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Credit Supply, Firms, and Earnings Inequality

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Abstract

We study the distributional consequences of monetary policy-induced credit supply in the German labor market. Firms in relationships with banks that are more exposed to the introduction of negative interest rates in 2014 experience a relative contraction in credit supply, associated with lower average wages. Within firms, initially lower-paid workers are more likely to leave employment, while initially higher-paid workers see a relative decline in wages. Between firms, wages fall by more at initially higher-paying employers. Our results suggest that credit affects the distribution of wages and employment both within and between firms.

JEL classification

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Keywords

wages, employment, distribution, credit supply, monetary policy, downward wage rigidity

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1 Introduction

Mounting empirical evidence highlights firms' important role in accounting for differences in labor market outcomes across workers (e.g., Card et al., 2013, 2018). To operate, firms commonly rely on external financing to fund working capital and produce value added, two-thirds of which accrues to labor (Schoefer, 2021). Consequently, firms' access to credit may affect workers' wages and employment. Motivated by these observations, we study the distributional effects of credit supply within and between firms. Our goal is to test the hypothesis that credit supply interacts with firm heterogeneity in determining workers' labor market outcomes.

To this end, we build a novel dataset that links workers' employment histories to firms' financials and banking relationships in Germany. We exploit banks' lending reaction to the introduction of negative monetary policy rates by the European Central Bank (ECB) as a credit supply shock. We show that firms in preexisting bank relationships that were more exposed to negative rates experience a relative reduction in credit supply. In turn, the relative reduction in credit supply is associated with lower firm-level average wages and employment. Our main result concerns the heterogeneous effects of credit supply on workers within and between firms. Within firms, initially lower-paid workers are more likely to leave employment, while higher-paid workers experience relative wage declines. Between firms, wages fall by more at initially higher-paying employers. Altogether, we find that a reduction in credit supply lowers wage inequality both within and between firms. This finding suggests that credit supply matters not only for average wages and employment, but also for their distribution across workers and firms.

To guide our empirical investigation, we develop a simple equilibrium model with search and credit frictions. Both frictions are necessary to understand the distributional consequences of credit supply: absent search frictions, workers receive identical pay across firms; absent credit frictions, credit supply has no role in determining pay. Because firms finance their labor expenses using debt, we show that those with more binding borrowing constraints have lower effective productivity, which reduces firm-level wages and employment. If wages are relatively rigid for low-skill workers, then credit tightening causes wages to decline more among high-skill workers and even more at firms with higher effective productivity. Consequently, a reduction in credit leads to lower within- and between-firm wage inequality.

We test these predictions by studying the ECB's introduction of negative deposit facility rates in 2014—a significant event in the European Economic and Monetary Union's history. As banks were reluctant or unable to pass on negative rates to their depositors, more deposit-reliant banks faced higher funding costs and lower net worth, resulting in lending cuts to their preexisting borrowers relative to less deposit-reliant banks (Heider et al., 2019; Ulate, 2021; Eggertsson et al., 2024). Our empirical strategy exploits heterogeneity in preexisting bank-firm relationships around the introduction of negative interest rates to generate variation in firm-level credit supply. An open question is how such a credit supply shock affects different workers and firms in the labor market.

To investigate the distributional effects of credit supply in the labor market, we proceed in two steps. In the first step, we show that negative interest rates lead firms in relationships with more deposit-reliant banks to experience a relative contraction in credit. To show that the reduction in firm borrowing is driven by credit supply rather than demand, we exploit between-bank variation in exposure to negative rates while controlling for time-varying unobserved firm heterogeneity (Khwaja and Mian, 2008). Furthermore, we find that firms cannot perfectly offset the drop in credit by switching to other banks or alternative sources of debt financing, with more affected firms reducing their leverage and building cash reserves. Firm-level deleveraging is relatively more pronounced at firms with lower cash buffers, less collateral, and more initial leverage.

In the second step, we study the effects of this credit supply shock on worker-level labor market outcomes. A reduction in credit leads to lower *average* firm-level wages. In terms of magnitudes, a one standard deviation increase in exposure to the negative credit supply shock is associated with a reduction in mean wages of up to 1.2 percent. Our estimates control for time-varying unobserved heterogeneity at the state and industry (and state by industry) levels as well as worker-firm match heterogeneity, which would be confounded with changes in worker composition absent individual-level microdata.

This average effect masks important heterogeneity across workers *within* firms, and our model suggests that it is also linked to heterogeneous unemployment risk. To test this, we estimate individual wage equations with controls for worker-firm match-specific and time-varying firm pay components (Engbom et al., 2023). We find that initially lower-paid workers are more likely to leave employment, while initially higher-paid workers see relative wage declines. A one standard deviation increase in exposure to the negative credit supply shock is associated with a reduction in top-tercile wages of around 0.8 percent relative to bottom-tercile workers. Consequently, within-firm wage inequality decreases. At the same time, the probability of leaving employment increases by around 0.6 percentage points per standard deviation of exposure among workers in the bottom relative to the top tercile.

We then extend our analysis to firm-level aggregates, including changes in employment and worker composition. In doing so, we account for worker separations and hires, which we previously held constant by including worker-firm match-specific controls. We continue to find that firms in preexisting relationships with high-deposit banks exhibit lower within-firm wage inequality than those in preexisting relationships with low-deposit banks following the introduction of negative rates. This also holds for the much smaller subsample of publicly listed firms.

For the largest publicly listed firms, which tend to offer variable compensation to their top management, we can further zoom in on the within-firm pay distribution by incorporating data on executive board members' compensation. Doing so, we confirm the inequality-reducing treatment effect. Importantly, consistent with the idea that, due to downward wage rigidities, wages are more likely to be rationed at the top, this treatment effect operates primarily through affected firms' adjustments to variable compensation. In addition, we show at the firm level that firms in preexisting relationships with high-deposit banks experience reduced employment, especially among nonmanagerial staff. As the latter are more likely from the bottom of the wage distribution, this is once again consistent with our worker-level results.

Credit supply also affects the distribution of wages and employment *between* firms. To this end, we estimate specifications that include an interaction term with firms' initial pay rank while controlling for worker-firm match heterogeneity and industry-state-year fixed effects. We find that among firms equally exposed to the credit supply shock, wages decline by more at initially higher-paying firms, while the probability of leaving employment rises by more at initially lower-paying firms. Consequently, between-firm wage inequality decreases.

In summary, we show that credit supply affects the distribution of wages and employment within and between firms. Our findings suggest that credit supply, firms, and earnings inequality are interlinked, in line with our equilibrium model with search and credit frictions. Finally, we provide evidence that our findings extend beyond the German setting. Based on identified US monetary policy shocks, contractionary surprises compress the state-level P90/P10 wage ratio in states with above-median minimum wages—with the effect driven by the top of the distribution and concentrated in more bank-dependent states. This pattern mirrors our results from Germany and supports the external validity of the distributional effects of credit supply, driven by two mechanisms: firms' dependence on bank credit and downward wage rigidity at the bottom of the wage distribution.

Related Literature. We contribute to an emerging literature on the distributional consequences of firms' external financing conditions. Specifically, we study the effects of a credit supply shock on the distribution of wages within and between firms by building on insights from related research on the pass-through of other firm-level shocks. Previous work has studied the pass-through of shocks to revenue productivity (Guiso et al., 2005; Lamadon et al., 2022; Friedrich et al., 2024), innovation (Van Reenen, 1996; Kline et al., 2019; Kogan et al., 2023), cash (Howell and Brown, 2022), and taxes (Fuest et al., 2018). Our estimates are consistent with previous pass-through studies in that we find positive wage sensitivity to credit and greater sensitivity among higher-ranked workers. A distinguishing feature of our paper is its focus on the heterogeneous pass-through of credit to high- versus low-paid workers at high- versus low-paying firms.

