourages the specific choice. The discount factor is set to $\beta_c = 0.999$, to match an annual discount rate of 5%. The parameter values for the separation rate s and the matching efficiency μ are taken from Michaillat (2012), who provides estimates based on the Job Opening and Labor Turnover Survey (JOLTS) for the time period between 2000 and 2009. The calibration targets for the matching elasticity η and for the bargaining power β are standard in the literature (cf. Petrongolo and Pissarides, 2001; Shimer, 2005). Following Michaillat (2012), who bases his estimates on studies by Barron et al. (1997) and Silva and Toledo (2009), the vacancy posting costs c are calibrated to 32% of the steady state wage. The probability of a pro-business ruling ρ is calibrated to match the average share of pro-business rulings in district courts in the Carp-Manning database. Finally, we calibrate the cost of a lost lawsuit $\mathbb{L}$ to 0.78 in order to match the average unemployment rate of 6.4% over the time period between 1978 and 2011.

Simulation Results We use the calibrated model to assess the theoretical effect of an increase in the share of pro-business rulings ρ. The results are summarized in Table 4.9. A ten percentage point increase in the probability of winning a lawsuit lowers the simulated unemployment rate by about 0.66 percentage points (compared to a decrease of 0.7 percentage points in the data). Given the simplicity of our model, these two numbers are surprisingly close.[21]

[21] Due to the increased labor market tightness, wage changes are significantly smaller compared to our empirical results. However, in the data wage decreases are likely magnified by the decreasing union coverage of workers

4.5 Conclusion

In this paper, we have documented substantial economic effects of ideological tendencies in court rulings. In a first stage, we have shown that the share of conservative rulings has increased in states with rather liberal district courts relative to states with rather conservative district courts following the shift of the Supreme Court towards the conservative end of the ideological spectrum since the late 1970s. In a second stage, we have exploited these differential effects on U.S. states in order to identify the economic impact of a conservative shift in ideological tendencies of the judiciary. We find that an increase in the share of conservative rulings substantially increases the employment rate and promotes labor market fluidity but also contributes to wage stagnation, job market polarization, deunionization, and rising income inequality.

which the model abstracts from. While an increase of 7.4 percentage points in the job-finding rate appears stark at first glance, keep in mind that the model is calibrated to quarterly frequency, whereas we look at weekly job-finding rates in the data.

5 Conclusion

This thesis provides three main conclusions, each related to one of the three preceding chapters.

First, I illustrate that a search and matching model with both financial frictions and wage rigidity is able to replicate important unemployment dynamics which are lacking in canonical models of the labor market. I find that wage rigidity is responsible for the steeper increase in the unemployment rate in recessions, whereas credit constraints ensure that the decrease in recovery periods is flatter. Joblessness occurs when recovery periods are characterized by concomitantly eroding credit conditions.

Second, Anna Hartmann and I provide empirical, analytical, and simulated results on the effects of routine-biased technical change on unionization rates. Employing a search and matching model with an occupational choice and endogenous union formation, we illustrate that manual workers, who benefit from the changing demand structure, are discouraged from voting in favor of a collective bargaining agreement, as unions aiming at equal wage growth would reallocate part of these benefits to routine workers. This result emphasizes routine-biased technical change as an important driver of both job market polarization and deunionization.

Third, Anna Hartmann, Christian Bredemeier, and I expose substantial effects of judicial ideology on the labor market. The economic impact is identified by exploiting differences in the effect of a shift in U.S. Supreme Court ideology on district court rulings across U.S. states. An increase in the share of conservative rulings increases the employment rate and promotes labor market fluidity but also contributes to wage stagnation, job market polarization, deunionization, and rising income inequality.

A Appendix to Chapter 2

A.1 Household's Maximization Problem

Having defined the state vectors in Section 2.3, the household maximization problem can be written as

$$\mathbb{H}_t(\omega_t^h;\Omega_t) = \max_{\{c_t,a_t\}} \ln(c_t) - \varphi n_t + \mathbb{E}_t\beta_h\mathbb{H}_{t+1}(\omega_{t+1}^h;\Omega_{t+1})$$

subject to

$$c_t + \frac{a_{t+1}}{R_t} + T_t \leq w_t n_t + a_t + u_t s. \tag{A.1}$$

A.2 Derivation of the Enforcement Constraint

The derivation of the enforcement constraint follows the derivation in Garín (2015). Referring to the respective optimization problem, the value of a firm can be written as

$$\mathbb{J}(\omega_t^c;\Omega_t) = d_t + \mathbb{E}_t\Lambda_{t+1|t}^c\mathbb{J}(\omega_{t+1}^c;\Omega_{t+1}).$$

With the possibility of default before the loan is due and after production is realized, the value of not defaulting is

$$\nu^{f,n} = \mathbb{E}_t\Lambda_{t+1|t}^c\mathbb{J}(\omega_{t+1}^c;\Omega_{t+1}).$$

In the case of default, firms and lenders renegotiate. If an agreement is reached, firms pay lenders a fraction ν_t of the continuation value. Therefore, the value of a successful renegotiation is

$$\nu^{f,s} = \mathbb{E}_t\Lambda_{t+1|t}^c\mathbb{J}(\omega_{t+1}^c;\Omega_{t+1}) + l_t - \nu_t,$$

where firms continue to produce, get another loan l_t, but have to pay a part of the continuation value to the lenders. As production cannot be seized by lenders in the case of default, the value of an unsuccessful renegotiation for the firm is simply $\nu^{f,u} = l_t$. Consequently, the net value of an agreement is given by

$$\nu^{f,net} = \mathbb{E}_t \Lambda^c_{t+1|t} \mathbb{J}(\omega^c_{t+1}; \Omega_{t+1}) - \nu_t.$$

From the perspective of a lender, the value of a successful renegotiation is

$$\nu^{l,s} = \nu_t + \frac{b_{t+1}}{R_t}.$$

In the case that no agreement is reached, lenders cannot seize production. As they do not value the stock of workers in the firm, the value of an unsuccessful renegotiation is

$$\nu^{l,u} = \eta_t q_{k,t} k_t$$

from the lender's point of view. This results in the net value of an agreement for the lender of

$$\nu^{v,net} = \nu_t + \frac{b_{t+1}}{R_t} - \eta_t q_{k,t} k_t.$$

The joint surplus of renegotiating is the sum of the net values of the firm and the lender. Since financial intermediaries have no bargaining power in the renegotiation of debt, the firm gets the value

$$\nu^{f,d} = \mathbb{E}_t \Lambda^c_{t+1|t} \mathbb{J}(\omega^c_{t+1}; \Omega_{t+1}) + l_t + \frac{b_{t+1}}{R_t} - \eta_t q_{k,t} k_t$$

in case of a default. This value is equal to its liquidity plus the joint surplus of renegotiating the debt. In order to rule out defaults, the value of not defaulting for the firm has to be at least as large as the value of defaulting. Using this inequality and rearranging terms results in the enforcement constraint

$$l_t + \frac{b_{t+1}}{R_t} \leq \eta_t q_{k,t} k_t, \tag{A.2}$$

which constrains a firm's ability to borrow below the value of the fraction of the physical capital stock that lenders can recuperate after default.

A.3 Firm's Maximization Problem

Using the assumptions made in Section 2.3, the optimization problem of the firm can be summarized by

$$\mathbb{J}_t(\omega^c_t; \Omega_t) = \max_{\{d_t, m_t, i_t, k_{t+1}, b_{t+1}\}} d_t + \mathbb{E}_t \Lambda^c_{t|t+1} \mathbb{J}_{t+1}(\omega^c_{t+1}; \Omega_{t+1})$$

subject to the budget constraint

$$z_t n_t^\alpha k_t^{1-\alpha} + \frac{b_{t+1}}{R_t} = d_t + b_t + w_t n_t + i_t + \frac{\kappa}{2} \left(\frac{m_t}{q(\theta_t)} \right)^2, \tag{A.3}$$

the law of motion for the capital stock

$$k_{t+1} = (1-\delta)k_t + \left[\frac{\delta^{\xi}}{1-\xi}\left(\frac{i_t}{k_t}\right)^{1-\xi} - \frac{\xi\delta}{(1-\xi)}\right]k_t, \tag{A.4}$$

the law of motion for employment,

$$n_t = (1-s)n_{t-1} + m_t, \tag{A.5}$$

and the borrowing constraint

$$d_t + w_t n_t + \frac{\kappa}{2}\left(\frac{m_t}{q(\theta_t)}\right)^2 + i_t + b_t \leq \eta_t q_{k,t} k_t, \tag{A.6}$$

where the loan l_t is replaced by the working capital requirements. Denoting the multipliers on the budget constraint, the law of motion for the capital stock, the law of motion for employment, and the borrowing constraint with $\mu_{c,t}$, $\mu_{k,t}$, $\mu e, t$, and $\mu_{b,t}$ and taking derivatives results in the following first order conditions:

$$\mu_{c,t} = 1 - \mu_{b,t} \tag{A.7}$$

$$\mu_{e,t} = \frac{\kappa m_t}{q(\theta_t)^2} \tag{A.8}$$

$$\mu_{k,t} = \frac{1}{\delta^{\xi}\left(\frac{k_t}{i_t}\right)^{\xi}} \tag{A.9}$$

$$\begin{aligned}\mu_{k,t} =& \mathbb{E}_t \Lambda^c_{t|t+1}\Bigg\{((1-\alpha)z_{t+1}n^{\alpha}_{t+1}k^{-\alpha}_{t+1})(1-\mu_{b,t+1}) - \frac{i_{t+1}}{k_{t+1}} \\ &+ \mu_{k,t+1}\left[(1-\delta) + \frac{\delta^{\xi}}{1-\xi}\left(\frac{i_t}{k_t}\right)^{1-\xi} - \frac{\xi\delta}{(1-\xi)}\right]\Bigg\} \\ &+ \mathbb{E}_t \Lambda^c_{t|t+1}\mu_{b,t+1}\eta_{t+1}q_{k,t+1}\end{aligned} \tag{A.10}$$

$$\frac{1}{R_t} = \mathbb{E}_t \Lambda^c_{t|t+1}\frac{1}{1-\mu_{b,t}}, \tag{A.11}$$

where q_k is the ratio of the Lagrange multipliers $\mu_{k,t}$ and $\mu_{c,t}$. As is standard in the literature, $q_{k,t}$ represents the value of the installed capital relative to its replacement costs.

A.4 Wage Bargaining

In order to derive the wage schedule, I first define the value functions $\mathbb{H}_{m,t}$, $\mathbb{H}_{n,t}$, and $\mathbb{H}_{u,t}$. $\mathbb{H}_{m,t}$ is the marginal value of having an additional member matched from the household perspective, $\mathbb{H}_{n,t}$ the value function associated with having an additional member employed, and $\mathbb{H}_{u,t}$ the value function associated with having an additional member unemployed. Using the law of motion for employment, $n_t = (1-x)n_{t-1} + f(\theta)u_t$, $\mathbb{H}_{n,t}$ can be written as

$$\mathbb{H}_{n,t} = -\varphi + \lambda_t w_t + \beta \mathbb{E}_t\{x[1-f(\theta_{t+1})]\mathbb{H}_{u,t+1} + [1-x+xf(\theta_{t+1})]\mathbb{H}_{n,t+1}\}$$

and $\mathbb{H}_{u,t}$ as

$$\mathbb{H}_{u,t} = \lambda_t s + \beta \mathbb{E}_t\{f(\theta_{t+1})]\mathbb{H}_{n,t+1} + [1-f(\theta_{t+1})]\mathbb{H}_{u,t+1}\},$$

where λ_t is the Lagrange multiplier on the household's budget constraint. The marginal value of a match, $\mathbb{H}_{m,t} = \frac{\mathbb{H}_{n,t}-\mathbb{H}_{u,t}}{\lambda_t}$ is therefore given by

$$\mathbb{H}_{m,t} = -\frac{\varphi}{\lambda_t} + w_t - s + (1-x)\mathbb{E}_t \Lambda^h_{t+1|t}[1-f(\theta_{t+1})]\mathbb{H}_{m,t+1}, \tag{A.12}$$

where $\Lambda^h_{t+1|t} = \beta_h \mathbb{E}_t \frac{\lambda_{t+1}}{\lambda_t}$ is the household's stochastic discount factor.

With both value functions defined, the wage that solves the bargaining problem can be expressed as

$$w_t^* = \text{arg} \max_{w_t} \ \mathbb{J}_{n,t}^{1-\phi}\mathbb{H}_{m,t}^{\phi},$$

where ϕ is the bargaining power of the worker. Taking the first derivative results in the standard first order condition

$$\phi \frac{\partial \mathbb{H}_{m,t}}{\partial w_t}\mathbb{J}_{n,t} + (1-\phi)\frac{\partial \mathbb{J}_{n,t}}{\partial w_t}\mathbb{H}_{m,t} = 0,$$

that can be rewritten as

$$\phi \mathbb{J}_{n,t} = (1-\phi)\mathbb{H}_{m,t}.$$

In the next step I define the total surplus of the match $\mathbb{S}_t$ as the sum of the firm's and the worker's surplus. This results in

$$\begin{aligned}\mathbb{S}_t = &(\alpha z_t n_t^{\alpha-1}k_t^{1-\alpha})(1-\mu_{b,t}) - s - \varphi c_t \\ &+ (1-x)\mathbb{E}_t\{\Lambda^c_{t+1|t}\mathbb{J}_{n,t+1}\Lambda^h_{t+1|t}[1-f(\theta_{t+1})]\mathbb{H}_{m,t+1}\}.\end{aligned}$$

Using $\mathbb{H}_{m,t}$ I can write $\mathbb{H}_{m,t} = \phi\mathbb{S}_t$ and $\mathbb{J}_{n,t} = (1-\phi)\mathbb{S}_t$. Multiplying the total surplus with $(1-\phi)$, using $\mathbb{J}_{n,t} = (1-\phi)\mathbb{S}_t$ and the first order condition

$$\begin{aligned}\mathbb{J}_{n,t} = &(1-\phi)(\alpha z_t n_t^{\alpha-1}k_t^{1-\alpha})(1-\mu_{b,t}) - (1-\phi)\left[s + \varphi c_t\right] \\ &+ (1-\phi)(1-x)\mathbb{E}_t\{\Lambda^c_{t+1|t}\mathbb{J}_{n,t+1}\} \\ &+ \phi(1-x)\mathbb{E}_t\{\Lambda^h_{t+1|t}[1-f(\theta_{t+1})]\mathbb{J}_{n,t+1}\}.\end{aligned}$$

In the last step I replace $\mathbb{J}_{n,t}$ by the value function of the firm and use that $\mathbb{J}_{n,t+1} = \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2}$.

Rearranging terms gives the wage equation

$$
\begin{aligned}
w_t^* = \phi &\left[\alpha z_t n_t^{\alpha-1} k_t^{1-\alpha}(1-\mu_{b,t}) + (1-s)\mathbb{E}_t \left\{ \Lambda^c_{t+1|t} \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2} \right\} \right] \\
&+ (1-\phi)\left[s + \varphi c_t\right] \\
&- \phi(1-x)\mathbb{E}_t \left\{ \Lambda^h_{t+1|t}[1-f(\theta_{t+1})] \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2} \right\}.
\end{aligned} \tag{A.13}
$$

A.5 Role of Investment Adjustment Costs

In this section I deviate from the conservative choice of investment adjustment costs in the main text. A parameter value at the lower end of the range found in the literature was chosen for the quantitative analysis because higher investment adjustment costs increase the incentive for firms to keep fluctuations in investment low. This increases the volatility of labor market variables and strengthens the asymmetric behavior of unemployment.

To asses the theoretical ability of the model mechanisms to generate positive skewness in the unemployment rate, the investment adjustment costs are increased by a factor ten to the value used in an early version of Perri and Quadrini (2018), $\xi = 0.5$. Under the new parametrization, the skewness of the unemployment rate increases to 0.92 in levels (0.72 in the data) and 0.70 in changes (1.30 in the data). All of the skewness in levels and over 50$\%$ of the skewness in changes is explained by financial frictions in conjunction with wage rigidity under high investment adjustment costs.

Even though investment adjustment costs can generate skewed unemployment dynamics in my model, I confirm that they are not the main driving force behind the results in this paper. Setting the adjustment costs to zero reduces the skewness of the simulated unemployment rate by no more than 5$\%$. This implies that a value of $\xi = 0.05$ is already small enough as not to blur the emphasized mechanism.

A.6 Benchmark Model

The model described in this section exhibits perfect credit markets and flexible wages. It acts as a benchmark in the quantitative analysis in order to gauge the performance of the complete model with financial frictions and wage rigidity. The benchmark model is based on the appendix to Garín (2015). The labor market and the household sector are unchanged by the introduction of financial frictions. Therefore, the first two subsections from Section 2.3 hold for this model as well.

A.6.1 Firms

Under the same assumptions as in Section 2.3, the maximization problem of the firm can be expressed as

$$\mathbb{J}_t(\omega_t^e;\Omega_t) = \max_{\{d_t,m_t,i_t,k_{t+1},b_{t+1}\}} d_t + \mathbb{E}_t\Lambda^c_{t|t+1}\mathbb{J}_{t+1}(\omega_{t+1}^e;\Omega_{t+1})$$

subject to the budget constraint

$$z_t n_t^{\alpha} k_t^{1-\alpha} + \frac{b_{t+1}}{R_t} = d_t + b_t + w_t n_t + i_t + \frac{\kappa}{2}\left(\frac{m_t}{q(\theta_t)}\right)^2, \tag{A.14}$$

the law of motion for the capital stock

$$k_{t+1} = (1-\delta)k_t + \left[\frac{\delta^{\xi}}{1-\xi}\left(\frac{i_t}{k_t}\right)^{1-\xi} - \frac{\xi\delta}{(1-\xi)}\right]k_t \tag{A.15}$$

and the law of motion for employment

$$n_t = (1-x)n_{t-1} + m_t. \tag{A.16}$$

Denoting the multipliers on the budget constraint, the law of motion for the capital stock, and the law of motion for employment, with $\mu_{c,t}$, $\mu_{k,t}$, and $\mu_{e,t}$ respectively, and taking derivatives results in the following first order conditions:

$$\mu_{c,t} = 1 \tag{A.17}$$

$$\mu_{e,t} = \frac{\kappa m_t}{q(\theta_t)^2} \tag{A.18}$$

$$\mu_{k,t} = \frac{1}{\delta^{\xi}\left(\frac{i_t}{k_t}\right)^{-\xi}} \tag{A.19}$$

$$\begin{aligned}\mu_{k,t} =& \mathbb{E}_t\Lambda^c_{t|t+1}\Bigg\{(1-\alpha)z_{t+1}n_{t+1}^{\alpha}k_{t+1}^{-\alpha} - \frac{i_{t+1}}{k_{t+1}} \\ &+ \mu_{k,t+1}\left[(1-\delta) + \frac{\delta^{\xi}}{1-\xi}\left(\frac{i_t}{k_t}\right)^{1-\xi} - \frac{\xi\delta}{(1-\xi)}\right]\Bigg\}\end{aligned} \tag{A.20}$$

$$\frac{1}{R_t} = \mathbb{E}_t\Lambda^c_{t|t+1}. \tag{A.21}$$

Under these specifications, the marginal value of an additional worker for the firm is given by

$$\mathbb{J}_{n,t} = \left[\alpha z_t n_t^{\alpha-1}k_t^{1-\alpha} - w_t\right] + (1-x)\mathbb{E}_t\Lambda^c_{t|t+1}\mathbb{J}_{n,t+1}. \tag{A.22}$$

The first term is the the net return of an additional worker for the firm. The second term is the discounted benefit of having an additional worker in the next period. Combining the marginal value with Equation (A.18) yields the job creation condition

$$\frac{\kappa m_t}{q(\theta_t)^2} = \left[\alpha z_t n_t^{\alpha-1} k_t^{1-\alpha} - w_t\right] + (1-x)\mathbb{E}_t \Lambda^c_{t|t+1} \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2}. \tag{A.23}$$

Again, the firm hires additional workers until the marginal costs of hiring equal the marginal benefits. In contrast to the model in Section 2.3, the job creation condition does not depend on financial conditions.

A.6.2 Wage Bargaining

Since the financial frictions introduced in Section 2.3 only affect the firm side, the marginal value of the match for the household is the same as in the model with financial frictions. With both value functions defined, the wage that solves the bargaining problem can be expressed as

$$w_t^* = \arg\max_{w_t} \ \mathbb{J}_{n,t}^{1-\phi} \mathbb{H}_{m,t}^{\phi},$$

where ϕ is a parameter that governs the bargaining power of the worker and the firm. Taking the first derivative results in the standard first order condition

$$\phi \frac{\partial \mathbb{H}_{m,t}}{\partial w_t} \mathbb{J}_{n,t} + (1-\phi) \frac{\partial \mathbb{J}_{n,t}}{\partial w_t} \mathbb{H}_{m,t} = 0$$

that can be rewritten as

$$\phi \mathbb{J}_{n,t} = (1-\phi)\mathbb{H}_{m,t}.$$

As a next step I define the total surplus of the match $\mathbb{S}_t$ as the sum of the firm's and the worker's surplus. This results in

$$\begin{aligned}\mathbb{S}_t = {} & \alpha z_t n_t^{\alpha-1} k_t^{1-\alpha} - s - \varphi c_t \\ & + (1-x)\mathbb{E}_t\{\Lambda^c_{t+1|t}\mathbb{J}_{n,t+1} + \Lambda^h_{t+1|t}[1 - f(\theta_{t+1})]\mathbb{H}_{m,t+1}\},\end{aligned}$$

where I use that $\lambda_t = u'(c_t)$. Next, I use Equation (A.12), $\mathbb{H}_{m,t} = \phi\mathbb{S}_t$ and $\mathbb{J}_{n,t} = (1-\phi)\mathbb{S}_t$. Multiplying the total surplus with $(1-\phi)$, using $\mathbb{J}_{n,t} = (1-\phi)\mathbb{S}_t$, $\phi\mathbb{J}_{n,t} = (1-\phi)\mathbb{H}_{m,t}$ and rearranging terms gives

$$\begin{aligned}\mathbb{J}_{n,t} = {} & (1-\phi)(\alpha z_t n_t^{\alpha-1} k_t^{1-\alpha}) - (1-\phi)\left[s + \varphi c_t\right] \\ & + (1-\phi)(1-x)\mathbb{E}_t\{\Lambda^c_{t+1|t}\mathbb{J}_{n,t+1}\} \\ & + \phi(1-x)\mathbb{E}_t\{\Lambda^h_{t+1|t}[1 - f(\theta_{t+1})]\mathbb{J}_{n,t+1}\}.\end{aligned}$$

In the last step I replace $\mathbb{J}_{n,t}$ by the value function and use that $\mathbb{J}_{n,t+1} = \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2}$. Rearranging terms gives the wage equation

$$
\begin{aligned}
w_t = \phi & \left[(\alpha z_t n_{c,t}^{\alpha-1} k_t^{1-\alpha}) + (1-x)\mathbb{E}_t \left\{ \Lambda^c_{t+1|t} \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2} \right\} \right] \\
& + (1-\phi)\left[s + \varphi c_t \right] \\
& - \phi(1-x)\mathbb{E}_t \left\{ \Lambda^h_{t+1|t} [1 - f(\theta_{t+1})] \frac{\kappa m_{t+1}}{q(\theta_{t+1})^2} \right\}.
\end{aligned} \tag{A.24}
$$

A.6.3 Equilibrium

With the model completetly described, I define the equilibrium.

Definition 3. *A recursive equilibrium is defined as a set of i) firm's policy functions* $d(\omega^c; \Omega)$, $k(\omega^c; \Omega)$, $i(\omega^c; \Omega)$, *and* $v(\omega^c; \Omega)$*; ii) household's policy function* $c(\omega^h; \Omega)$*; iii) a lump-sum tax* $T(\Omega)$*, iv) prices* $w(\Omega)$ *and* $R(\Omega)$*; and v) a law of motion for the aggregate states,* $\Omega' = \Psi(\Omega)$*, such that: i) firm's policies satisfy the firm's first order conditions (Equations (A.17)–(A.21)) and the job creation condition (Equation (A.23)); ii) household's policy function satisfies the household's first order condition (Equation (2.2)), iii) the wage is determined by Equation (A.24); iv) the law of motion* $\Psi(\Omega)$ *is consistent with individual decisions and with the stochastic process for technology z, and v) the government has a balanced budget such that* $s(1-n) = T$*.*

A.6.4 Simulation Results

The model without financial frictions and without wage rigidity is calibrated to match the same steady state values as the complete model. Simulation results are given in Table A.1 and are strikingly close to the results obtained by Shimer (2005) in the simulation of a model with only labor productivity shocks. While the model performance is good along several important dimensions, it is unable to match the high volatility of the key labor market variables unemployment, vacancies, and labor market tightness. The volatiliy of vacancies and labor market tightness is even lower than in Shimer (2005), as the presence of capital and bonds in the benchmark model gives firms more possibilities to adjust to technology shocks.

A.7 Private Efficiency of Wages

Proof. The proof draws from the proof of private efficiency in the online appendix to Michaillat (2012). Note that wages are privately efficient if neither firms nor workers have any incentive to separate as long as there are positive bilateral gains from the match.

The first part of the proof is relatively simple. For the household side private efficiency implies that the marginal value of an additional matched worker has to be positive:

$$
\mathbb{H}_{m,t} = -\frac{\varphi}{\lambda_t} + w_t - s + (1-x)\mathbb{E}_t \Lambda^h_{t+1|t} [1 - f(\theta_{t+1})] \mathbb{H}_{m,t+1} \geq 0.
$$

This equation can be rearranged to give

Table A.1. Simulated Moments – Flexible Wages and Perfect Credit Markets

	u	v	θ	w	y	z
Standard deviation	0.009	0.014	0.023	0.014	0.016	0.015
	(0.002)	(0.003)	(0.005)	(0.003)	(0.004)	(0.003)
Autocorrelation	0.940	0.792	0.865	0.839	0.845	0.830
	(0.028)	(0.078)	(0.058)	(0.070)	(0.067)	(0.069)
Correlation	1	-0.891	-0.958	-0.941	-0.947	-0.933
		(0.035)	(0.016)	(0.024)	(0.022)	(0.026)
	—	1	0.983	0.970	0.974	0.994
			(0.004)	(0.009)	(0.006)	(0.001)
	—	—	1	0.985	0.991	0.996
				(0.007)	(0.004)	(0.001)
	—	—	—	1	0.999	0.990
					(0.000)	(0.005)
	—	—	—	—	1	0.992
						(0.003)
	—	—	—	—	—	1

Note: Results from simulating the model with stochastic technology with a second-order perturbation method. All variables are log deviations from an HP trend with smoothing parameter 10^5. Simulated standard errors (standard deviations across 500 simulations) are reported in parentheses.

$$w_t + (1-s)\mathbb{E}_t\Lambda^h_{t+1|t}[1 - f(\theta_{t+1})]\mathbb{H}_{m,t+1} \geq s + \frac{\varphi}{\lambda_t}.$$

The household has no incentive to have the last worker separated from the match if the wage plus the continuation value of the match is larger than unemployment benefits plus the utility value of leisure. Since I focus only on symmetric equilibria, all firms pay equal wages and no worker has an incentive to switch firms.

For the second part of the proof, let the marginal revenue of an additional worker be defined by

$$\hat{v}_t \equiv \alpha z_t[(1-x)n_{t-1}]^{\alpha-1}k_t^{1-\alpha}(1-\mu_{b,t}),$$

which is the highest marginal product the firm can receive in a given period without laying off workers. Assume that there exist marginal costs $\hat{v}_t^H > \hat{v}_t^L$ such that

(i) if $\hat{v}_t < \hat{v}_t^L$, the firm lays off workers[1];

(ii) if $\hat{v}_t \in [\hat{v}_t^L, \hat{v}_t^H]$, the firm freezes hiring;

(iii) if $\hat{v}_t > \hat{v}_t^H$, the firm hires workers.

Now define as L_t the value function of the firm accounting for the possibility of layoffs. This function is given by

$$L_t = \max_{\{d_t, h_t, i_t, k_{t+1}, b_{t+1}\}} d_t + \mathbb{E}_t \Lambda^c_{t|t+j} \mathbb{J}_{t+1}(\omega^e_{t+1}; \Omega_{t+1})$$

subject to

$$\begin{aligned} z_t n_t^\alpha k_t^{1-\alpha} + \frac{b_{t+1}}{R_t} &= d_t + b_t + w_t n_t + i_t \\ &+ \mathbb{1}\{n_t > (1-x)n_{t-1}\} \frac{\kappa}{2} \left(\frac{m_t}{q(\theta_t)} \right)^2 [n_t - (1-x)n_{t-1}], \end{aligned}$$

$$k_{t+1} = (1-\delta)k_t + \left[\frac{\delta^\xi}{1-\xi} \left(\frac{i_t}{k_t} \right)^{1-\xi} - \frac{\xi\delta}{(1-\xi)} \right] k_t,$$

$$n_t = (1-x)n_{t-1} + m_t,$$

and

$$d_t + w_t n_t + \mathbb{1}\{n_t > (1-x)n_{t-1}\} \frac{\kappa}{2} \left(\frac{m_t}{q(\theta_t)} \right)^2 [n_t - (1-x)n_{t-1}] + i_t + b_t \leq \eta_t q_{k,t} k_t,$$

where $\mathbb{1}\{n_t > (1-x)n_{t-1}\}$ is the indicator function that is equal to one if and only if the firm hires workers and equal to zero otherwise. The marginal costs $\hat{v}_t^H$ and $\hat{v}_t^L$ are defined as follows:

$$\hat{v}_t^L = w_t - \Lambda^c_{t|t+j} \mathbb{E}_t \left[\frac{\partial L_{t+1}}{\partial n_t} \right]$$

and

$$\hat{v}_t^H = w_t + \frac{\kappa m_t}{q(\theta_t)^2} - \Lambda^c_{t|t+j} \mathbb{E}_t \left[\frac{\partial L_{t+1}}{\partial n_t} \right],$$

where L_{t+1} is the value function of the firm as seen from period $t+1$. $\hat{v}_t^L$ are the lowest marginal costs a firm can achieve by keeping its workforce, while $\hat{v}_t^H \geq \hat{v}_t^L$ are the lowest marginal costs a firm can achieve by hiring an infinitesimal amount of workers. Now let $\mathcal{F}$ be the σ-algebra generated by future realizations of the stochastic processes z and η, taking as given the information set at time t. $\mathcal{F}$ can be partitioned in

[1] Here I allow for $m_t < 0$.

$$\mathcal{F} = \mathcal{F}^+ \cup \mathcal{F}^- \cup_{h=1}^{+\infty} \mathcal{F}^h,$$

where $\mathcal{F}^+$ is the subset of all future realizations of z and η such that the firm is hiring next period, $\mathcal{F}^-$ is the subset such that there are layoffs and $\mathcal{F}^h$ is the subset such that there is a hiring freeze for the next h periods. Let $p^+ = \mathbb{P}(\mathcal{F}^+)$, $p^- = \mathbb{P}(\mathcal{F}^-)$, and $p^h = \mathbb{P}(\mathcal{F}^h)$ be the measures of these subsets, then it holds that

$$\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}\right] = p^+\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^+\right] + p^-\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^-\right] + \sum_{h=1}^{+\infty} p^h\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^h\right].$$

Using the value function, it follows that

$$\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^+\right] = (1-x)\mathbb{E}_t\left[\frac{\kappa m_t}{q(\theta_t)}|\mathcal{F}^+\right],$$

$$\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^-\right] = 0,$$

and

$$\begin{aligned}\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^h\right] = \mathbb{E}_t\Bigg[&\sum_{j=t+1}^{t+h} \Lambda^c_{0|j-(t+1)}(1-s)^{j-t} \\ &\left\{\alpha z_j[(1-x)^{j-t}n_t]^{\alpha-1}k_j^{1-\alpha}(1-\mu_{b,j}) - w_j\right\} \\ &+ \Lambda^c_{0|h}(1-x)^{h+1}\frac{\kappa m_{t+h+1}}{q(\theta_{t+h+1})^2}|\mathcal{F}^+\Bigg].\end{aligned}$$

Next, note that in a symmetric environment hiring freezes occur with a probability of zero. As the environment is symmetric, if one firm decides to freeze hiring all firms will do so. However, when all firms freeze hiring, θ is equal to zero, as there are no vacancies. This implies $\frac{\kappa m_t}{q(\theta_t)^2} = 0$ and thus $v_t^L = v_t^H$. I have already shown that a necessary and sufficient condition to avoid layoffs is $\hat{v}_t \geq v_t^L$. Now since $p^h = 0$ and $\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^-\right] = 0$ it holds that

$$\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}\right] = p^+\mathbb{E}_t\left[\frac{\partial L_{t+1}}{\partial n_t}|\mathcal{F}^+\right] = (1-s)\mathbb{E}_t\left[\frac{\kappa m_{t+1}}{q(\theta_{t+1})^2}\right].$$

Using this equation, a necessary and sufficient condition to avoid layoffs is

$$\begin{aligned}&s + \varphi c_t - (1-s)\mathbb{E}_t\Lambda^h_{t+1|t}[1-f(\theta_{t+1})]\mathbb{H}_{m,t+1} \\ &\leq w_t \leq \alpha z_t[(1-x)n_{t-1}]^{\alpha-1}k_t^{1-\alpha}(1-\mu_{b,t}) \\ &+ (1-x)\mathbb{E}_t\Lambda^h_{t+1|t}\frac{\kappa m_{t+1}}{q(\theta_{t+1})^2},\end{aligned}$$

which is equal to the equation in Proposition 2. □

A.8 Staggered Nash Bargaining

In this section I derive the wage schedule arising from a staggered Nash bargaining setup as in Gertler and Trigari (2009) in a model with financial frictions.[2]

Going back to the idea of Calvo (1983), only a fraction τ of all firms is able to renegotiate wages with its workforce in each period. Workers hired in between wage renegotiations receive the current wage. In order to ensure that a determinate equilibrium exists, the model presented in Section 2.3 needs to be adjusted. First, I assume quadratic costs of adjusting employment. Second, workers hired in this period do not become productive immediately but only in the next period. Under these assumptions the job creation condition for firm i is given by

$$\begin{aligned}\frac{\kappa}{2}x_t(i) = \Lambda^c_{t+1|t}[&\alpha z_{t+1} n^{\alpha-1}_{t+1} k^{1-\alpha}_{t+1}(1-\mu_{b,t+1}) - w_{t+1}(i) \\ &+ \frac{\kappa}{4}x^2_{t+1}(i) + (1-x)\frac{\kappa}{2}x_{t+1}(i)],\end{aligned},$$

where $x_t(i)$, the hiring rate in firm i, is defined as $x_t(i) = \frac{q_t v_t(i)}{n_t(i)}$.