By studying the link between credit supply and wage inequality, our work highlights the role of firm heterogeneity in the labor market. Firms have been shown to be the natural unit of analysis for wage-setting and employment decisions (Card et al., 2013, 2018). Related work by Michelacci and Quadrini (2009) and Guiso et al. (2013) shows that credit matters for new hires' initial levels and subsequent growth of average wages. Our work complements theirs by highlighting the distributional effects of credit in the labor market, in that credit has heterogeneous impacts on wages and employment within and between firms.

Firm credit has been the focus of a tradition of work based on frictionless labor markets (e.g., Neumeyer and Perri, 2005; Jermann and Quadrini, 2012). We study credit supply at a more granular level by empirically tracing its effects on the distribution of workers' wages and employment at differentially affected firms. To make this possible, we build a novel dataset that tracks the complete credit chain—from banks to firms to workers—in Germany. This allows us to shed light on the distributional consequences of credit at the worker level that would remain hidden in more aggregate data due to changes in the underlying worker composition.

Conceptually, our paper highlights credit supply as a source of wage inequality. This complements studies of the effects of credit on employment (Chodorow-Reich, 2014; Berton et al., 2018; Benmelech et al., 2019; Caggese et al., 2019; Barbosa et al., 2020; Bergman et al., 2022; Coglianese et al., 2023).[1] While employment significantly responds to credit supply, the effect of credit on wages matters for the majority of workers who remain employed. Consequently, it matters to what extent

[1]In other settings, credit supply has been shown to affect productivity (Gilchrist et al., 2013), employment (Chodorow-Reich, 2014), product pricing (Gilchrist et al., 2017), investment (Amiti and Weinstein, 2018), innovation (Huber, 2018), household demand (Mian et al., 2020), organizational structure (Sforza, 2020), aggregate output (Herreño, 2023), and firm dynamics (Acabbi et al., 2025).

different workers are heterogeneously affected by credit. A recent strand of work has examined the response of wages to firm-level credit shocks (Arabzadeh et al., 2020; Jasova et al., 2021; Fonseca and Van Doornik, 2022; Adamopoulou et al., 2023). Relatedly, Broer et al. (2026) quantify the effects of monetary policy shocks on workers' job-finding and separation rates in Germany. Our work adds to this recent strand by documenting the distributional effects of credit supply on pay within and between firms, which we find are consistent with our simple equilibrium model.

2 Equilibrium Model

We would like a conceptual framework that links credit supply to the distribution of wages and employment within and between firms facing frictions in both credit and labor markets. Credit frictions imply that firms finance their wage bill and recruiting costs subject to idiosyncratic borrowing constraints (Neumeyer and Perri, 2005; Jermann and Quadrini, 2012). Our framework incorporates a role for firms' idiosyncratic credit constraints to affect both employment and wages. Labor market frictions give rise to market power, resulting in identical workers receiving different pay across employers, as in the seminal framework by Burdett and Mortensen (1998).

2.1 Workers

Workers are infinitely lived, risk-neutral, and discount time at a rate ρ. Ex ante, they differ in their skill $a \in \{a_L, a_H\}$, where $0 < a_L < a_H$, which we refer to as low-skill and high-skill, with population shares μ_a. Ex post, workers are either employed at some wage w or unemployed.

Job Search. Unemployed workers enjoy flow utility b_a, where $b_{a_L} \leq b_{a_H}$, engage in random job search within labor markets segmented by skill a, and receive job offers at rate λ_a^u. Employed workers enjoy flow utility equal to their wage w and also receive job offers at a rate $\lambda_a^e = s_a^e \lambda_a^u$, with relative on-the-job search intensity satisfying $s_{a_L}^e = 0 < s_{a_H}^e \leq 1$. A job offer entails a wage w drawn from the endogenous offer distribution $F_a(w)$. Jobs end exogenously at a rate δ_a.

Value Functions. The value of an employed worker of skill a in a job with wage w is

$$\rho S_a(w) = w + \lambda_a^e \int\limits_{w' > w} \left[S_a(w') - S_a(w) \right] dF_a(w') + \delta_a \left[W_a - S_a(w) \right], \quad \forall a. \tag{1}$$

The value of an unemployed worker of skill a is

$$\rho W_a = b_a + \lambda_a^u \int\limits_{w'} \max\left\{S_a\left(w'\right) - W_a, 0\right\}\, dF_a\left(w'\right), \quad \forall a. \tag{2}$$

Policy Functions. Employed workers accept any higher wage. Unemployed workers have a reservation wage ϕ_a, which we assume is low enough so that all firms hire both skill types.

2.2 Firms

A mass E of firms differ in their productivity $p > 0$ and credit limit $\xi > 0$, where (p, ξ) follows a continuous cumulative distribution function $\Gamma(\cdot)$.

Wages and Job Vacancies. Firms post, for each worker skill a, a market-specific wage w_a and vacancies v_a subject to strictly convex increasing recruiting costs $c_a(v_a)$: $c_a'(\cdot) > 0, c_a''(\cdot) > 0$.

Production. A firm with productivity p_j employing $\{l_a\}_{a\in\{a_L,a_H\}}$ workers of each skill level produces output according to the linear production function $y(p_j, \{l_a\}_{a\in\{a_L,a_H\}}) = p_j \sum_{a\in\{a_L,a_H\}} a l_a$.

Credit Constraint. Before production occurs, firms take up debt $D \geq 0$ to finance their working capital, consisting of their wage bill $\sum_a w_a l_a(w_a, v_a)$ and recruiting costs $\sum_a c_a\left(v_a\right)$. Given the interest rate $r > 0$, firms face idiosyncratic credit limits, net of any cash reserves and self-financing, given by $rD \leq \xi_j$.

Value Function. The value of a firm of type (p_j, ξ_j) is the net present value of revenues minus the wage bill minus recruiting costs minus the cost of servicing debt, which can be written as

$$r\Pi(p_j, \xi_j) = \max_{\{w_a, v_a\}_a} \left\{ \sum_a \left[\left(p_j a - (1+r) w_a\right) l_a(w_a, v_a) - (1+r) c_a(v_a) \right] \right\} \tag{3}$$

$$\text{s.t.} \quad r \sum_a \left[w_a l_a(w_a, v_a) + c_a\left(v_a\right) \right] \leq \xi_j. \tag{4}$$

Firm Ranks. As a result of firms' optimization problem in equations (3)–(4), each firm j is assigned a utility rank $r(j) \in [0, 1]$, with higher values of r representing higher worker utility levels.

2.3 Matching and Firm Sizes

A Cobb-Douglas matching function with constant returns to scale combines the effective job searchers $U_a = \mu_a \left[u_a + s_a^e(1-u_a)\right]$ with the aggregate vacancies $V_a = E \int_j v_a(j)\, d\Gamma(j)$ to produce, for each a, matches $m_a = \chi_a V_a^\alpha U_a^{1-\alpha}$ with matching efficiency $\chi_a > 0$ and elasticity $\alpha \in (0,1)$.

2.4 Equilibrium Pay and Employment Decisions under Credit Constraints

We define a *stationary equilibrium* of the economy in Appendix B.1. A firm's optimal wage and vacancy policies depend on both its productivity and its credit constraint, as characterized by the following first-order conditions (FOCs):

$$[\partial w_a]: \quad p_j a \frac{\partial l_a(w_a, v_a)}{\partial w_a} - \left(1 + (1+\psi_j) r\right) \left[l_a(w_a, v_a) + w_a \frac{\partial l_a(w_a, v_a)}{\partial w_a} \right] = 0, \quad \forall a, \tag{5}$$

$$[\partial v_a]: \quad p_j a \frac{\partial l_a(w_a, v_a)}{\partial v_a} - \left(1 + (1+\psi_j) r\right) \left[w_a \frac{\partial l_a(w_a, v_a)}{\partial v_a} + \frac{\partial c_a(v_a)}{\partial v_a} \right] = 0, \quad \forall a, \tag{6}$$

where $\psi_j \geq 0$ is the Lagrange multiplier on firm j's credit constraint (4). For unconstrained firms, $\psi_j = 0$, while for constrained firms, $\psi_j > 0$.[2]

The FOCs in equations (5) and (6) are identical to those of a firm with *effective productivity*

$$\tilde{p}_j = p_j \frac{1+r}{1+(1+\psi_j)r}. \tag{7}$$

Note that $\tilde{p}_j = p_j$ for unconstrained firms with $\psi_j = 0$, while $\tilde{p}_j < p_j$ for credit constrained firms with $\psi_j > 0$. For a given level of physical productivity p_j, firms facing a tighter credit limit—as reflected in a higher Lagrange multiplier ψ_j on the credit constraint in equation (4)—have lower effective productivity due to their higher shadow cost of resources.

2.5 The Effect of Credit Supply on the Distribution of Wages and Employment

The following proposition provides comparative statics for firms' wage and employment policies with respect to changes in idiosyncratic credit supply.

Proposition 1 (Comparative statics of wages and employment with respect to credit supply). *For any firm j, a decrease in their idiosyncratic credit limit ξ_j leads to*

[2] From here on, we ignore the knife-edge case in which the credit constraint (4) holds with equality but $\psi_j = 0$ for some firm j. Thus, we assume that (4) holds with strict inequality whenever $\psi_j = 0$.

(i) lower firm-level average wages for the same workers at firm j,

(ii) lower firm-level employment at firm j,

(iii) lower within-firm wage inequality, in the sense of a relatively greater reduction in wages among initially higher-paid workers at firm j, and

(iv) lower between-firm wage inequality, measured by a lower distance between the average wage at firm j and that at firm j' with rank $r(j') = 0$.

The comparative statics for any firm j with a binding credit constraint ($\psi_j > 0$) involve

- *strictly lower wage outcomes at interior firm ranks ($r(j) > 0$), as in (i), (iii), and (iv) above, and*
- *strictly lower employment, as in (ii) above.*

Any firm j with a strictly slack credit constraint ($\psi_j = 0$) experiences no changes in wages or employment.

Proof. See Appendix B.2. □

Intuitively, Proposition 1 states that a tighter credit constraint lowers a firm's effective productivity. This leads to a reduction in the relative wages of high-skill workers, whose wages are flexible downward. In the presence of a binding wage floor for low-skill workers, any adjustment for them must occur in the form of employment losses. Similar insights would obtain in a setting with Nash-bargained wages subject to a binding minimum wage (Flinn, 2006). More broadly, our analysis aligns well with recent papers studying nominal wage rigidity (e.g., Guerrieri et al., 2022; Hazell and Taska, 2025; Blanco et al., 2026; Rodríguez-Clare et al., 2026).

Our simple theory can accommodate firms' self-financing of working capital (Moll, 2014), which we can view as an additional positive term on the right-hand side of the firms' credit constraint in equation (4). Such self-financing would reduce the bindingness of the firm's credit constraint, with strictly self-financed firms being locally unaffected by their nominal credit limit.

The equilibrium model presented above has sharp predictions for the effects of credit supply on the distribution of wages and employment within and between firms in this environment. The timing and magnitude of the predicted effects of credit supply on the distribution of wages and employment are ultimately an empirical question.[3] Therefore, we test these predictions using an identified credit supply shock in the data.

[3]Our comparative statics results pertain to steady states and are silent on the speed of the transition. Real wages may either adjust immediately through nominal wage cuts or adjust slowly over time by staying constant in nominal terms in the wake of inflation. Analogously, employment may either adjust immediately through the firing of existing workers or adjust slowly over time as new hires are reduced following worker separations.

3 Empirical Strategy

Based on the theoretical predictions from Section 2, we now set out to empirically identify the effects of credit supply on the distribution of wages and employment within and between firms.

3.1 Identifying Credit Supply

To assess the distributional effects of credit in the labor market, the ideal experiment would involve manipulating the credit supply to a known subset of firms in a "macroeconomic laboratory." Absent such an experiment, we exploit quasi-natural variation in firm-level credit supply. Specifically, we study the heterogeneous transmission of monetary policy to bank lending following the historically notable implementation of negative deposit facility rates in the euro area.

The deposit facility rate is the rate at which banks may make overnight deposits with the Eurosystem. It is one of three main policy rates set by the Governing Council of the ECB. Through its transmission to banks' funding costs, the deposit facility rate is a key determinant of banks' lending activity. In June 2014, for the first time in the history of the euro, the deposit facility rate was set to negative. This was an important event in the history of the European Economic and Monetary Union and has sparked a lot of attention among academics and policy makers alike. There is broad consensus that this unprecedented step came as a surprise to financial institutions and firms, as evidenced by the sharp market reaction (Hirst, 2014). Since then, the deposit facility rate remained negative for over eight years, reaching beyond the time period studied in this paper.

In *non-negative* territory, lower interest rates decrease banks' funding costs independent of their funding structure, which induces them to increase lending to firms, in line with classical monetary theory (Gertler and Kiyotaki, 2010). Our identification strategy exploits cross-sectional variation in banks' exposure to the important episode of *negative* rates. Negative rates have been shown to affect bank lending through two channels. The first channel is due to the imperfect pass-through of negative monetary policy rates to deposit rates. With the exception of some corporate deposit accounts (Heider et al., 2021; Altavilla et al., 2022), banks have been reluctant—or unable—to charge negative rates to their retail depositors, as opposed to rates on other types of short-term debt, e.g., interbank funding.[4] As a result, deposit spreads are squeezed, so banks with greater reliance on deposit funding experience increased funding costs (Eggertsson et al., 2024; Ulate, 2021). The second channel is due to the effect of negative rates on banks' net worth or equity

[4]In this sense, our work is related to Drechsler et al. (2017) who study market power in deposit markets.

value, which falls in line with the decline in profitability of banks that are more reliant on deposit funding (Ampudia and Van den Heuvel, 2022). The decline in their net worth leads banks to reach for yield by channeling credit away from existing and/or safe borrowers toward new, and potentially riskier, projects (Heider et al., 2019).

Both of these channels lead to a relative reduction in credit supply to *existing* borrowers from banks that are more exposed to negative rates because of their deposit reliance. Therefore, to the extent that banking relationships are sticky (Chodorow-Reich, 2014; Darmouni, 2020), firms in preexisting relationships with more deposit-reliant banks should experience a relative contraction in credit supply.

This allows us to identify variation in firms' credit access using information on their relationship banks, akin to Huber (2018). For this purpose, we combine data on firms' self-reported banking relationships with bank-level balance sheet information. Specifically, let $\textit{Deposit ratio}_j$ denote the average deposit ratio, that is the ratio of deposits to assets, across all euro-area banks that firm j reports to be in a banking relationship with during the preperiod from 2010 to 2013. Let $\textit{After}(2014)_t$ denote a dummy variable for the years 2014–2017. In a difference-in-differences setting, we define as our credit supply shock

$$\textit{Deposit ratio}_j \times \textit{After}(2014)_t,$$

which captures the idea that firms in relationships with euro-area banks that rely more on deposit funding experience a negative credit supply shock after June 2014.

3.2 Measuring the Effects of Credit Supply Within and Between Firms

To test for distributional effects of a monetary policy-induced credit supply shock on workers and firms, we consider a panel of workers indexed by i across German firms j and years t. We want to track wages and employment of workers in the years around the firm-level credit supply shock.

Mean Wage Effects. While the credit supply shock is at the firm-year level, we study individual wages at the level of the worker-firm-year ijt, which simplifies to the worker-year level as we keep only the main job j in a given worker-year it. We estimate the following specification:

$$y_{ijt} = \beta Deposit\ ratio_j \times After(2014)_t + \theta_{ij} + \zeta_{k(j)t} + \varepsilon_{ijt}, \tag{8}$$

where y_{ijt} is the natural logarithm of the wage for worker i employed at firm j in year t, *Deposit ratio*$_j$ is the mean deposits-to-assets ratio, measured in 2013, across all (typically German) banks that firm j reports to be in a banking relationship with from 2010 to 2013, *After*(2014)$_t$ is a dummy variable for the years 2014–2017, and θ_{ij} and $\zeta_{k(j)t}$ denote, respectively, worker-firm and industry-year, state-year, or industry-state-year fixed effects (corresponding to the state that firm j is located in and/or its one-digit industry code). We cluster standard errors at the firm level since we exploit variation in firm-level exposure to a bank-specific lending shock.