Next, I define the value functions. The worker surplus at firm i is $H_t(i) = V_t(i) - U_t(i)$, where $V_t(i)$ is the value of a worker being employed at firm i. U_t is the value of being unemployed. This value can be calculated as:

$$H_t(i) = -\frac{\varphi}{\lambda_t} + w_t(i) - s + \mathbb{E}_t \Lambda^h_{t+1|t}[(1 - x + x f(\theta_{t+1})] H_{t+1}(i) - f(\theta_{t+1}) H_{x,t+1},$$

where $H_{x,t+1}$ is the average worker surplus conditional on being a new hire.

Under the assumption of Nash bargaining, the wage schedule solves

$$\max_{w_t} H^{\phi}_t(r) J^{1-\phi}_t(r),$$

where $H_t(r)$ and $J_t(r)$ are the value functions of renegotiating workers and firms, respectively. Due to the multi-period setup of the wage schedule, firms need to take into account the discounted expected sum of future wage payments and workers the discounted future sum of expected wages. The corresponding equations are given by

$$W^f_t(r) = \Sigma_t(r) w^*_t + \tau \mathbb{E}_t \sum_{j=1}^{\infty} \frac{n_{t+j}}{n_t}(r) \Lambda^c_{t+j|t} \Sigma_{t+j}(r) w^*_{t+j}$$

and

$$W^w_t(r) = \Delta w^*_t + \tau \mathbb{E}_t \sum_{j=1}^{\infty} (1-x)^j \Lambda^h_{t+j|t} \Delta_{t+j} w^*_{t+j},$$

[2]For a more detailed derivation of the wage schedule in a model without financial frictions see Gertler and Trigari (2009).

where $\Sigma_t(r) = \mathbb{E}_t \sum_{j=1}^{\infty} \frac{n_{t+j}}{n_t}(r)(1-\tau)^j \Lambda^c_{t+j|t}$ and $\Delta_t = \mathbb{E}_t \sum_{j=1}^{\infty}(1-x)^j(1-\tau)^j \Lambda^h_{t+j|t}$ are the cumulative discount factors of the firm and the worker, respectively.

The solution to the Nash bargaining problem is

$$\chi_t(r) J_t(r) = (1-\chi_t(r)) H_t(r)$$

with $\chi_t(r) = \frac{\phi}{\phi + (1-\phi)\frac{\Sigma_t(r)}{\Delta_t}}$.

Combining equations yields the difference equation

$$\begin{aligned}\Delta_t w_t^* = {} & w_t^0(r) + \chi_t(r)(1-\gamma)(1-x)\mathbb{E}_t \Lambda^c_{t+1|t}\Delta_{t+1} w^*_{t+1} \\ & + (1-\chi_t(r))(1-\gamma)(1-x)\mathbb{E}_t \Lambda^h_{t+1|t}\Delta_{t+1} w^*_{t+1}\end{aligned}$$

for the wage, where the the target wage is given by

$$\begin{aligned}w_t^0(r) = {} & \chi_t(r)\left[\alpha z_t n_t^{\alpha-1} k_t^{1-\alpha}(1-\mu_{b,t}) + \frac{\kappa}{4} x_t^2(i)\right] \\ & + (1-\chi_t(r)\left[-\frac{\varphi}{\lambda_t} + s + f(\theta_t)\mathbb{E}_t \Lambda^h_{t+1|t} H_{x,t}\right].\end{aligned}$$

This wage schedule collapses to the one derived in Gertler and Trigari (2009) in the absence of financial frictions (i.e. setting $\mu_{b,t} = 0$), when setting $\beta_h = \beta_c$ and $\frac{\kappa}{2} = \kappa$.

In the remainder of this section I discuss the results obtained by simulating the model with financial frictions using the wage schedule derived above. The model is calibrated to match the same steady state values as the model in the main text. The simulated moments are given in Table A.2.

The second moments of key labor market variables are close to the results in Section 2.4. The amplification of shocks is again close to the data. In the model, a 1% decrease in productivity increases unemployment by 4.9% (4.2% in the data), decreases vacancies by 6.7% (4.5% in the data), and decreases labor market tightness by 12.2% (8.6% in the data).

The findings concerning the skewness of the unemployment rate in levels and in changes also carry over to this version of the model. However, the wage rate arising from this staggered bargaining setup generates skewness at a higher frequency than the ad-hoc wage schedule. Before filtering the data, the skewness results for the two models are qualitatively the same and quantitatively very close. After filtering the skewness is smaller, but the qualitative results still hold.

A.9 Further Business Cycle Statistics

In this section I evaluate the business cycle statistics for the benchmark model with wage rigidity and for the model with financial frictions but without wage rigidity. This robustness exercise emphasizes the importance of the interaction between wage rigidity and financial frictions, not only for explaining unemployment dynamics, but also for explaining business

Table A.2. Simulated Moments – Alternative Wage Schedule and $\tau = 0.25$

	u	v	θ	w	y	z
Standard deviation	0.096	0.115	0.204	0.012	0.022	0.015
	(0.024)	(0.020)	(0.052)	(0.002)	(0.005)	(0.003)
Autocorrelation	0.920	0.813	0.912	0.717	0.910	0.829
	(0.030)	(0.065)	(0.031)	(0.115)	(0.039)	(0.070)
Correlation	1	-0.644	-0.863	-0.579	-0.882	-0.772
		(0.228)	(0.085)	(0.105)	(0.044)	(0.082)
	—	1	0.921	0.402	0.810	0.887
			(0.066)	(0.150)	(0.209)	(0.189)
	—	—	1	0.523	0.913	0.907
				(0.092)	(0.066)	(0.054)
	—	—	—	1	0.664	0.559
					(0.130)	(0.135)
	—	—	—	—	1	0.963
						(0.010)
	—	—	—	—	—	1

Note: Results from simulating the model with stochastic technology with a second-order perturbation method. All variables are log deviations from an HP trend with smoothing parameter 10^5. Simulated standard errors (standard deviations across 500 simulations) are reported in parentheses.

cycle statistics. I strengthen this point by showing that neither the benchmark model with wage rigidity nor the model with financial frictions and without wage rigidity is able to generate volatility in labor market variables comparable to the data.

Without wage rigidity, the volatility of the key labor market variables drops sharply. The volatility of unemployment decreases by 39%, the volatility of vacancies by 30%, and the volatility of labor market tightness by 32%. The response of unemployment, vacancies and labor market tightness to a 1% percent decrease in productivity is lower: unemployment increases by 1.5%, vacancies decrease by 2.5%, and labor market tightness decreases by 3.9%.

For the benchmark model, the introduction of wage rigidity increases the volatility of unemployment by 45%, the volatility of vacancies by 71%, and the volatility of labor market tightness by 61%. Despite the large relative increases, the absolute values remain small. Even with larger wage rigidity than in the model with financial frictions, the benchmark model does not generate enough amplification to match the empirical volatility of the key labor market variables or the amplification of shocks in the data.

Table A.3. Simulated Moments – Financial Frictions and $\tau = 0$

	u	v	θ	w	y	z
Standard deviation	0.049	0.114	0.154	0.013	0.018	0.015
	(0.009)	(0.011)	(0.015)	(0.002)	(0.004)	(0.003)
Autocorrelation	0.758	0.257	0.408	0.841	0.855	0.829
	(0.055)	(0.077)	(0.083)	(0.066)	(0.065)	(0.070)
Correlation	1	-0.772	-0.876	-0.482	-0.631	-0.443
		(0.018)	(0.015)	(0.192)	(0.142)	(0.202)
	—	1	0.979	0.357	0.454	0.336
			(0.003)	(0.116)	(0.093)	(0.124)
	—	—	1	0.418	0.534	0.388
				(0.142)	(0.109)	(0.151)
	—	—	—	1	0.980	0.973
					(0.016)	(0.012)
	—	—	—	—	1	0.958
						(0.020)
	—	—	—	—	—	1

Note: Results from simulating the model with stochastic technology with a second-order perturbation method. All variables are log deviations from an HP trend with smoothing parameter 10^5. Simulated standard errors (standard deviations across 500 simulations) are reported in parentheses.

Table A.4. Simulated Moments – Benchmark Model and $\tau = 0.25$

	u	v	θ	w	y	z
Standard deviation	0.012	0.024	0.037	0.012	0.017	0.015
	(0.003)	(0.003)	(0.005)	(0.003)	(0.004)	(0.003)
Autocorrelation	0.888	0.610	0.726	0.949	0.849	0.830
	(0.036)	(0.091)	(0.077)	(0.028)	(0.065)	(0.069)
Correlation	1	-0.837	-0.929	-0.798	-0.947	-0.967
		(0.026)	(0.016)	(0.034)	(0.013)	(0.008)
	—	1	0.979	0.499	0.817	0.865
			(0.001)	(0.065)	(0.016)	(0.009)
	—	—	1	0.630	0.899	0.938
				(0.056)	(0.014)	(0.007)
	—	—	—	1	0.904	0.857
					(0.035)	(0.038)
	—	—	—	—	1	0.993
						(0.003)
	—	—	—	—	—	1

Note: Results from simulating the model with stochastic technology with a second-order perturbation method. All variables are log deviations from an HP trend with smoothing parameter 10^5. Simulated standard errors (standard deviations across 500 simulations) are reported in parentheses.

B Appendix to Chapter 3

B.1 First Order Conditions of Firms

Defining the value of a marginal worker in an abstract non-routine cognitive occupations for a firm as J_a, the first order conditions for hiring and for vacancy posting are given by

$$c_a = \mu_a q_a$$
$$\mu_a = \beta J_{a,+1},$$

where μ_a is the Lagrange-multiplier on the employment constraint for workers in abstract occupations. The corresponding value of a marginal worker in abstract non-routine cognitive occupations is

$$J_a = p_{Z_a} - \mathbb{1}_u w_a^u - (1 - \mathbb{1}_u) w_a^n + (1 - s_a)\beta J_{a,+1}.$$

Defining the value of a marginal worker with ability η in a routine occupation for a firm as $J_r(\eta)$, the first order conditions for hiring workers in routine tasks and for vacancy posting are given by

$$c_r = \mu_r q_r$$
$$\mu_r = \beta J_{r,+1},$$

where μ_r is the Lagrange-multiplier on the employment constraint for a worker in routine occupations. The corresponding value of a marginal worker with ability η in routine occupations is

$$J_r = p_{Z_r}\overline{y_r} - \mathbb{1}_u \overline{w_r^u} - (1 - \mathbb{1}_u)\overline{w_r^n} + (1 - s_r)\beta J_{r,+1},$$
$$\text{with } y_r(\eta) = \frac{\partial Z_r}{\partial L_r(\eta)} = \eta(1-\mu)^\sigma \left[(1-\mu)^\sigma + (\mu k)^\sigma\right]^{\frac{1}{\sigma}-1} \text{ and } k \equiv \frac{K}{\int_{\eta_m}^{\bar{\eta}} \eta L_r(\eta)},$$

where $\overline{y_r}$ is the expected marginal product of a routine worker, $\overline{w_r^u}$ is the expected union wage, and $\overline{w_r^n}$ the expected non-union wage. The average marginal product and the average wages are used here, as firms are unable to condition their job search on the ability level η.

Defining the value of a marginal worker with ability η in a non-routine manual occupation for a firm as J_m, the first order conditions for hiring workers in manual tasks and for vacancy posting are given by

$$\begin{aligned} c_m &= \mu_m q_m \\ \mu_m &= \beta J_{m,+1}, \end{aligned}$$

where μ_m is the Lagrange-multiplier on the employment constraint for worker in manual occupations. The corresponding value of a marginal worker with ability η in manual occupations is

$$J_m = p_{Z_m} - \mathbb{1}_u w_m^u - (1 - \mathbb{1}_u) w_m^n + (1 - s_m)\beta J_{m,+1}.$$

B.2 Job Creation Conditions

The job creation conditions are given by

$$\frac{c_i}{q_i} = \beta J_{i,+1}$$
$$\text{with } i = a, r, m.$$

Together with the values of marginal workers for firms, it follows that

$$\begin{aligned} \frac{c_a}{q_a} &= \beta \left[p_{Z_a} - \mathbb{1}_{u,+1} w_a^u - (1 - \mathbb{1}_{u,+1})\, w_a^n + (1 - s_a)\frac{c_a}{q_{a,+1}} \right] \\ \frac{c_r}{q_r} &= \beta \left[p_{Z_r}\overline{y_r} - \mathbb{1}_{u,+1}\overline{w_r^u} - (1 - \mathbb{1}_{u,+1})\, \overline{w_r^n} + (1 - s_r)\frac{c_r}{q_{r,+1}} \right] \\ \frac{c_m}{q_m} &= \beta \left[p_{Z_m} - \mathbb{1}_{u,+1} w_m^u - (1 - \mathbb{1}_{u,+1})\, w_m^n + (1 - s_m)\frac{c_m}{q_{m,+1}} \right]. \end{aligned}$$

As we are mainly interested in the long-run effect of routine-biased technical change on the economy and on the wage bargaining regimes, we focus on the steady state of the economy. The steady state job creation conditions are given by

$$\frac{c_a}{q_a} = \beta\left[p_{Z_a} - \mathbb{1}_u w_a^u - (1-\mathbb{1}_u)\, w_a^n + (1-s_a)\frac{c_a}{q_a}\right] \tag{B.1}$$

$$\frac{c_r}{q_r} = \beta\left[p_{Z_r}\overline{y_r} - \mathbb{1}_u \overline{w_r^u} - (1-\mathbb{1}_u)\, \overline{w_r^n} + (1-s_r)\frac{c_r}{q_r}\right] \tag{B.2}$$

$$\frac{c_m}{q_m} = \beta\left[p_{Z_m} - \mathbb{1}_u w_m^u - (1-\mathbb{1}_n)\, w_m^n + (1-s_m)\frac{c_m}{q_m}\right]. \tag{B.3}$$

A firm hires workers of each type and each ability level η until the costs of labor are equal to the discounted expected marginal product. Here the costs consist of the vacancy posting costs and the discounted expected wage minus the discounted cost of hiring next period.

B.3 Derivation of Wages

In this section we derive the non-union wages in the model. The first order conditions are given by

$$W_i^n(\eta) - U_i(\eta) = \frac{\gamma^i}{1-\gamma^i} J_i^n(\eta),$$
$$\text{with } i = a, r, m.$$

Abstract Workers

After replacing the value function, the Nash sharing rule for abstract workers is

$$\begin{aligned} &w_a^n + \beta\left[(1-s_a)W_a^n + s_a U_a\right] - z_a - \beta[(1-f_a)U_a^n + f_a W_a^n\} \\ &= \frac{\gamma^a}{1-\gamma^a}\left[p_{Z_a} - w_a^n + (1-s_a)\beta J_a^n\right]. \end{aligned}$$

After some rearrangement, we get

$$\begin{aligned} w_a^n &= \gamma^a p_{Z_a} + (1-\gamma^n) z_a + \gamma^a (1-s_a)\beta J_a^n \\ &\quad + (1-\gamma^a)\beta\left[f_a\left(W_a^n - U_a^n\right) - (1-s_a)\left(W_a^n - U_a^n\right)\right]. \end{aligned}$$

By using the job creation condition (B.1), $\frac{c_a}{q_a} = \beta J_{a,+1}^n$, and the first order condition resulting from the Nash sharing rule

$$(1-\gamma^a)\left(W_a^n - U_a^n\right) = \gamma^a J_a^n = \gamma^a \frac{c_a}{\beta q_a},$$

we obtain the wage equation

$$w_a^n = \gamma^a p_{Z_a} + \gamma^a c_a \theta_a + (1-\gamma^a) z_a.$$

Routine Workers

After replacing the value function, the Nash sharing rule for routine workers of ability level η is

$$\begin{aligned} & w_r^n(\eta) + \beta\left[(1-s_r)W_r^n(\eta) + s_r U_r(\eta)\right] - z_r(\eta) - \beta[(1-f_r)U_r^n(\eta) + f_r W_r^n(\eta)\} \\ & = \frac{\gamma^r}{1-\gamma^r}\left[p_{Z_r} y_r(\eta) - w_r^n(\eta) + (1-s_r)\beta J_r^n\right]. \end{aligned}$$

After some rearrangement, we arrive at

$$\begin{aligned} w_r^n(\eta) = {} & \gamma^r p_{Z_r} y_r(\eta) + (1-\gamma^r) z_r(\eta) + \gamma^r (1-s_r)\beta J_r^n \\ & + (1-\gamma^r)\beta\left[f_r\left(W_r^n(\eta) - U_r^n(\eta)\right) - (1-s_r)\left(W_r^n(\eta) - U_r^n(\eta)\right)\right]. \end{aligned}$$