The coefficient of interest in equation (8) is β, which measures the average wage response to variation in credit supply. The inclusion of worker-firm match fixed effects means that we identify this coefficient off the effect on workers that were already employed at the same firm prior to the credit supply shock. By first excluding and then including controls for worker-firm match heterogeneity, our estimates shed light on the different margins of labor market adjustments, specifically changes in worker composition through employment transitions. By additionally controlling for industry-year and state-year fixed effects—or, more granularly, industry-state-year fixed effects—we absorb time-varying factors at the industry level as well as aggregate trends and regional business cycle fluctuations that equally affect all workers in a given state each year.

Within-Firm Heterogeneity. To estimate within-firm heterogeneity in the effect of credit, we interact the credit supply shock with a worker's pay rank within the firm:

$$
\begin{aligned}
y_{ijt} = {} & \beta_1 Deposit\ ratio_j \times After(2014)_t \times Bottom\ 33\%\ within\ firm_i \\
& + \beta_2 Deposit\ ratio_j \times After(2014)_t \times Middle\ 33\%\ within\ firm_i \\
& + \beta_3 Deposit\ ratio_j \times Bottom\ 33\%\ within\ firm_i + \beta_4 Deposit\ ratio_j \times Middle\ 33\%\ within\ firm_i \\
& + \beta_5 After(2014)_t \times Bottom\ 33\%\ within\ firm_i + \beta_6 After(2014)_t \times Middle\ 33\%\ within\ firm_i \\
& + \theta_{ij} + \eta_{jt} + \varepsilon_{ijt},
\end{aligned} \tag{9}
$$

where y_{ijt} is either the natural logarithm of the wage or an indicator for unemployment next year for worker i employed at firm j in year t, *Bottom* 33% *within firm*$_i$ (*Middle* 33% *within firm*$_i$) is an indicator variable for whether worker i's wage is in the bottom (middle) tercile of the wage distribution of the firm where worker i was employed in the last available year during the preperiod from 2010 to 2013, and θ_{ij} and η_{jt} denote worker-firm and firm-year fixed effects, respectively.[5]

The coefficients of interest in equation (9) are β_1 and β_2. They capture the extent to which firms' exposure to negative rates differentially affects workers within the bottom tercile and middle tercile of the wage distribution relative to workers in the top tercile, while β_5 and β_6 capture any potential mean reversion in the wage ranking of workers that may occur over time. In addition to worker-firm fixed effects, we also add a set of firm-year fixed effects that control for time-varying unobserved heterogeneity at the firm level. This powerful control absorbs any aggregate trends and idiosyncratic firm innovations that equally affect all workers within a firm.

Between-Firm Heterogeneity. To estimate between-firm heterogeneity in the effect of credit, we interact the credit supply shock with a firm's mean pay rank:

$$\begin{aligned} y_{ijt} = &\ \beta_1 \mathit{Deposit\ ratio}_j \times \mathit{After}(2014)_t \times \mathit{Firm\ pay\ rank}_j \\ &+ \beta_2 \mathit{Deposit\ ratio}_j \times \mathit{After}(2014)_t + \beta_3 \mathit{After}(2014)_t \times \mathit{Firm\ pay\ rank}_j \\ &+ \theta_{ij} + \zeta_{k(j)t} + \varepsilon_{ijt}, \end{aligned} \tag{10}$$

where $\mathit{Firm\ pay\ rank}_j$ is firm j's mean wage rank among all firms in 2013, with 0 being the lowest rank and 1 being the highest rank, and θ_{ij} and $\zeta_{k(j)t}$ denote, respectively, worker-firm and industry-year, state-year, or industry-state-year fixed effects (corresponding to the state that firm j is located in and/or its one-digit industry code). The coefficient of interest in equation (10) is β_1, which captures the extent to which initially higher-paying firms respond differentially to the credit supply shock due to the introduction of negative interest rates.

4 Data

4.1 Data Sources

For the first time, this paper spans the complete credit chain in Germany: starting from banks' balance sheet exposure to monetary policy, to bank-firm relationships and loan transactions, to firm financials, and finally to worker-level outcomes. Building this data infrastructure requires us to combine microdata from several different data providers, including private and restricted public data sources.

[5] Of course, the (nuisance) parameters β_3 and β_4 will not be separately identified from worker-firm match fixed effects, when included.

Employment Histories (IAB). At the heart of our analysis lie the administrative linked employer-employee data hosted at Germany's Institute for Employment Research (IAB). These restricted public data contain employment histories based on social security records for essentially the universe of workers and establishments in Germany, excluding civil servants and the self-employed. The linked employer-employee nature of the data means that we observe all workers within each establishment and that we can track workers across establishments and over time.

Firm Financials (Orbis). We use firm financial data comprising balance sheet information for private and public firms. These private data can be purchased from Bureau van Dijk (BvD) and are distributed as part of the Orbis Historical data product. The merge between the IAB's linked employer-employee data and the firm financials data forms part of the IAB-internal data product *Orbis-ADIAB* (Schild, 2016; Antoni et al., 2018). We extend the preexisting record linkages beyond 2013 to cover our complete sample period from 2010–2017. This merge allows us to link individual establishments in the IAB's data at the firm level.

Board Compensation (BoardEx). We supplement the IAB data's worker earnings records with small-sample information on compensation—including salary and bonus components—of board members at companies listed on the German stock market index (DAX) from 2010 to 2016. We source this information from *BoardEx*, which we access via Wharton Research Data Services (WRDS) and merge with the other datasets via consistent BvD identifiers.

Bank-Firm Relationships (Creditreform). To capture German firms' bank credit relationships, we primarily use their self-reported bank relationships in *Creditreform* (Huber, 2018). These data identify firms' principal and other bank affiliations, which we merge using BvD identifiers.

Loan Transactions (DealScan). As an additional source of information on firms' bank credit relationships, we use data from Thomson Reuters *DealScan* on (typically large, public) firms' transactions in the syndicated loans market based on public filings, company statements, and media reports. We hand-match data from DealScan to firms in the other datasets using a combination of firm name, industry, and address, similar to Acharya et al. (2019) and Heider et al. (2019). To conform as closely as possible with the Orbis-ADIAB sample that we use for identifying heterogeneous worker effects, we limit our analysis to German firms in Orbis with data coverage throughout 2010–2017 and at least ten employees. Furthermore, we drop a small number of firms that, according to the Orbis data, have ratios of the sum of long-term debt and short-term loans over assets of 0.05 and less, as those firms are unlikely to be affected by financial shocks.

Bank Balance Sheets (SNL Financial). To measure banks' exposure to negative rates, we take balance sheet data from *SNL Financial* (now S&P Global Market Intelligence), a financial news and data services provider, for all banks that appear in the other datasets.

4.2 Variables and Sample Selection

The main variables of interest for our analysis are the deposit ratios of firms' relationship banks as well as workers' wages and employment status. We measure a firm's exposure to negative rates through the mean ratio of deposits to assets across all euro-area banks (typically in Germany) that a firm reports to be in a banking relationship with during the preperiod from 2010 to 2013.[6] Two-thirds of all firms indicate that they are in a relationship with a single bank. Only 9 percent of all German firms report to be in a relationship with more than two banks.

Wages are defined as the mean (log) daily earnings of full-time employees as reported in the IAB's linked employer-employee data.[7] Since these data are based on social security records and reporting is subject to statutory contribution limits, earnings are winsorized around the 90$^{\text{th}}$ percentile of the population. Finally, unemployment is defined as a worker leaving our sample of employment records in a given year, excluding temporary leaves and recalls within one year.