By using the job creation condition (B.2), $\frac{c_r}{q_r(\eta)} = \beta J_r^n(\eta)$, and the first order condition resulting from the Nash sharing rule

$$(1-\gamma^r)\left(W_r^n(\eta) - U_r^n(\eta)\right) = \gamma^r J_r^n(\eta) = \gamma^r \frac{c_r}{\beta q_r},$$

we obtain the wage equation

$$w_r^n(\eta) = \gamma^r p_{Z_r} y_r(\eta) + \gamma^r c_r \theta_r + (1-\gamma^r) z_r(\eta).$$

Manual Workers

After replacing the value function, the Nash sharing rule for manual workers is

$$\begin{aligned} & w_m^n + \beta\left[(1-s_m)W_m^n + s_m U_m\right] - z_m(\eta) - \beta[(1-f_m)U_m^n + f_m W_m^n\} \\ & = \frac{\gamma^m}{1-\gamma^m}\left[p_{Z_m} - w_m^n + (1-s_m)\beta J_m^n\right]. \end{aligned}$$

After some rearrangement, we get

$$\begin{aligned} w_m^n = {} & \gamma^m p_{Z_m} + (1-\gamma^m) z_m(\eta) + \gamma^m (1-s_m)\beta J_m^n \\ & + (1-\gamma^m)\beta\left[f_m\left(W_m^n - U_m^n\right) - (1-s_m)\left(W_m^n - U_{m,+1}^n\right)\right]. \end{aligned}$$

By using the job creation condition (B.3), $\frac{c_m}{q_m)} = \beta J_m^m$, and the first order condition resulting from the Nash sharing rule

$$(1-\gamma^m)\left(W_m^n - U_m^n\right) = \gamma^m J_m^n = \gamma^m \frac{c_m}{\beta q_m},$$

we obtain the wage equation

$$w_m^n = \gamma^m p_{Z_m} + \gamma^m c_m \theta_m + (1-\gamma^m) z_m(\eta).$$

B.4 Union Surplus

In this section we derive the industrial union surplus. The derivation of the craft union surplus proceeds analogously. The first order condition in the collective bargaining problem is given by

$$\begin{aligned}
&\sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta)\left[W_i^u(\eta) - W_i^{u,s}(\eta)\right] \mathrm{d}\,\eta \\
&= \frac{\gamma^u}{1-\gamma^u} \sum_i \Big\{ p_{Z_i} Z_i - p'_{Z_i} Z'_i - \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^u(\eta) \,\mathrm{d}\,\eta \Big\}, \\
&\text{with } i = r, m.
\end{aligned}$$

After replacing the value function and using the job creation conditions (B.2) and (B.3), the Nash sharing rule is

$$\begin{aligned}
&\sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta)\left[w_i^u(\eta) - w_i^{u,s}(\eta)\right] \mathrm{d}\,\eta \\
&= \frac{\gamma^u}{1-\gamma^u} \sum_i \Big\{ p_{Z_i} Z_i - p'_{Z_i} Z'_i - \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^u(\eta) \,\mathrm{d}\,\eta \Big\}.
\end{aligned}$$

After some rearrangement, we have

$$\begin{aligned}
&\gamma^u \sum_i \left(p_{Z_i} Z_i - p'_{Z_i} Z'_i\right) + (1-\gamma^u) \sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^{u,s}(\eta) \,\mathrm{d}\,\eta \\
&= \gamma^u \sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^u(\eta) \,\mathrm{d}\,\eta + (1-\gamma^u) \sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^u(\eta) \,\mathrm{d}\,\eta.
\end{aligned}$$

Thus, the total union surplus is given by

$$\begin{aligned}
S^u &= \sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^u(\eta) \,\mathrm{d}\,\eta \\
&= \gamma^u \sum_i \left(p_{Z_i} Z_i - p'_{Z_i} Z'_i\right) + (1-\gamma^u) \sum_i \int_{\underline{\eta}}^{\bar{\eta}} L_i(\eta) w_i^{u,s}(\eta) \,\mathrm{d}\,\eta \\
&\text{with } i = r, m.
\end{aligned}$$

B.5 Theoretical Evaluation of the Main Mechanisms

The arguments in this section proof Propositions 3 and 4, which state the main mechanisms in our paper.

B.5.1 Polarization

Routine-biased technical change is modeled as a drop in p_k, the relative price of computer capital. As we are concerned with the incentives of previous routine workers to switch to manual occupations, we consider the effects of a decrease in p_k before any employment adjustment occurs. Thus, L_a, L_r, and L_m are constant.

Note that the decrease in the relative price only affects the intermediate firm producing Z_r directly. From the first order condition with respect to computer capital

$$\frac{\partial Z_r}{\partial K} = \mu^\sigma \left[\left(\frac{1-\mu}{k} \right)^\sigma + \mu^\sigma \right]^{\frac{1}{\sigma}-1}$$

it follows that K increases if and only if computer capital and workers performing routine tasks are substitutes, i.e, if $\sigma > 0$.[1] The increasing computer capital stock increases the production of the intermediate good Z_r.

Keep in mind that a unemployed routine worker switches occupations if $U_m(\eta) > U_r(\eta)$. Thus, given that unemployment benefits and separation rates are not affected by the drop in capital prices, the two variables driving changes in the incentives are wages and job-finding rates. From the wage equations and job creation conditions for both types of occupations it immediately follows that both variables of interest are driven by changes in the marginal productivity of the respective workers.

As the relevant elasticities (the elasticity of the wage with respect to productivity and labor market tightness and the elasticity of the job-finding rate with respect to productivity and wages) are identical for both types of occupations, it remains to show that the marginal productivity of manual workers increases by more compared to the marginal productivity of routine workers due to routine-biased technical change.

The relative marginal productivity of routine workers compared to manual workers is given by

$$\frac{p_{Z_r} y_r(\eta)}{p_{Z_m}} = \eta(1-\alpha)(1-\mu)^\sigma \left(\frac{A^{1+\frac{1}{\rho}}}{A_m} \right)^\rho \left(\frac{Z_a^{\frac{\alpha\rho}{\rho-1}}}{Z_m} \right)^{\rho-1}$$
$$\left((1-\mu) \int_{\eta_m}^{\bar{\eta}} \eta L_r(\eta) \, \mathrm{d}\, \eta \right)^{\sigma-1} Z_r^{(1-\alpha)\rho-\sigma}.$$

[1] Since the computer capital stock can be adjusted instantaneously and without frictions, an increase in K before occupational switches occur is in line with the model setup.

Thus, the relative productivity of routine workers decreases in Z_r, if $\sigma > (1-\alpha)\rho$, which proofs Proposition 3. Intuitively, in order for routine-biased technical change to increase the incentives for occupational switches, capital and routine tasks need to be substitutes and they need to be better substitutes than routine and manual tasks in the production of the final good.

B.5.2 Voting Incentives

A manual worker inside a unionized firm votes in favor of collective bargaining coverage, if the value of being a manual worker in a unionized firm is larger than the value of being a worker in a non-unionized firm, i.e., if $W_m^u > W_m^n$. As in Appendix B.5.1, the relevant variables are again the wages and the job-finding rates. As the marginal productivity of a manual worker is independent of the union status of the firm, relative changes in the job-finding rates are entirely driven by relative wage changes. Thus, it suffices to show that the non-union wage rate for manual workers increases relative to the union wage rate.[2]

Using the equation for the union surplus (3.8), the union wage schedule (3.6), and the non-union wage for manual workers (3.5), the relative union wage for a manual worker is given by[3]

$$\frac{w_m^u}{w_m^n} = \frac{\left[\gamma^u(p_{Z_m}Z_m - p'_{Z_m}Z'_m) + \gamma^u(p_{Z_r}Z_r - p'_{Z_r}Z'_r)\right]/(L_m+L_r)}{\gamma^m p_{Z_m} + \gamma^m c_m \theta_m^n}.$$

Using the production functions, this expression can be rewritten as

$$\frac{w_m^u}{w_m^n} = \frac{\left[\gamma^u p_{Z_m}Z_m\right]/(L_m+L_r)}{\gamma_m p_{Z_m} + \gamma_m c_m \theta_m^n} + \frac{\left[\gamma^u(p_{Z_r}Z_r - p'_{Z_r}Z'_r)\right]/(L_m+L_r)}{\gamma_m p_{Z_m} + \gamma_m c_m \theta_m^n}. \tag{B.4}$$

First, following the arguments in Appendix B.5.1, routine-biased technical change implies an increase in Z_r and thus an increase in the marginal productivity of manual workers, p_{Z_m}. Second, note that the effect of routine-biased technical change on the first term only depends on the elasticity of this term with respect to p_{Z_m}. Combining the job creation condition (B.3) and the wage for manual workers (3.5) yields

$$((1/\beta) - 1 + s_m)c_m\Psi_m(\theta_m^n)^\eta + c_m\gamma^m\theta_m^n = (1-\gamma^m)p_{Z_m}.$$

From this expression it is easy to see that the elasticity of θ_m^n with respect to p_{Z_m} is larger than one. Next, we use that for two functions f and g the elasticity of $(g+f)$ is given by $\epsilon_{f+g} = \frac{f\epsilon_f + g\epsilon_g}{f+g}$ to establish that the elasticity of the non-union wage of manual workers is larger than one. This directly implies that the first term of equation (B.4) decreases in p_{Z_m}.

[2] Note that the positive effect of a wage increase on the value function is not offset by a decrease in the job-finding rate.

[3] Since w_i^u and $z_i(\eta)$ are both unaffected by routine-biased technical change and set to zero in the simulation, they are left out in order to facilitate representation.

Intuitively, routine-biased technical change increases the productivity of and therefore the demand for manual workers. The non-union wage for manual workers increases as both the productivity and the labor market tightness increase. The union wage for manual workers increases by less, as the different outside options in the two bargaining regimes imply that the greater labor market tightness does not affect the collective bargaining.

For the second term in equation (B.4) it holds that

$$\frac{Z_r}{Z_r'} = \left[1 + \left(\frac{(1-\mu)\int_{\eta_m}^{\bar{\eta}} \eta L_r(\eta)\,\mathrm{d}\,\eta}{\mu K}\right)^{\sigma}\right]^{\frac{1}{\sigma}}.$$

Thus, an increase in K due to routine-biased technical change reduces $\frac{Z_r}{Z_r'}$. After some rearrangement, $\frac{p_{Z_r} Z_r}{p'_{Z_r} Z' r}$ is given by

$$\frac{p_{Z_r} Z_r}{p'_{Z_r} Z_r'} = \frac{[(A Z_a^{\alpha} Z_r^{1-\alpha})^{\rho} + (A_m Z_m)^{\rho}]^{1/\rho} - 1}{[(A Z_a^{\alpha} (Z_r')^{1-\alpha})^{\rho} + (A_m Z_m)^{\rho}]^{1/\rho - 1}} \left(\frac{Z_r}{Z_r'}\right)^{(1-\alpha)\rho}.$$

Using that $\frac{Z_r}{Z_r'}$ decreases with K, it is straightforward to show that an increase in K reduces $\frac{p_{Z_r} Z_r}{p'_{Z_r} Z' r}$ if routine and manual tasks are substitutes, i.e, if $\rho > 0$.

Taken together, routine-biased technical change reduces the union wage of manual workers relative to the non-union wage of manual workers, if $\rho > 0$. This proofs Proposition 4. This result does not depend on our choice of the union wage schedule, as the proof also holds if we exchange the union wage of manual workers for the union surplus.

B.6 Robustness Checks

In this section we present several robustness checks: regressions using the average routine share instead of the initial routine share in our instrument, unweighted regressions, and regressions using union coverage as the dependent variable,. The results are summarized in Tables B.1–B.3. Our instrument remains highly statistically significant across all alternative specifications. As was to be expected, union coverage reacts less to falling relative prices for investment goods.

Table B.1. Regression Results for Unionization Rates – Average Routine Share

	(1)	(2)	(3)	(4)
Capital prices	0.4116***	0.6254***	0.4054***	0.6099***
× routine employment share	(0.0743)	(0.0724)	(0.0702)	(0.0682)
Observations	1116	1116	1116	1116
R^2	0.9870	0.9834	0.9863	0.9822
Industry and occupation controls	yes	yes	no	no
State policy controls	yes	no	yes	no
State legislation controls	yes	no	yes	no
State demographic controls	yes	no	yes	no
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: Observations are weighted by the average state population over our sample period. The standard errors are reported in parentheses. *** indicates significance at the 1 percent level, ** indicates significance at the 5 percent level, and * indicates significance at the 10 percent level.

Table B.2. Regression Results for Unionization Rates – Unweighted

	(1)	(2)	(3)	(4)
Capital prices	0.1961***	0.1982***	0.1508***	0.1522***
× routine employment share	(0.0422)	(0.0416)	(0.0379)	(0.0403)
Observations	1116	1116	1116	1116
R^2	0.9765	0.9727	0.9753	0.9721
Industry and occupation controls	yes	yes	no	no
State policy controls	yes	no	yes	no
State legislation controls	yes	no	yes	no
State demographic controls	yes	no	yes	no
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: Observations are weighted by the average state population over our sample period. The standard errors are reported in parentheses. *** indicates significance at the 1 percent level, ** indicates significance at the 5 percent level, and * indicates significance at the 10 percent level.

Table B.3. Regression Results for Union Coverage Rates

	(1)	(2)	(3)	(4)
Capital prices	0.2031***	0.3468***	0.1913***	0.2817***
× routine employment share	(0.0583)	(0.0578)	(0.0536)	(0.0514)
Observations	1116	1116	1116	1116
R^2	0.9839	0.9802	0.9828	0.9784
Industry and occupation controls	yes	yes	no	no
State policy controls	yes	no	yes	no
State legislation controls	yes	no	yes	no
State demographic controls	yes	no	yes	no
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: Observations are weighted by the average state population over our sample period. The standard errors are reported in parentheses. *** indicates significance at the 1 percent level, ** indicates significance at the 5 percent level, and * indicates significance at the 10 percent level.

B.7 Data Appendix

In Table B.4 we provide the complete list of control variables used in the regressions in Section 3.2.

Table B.4. List of Control variables

State demographics	
Share of population living in a central city (urban density)	CPS
Share of population living in a city (urban density)	CPS
Share of black population (ethnic composition)	CPS
Share of white population (ethnic composition)	CPS
Shares of population in age groups 16-24; 25-44; 45-54; >55	CPS
Share of workers with each educational level:	CPS
less than high school; high school; some college; college or more	
Share of male population	CPS
Industry-occupation controls	
Shares of workers employed in industry groups:	CPS
construction; manufacturing; transportation, communcations, and other public utilities;	
wholesale and retail trade; services; finance, insurance, and real estate	
Shares of workers employed in occupational groups:	CPS + AD
abstract; routine; manual	
State policy controls	
Minimum wage rate	FRED
Total federal intergovernmental revenue	SLGFD
Total tax burden	TF
State gov. and leg. controls	
State senate majority party	NCLS
(1: Rep., -1: Dem., 0: other/indep./no majority)	
State house majority party	NCLS
(1: Rep., -1: Dem., 0: other/indep./no majority)	
Political party of the governor	NCLS
(1: Rep., -1: Dem., 0: other/indep./no majority)	

Note: AD: Autor and Dorn (2013); CPS: Current Population Survey; FRED: Federal Reserve Economic Data; NCLS: National Conference of State Legislatures; SLGFD: State and Local Government Finance Dataset; TF: Tax Foundation.

C Appendix to Chapter 4

C.1 Ideological Leanings in the District Courts

In Figures C.1–C.6 we provide evidence on the evolution of the average ideology score in each of the 90 U.S. district courts that have been active over our entire sample period from 1978 to 2011. While the ideology score did not experience a strong conservative shift in most district courts, there is some evidence of liberal (conservative) district courts becoming more conservative (liberal) over time. These slight tendencies towards the middle motivate us to include the lagged dependent variable in our regressions for district court rulings to account for mean-reverting dynamics which may also be present in rulings.

Figure C.1. Average Ideology Score in the District Courts (1/6)

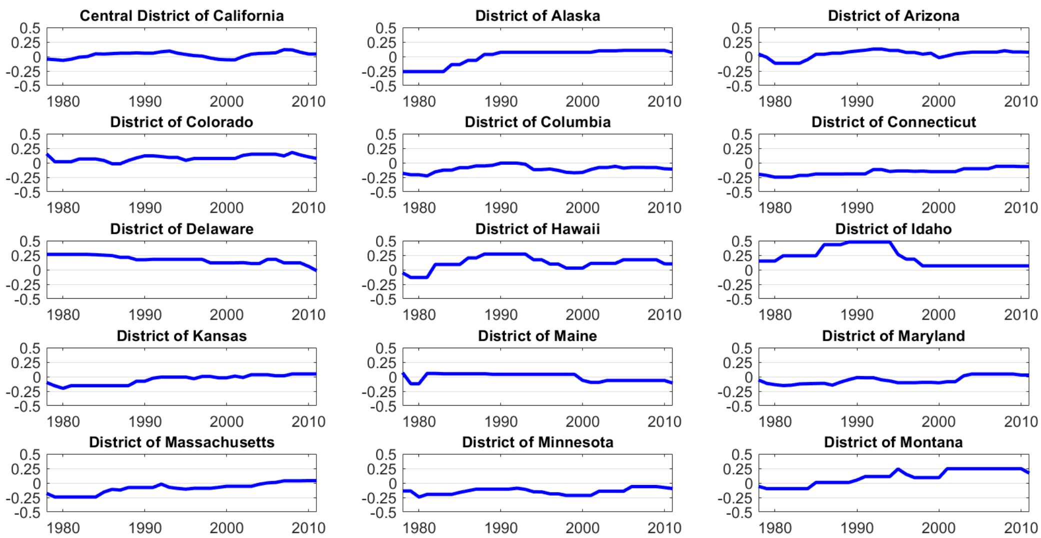


Figure C.2. Average Ideology Score in the District Courts (2/6)

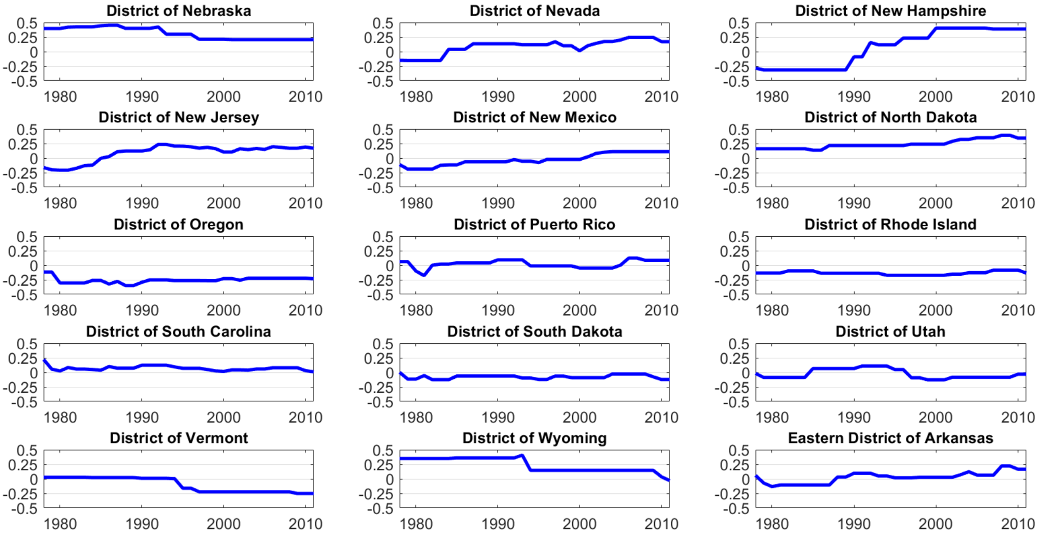


Figure C.3. Average Ideology Score in the District Courts (3/6)

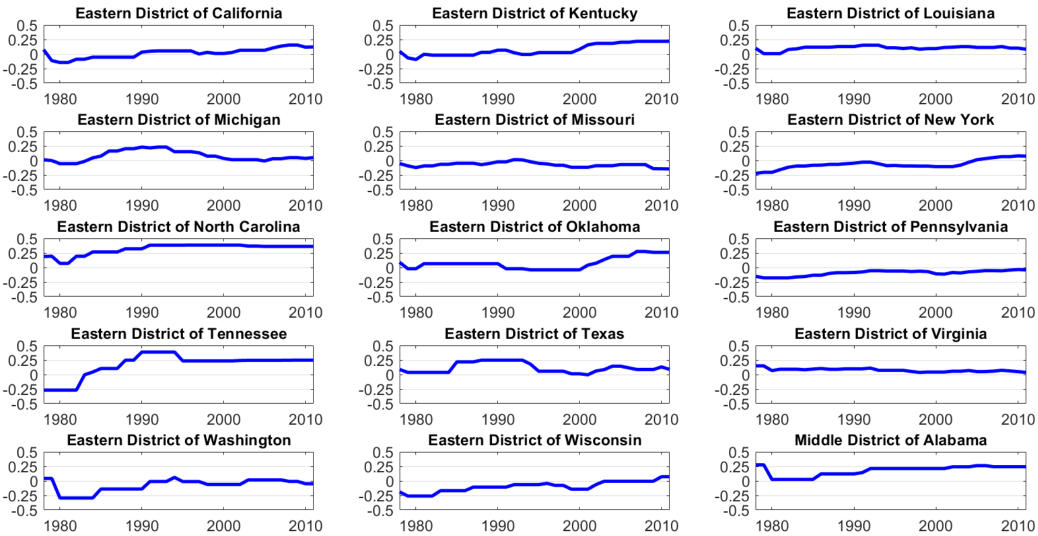


Figure C.4. Average Ideology Score in the District Courts (4/6)

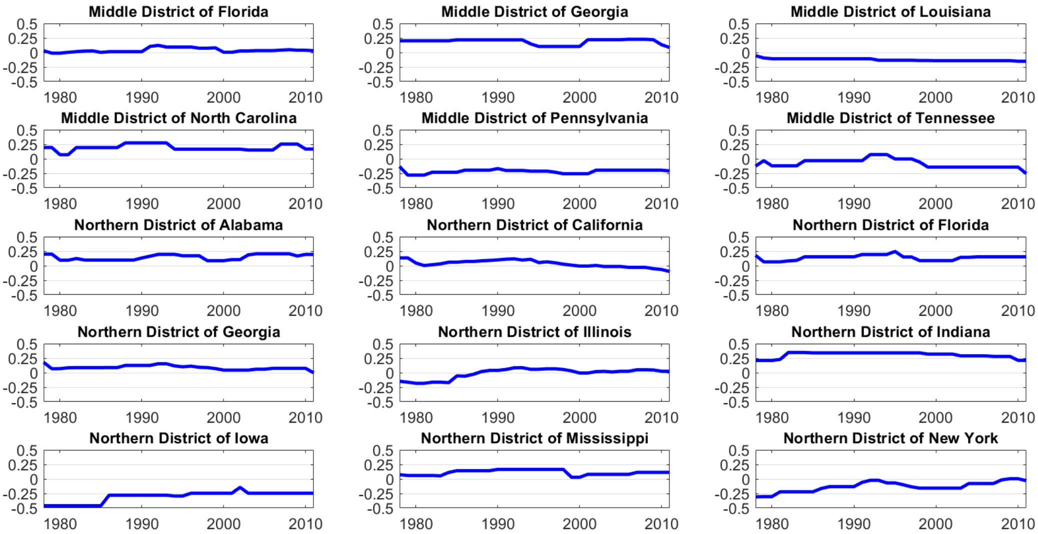


Figure C.5. Average Ideology Score in the District Courts (5/6)

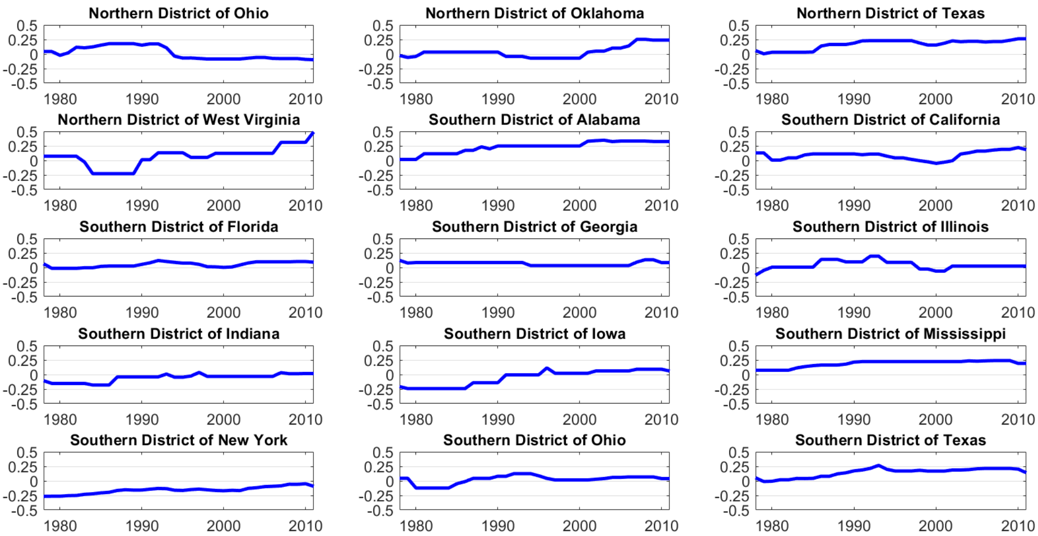
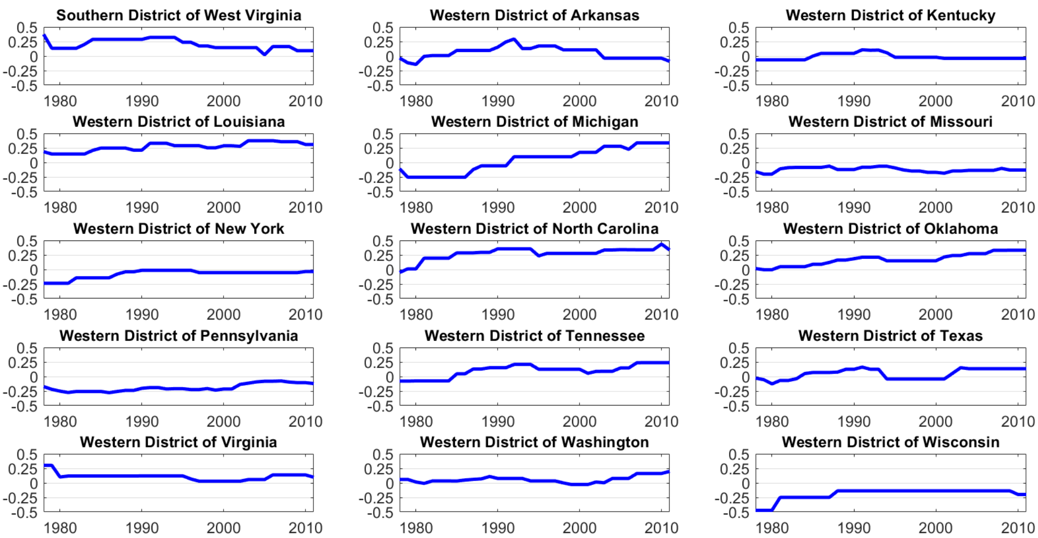


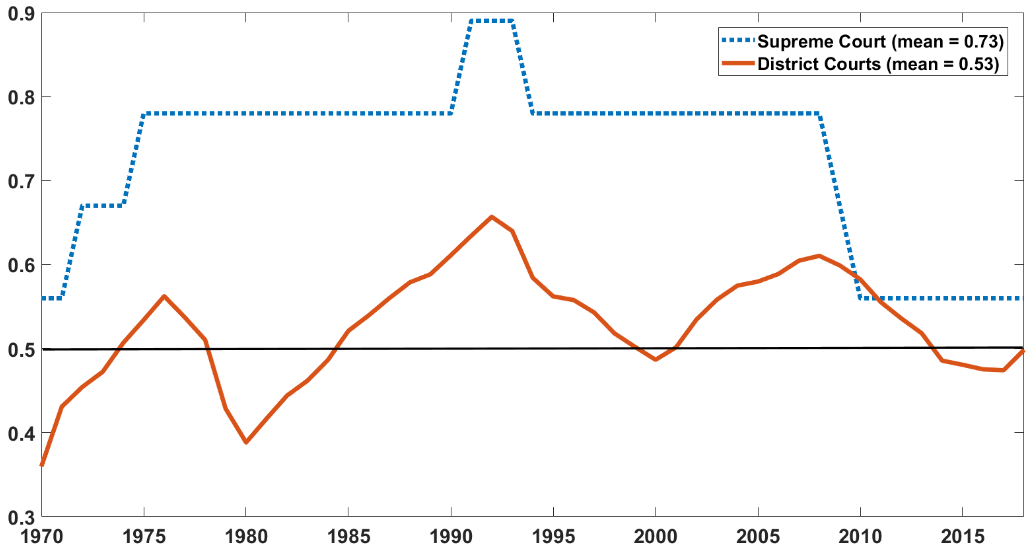
Figure C.6. Average Ideology Score in the District Courts (6/6)



C.2 Share of Judges Appointed by a Republican President

In this section we compare the share of district court judges appointed by Republican presidents to the share of Supreme Court justices appointed by Republican presidents. As depicted in Figure C.7, five of the nine Supreme Court Justices serving in 1970 have been appointed by a Republican president. This ratio increased to eight out of nine justicesin the early 1990s and only reverted back in 2010. In contrast, the share of district court judges appointed by a Republican president has remained close to 50% over the entire time period. Thus, it is unlikely that the relative increase in the share of conservative rulings is driven by increasingly conservative district court judges. Furthermore, effects of changes in the national composition of district court judges are absorbed in the time fixed effects in our regressions.

Figure C.7. Share of Justices and Judges Appointed by a Republican President



Note: This graph depicts the share of Supreme Court justices and the share of district court judges appointed by a Republican president. The black line indicates parity between the number of justices and judges appointed by a Republican president and the number of justices and judges appointed by a Democratic president.

We additionally provide evidence on the evolution of the share of district court judges that were appointed by a Republican president in each district court in Figures C.8–C.13. The majority of the district courts did not experience a strong conservative shift between 1978 and 2011. However, as there is some evidence that at district courts where many judges were appointed by a Republican (Democratic) president the share of Republican (Democratic) appointees declines over time, we include the share of district court judges appointed by a Republican president as a control variable in our regressions.

Figure C.8. Share of District Court Judges Appointed by a Republican President (1/6)

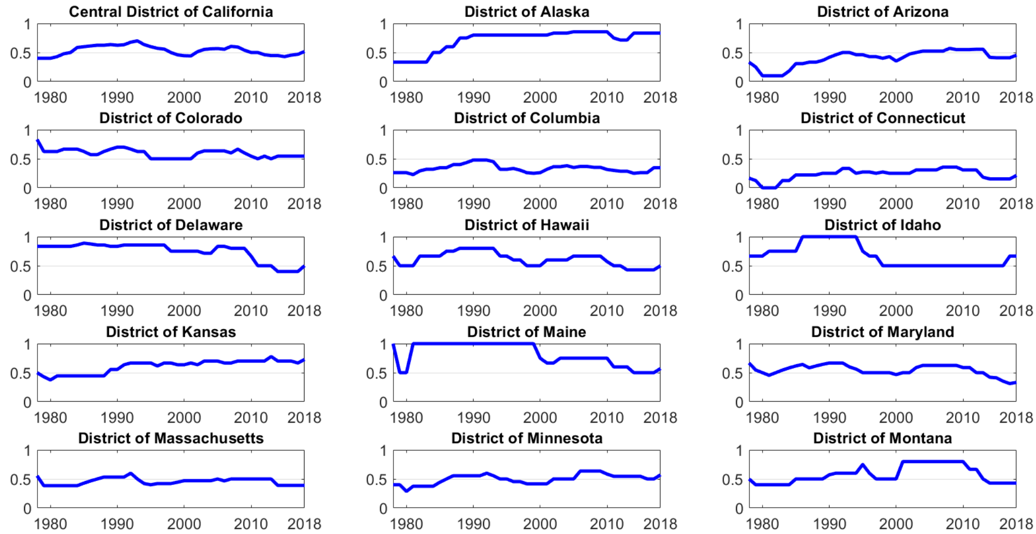


Figure C.9. Share of District Court Judges Appointed by a Republican President (2/6)

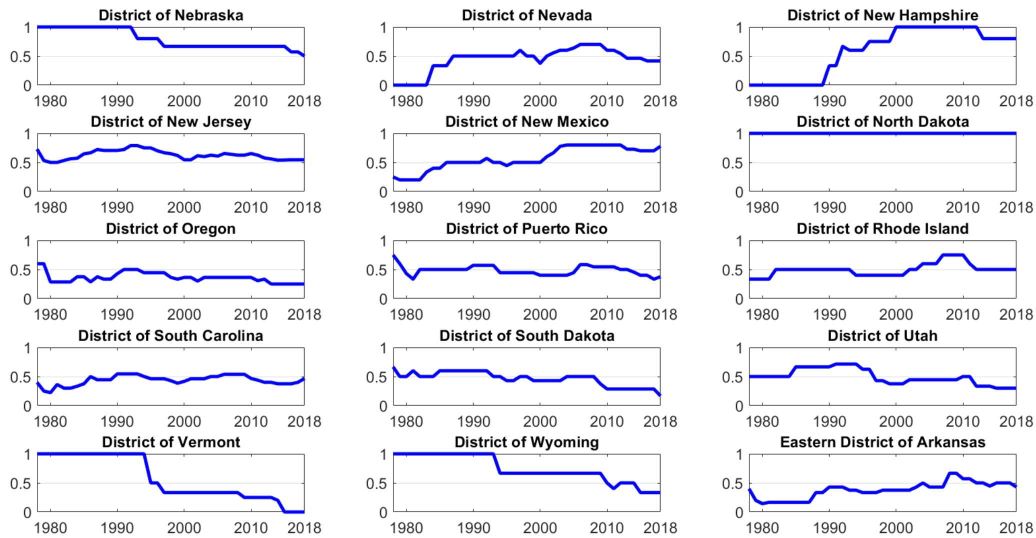


Figure C.10. Share of District Court Judges Appointed by a Republican President (3/6)