We use data from 2010 to 2017 around the introduction of negative monetary policy rates in 2014. Exploiting the matched employer-employee dimension of the merged data, we build a panel of workers indexed by i across firms j and years t. Within a given worker-year it, we retain the main job j, defined as the highest-paid full-time job held by worker i in year t. We then limit the sample to firms with information on bank relationships from Creditreform, which we use to construct the credit supply shock exposure variable, $\textit{Deposit ratio}_j$.

Our final sample covers approximately 36 percent of all full-time workers in Germany, which constitutes a large subset of the German labor force. Table 1 presents summary statistics for this sample and key variables from the merged dataset at all relevant levels of our analysis. In Panel A, we start out with German firms' activities in the syndicated loans market, and build a panel

[6]We construct the unweighted mean ratio of deposits to assets across all euro-area banks since the Creditreform data do not quantify the intensities of bank-firm relationships.

[7]We separately study the part-time employment share as an outcome in our firm-level analysis.

Table 1: Summary Statistics

	Mean	Std. dev.	P5	P50	P95	No. of observations
Panel A: Firm-bank-half-year level						
Deposit ratio$_k$	0.370	0.117	0.241	0.337	0.552	16,266
Deposit ratio$_j$	0.489	0.127	0.257	0.498	0.756	21,274
Any loan share	0.141	0.348	0	0	1	22,016
Total loan amount (bn euros)	0.069	0.194	0.008	0.035	0.152	3,068
Panel B: Worker-year level						
Annualized wage (euros)	37,278	18,553	8,238	35,238	70,941	72,161,941
Unemployed next year	0.147	0.354	0	0	1	67,981,002
Panel C: Firm-year level						
Deposit ratio	0.654	0.153	0.257	0.693	0.837	2,810,558
Wage P90/P10	4.374	216.636	1.000	2.093	9.529	2,779,570
Wage P90/P10 at public firms	2.589	3.267	1.171	2.006	4.352	1,340
Board total P50/Wage P5	189.077	852.580	28.666	60.360	298.299	266
Board salary P50/Wage P5	64.158	294.934	12.883	25.318	85.895	266
Board bonus P50/Wage P5	126.299	580.321	11.868	35.884	213.693	263
No. of employees	42.841	528.685	1	11	133	2,744,131
No. of nonmanagerial employees	40.185	483.374	1	10	125	2,744,131
No. of part-time employees	14.809	188.805	0	3	43	2,744,131

Notes: The summary statistics in Panel A refer to the firm-bank-half-year level for syndicated loans granted to German firms in DealScan, and correspond to the respective descriptions and samples in Tables 3 and 4. The total loan amount is conditional on having any loan. The summary statistics in Panel B refer to the dependent variables at the worker-year level, and correspond to the respective descriptions in Tables 5, 6, and 9. The variables in Panel C correspond to the respective descriptions in Tables 7 and 8.

at the firm-bank-half-year level for syndicated loans granted to German firms in DealScan. Panel B shows summary statistics at the worker-year level based on the merged data. Altogether, our sample covers over 72 million worker-year observations, or an average of 9 million observations per year. Finally, Panel C summarizes key variables at the firm-year level based on the merged data. The average deposit ratio is around 0.65. The mean P90/P10 wage percentile ratio is around 4.4 for all firms and around 2.6 for the subset of publicly listed firms. Using small-sample evidence on compensation of board members at public firms, we find a large pay gap between board members and regular workers. While the average firm in our sample has 42.8 full-time employees, the firm size distribution is positively skewed and fat-tailed. The mean number of nonmanagerial employees who work full-time is 40.2, while the mean number of part-time employees is 14.8.

Table 2 presents firm-level summary statistics separately for firms in relationships with banks in the top and bottom quartiles of the distribution of deposit ratios. Firms in relationships with high-deposit banks, which have greater exposure to negative rates, and firms in relationships with low-deposit banks are similar along several observable characteristics (e.g., composition of work-

Table 2: Summary Statistics for Firms with High versus Low Exposure to Negative Rates

	Mean	Std. dev.	P5	P50	P95	No. of firms
Panel A: German firms related to banks in the highest quartile of the deposit ratio distribution						
No. of all employees	23.621	100.783	1	9	76	89,866
Average annualized wage (euros)	27,473	11,306	11,614	26,065	48,452	89,866
Proportion female	0.251	0.318	0.000	0.111	1	89,866
Proportion foreigner	0.070	0.181	0.000	0.000	0.477	89,866
Proportion university	0.109	0.235	0.000	0.000	0.667	89,866
Leverage	0.201	0.244	0.000	0.098	0.730	34,224
ROA	0.113	0.127	0.005	0.071	0.368	8,191
ROA volatility	0.062	0.064	0.006	0.041	0.188	4,379
Cash/Assets	0.192	0.207	0.001	0.117	0.635	59,711
Investment/Assets	0.070	0.101	0.000	0.033	0.272	25,585
Panel B: German firms related to banks in the lowest quartile of the deposit ratio distribution						
No. of all employees	70.515	917.632	1	11	216	88,077
Average annualized wage (euros)	33,050	14,015	12,505	31,415	58,492	88,077
Proportion female	0.297	0.316	0.000	0.200	1	88,077
Proportion foreigner	0.080	0.184	0.000	0.000	0.500	88,077
Proportion university	0.190	0.285	0.000	0.036	1	88,077
Leverage	0.158	0.228	0.000	0.031	0.675	37,468
ROA	0.125	0.131	0.007	0.085	0.388	13,557
ROA volatility	0.071	0.066	0.009	0.052	0.200	9,636
Cash/Assets	0.194	0.214	0.001	0.113	0.650	59,007
Investment/Assets	0.065	0.105	0.000	0.025	0.271	25,173

Notes: This table shows firm-level summary statistics for 2013, the last year before the introduction of negative rates, for German corporations in the top (Panel A) and bottom (Panel B) quartile of the distribution of $\textit{Deposit ratio}_j$, which is the average deposit ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013.

force, return on assets, volatility of return on assets, cash/assets, and investment/assets), but there are also some notable differences between the two groups. The average firm in relationships with high-deposit banks has 23.6 employees, compared to 70.5 employees for firms in relationships with low-deposit banks. Mean pay at firms in relationships with high-deposit banks is 27,473 euros, slightly less than at firms in relationships with low-deposit banks, which is 33,050 euros. These differences are relatively smaller, however, when comparing median values.

It is important to note that baseline differences between firms in relationships with high- versus low-deposit banks are not a threat to our identification. By including firm fixed effects in all worker-level regression specifications, we control for permanent (unobserved) firm heterogeneity. We also explicitly address nonrandom matching between firms and banks by including bank-firm match fixed effects in all credit-related specifications. In our analysis of within-firm inequality, we also include firm-year fixed effects, which account for both permanent and time-varying (unobserved) employer differences, thereby subsuming any firm-specific trends.

5 Empirical Results

5.1 Effects of Negative Interest Rates on Credit Supply

We start by estimating the extent to which German firms in relationships with high-deposit rather than low-deposit banks experience a relative reduction in credit supply following the introduction of negative interest rates in June 2014. For this purpose, we use transaction-level data on German firms' syndicated loans based on DealScan. While only a subset of German firms in our sample are active in the syndicated loans market, the granularity of these data allows us to control for a rich set of codeterminants of firms' access to credit. We focus on banks that act as lead arrangers in the syndication process. Lead arrangers are members of a syndicate typically responsible for traditional bank duties, including due diligence, payment management, and monitoring of the loan (Ivashina and Scharfstein, 2010).

Based on all lead banks' shares of completed syndicated loans of German corporations between January 1, 2010, and December 31, 2017, we extend the sample to a balanced panel of borrowers j and banks k over time t at semi-annual frequency. Following Heider et al. (2019) and Eggertsson et al. (2024), we use bank k's deposit ratio as the exposure variable and limit the sample to lead banks in negative-rate currency areas from which firm j borrowed anytime in the preperiod. We then estimate the following difference-in-differences specification at the firm-bank-time level jkt, where time therefore refers to the semi-annual level:

$$y_{jkt} = \beta \mathit{Deposit\ ratio}_k \times \mathit{After}(06/2014)_t + \kappa_{jk} + \lambda_{jt} + \varepsilon_{jkt}, \tag{11}$$

where y_{jkt} is an outcome associated with lending by bank k to firm j at time t, $\mathit{Deposit\ ratio}_k$ is bank k's deposits-to-assets ratio, measured in 2013, $\mathit{After}(06/2014)_t$ is an indicator for whether the date falls on or after June 2014, and κ_{jk} and λ_{jt} denote firm-bank and firm-time fixed effects, respectively. We cluster standard errors at the bank level.

Our interest lies in estimates of the coefficient β in equation (11). In the presence of firm-time fixed effects that absorb time-varying unobserved heterogeneity at the firm level, including loan demand (Khwaja and Mian, 2008), β captures changes in bank k's credit supply to its existing borrowers as a result of greater exposure to negative rates.

Table 3 presents the results of estimating (11). In columns 1–3, the dependent variable is an indicator for any non-zero share of firm j's syndicated loans held by bank k at date t. Column

Table 3: Impact of Negative Policy Rates on German Firms' Preexisting Banking Relationships

	Any loan share $\in \{0,1\}$			$\ln(1 + \text{total loan volume})$		
Sample	2010–2017		2013–2015	2010–2017		2013–2015
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio$_k$ × After(06/2014)	-0.085*	-0.122**	-0.158**	-1.475*	-2.099*	-2.630*
	(0.048)	(0.061)	(0.076)	(0.852)	(1.108)	(1.382)
Deposit ratio$_k$ × After(07/2012)		0.066			1.113	
		(0.089)			(1.611)	
Bank-firm FE	Y	Y	Y	Y	Y	Y
Firm-time FE	Y	Y	Y	Y	Y	Y
N	15,554	15,554	6,508	15,554	15,554	6,508

Notes: Based on all lead banks' shares of completed syndicated loans of German corporations j anytime from January 2010 to June 2014, the sample is extended so as to represent a balanced panel of all borrower-bank pairs at the semi-annual frequency from 2010 to 2017. Time, therefore, refers to the semi-annual level. Furthermore, the sample is limited to banks in currency areas with negative monetary policy rates that lend to German firms at any point in the preperiod from January 2010 to June 2014. In columns 3 and 6, the sample runs from the first half of 2013 to the second half of 2015. All singletons are dropped from the total number of observations N. In columns 1–3, the dependent variable is an indicator for any nonzero share of firm j's loans held by bank k in t. In columns 4–6, the dependent variable is the natural logarithm of one plus the total loan volume granted to firm j by bank k in t. *Deposit ratio*$_k$ $\in [0,1]$ is bank k's ratio of deposits over total assets in 2013. *After*(06/2014)$_t$ is a dummy variable for the period from June 2014 onwards. *After*(07/2012)$_t$ is a dummy variable for the period from July 2012 onwards. Energy and financial services borrower firms are dropped. Robust standard errors (clustered at the bank level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

1 shows that high-deposit banks experience a relative reduction in their credit supply after the introduction of negative rates. A one standard deviation increase in bank-level deposit ratios implies a lower relative likelihood of granting any loans through syndication by $0.117 \times 0.085 = 1.0$ percentage points.

Our identification rests on the assumption that banks' funding structure (i.e., deposit reliance) matters for their credit supply when monetary policy rates are disconnected from deposit rates. Ulate (2021) argues that this is the case for negative interest rates, as the returns on deposits are constrained to being nonnegative. During our sample period, banks are indeed reluctant or unable to pass negative rates on to most of their (typically household) depositors (Heider et al., 2019; Eggertsson et al., 2024; Heider et al., 2021; Balloch et al., 2022; Kwan et al., 2025).

To corroborate this assumption, in column 2 we interact the deposit ratio with an indicator for the period starting in July 2012, which is when the ECB reduced the deposit facility rate from 0.25 percent to 0 percent, the lowest nonnegative interest rate. We find that high-deposit and low-deposit banks do not respond differently to this cut in positive rates.[8] Instead, we continue to find

[8] Our results are robust to using the mean deposit ratio over the preperiod before 2012, likely reflecting the stability of firms' relationship banks in our sample.

Table 4: Impact of Negative Policy Rates on Lending to German Firms

	Any loan share $\in \{0,1\}$		$\ln(1 + \text{total loan volume})$	
	(1)	(2)	(3)	(4)
Deposit ratio$_j$ $\times$ After(06/2014)	-0.084***	-0.101***	-1.254**	-1.559***
	(0.030)	(0.030)	(0.511)	(0.514)
Bank-firm FE	Y	Y	Y	Y
Time FE	Y	N	Y	N
Bank-time FE	N	Y	N	Y
N	21,274	21,158	21,274	21,158

Notes: Based on all lead banks' shares of completed syndicated loans of German corporations j anytime from January 2010 to December 2017, the sample is extended so as to represent a balanced panel of all borrower-bank pairs at the semi-annual frequency. Time, therefore, refers to the semi-annual level. All singletons are dropped from the total number of observations N. In the columns 1–2, the dependent variable is an indicator for any nonzero share of firm j's loans held by bank k in t. In the columns 3–4, the dependent variable is the natural logarithm of one plus the total loan volume granted to firm j by bank k in t. *Deposit ratio*$_j$ $\in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. *After*(06/2014)$_t$ is a dummy variable for the period from June 2014 onwards. Energy and financial services borrower firms are dropped. Robust standard errors (clustered at the bank level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

that high-deposit banks reduce credit supply relative to low-deposit banks after the introduction of negative policy rates in June 2014.

In column 3, we estimate the same specification as in column 1, using a shorter three-year time window centered around the introduction of negative rates in June 2014. This reduces the likelihood that other events, including monetary policy decisions, could interfere with our identification. The resulting estimate becomes somewhat larger and is significant at the 5 percent level.

All of these results also hold when we replace the dependent variable with the natural logarithm of 1 plus the total loan volume granted to firm j by bank k in t, as shown in columns 4–6. For each syndicated loan, we use DealScan data on each lead bank's share to compute the total loan amount granted to a firm in a given time period.[9]

Our results indicate that high-deposit banks reduce their credit supply relative to low-deposit banks in response to the introduction of negative rates, consistent with both bank-firm-level and bank-level evidence in Heider et al. (2019) and Eggertsson et al. (2024). In the next step, we establish that firms in existing relationships with affected high-deposit banks not only receive less credit but also cannot fully offset the credit contraction by switching to other banks. For this purpose, we extend our balanced panel to include all lending relationships, including those with banks

[9] Whenever available, we use loan shares as reported in DealScan. Otherwise, as in Chodorow-Reich (2014), we set the total loan share held by lead arrangers in the syndicate equal to the sample mean and divide it equally among all lead arrangers.

outside the euro area, from 2010 to 2017. Furthermore, we replace *Deposit* $ratio_k$ in (11) by *Deposit* $ratio_j$, which captures a firm's exposure to the introduction of negative rates and is computed as the mean deposit ratio in 2013 of its relationship banks in the preperiod from 2010 to 2013. As a consequence, we can no longer control for firm-time fixed effects, as they would match the level of our identifying variation.

The results from analyzing the extended set of lending relationships are shown in Table 4. We find a significant relative reduction in credit for firms in relationships with high-deposit banks, regardless of the credit-granting bank, after the introduction of negative rates in 2014. In column 1, we include only bank-firm and time fixed effects and find that a one standard deviation increase in *Deposit* $ratio_j$ is associated with a $0.127 \times 0.084 = 1.1$ percentage point lower likelihood of obtaining any loan, very close to our previous estimate surrounding Table 3.