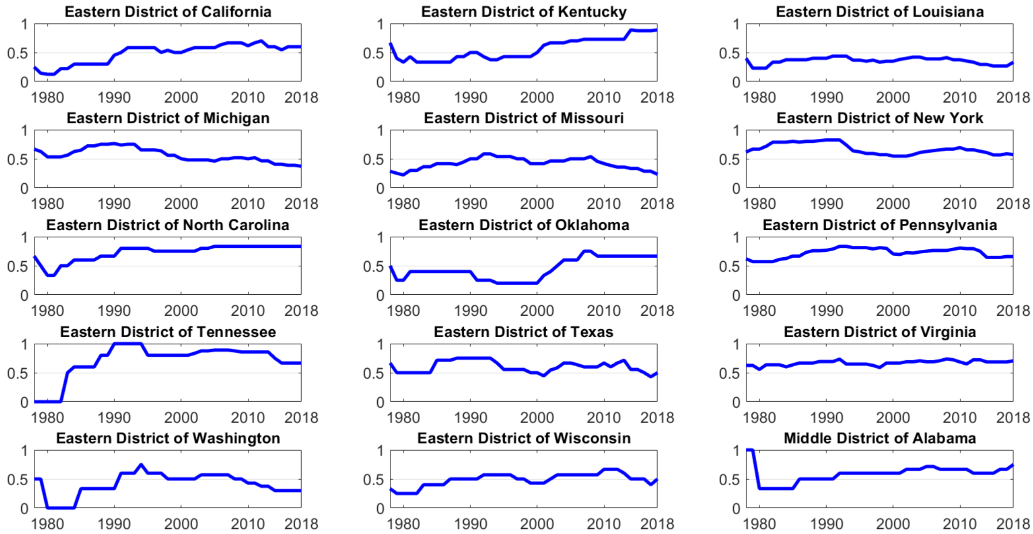


Figure C.11. Share of District Court Judges Appointed by a Republican President (4/6)

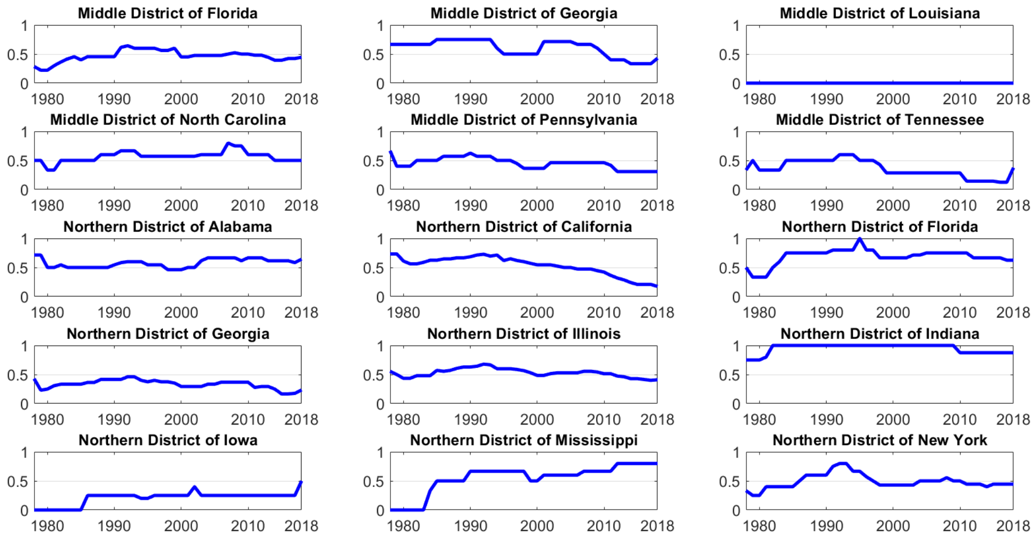


Figure C.12. Share of District Court Judges Appointed by a Republican President (5/6)

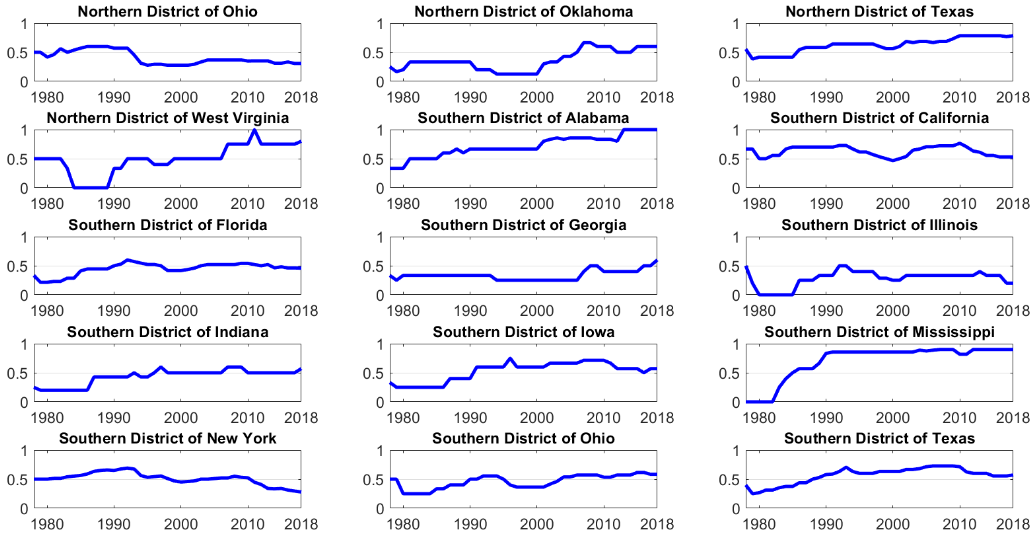
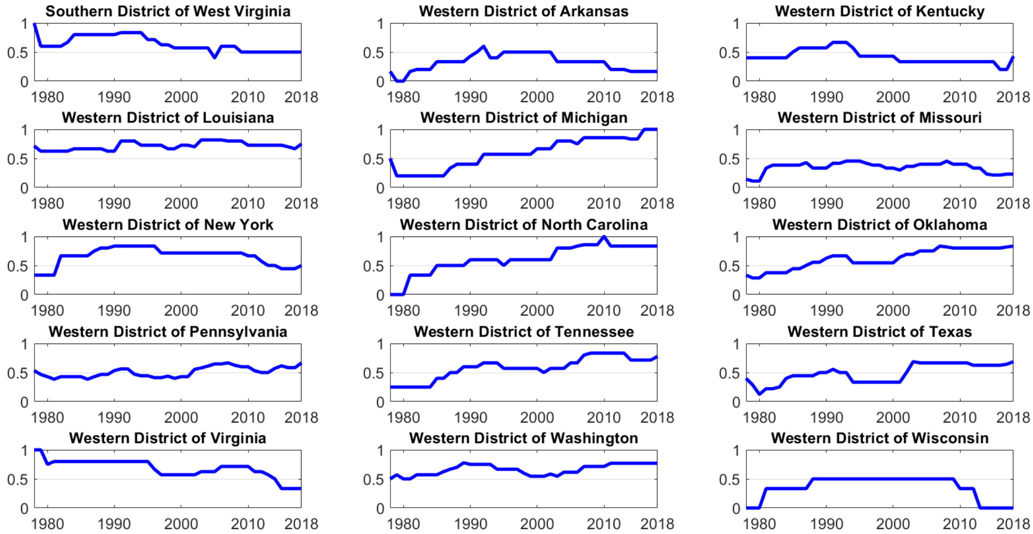


Figure C.13. Share of District Court Judges Appointed by a Republican President (6/6)



C.3 Further Rulings Regressions

In Table C.1 we present the results for several robustness checks. In Column (1), we control for local labor market conditions by including the unemployment rate and the real state GDP growth rate. This evaluation is motivated by the evidence that court rulings can be affected by economic conditions, see Section 4.2. In Column (2), we use a moving average over a four year window of our measure of Supreme Court ideology sci_t, taking into account the possibility that district court judges orientate themselves partly on past Supreme Court ideology. In Column (3), we include four (instead of one) lags of the dependent variable, allowing us to capture more general mean-reverting tendencies in district court rulings. Finally, in Column (4), we weigh rulings by the number of rulings per state population which reduces the importance of observations where unusually few rulings are published. In all of these specifications the coefficient on the interaction term between Supreme Court ideology and district court ideology remains negative and statistically siginificant.

Table C.1. Regression Results for District Court Rulings – Robustness Checks

	(1)	(2)	(3)	(4)
Supreme Court ideology	-2.0298		-2.0738	-1.7567
× district court ideology	(0.7114)		(0.7155)	(0.7227)
	p=0.0044		*p=0.0038*	*p=0.0152*
Supreme Court ideology (MA)		-1.7085		
× district court ideology		(0.9320)		
		p=0.0670		
Observations	1499	1499	1499	1499
R^2	0.2631	0.2592	0.2734	0.2748
Lagged dependent variable	yes	yes	yes	yes
Additional lags	no	no	yes	no
Weights	no	no	no	yes
State demographics	yes	yes	yes	yes
Court, judge, and case characteristics	yes	yes	yes	yes
State gov. and leg. controls	yes	yes	yes	yes
State GDP and unemployment	yes	no	no	no
Year fixed effects	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors. Supreme Court ideology (MA) = $1/4 \cdot \sum_{\tau=0}^{3} sci_{t-\tau}$.

C.4 Further Labor Market Regressions

In this section we present regressions for additional labor market outcomes and alternative specifications for the regressions in Section 4.4.

C.4.1 Additional Outcome Variables

First, we run the regression in Equation (4.6) for additional industry groups and for additional inequality measures. Table C.2 shows results for further industry groups. We find that conservative court rulings increase employment in financial industries disproportionately, while there is no discernible change in the trade and transportation employment shares. Table C.3 shows results for additional inequality measures, which support our findings of increasing inequality in response to rising judicial coservatism documented in Section 4.4.

Table C.2. Regression Results for Industry Employment Shares – Additional Industry Groups

	(1)	(2)	(3)
Dependent variable	Trade emp. share	Transport emp. share	Finance emp. share
Supreme Court ideology sci_t	0.0207	0.0206	-0.0287
$\times$ district court ideology dci_s	(0.0220)	(0.0132)	(0.0117)
	p= 0.3477	*p= 0.1177*	*p= 0.0148*
Effect of conservatism (γ^z/γ)	-0.0105	-0.0105	0.0146
Observations	1499	1499	1499
R^2	0.5962	0.6122	0.7606
Industry and occupation controls	no	no	no
State policy controls	yes	yes	yes
State gov. and leg. controls	yes	yes	yes
Year fixed effects	yes	yes	yes
State fixed effects	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Table C.3. Regression Results for Inequality – Additional Percentiles

	(1)	(2)	(3)
Dependent variable	80/20 percentiles	80/50 percentiles	50/20 percentiles
Supreme Court ideology sci_t	-0.2235	-0.1157	-0.1078
$\times$ district court ideology dci_s	(0.1037)	(0.0584)	(0.0807)
	p= 0.0313	*p= 0.0476*	*p= 0.1816*
Effect of conservatism (γ^z/γ)	0.1136	0.0588	0.0548
Observations	1499	1499	1499
R^2	0.8427	0.8077	0.7182
Industry and occupation controls	yes	yes	yes
State policy controls	yes	yes	yes
State gov. and leg. controls	yes	yes	yes
Year fixed effects	yes	yes	yes
State fixed effects	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

C.4.2 Alternative Specifications

Second, we present the results of several robustness checks. For simplicity, we concentrate on five dependent variables which represent our main results that conservative court rulings promote labor market fluidity but also contribute to wage stagnation, job market polarization, deunionization, and rising inequality. Specifically, we show results for the unemployment rate, the average hourly wage rate, the employment share in routine occupations, the union coverage rate, and the 90/10 income ratio.

In Table C.4, we additionally control for state demographics, which are also included in the regressions for district court rulings. Results are similar to the baseline case presented in Section 4.4.

In Table C.5 we take into account the possibility that the economic effects of national executive policy vary by state political orientation. Specifically, we additionally include the interaction between a dummy for a Republican president and the state-specific 2008 presidential-election voting share for John McCain as an indicator for the state's Republican orientation. The results are only slightly affected by this inclusion. We have also considered other indicators of state political orienations such as the number of years with a Republican governor or a Republican state legislative majority in our sample which led to similar results (not shown).

In Tables C.6–C.8 we leave out sets of control variables one after another. These exercises serve two purposes. First, they reveal whether our results rely on specific control variables. Second, they are informative about endogenous responses of the control variables to changing Supreme Court ideology and their effects on our variables of interest. These indirect effects allow for a more complete picture of the effects of changing Supreme Court ideology but are not part of the direct effects of ideological leanings in court rulings which we are primarily interested in.

In Table C.6, we leave out control variables for state politics. This has no effect on the direction or the significance of the effects but changes the size of some coefficients visibly. For example, the effect on wage is strengthened, suggesting that increasing Supreme Court conservatism induces changes in state governments and legislatures which further weaken workers' bargaining power. In Table C.7, we leave out control variables for state-specific policies. The effects on the results are negligible. Finally, we refrain from controlling for the state's industry-occupation composition in Table C.8. By construction, the specification for the routine employment share is unchanged because it did not control for the industry-occupation composition in the first place. Most of the results are barely affected, only for the log hourly wage rate there seem to be some counteracting composition effects which weaken the precision of the estimate.

Table C.4. Regression Results – Controlling for State Demographics

	(1)	(2)	(3)	(4)	(5)
Dependent variable	Unemployment rate	Avg. hourly wage rate	Routine emp. share	Union coverage	90/10 percentiles
Supreme Court ideology sci_t	0.0706	0.1283	0.0706	0.1095	-0.4061
$\times$ district court ideology dci_s	(0.0209)	(0.0679)	(0.0251)	(0.0256)	(0.1604)
	p= 0.0008	*p= 0.0590*	*p= 0.0049*	*p= 0.0000*	*p= 0.0115*
Effect of conservatism (γ^z/γ)	-0.0359	-0.0652	-0.0359	-0.0557	0.2064
Observations	1499	1499	1499	1499	1499
R^2	0.7617	0.9940	0.8335	0.9705	0.8266
Industry and occupation controls	yes	yes	no	yes	yes
State policy controls	yes	yes	yes	yes	yes
State gov. and leg. controls	yes	yes	yes	yes	yes
State demographics	yes	yes	yes	yes	yes
Year fixed effects	yes	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Table C.5. Regression Results – Accounting for Potentially Heterogenous Effects of National Executive Policy

	(1)	(2)	(3)	(4)	(5)
Dependent variable	Unemployment rate	Avg. hourly wage rate	Routine emp. share	Union coverage	90/10 percentiles
Supreme Court ideology sci_t	0.0635	0.1743	0.0631	0.1329	-0.4065
$\times$ district court ideology dci_s	(0.0205)	(0.0708)	(0.0251)	(0.0267)	(0.1581)
	p= 0.0020	*p= 0.0139*	*p= 0.0120*	*p= 0.0000*	*p= 0.0103*
Effect of conservatism (γ^z/γ)	-0.0323	-0.0886	-0.0323	-0.0676	0.2066
Observations	1499	1499	1499	1499	1499
R^2	0.7638	0.9933	0.8267	0.9667	0.8257
Industry and occupation controls	yes	yes	no	yes	yes
State policy controls	yes	yes	yes	yes	yes
State gov. and leg. controls	yes	yes	yes	yes	yes
Republican president $\times$ state ideology	yes	yes	yes	yes	yes
Year fixed effects	yes	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Table C.6. Regression Results – Not Controlling for State Politics

	(1)	(2)	(3)	(4)	(5)
Dependent variable	Unemployment rate	Avg. hourly wage rate	Routine emp. share	Union coverage	90/10 percentiles
Supreme Court ideology sci_t	0.0736	0.2083	0.0659	0.1420	-0.3412
$\times$ district court ideology dci_s	(0.0208)	(0.0719)	(0.0249)	(0.0269)	(0.1589)
	p= 0.0004	*p= 0.0038*	*p= 0.0084*	*p= 0.0000*	*p= 0.0320*
Effect of conservatism (γ^z/γ)	-0.0374	-0.1059	-0.0335	-0.0722	0.1734
Observations	1499	1499	1499	1499	1499
R^2	0.7535	0.9930	0.8262	0.9658	0.8217
Industry and occupation controls	yes	yes	no	yes	yes
State policy controls	yes	yes	yes	yes	yes
State gov. and leg. controls	no	no	no	no	no
Year fixed effects	yes	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Table C.7. Regression Results – Not Controlling for State Policies

	(1)	(2)	(3)	(4)	(5)
Dependent variable	Unemployment rate	Avg. hourly wage rate	Routine emp. share	Union coverage	90/10 percentiles
Supreme Court ideology sci_t	0.0824	0.1412	0.1151	0.1447	-0.4649
$\times$ district court ideology dci_s	(0.0198)	(0.0663)	(0.0237)	(0.0255)	(0.1494)
	p= 0.0000	*p= 0.0333*	*p= 0.0000*	*p= 0.0000*	*p= 0.0019*
Effect of conservatism (γ^z/γ)	-0.0419	-0.0718	-0.0585	-0.0736	0.2363
Observations	1499	1499	1499	1499	1499
R^2	0.7453	0.9932	0.8202	0.9650	0.8208
Industry and occupation controls	yes	yes	no	yes	yes
State policy controls	no	no	no	no	no
State gov. and leg. controls	yes	yes	yes	yes	yes
Year fixed effects	yes	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

Table C.8. Regression Results – Not Controlling for the Industry-Occupation Composition

	(1)	(2)	(3)	(4)	(5)
Dependent variable	Unemployment rate	Avg. hourly wage rate	Routine emp. share	Union coverage	90/10 percentiles
Supreme Court ideology sci_t	0.0658	0.0927	0.0633	0.1405	-0.5037
$\times$ district court ideology dci_s	(0.0209)	(0.0726)	(0.0250)	(0.0270)	0.1637
	p= 0.0017	*p= 0.2017*	*p= 0.0116*	*p= 0.0000*	*p= 0.0021*
Effect of conservatism (γ^z/γ)	-0.0334	-0.0471	-0.0322	-0.0714	0.2560
Observations	1499	1499	1499	1499	1499
R^2	0.7430	0.9927	0.8267	0.9647	0.8057
Industry and occupation controls	no	no	no	no	no
State policy controls	yes	yes	yes	yes	yes
State gov. and leg. controls	yes	yes	yes	yes	yes
Year fixed effects	yes	yes	yes	yes	yes
State fixed effects	yes	yes	yes	yes	yes

Note: The standard errors are reported in parantheses. The p-values are reported below the standerd errors.