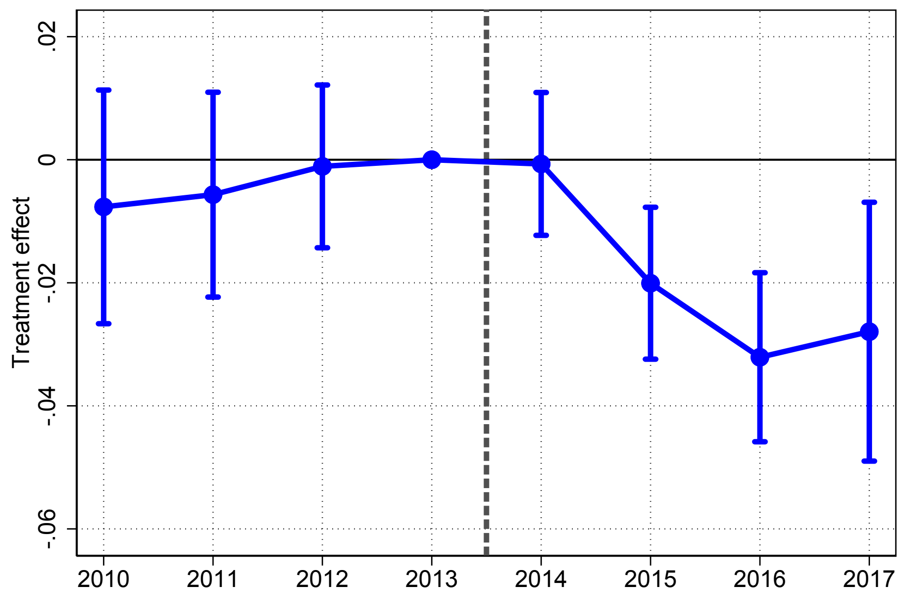
This estimate becomes even larger in column 2, which adds bank-time fixed effects to control for bank-wide shocks, such as regulatory changes, that affect bank lending across all clients. Crucially, the coefficient of interest, β, is now estimated based on firms in relationships with the same bank in a given year. Among these firms, β captures the effect of differential exposure to high- versus low-deposit banks in the preperiod on current lending by new bank relationships. As before, all of these results hold when we replace the dependent variable with the actual loan amounts granted by lead banks through syndication, as shown in columns 3 and 4 of Table 4.

These findings imply that, on average, firms in relationships with high-deposit banks receive less credit from *any* bank, including those outside the euro area, following the introduction of negative rates. As our matched employer-employee data comprise many small, bank-dependent firms without access to capital markets, these firms cannot readily compensate for a loss of bank credit access by switching to other sources of debt financing either. To bolster this assumption at least among German firms with balance-sheet data, we use the firm-level panel in Orbis and estimate the following firm-year-level regression:

$$y_{jt} = \sum_{\tau=2010}^{2017} \beta_\tau \textit{Deposit ratio}_j \times \mathbf{1}[t = \tau] + \psi_j + \delta_t + \varepsilon_{jt}, \tag{12}$$

where y_{jt} is the dependent variable of interest at the firm-year level, where t represents the respective year-end, *Deposit* $ratio_j$ is the mean deposits-to-assets ratio, measured in 2013 across all (typically German) banks that firm j reports to be in a banking relationship with from 2010 to 2013, $\mathbf{1}[t = \tau]$ is a dummy variable for the year t being equal to τ, and ψ_j and δ_t denote firm and year

Figure 1: Impact of Negative Policy Rates on Firms' Leverage



Notes: This figure plots estimates of β_τ, alongside 95 percent confidence bands, over time (each year represents the respective year-end) based on the event study specification in (12), using as dependent variable firm j's leverage ratio, estimated on the sample of German firms in the administrative linked employer-employee data merged with Orbis from 2010 to 2017.

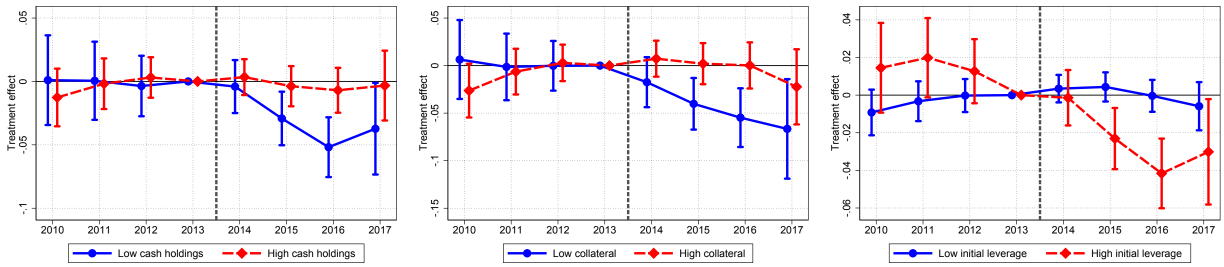
fixed effects, respectively. Standard errors are clustered at the firm level.

Figure 1 plots estimates of β_τ from equation (12) alongside 95 percent confidence intervals, relative to the year 2013 and using as dependent variable $Leverage_{jt}$, which we define as the ratio of the sum of long-term debt and short-term loans (in Orbis) to firm j's assets in year-end t. Firm leverage has been shown to be relevant for the transmission of other macroeconomic shocks (e.g., Giroud and Mueller, 2017) and is associated with both credit risk and labor compensation (Favilukis et al., 2020). The coefficient is statistically insignificantly different from zero for the preperiod from 2010–2013 and becomes negative and significant at the 5 percent level starting with the first full year of negative rates in 2015. The relatively greater deleveraging among more-affected firms in the data corresponds to a tightening of their idiosyncratic credit constraint, ξ_j, in our theoretical model in Section 2.

Figure 2 explores the heterogeneity underlying this mean effect on leverage across firm characteristics that economic theory predicts should govern exposure to credit supply shocks. Panel A splits firms by cash holdings: firms with below-median cash, which hold less internal liquidity to substitute for external finance, delever substantially more than cash-rich firms, whose response

Figure 2: Heterogeneity in the Impact of Negative Policy Rates on Firms' Leverage

(A) By Cash Holdings **(B) By Collateral** **(C) By Initial Leverage**



Notes: This figure plots estimates of β_τ, alongside 95 percent confidence bands, over time (each year represents the respective year-end) based on the event study specification in (12), using as dependent variable firm j's leverage ratio, estimated on the sample of German firms in the administrative linked employer-employee data merged with Orbis from 2010 to 2017. The specification is separately estimated by, and the results are shown for, subgroups below and above median cash holdings (Panel A), collateral defined as the capital intensity or capital per worker (Panel B), and initial leverage (Panel C).

is statistically indistinguishable from zero. Panel B splits firms by collateral, proxied by capital intensity or capital per worker: firms with less collateral, and hence tighter borrowing limits, exhibit a markedly larger contraction in leverage than collateral-rich firms. Panel C splits firms by preperiod leverage: highly leveraged firms—those closest to their debt-capacity limit and thus most vulnerable to a contraction in credit supply—reduce leverage significantly more than less indebted firms. Taken together, these patterns are consistent with negative policy rates operating through a bank credit-supply channel that bites hardest where economic theory guides us to look for its effects. This cross-sectional variation maps directly into heterogeneity in the firm-specific tightness of credit constraints net of cash holdings, $\tilde{\xi}_j$, in our model.

Supporting our interpretation of the introduction of negative rates as a credit supply shock, Appendix Figure A.1 shows that firms in relationships with high-deposit banks start accumulating significantly more cash following the introduction of negative rates. This is in line with an increased self-financing motive (Almeida et al., 2004). Appendix Figure A.2 also shows heterogeneity in the accumulation of cash, which is more emphasized, for instance, for firms with below-median cash that delever sharply, freeing up retained earnings for cash accumulation.