C.5 Data Appendix

In Tables C.9–C.11 we provide a complete list of dependent variables and a complete list of control variables that were used in our regressions. All regressions include state and year fixed effects. GDP growth by state is calculated using data from the Bureau of Economic Analysis. The term publishing judge refers to the judge publishing the opinion in the Federal Supplement. Concerning the state legislature controls, Nebraska constitutes an exception in the sense that it is both unicameral and officially non-partisan. We decide to use the de facto majority in the Nebraska legislature for both the state house and state senate variable. However, as the state legislature in Nebraska has featured a de facto Republican majority in all years of our sample, independent of how we handle Nebraska, effects of the state legislature will be absorbed in the state fixed effect.

Table C.9. Dependent Variables

Dependent variable in Section 3 and Appendix C	
Average idelogy score (1: conservative, -1: liberal) of rulings in Economic Regulation and/or Labor Cases in federal district courts by state and year	CM

Dependent variables in Section 4 and Appendix D.2	
Unemployment rate (number unemployed divided by labor force)	CPS
Job-finding rate (inverse of average duration of unemployment in weeks)	CPS
Employment rate (number employed divided by adult population)	CPS
Avg. hourly wage rate (log of the wage rate)	CPS
Vol. PT share (log of number voluntary part-time employed divided by all employed)	CPS
PT/FT wage rate (log of voluntary part-time wages divided by full-time wages)	CPS
Employment share in construction industries	CPS
Employment share in manufacturing industries	CPS
Employment share in service industries	CPS
Employment share in abstract-intensive occupations	CPS + AD
Employment share in routine-intensive occupations	CPS + AD
Employment share in manual task intensive occupations	CPS + AD
90/10 percentiles (log of 90th percentile family income divided by 10th percentile)	CPS
90/50 percentiles (log of 90th percentile family income divided by 50th percentile)	CPS
50/10 percentiles (log of 50th percentile family income divided by 10th percentile)	CPS

Dependent variables in Appendix D.1	
Employment share in wholesale and retail trade industries	CPS
Employment share in transportation, communcations, and other public utilities industries	CPS
Employment share in finance, insurance, and real estate industries	CPS
80/20 percentiles (log of 80th percentile family income divided by 20th percentile)	CPS
80/50 percentiles (log of 80th percentile family income divided by 50 percentile)	CPS
50/20 percentiles (log of 50th percentile family income divided by 20th percentile)	CPS

Note: AD: Autor and Dorn (2013); CM: Carp and Manning (2016); CPS: Current Population Survey.

Table C.10. Independent Variables (1/2)

Regressor of interest	
Median ideology score of Supreme Court justices by year ×	Boyd + Bailey
pre-sample average ideology score of district court judges by state	
Court, Judge, and Case Characteristics	
Share of judges appointed by a Republican president	CM+FJC
Average ideology score at the responsible court of appeals	CM+FJC
Share of cases in each case type category in the U.S. District Court Database	CM
(union v. company; member v. union; employee v. employer; commercial regulation; environmental protection local/state economic; labor dispute – govt v. union/employer; rent control; excess profits)	
Average age of district court judges	FJC
Share of white district court judges	FJC
Share of male district court judges	FJC
Share of publishing judges with Republican Party affiliation	CM
Share of publishing judges with Democrat Party affiliation	CM
Share of white publishing judges	CM
Share of male publishing judges	CM
Shares of publishing judges appointed by each president	CM
Experience of publishing judges (years of service at current court, shares)	CM
State demographics	
Total adult state population	CPS
Share of population living in a central city (urban density)	CPS
Share of population living in a city (urban density)	CPS
Share of black population (ethnic composition)	CPS
Share of white population (ethnic composition)	CPS
Shares of population in age groups 16-24; 25-44; 45-54; >55	CPS
Industry-occupation controls	
Shares of workers employed in industry groups:	CPS
construction; manufacturing; transportation communcations, and other public utilities;	
wholesale and retail trade; services; finance, insurance, and real estate	
Shares of workers employed in occupational groups:	CPS + AD
abstract; manual; routine	

Note: AD: Autor and Dorn (2013); Bailey: Bailey (2013); Boyd: Boyd (2015a); CM: Carp and Manning (2016); CPS: Current Population Survey; FJC: Federal Judicial Center: Biographical Directory of Article III Federal Judges.

Table C.11. Independent Variables (2/2)

State policy controls	
Minimum wage rate	FRED
Total federal intergovernmental revenue	SLGFD
Total tax burden	TF
Public policy exception to employment at-will	ADS
Implied contract exception to employment at-will	ADS
Good faith exception to employment at-will	ADS
State gov. and leg. controls	
State senate majority party	NCLS
(1: Rep., -1: Dem., 0: other/indep./no majority)	
State house majority party	NCLS
(1: Rep., -1: Dem., 0: other/indep./no majority)	
Political party of the governor	NCLS
(1: Rep., -1: Dem., 0: other/indep./no majority)	
Additional control variables in robustness checks	
State unemployment rate	FRED
Growth rate of real state GDP	BEA
Voting share for John McCain in 2008 presidential election	FEC
× Republican president	

Note: ADS: Autor et al. (2006a); BEA: Bureau of Economic Analysis; FEC: Federal Election Commission; FRED: Federal Reserve Economic Data; NCLS: National Conference of State Legislatures; SLGFD: State and Local Government Finance Dataset; TF: Tax Foundation.

Bibliography

Abadie, A., Athey, S., Imbens, G.W., Wooldridge, J., 2017. When should you adjust standard errors for clustering? NBER Working Paper 24003.

Acemoglu, D., Aghion, P., Violante, G.L., 2001. Deunionization, technical change and inequality. Carnegie-Rochester Conference Series on Public Policy 55, 229–264.

Açıkgöz, O.T., Kaymak, B., 2014. The rising skill premium and deunionization. Journal of Monetary Economics 63, 37–50.

Aghion, P., Algan, Y., Cahuc, P., 2011. Civil society and the state: The interplay between cooperation and minimum wage regulation. Journal of the European Economic Association 9, 3–42.

Aizer, A., Doyle Jr., J.J., 2015. Juvenile incarceration, human capital, and future crime: Evidence from randomly assigned judges. Quarterly Journal of Economics 130, 759–803.

Albertini, J., Hairault, J.O., Langot, F., Sopraseuth, T., 2017. A tale of two countries: A story of the French and US polarization. IZA Discussion Papers No. 11013.

Ashworth, J., Ransom, T., 2019. Has the college wage premium continued to rise? Evidence from multiple U.S. surveys. Economics of Education Review 69, 149–154.

Atolia, M., Gibson, J., Marquis, M., 2018. Asymmetry and the amplitude of business cycle fluctuations: A quantitative investigation of the role of financial frictions. Macroeconomic Dynamics 22, 279–306.

Autor, D.H., Donohue III, J.J., Schwab, S.J., 2006a. The costs of wrongful-discharge laws. Review of Economics and Statistics 88, 211–231.

Autor, D.H., Dorn, D., 2013. The growth of low-skill service jobs and the polarization of the US labor market. American Economic Review 103, 1553–1597.

Autor, D.H., Dorn, D., Hanson, G.H., 2013. The China syndrome: Local labor market effects of import competition in the United States. American Economic Review 103, 2121–2168.

Autor, D.H., Dorn, D., Hanson, G.H., 2015. Untangling trade and technology: Evidence from local labour markets. Economic Journal 125, 621–646.

Autor, D.H., Katz, L.F., Kearney, M.S., 2006b. The polarization of the U.S. labor market. American Economic Review 96, 189–194.

Autor, D.H., Katz, L.F., Krueger, A.B., 2006c. Computing inequality: Have computers changed the labor market? Quarterly Journal of Economics 96, 863–876.

Autor, D.H., Levy, F., Murnane, R.J., 2003. The skill content of recent technological change: An empirical exploration. Quarterly Journal of Economics 118, 1279–1333.

Baccaro, L., Locke, R., 1998. The end of solidarity? The decline of egalitarian wage policies in Italy and Sweden. European Journal of Industrial Relations 4, 283–308.

Bailey, M.A., 2013. Is today's court the most conservative in sixty years? Challenges and opportunities in measuring judicial preferences. Journal of Politics 75, 821–834.

Baldwin, R., 2003. The decline of US labor unions and the role of trade. Peterson Institute for International Economics.

Barnichon, R., 2010. The lumpy job separation rate: Reconciling search models with the ins and outs of unemployment. URL: http://sites.google.com/site/regisbarnichon/cv/Lumpyjs.pdf. Last accessed on 03.12.2018.

Barro, R.J., 1977. Long-term contracting, sticky prices, and monetary policy. Journal of Monetary Economics 3, 305–316.

Barron, J.M., Berger, M.C., Black, D.A., 1997. Employer search, training and vacancy duration. Economic Inquiry 35, 167–192.

Beaudry, P., Green, D.A., Sand, B.M., 2016. The great reversal in the demand for skill and cognitive tasks. Journal of Labor Economics 34, 199–247.

Benesh, S.C., Reddick, M., 2002. Overruled: An event history analysis of lower court reaction to Supreme Court alterations of precedent. Journal of Politics 64, 534–550.

Bernstein, S., Colonnelli, E., Giroud, X., Iverson, B., 2019a. Bankruptcy spillovers. Journal of Financial Economics 133, 608–633.

Bernstein, S., Colonnelli, E., Iverson, B., 2019b. Asset allocation in bankruptcy. Journal of Finance 74, 5–53.

Blanchard, O., Diamond, P., 1990. The cyclical behavior of the gross flows of U.S. workers. Brookings Papers on Economic Activity 1990, 85–156.

Blanchard, O., Galí, J., 2010. Labor markets and monetary policy: A new keynesian model with unemployment. American Economic Journal: Macroeconomics 2, 1–30.

Boal, W.M., Pencavel, J., 1994. The effects of labor unions on employment, wages, and days of operation: Coal mining in west virginia. Quarterly Journal of Economics 109, 267–298.

Booth, A., Bryan, M.L., 2004. The union membership wage-premium puzzle: Is there a free rider problem? International Labor Relations Review 57, 402–421.

Boyd, C.L., 2015a. Federal district court judge ideology data. Available at: http://cLboyd.net/ideology.html.

Boyd, C.L., 2015b. The hierarchical influence of courts of appeals on district courts. Journal of Legal Studies 44, 113–141.

Boyd, C.L., Epstein, L., Martin, A.D., 2010. Untangling the causal effects of sex on judging. American Journal of Political Science 54, 389–411.

Bradley, D., Kim, I., Tian, X., 2017. Do unions affect innovation? Management Science 63, 2251–2271.

Brändle, T., Goerke, L., 2018. The one constant: A causal effect of collective bargaining on employment growth? Evidence from German linked-employer-employee data. Scottish Journal of Political Economy 65, 445–478.

von Brasch, T., Gjelsvik, M.L., Sparrman, V., 2018. Deunionization and job polarization – a macroeconomic model analysis for a small open economy. Economic Systems Research 30, 380–399.

Breda, T., 2015. Firms' rents, workers' bargaining power and the union wage premium. Economic Journal 125, 1616–1652.

Brügemann, B.A., 2017. Privately efficient wage rigidity under diminishing returns. 2017 Meeting Papers 978, Society for Economic Dynamics.

Bryson, A., 2002. The union membership wage premium: An analysis using propensity score matching. CEP Discussion Papers No. 530.

Bryson, A., Dale-Olsen, H., Nergard, K., 2016. Gender differences in the union membership wage premium? A comparative case study. European Journal of Industrial Relations, forthcoming.

Buera, F.J., Shin, Y., 2013. Financial frictions and the persistence of history: A quantitative exploration. Journal of Political Economy 121, 221–272.

Cahuc, P., Carcillo, S., Patault, B., 2019. Are pro-worker judges detrimental to firm survival and employment? Working Paper.

Cahuc, P., Wasmer, E., 2001. Does intrafirm bargaining matter in the large firms matching model. Macroeconomic Dynamics 5, 742–747.

Calmfors, L., Booth, A., Burda, M., Checchi, D., Naylor, R., Visser, J., 2001. The Role of Unions in the Twenty-First Century: A Report for the Fondazione Rodolfo Debenedetti. Oxford University Press. chapter 1. pp. 1–158.

Calvo, G.A., 1983. Staggered prices in a utility-maximizing framework. Journal of Monetary Economics 12, 383–398.

Calvo, G.A., 2015. The liquidity approach to bubbles, crises, jobless recoveries, and involuntary unemployment. Journal Economía Chilena 18, 4–27.

Calvo, G.A., Coricelli, F., Ottonello, P., 2014. Labor market, financial crises and inflation: Jobless and wageless recoveries. NBER Working Paper No. 18480.

Cannon, B.C., Johnson, C.A., 1984. Judicial policies: Implementation and impact. Washington, DC: Congressional Quarterly Press.

Card, D., Lemieux, T., Riddel, C.W., 2004. Unions and wage inequality. Journal of Labor Research 25, 519–559.

Carp, R.A., Manning, K.L., 2016. U.S. district court database. 2016 version n=110977. URL: http://districtcourtdatabase.org.

Checchi, D., Visser, J., van de Werfhorst, H.G., 2010. Inequality and union membership: The influence of relative earnings and inequality attitudes. British Journal of Industrial Relations 48, 84–108.

Chew, P.K., Robert E, K., 2009. Myth of the color-blind judge: An empirical analysis of racial harassment cases. Washington University Law Review 86, 1117–1166.

Choi, S.J., Gulati, M., Posner, E.A., 2012. What do federal district judges want? An analysis of publications, citations, and reversals. Journal of Law, Economics, and Organization 28, 518–549.

Chugh, S.K., 2013. Costly external finance and labor market dynamics. Journal of Economic Dynamics and Control 37, 2882–2912.

Cohen, A., Yang, C.S., 2019. Judicial politics and sentencing decisions. American Economic Journal: Economic Policy 11, 160–191.

Connolly, R.A., Hirsch, B.T., Hirschey, M., 1986. Union rent seeking, intangible capital, and market value of the firm. Review of Economics and Statistics 68, 567–577.

Cortes, G.M., Gallipoli, G., 2017. The cost of occupational mobility: An aggregate analysis. Journal of the European Economic Association 16, 275–315.

Dahl, G.B., Kostøl, A.R., Mogstad, M., 2014. Family welfare cultures. Quarterly Journal of Economics 129, 1711–1752.

DiNardo, J., Fortin, N.M., Lemieux, T., 1996. Labor market institutions and the distribution of wages, 1973-1992: A semiparametric approach. Econometrica 64, 1001–1044.

DiNardo, J., Lee, D.S., 2004. Economic impacts of new unionization on private sector employers: 1984-2001. Quarterly Journal of Economics 119, 1383–1441.

Dinlersoz, E., Greenwood, J., 2016. The rise and fall of unions in the United States. Journal of Monetary Economics 83, 129–146.

Dobbie, W., Goldin, J., , Yang, C.S., 2018. The effects of pretrial detention on conviction, future crime, and employment: Evidence from randomly assigned judges. American Economic Review 108, 201–240.

Duclos, J.Y., Esteban, J., Ray, D., 2004. Polarization: Concepts, measurement, estimation. Econometrica 72, 1737–1772.

Dunn, M., Walker, J., 2016. BLS spotlight on statistics: Union membership in the United States. U.S. Department of Labor, Bureau of Labor Statistics.

Dustmann, C., Schönberg, U., Stuhler, J., 2017. Labor supply shocks, native wages, and the adjustment of local employment. Quarterly Journal of Economics 132, 435–483.

Edwards, B.C., 2019. Why appeals courts rarely reverse lower courts: An experimental study to explore affirmation bias. Emory Law Journal Online 68, 1036–1058.

Eisenberg, T., 2004. Appeal rates and outcomes in tried and nontried cases: Further exploration of anti-plaintiff appellate outcomes. Cornell Law Faculty Publications. Paper 359.

Elsby, M.W., Michaels, R., Solon, G., 2009. The ins and outs of cyclical unemployment. American Economic Journal: Macroeconomics 1, 84–110.

Epstein, L., Landes, W.M., Posner, R.A., 2013. How business fares in the Supreme Court. Minnesota Law Review 97, 1431–1472.

Epstein, L., Martin, A.D., Segal, J.A., Westerland, C., 2007. The judical common space. Journal of Law, Economics & Organization 23, 303–325.

Eren, O., Mocan, N., 2018. Emotional judges and unlucky juveniles. American Economic Journal: Applied Economics 10, 171–205.

Farber, H.S., Herbst, D., Kuziemko, I., Naidu, S., 2018. Unions and inequality over the twentieth century: New evidence from survey data. NBER Working Papers No. 24587.

Farber, H.S., Krueger, A.B., 1992. Union membership in the United States: The decline continues. NBER Working Paper 4216.

Federal Judicial Center, 2020. History of the federal judiciary. http://www.fjc.gov. Web site of the Federal Judicial Center, Washington, DC.

Feng, A., Graetz, G., 2015. Rise of the machines: The effects of labor-saving innovations on jobs and wages. CEP Discussion Papers No. dp1330.

Fitzenberger, B., Kohn, K., Lembcke, A., 2006. Union density and varieties of coverage: The anatomy of union wage effects in Germany. International Labor Relations Review 66, 169–197.

Flood, S., King, M., Rodgers, R., Ruggles, S., Warren, J.R., 2020. Integrated Public Use Microdata Series, Current Population Survey: Version 7.0 [dataset]. Minneapolis, MN: IPUMS.

Frandsen, B.R., 2012. Why unions still matter: The effects of unionization of the distribution of employee earnings. Working Paper.

French, E., Song, J., 2014. The effect of disability insurance receipt on labor supply. American Economic Journal: Economic Policy 6, 291–337.

Fujita, S., Ramey, G., 2006. The cyclicality of job loss and hiring. Working Papers 06-17, Federal Reserve Bank of Philadelphia.

Galí, J., Smets, F., Wouters, R., 2012. Slow recoveries: A structural interpretation. Journal of Money, Credit and Banking 44, 9–30.

Garín, J., 2015. Borrowing constraints, collateral fluctuations, and the labor market. Journal of Economic Dynamics and Control 57, 112–130.

Gathmann, C., Schönberg, U., 2010. How general is human capital? A task-based approach. Journal of Labor Economics 28, 1–49.

Gennaioli, N., Shleifer, A., 2008. Judicial fact discretion. Journal of Legal Studies 37, 1–35.

Gertler, M., Trigari, A., 2009. Unemployment fluctuations with staggered nash bargaining. Journal of Political Economy 117, 38–86.

Giles, M.W., Hettinger, V.A., Peppers, T., 2001. Picking federal judges: A note on policy and partisan selection agendas. Political Research Quarterly 54, 623–641.

Gilman, M.E., 2014. A court for the one percent: How the Supreme Court contributes to economic inequality. Utah Law Review 2014, 389–463.

Glynn, A.N., Sen, M., 2015. Identifying judicial empathy: Does having daughters cause judges to rule for women's issues? American Journal of Political Science 59, 37–54.

Goldsmith-Pinkham, P., Sorkin, I., Swift, H., 2018. Bartik instruments: What, when, why, and how. NBER Working Paper 24408.

Goos, M., Manning, A., Salomons, A., 2009. Job polarization in Europe. American Economic Review 99, 58–63.

Gottschalk, P., 2005. Downward nominal wage flexibility: Real or measurement error. Review of Economics and Statistics 87, 556–568.

Groshen, E., Potter, S., 2003. Has structural change contributed to a jobless recovery? Current Issues in Economics and Finance - Federal Reserve Bank of New York 9, 1–7.

Haefke, C., Sonntag, M., van Rens, T., 2013. Wage rigidity and job creation. Journal of Monetary Economics 60, 887–899.

Hall, R.E., 2005. Employment fluctuations with equilibrium wage stickiness. American Economic Review 95, 50–65.

Heyes, A., Saberian, S., 2019. Temperature and decisions: Evidence from 207,000 court cases. American Economic Journal: Applied Economics 11, 238–265.

Hirsch, B.T., Link, A.N., 1987. Labor union effects on innovative activity. Journal of Labor Research 8, 323–331.

Hirsch, B.T., Macpherson, D.A., 2003. Union membership and coverage database from the current population survey: Note. Industrial and Labor Relations Review 56, 349–354.

Hirsch, B.T., Schumacher, E.J., 2004. Match bias in wage gap estimates due to earnings imputation. Journal of Labor Economics 22, 689–722.

Ichino, A., Polo, M., Rettore, E., 2003. Are judges biased by labor market conditions? European Economic Review 47, 913–944.

Jaeger, D., Ruist, J., Stuhler, J., 2018. Shift-share instruments and the impact of immigration. NBER Working Paper 24285.

Jaimovich, N., Siu, H.E., 2018. The trend is the cycle: Job polarization and jobless recoveries. *Review of Economics and Statistics*, forthcoming.

Jermann, U.J., Quadrini, V., 2012. Macroeconomic effects of financial shocks. American Economic Review 102, 238–271.

Kaas, L., Kircher, P., 2015. Efficient firm dynamics in a frictional labor market. American Economic Review 105, 3030–3060.

Kambourov, G., Manovskii, I., 2009. Occupational specificity of human capital. International Economic Review 50, 63–115.

Kastellec, J.P., 2013. Racial diversity and judicial influence on appellate courts. American Journal of Political Science 57, 167–183.

Kiyotaki, N., Moore, J., 1997. Credit cycles. Journal of Political Economy 105, 211–248.

Kling, J.R., 2006. Incarceration length, employment, and earnings. American Economic Review 96, 863–876.

Knepper, M., 2017. When the shadow is the substance: Judge gender and the outcomes of workplace sex discrimination cases. Journal of Labor Economics 36, 623–664.

Krause, M.U., Lubik, T., 2007. The (ir)relevance of real wage rigidity in the new keynesian model with search frictions. Journal of Monetary Economics 54, 706–727.

Kugler, A.D., Saint-Paul, G., 2004. How do firing costs affect worker flows in a world with adverse selection? Journal of Labor Economics 22, 553–584.

Lee, W., Roemer, J.E., 2005. The rise and fall of unionised labour markets: A political economy approach. Economic Journal 115, 28–67.

Lewis, P., 1989. The unemployed and trade union membership. Industrial Relations Journal 20, 271–279.

Marinescu, I., 2011. Are judges sensitive to economic conditions? Evidence from UK employment tribunals. International Labor Relations Review 64, 673–698.

Martin, A., Quinn, K., 2002. Dynamic ideal point estimation via markov chain monte carlo for the U.S. Supreme Court, 1953-1999. Political Analysis 10, 134–153.

Mayer, G., 2004. Union membership trends in the United States. Congressial Research Service.

McKay, A., Reis, R., 2008. The brevity and violence of contractions and expansions. Journal of Monetary Economics 55, 738–751.

Meltzer, A., 2003. Jobless recovery. The Wallstreet Journal.

Merz, M., Yashiv, E., 2007. Labor and the market of the firms. American Economic Review 97, 1419–1431.

Meyer, B., 2019. Financialization, technological change, and trade union decline. Socio-Economic Review 17, 477–502.

Miceli, T.J., Coşgel, M.M., 1994. Reputation and judicial decision-making. Journal of Economic Behavior and Organization 23, 31–51.

Michaels, G., Natraj, A., Reenen, J.V., 2014. Has ICT polarized skill demand? Evidence from eleven countries over twenty-five years. Review of Economics and Statistics 96, 60–77.

Michaillat, P., 2012. Do matching frictions explain unemployment? Not in bad times. American Economic Review 102, 1721–1750.

Mincer, J., 1974. Schooling, experience, and earnings. Columbia University Press, NY.

Moiseeva, Y., 2018. The interaction between credit and labour market frictions. Working Paper.

Montgomery, E., 1989. The employment and unemployment effects of unions. Journal of Labor Economics 7, 170–190.

Mortensen, D.T., Pissarides, C.A., 1994. Job creation and job destruction in the theory of unemployment. The Review of Economic Studies 61, 397–415.

Nickell, S., Layard, R., 1999. Labor market institutions and economic performance. Handbook of Economics 3, 3029–2084.

Oh, W., 1989. Craft versus Industrial Unions: Union Organization Within the Work Place. Ph.D. thesis. The Ohio State University.

Ortigueira, S., 2013. The rise and fall of centralized wage bargaining. Scandinavian Journal of Economics 115, 825–855.

Perri, F., Quadrini, V., 2018. International recessions. American Economic Review 108, 935–984.

Petrongolo, B., 2001. Reemployment probabilities and returns to matching. Journal of Labor Economics 19, 716–741.

Petrongolo, B., Pissarides, C.A., 2001. Looking into a black box: A survey of the matching function. Journal of Economic Literature 39, 390–431.

Petrosky-Nadeau, N., 2014. Credit, vacancies and unemployment fluctuations. Review of Economic Dynamics 17, 191–205.

Petrosky-Nadeau, N., Wasmer, E., 2015. Macroeconomic dynamics in a model of goods, labor, and credit market frictions. Journal of Monetary Economics 72, 97–113.

Pfeifer, J., 2016. Macroeconomic effects of financial shocks: Comment. Dynare Working Papers, 50, CEPREMAP.

Philippe, A., Ouss, A., 2018. "no hatred or malice, fear or affection": Media and sentencing. Journal of Political Economy 126, 2134–2178.

Posner, R.A., 2005. Judicial behavior and performance: An economic approach. 32 Florida State University Law Review 1259.

Post, C., 2017. The roots of Trumpism. Cultural Dynamics 29, 100–108.

Randazzo, K.A., 2008. Strategic anticipation and the hierarchy of justice in U.S. District Courts. American Politics Research 36, 669–693.

Rogers, J., Streeck, W., 1995. Works councils: Consultation, representation, and cooperation in industrial relations. National Bureau of Economic Research, Inc.

Rueda, D., Pontusson, J., Way, C.R., 2002. Comparative political economy of wage distribution: The role of partisanship and labour market institutions. British Journal of Political Science 32, 281–308.

Schanzenbach, M.M., Tiller, E.H., 2007. Strategic judging under the U.S. sentencing guidelines: Positive political theory and evidence. Journal of Law, Economics, and Organization 23, 24–56.

Schanzenbach, M.M., Tiller, E.H., 2008. Reviewing the sentencing guidelines: Judicial politics, empirical evidence, and reform. Journal of Law, Economics, and Organization 75, 715–760.

Schmitt-Grohé, S., Uribe, M., 2017. Liquidity traps and jobless recoveries. American Economic Journal: Macroeconomics 9, 165–204.

Schoefer, B., 2016. The financial channel of wage rigidity. 2016 Meeting Papers 1605, Society for Economic Dynamics.

Schott, I., 2013. Startups, credit, and the jobless recovery. 2013 Meeting Papers 340, Society for Economic Dynamics.

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

Shimer, R., 2012. Wage rigidities and jobless recoveries. Journal of Monetary Economics 59, 65–77.

Sichel, D.E., 1993. Business cycle asymmetry: A deeper look. Economic Inquiry 31, 224–236.

Silva, J.I., Toledo, M., 2009. Labor turnover costs and the cyclical behavior of vacancies and unemployment. Macroeconomic Dynamics 13, 76–96.

Smith, C.L., 2013. The dynamics of labor market polarization. Finance and Economics Discussion Series 2013-57, Board of Governors of the Federal Reserve System (U.S.).

Songer, D.R., Segal, J.A., Cameron, C.M., 1994. The hierachy of justice: Testing a principal-agent model of Supreme Court-circuit court interactions. American Journal of Political Science 38, 673–696.

Stock, J.H., Watson, M.W., 1999. Forecasting inflation. Journal of Monetary Economics 44, 293–335.

Stole, L.A., Zwiebel, J., 1996. Intra-firm bargaining under non-binding contracts. Review of Economic Studies 63, 375–410.

Streeck, W., 2005. The Sociology of Labor Markets and Trade Unions. In: N.J. Smelser and R.Swedberg *The Handbook of Economic Sociology*. Princeton University Press. pp. 254–283.

Swenson, K., 2004. Federal district court judges and the decision to publish. The Justice System Journal 25, 121–142.

Taha, A.E., 2004. Publish or paris? Evidence of how judges allocate their time. American Law and Economics Review 6, 1–27.

Taschereau-Dumouchel, M., 2017. The union threat. Working Paper.

Taylor, J.B., 1999. Staggered price and wage setting in macroeconomics. Handbook of Macroeconomics 1, 1009–1050.

Tiede, L., Carp, R., Manning, K.L., 2010. Judicial attributes and sentencing-deviation cases: Do sex, race, and politics matter? Justice System Journal 31, 249–272.

Traxler, F., 1994. Collective bargaining: Levels and coverage. Employment Outlook 1994, OECD, 167–194.

Troy, L., Sheflin, N., 1985. U.S. union sourcebook. Industrial Relations Data and Information Services, West Orange, New Jersey.

Tüzemen, D., Willis, J.L., 2013. The vanishing middle: Job polarization and workers' response to the decline in middle skill jobs. Economic Review Q I, 5–32.

Waddington, J., 2005. Trade union membership in Europe: The extent of the problem and range of union responses. Host publication.

Wasby, S.L., 1970. The impact of the United States Supreme Court: Some perspectives. Homewood, Il.: Dorsey Press.

Wasmer, E., 2006. General versus specific skills in labor markets with search frictions and firing costs. American Economic Review 96, 811–831.

Wasmer, E., Weil, P., 2004. The macroeconomics of credit and labor market imperfections. American Economic Review 94, 944–963.

Welch, S., Combs, M., Gruhl, J., 1988. Do black judges make a difference. American Journal of Political Science 32, 126–136.

Wesselbaum, D., 2019. Jobless recoveries: The interaction between financial and search frictions. Journal of Macroeconomics 61, 103126.

West Publishing Company, 1994. Publication Guide for Judges of the United States District Courts. St. Paul: West Publishing Company.

Yang, C.S., 2015. Free at last? judicial discretion and racial disparities in federal sentencing. Journal of Legal Studies 44, 75–111.

Zorn, C., Bowie, J., 2010. Ideological influences on decision making in the federal judicial hierarchy: An empirical assessment. The Journal of Politics 72, 1212–1221.

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Forschungsprojekte und Veröffentlichungen

Die Polarisierung des Arbeitsmarktes und die Bedeutung von Gewerkschaften, *mit Anna Hartmann, Theresa Markefke und Rebekka Rehm*
Veröffentlicht in *Kölner Impulse zur Wirtschaftspolitik*, Ausgabe 05/2019

Collateral Constraints, Wage Rigidity, and Jobless Recoveries

A Joint Theory of Polarization and Deunionization, *mit Anna Hartmann*

Out-Lawed: Estimating the Effects of Ideological Leanings of U.S. Supreme Court Justices on Labor Market Prospects, *mit Christian Bredemeier und Anna Hartmann*

Konferenzen und Summer Schools

Konferenzvorträge

2019 Royal Economic Society Annual Conference (Warwick), Midwest Macroeconomics Meeting (Athens, GA), European Association of Labour Economists Annual Conference (Uppsala), Jahrestagung Verein für Socialpolitik (Leipzig), IAB/IWH Workshop zur Arbeitsmarktpolitik (Halle)

2018 Ruhr Graduate School Doctoral Conference (Essen), Search and Matching Annual Conference (Cambridge), European Economic Association Annual Conference (Köln), European Association of Labour Economists Annual Conference (Lyon)

2017 Jahrestagung Verein für Socialpolitik (Wien), Rhineland Workshop (Bonn)

2016 IZA-CMR PhD Workshop (Bonn)

Summer Schools

08/2015 LSE Methods Summer Programme, *Tools for Macroeconomists: Advanced Tools*, Kursleiter: Prof. Wouter den Haan, PhD

06/2015 – 07/2015 Barcelona GSE Summer School of Economics, *Labor Market Outcomes*, Kursleiter: Prof. Robert Shimer, PhD

Lehrerfahrung

Seit 10/2019 Fieldseminar in *Growth, Education and Inequality in the Global Economy*, Masterkurs, für Prof. Michael Krause, PhD

Seit 04/2019 Bachelorseminar *Volkswirtschaftslehre (CMR)*, Bachelorkurs

Seit 10/2015 Übung zu *Wirtschaftspolitik II: Arbeitsmarkt- und Konjunkturpolitik*, Bachelorkurs, für Prof. Michael Krause, PhD

10/2015 – 09/2018 Übung zu *Topics in Macroeconomics, Money and Financial Markets*, Bachelorkurs, für Jun.-Prof. Dr. Martin Scheffel, Jun.-Prof. Dr. Thomas Schelkle und Jun.-Prof. Dr. Paul Schempp

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ILIAS Regelmäßige Anwendung seit 5 Jahren (sehr gute Kenntnisse)

LaTeX Tägliche Anwendung seit 5 Jahren (sehr gute Kenntnisse)

Matlab Regelmäßige Anwendung seit 5 Jahren (gute Kenntnisse)

Microsoft Office Regelmäßige Anwendung seit 5 Jahren (gute Kenntnisse)

Stata Anwendung während eines Forschungsprojekts (gute Kenntnisse)

TYPO3 Bearbeitung der Inhalte der CMR-Homepage (gute Kenntnisse)

MySQL Anwendung als Werkstudent bei Adcloud (Grundkenntnisse)

R Anwendung als Werkstudent bei Adcloud (Grundkenntnisse)

Sprachen

Deutsch Muttersprache

Englisch Verhandlungssicher in Wort und Schrift (GER C1)

Französisch Grundkenntnisse (GER A2)

Bergheim, 09.05.2020

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