5.2 Effects on the Distribution of Wages and Employment

So far, we have established that firms in relationships with high-deposit banks receive relatively less credit after the introduction of negative monetary policy rates. Next, we estimate how this

Table 5: **Effects of Monetary Policy-Induced Credit Supply on Mean Wages**

	(1)	(2)	(3)	(4)
Deposit ratio × After(2014)	-0.020**	-0.077***	-0.045***	-0.044***
	(0.009)	(0.010)	(0.007)	(0.007)
Worker FE	Y	N	N	N
Firm FE	Y	N	N	N
Worker-firm FE	N	Y	Y	Y
Year FE	Y	Y	N	N
Industry-year FE	N	N	Y	N
State-year FE	N	N	Y	N
Industry-state-year FE	N	N	N	Y
N	70,169,491	67,763,431	67,745,191	67,745,191

Notes: The sample consists of full-time employees i at German corporations j in year t from 2010 to 2017. The dependent variable is the natural logarithm of the wage of individual i at firm j in year t. *Deposit ratio*$_j \in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. *After*(2014)$_t$ is a dummy variable for the years 2014–2017. Industry-year fixed effects are based on the one-digit industry code of firm j's industry. State-year fixed effects are based on the modal location (state) of firm j's establishments. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

credit supply shock affects the distribution of wages and employment across workers and firms.

Mean Wage Effects. We first consider the effects of credit supply on mean wages. Table 5 shows results from estimating variants of equation (8) using worker i's log wage at firm j as the dependent variable. When including worker, firm, and year fixed effects, we find that workers at more-exposed firms experience, on average, a relative reduction in wages (column 1). This effect becomes stronger when including worker-firm fixed effects (column 2), so the coefficient of interest, β, is estimated based on workers employed at the same firm both before and after 2014. A one standard deviation increase in firms' exposure, captured by *Deposit ratio*$_j$, translates into $0.153 \times 0.077 = 1.2$ percent lower wages. The estimate is fairly robust to the inclusion of industry-year and state-year fixed effects (in column 3), which control for time-varying unobserved heterogeneity across industries and states, or even more granular industry-state-year fixed effects (in column 4). These findings are in line with Part (i) of Proposition 1 of our theoretical model.

Within-Firm Heterogeneity. The estimated mean effect on wages may mask important heterogeneity across worker groups within firms. In addition, consistent with Part (ii) of Proposition 1, heterogeneous wage effects may be linked to heterogeneous unemployment risk. To investigate this, we estimate specification (9) of our empirical strategy, which adds an interaction term indicating a worker's position in the within-firm wage distribution and also uses workers' employment status the following year as a dependent variable.

Table 6 presents the results. We always include worker fixed effects to control for time-invariant heterogeneity at the worker level. In column 1, we include firm and year fixed effects; in column 2, we replace them with firm-year fixed effects, which control for time-varying heterogeneity at the firm level, e.g., firm-wide developments that may be correlated with firms' heterogeneous exposure to negative rates through their banking relationships. As such, they subsume industry-year, state-year, and industry-state-year fixed effects, which we included in Table 5. By controlling for firm-year fixed effects, we also address a potential weakness of our identification strategy, namely that confounding firm characteristics could affect firms' wage-setting and employment behavior around the introduction of negative rates, including firm-specific pretrends.

We find that individuals who used to earn wages in the bottom tercile of their respective firms' wage distributions see their wages grow more at more-exposed firms after the introduction of negative rates than do those in the top tercile (i.e., the omitted category). The negative middle-tercile coefficient in column 1 reflects between-firm variation in mean wages (cf. Table 5) rather than a within-firm rank effect. Once firm-year fixed effects in columns 2 and 3 absorb the firm-level mean wage response, the within-firm pattern emerges cleanly: the bottom tercile gains relative to the top, while the middle and top terciles are statistically indistinguishable. This reflects the decomposition predicted by Proposition 1, in which firm-level wages decline on average (Part (i)) and within-firm inequality compresses through differential effects at the top (Part (iii)).

After adding worker-firm fixed effects in column 3, a one standard deviation increase in firms' exposure as captured by *Deposit ratio*$_j$ translates into a $0.153 \times 0.050 = 0.8$ percent relative reduction in wages of workers in the top tercile versus those in the bottom tercile of the within-firm wage distribution. Since the coefficient of interest in the wage regression is now estimated for workers who remain with the same employer before and after the introduction of negative rates, these results are driven by wage effects on incumbents rather than on new hires.

In columns 4 to 6, we estimate specification (9) with the dependent variable replaced by an indicator for whether worker i is unemployed in year $t+1$. Across all three specifications, we find that workers outside the top tercile of the within-firm wage distribution face a higher risk of layoff following the negative credit supply shock. Quantitatively, in our preferred specification (column 6), the additional risk of leaving employment per standard deviation of exposure is $0.153 \times 0.028 =$

0.4 percentage points for middle-tercile workers and $0.153 \times 0.040 = 0.6$ percentage points for bottom-tercile workers, both relative to the top tercile. The inclusion of worker-firm match fixed effects implies that we identify the effect in column 6 based on workers who were either employed at the same firm before and after 2014 or left employment after 2014.

The empirical observation that wages are more rigid for lower-paid workers may partly reflect the introduction of a federal minimum wage in Germany in 2015. To the extent that workers near the bottom of the within-firm wage distribution find themselves at or near this threshold, their wages are downwardly rigid. Stronger downward wage rigidity of low-paid workers could also rationalize our finding that these workers are relatively more likely to become unemployed following the credit supply shock. This finding is consistent with the prediction from our theoretical model that larger firms initially pay a relative premium for high-skill workers, which is subsequently reduced due to tighter credit constraints.

As alluded to earlier, the German administrative earnings data are winsorized at around the 90$^{\text{th}}$ percentile of the population. We argue here that this is not a major concern for our identification, since it actually works *against* our empirical results. A comparison of Panels A and B of Table 2 reveals that firms in relationships with low-deposit (high-deposit) banks have relatively higher (lower) average wages, which is also reflected by their winsorized share of worker-years being 15.5 (5.0) percent. Thus, firms in relationships with low-deposit banks are relatively more likely to have high earnings winsorized. It is useful to think of our empirical setting as a combination of aggregate wage growth and differential firm-level wage growth arising from firms' varying exposure to the credit supply shock. The aggregate wage growth component alone pushes more employees at firms with relationships to low-deposit banks into the winsorizing range, leading to a mechanical decrease in relative within-firm inequality at these firms. This is the exact opposite of what we find: within-firm inequality declines more at firms in relationships with high-deposit banks. This suggests that our results are not driven by this statistical artifact.

In addition, our results are robust to considering a quintile-based grouping (bottom 20 percent, middle 60 percent, and top 20 percent) or just two groups, below versus above the median wage within a firm—see Tables A.1 and A.2 in Appendix A.

Table 6: Effects of Monetary Policy-Induced Credit Supply on Wages and Employment, by Workers' Within-Firm Pay Rank

	ln(wage)			Unemployed next year $\in \{0,1\}$		
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio × After(2014) × Bottom 33% within firm	0.055***	0.067***	0.050***	0.032***	0.027***	0.040***
	(0.013)	(0.013)	(0.012)	(0.006)	(0.005)	(0.006)
Deposit ratio × After(2014) × Middle 33% within firm	-0.015**	-0.010	-0.010	0.024***	0.025***	0.028***
	(0.007)	(0.007)	(0.008)	(0.004)	(0.003)	(0.004)
Deposit ratio × After(2014)	-0.024***			-0.010		
	(0.008)			(0.009)		
Deposit ratio × Bottom 33% within firm	-0.149***	-0.142***		-0.002	0.010**	
	(0.020)	(0.016)		(0.005)	(0.005)	
Deposit ratio × Middle 33% within firm	-0.101***	-0.098***		-0.000	0.004	
	(0.014)	(0.013)		(0.004)	(0.004)	
After(2014) × Bottom 33% within firm	0.063***	0.059***	0.016**	0.007	0.014***	0.025***
	(0.009)	(0.009)	(0.007)	(0.004)	(0.004)	(0.005)
After(2014) × Middle 33% within firm	0.005	0.003	-0.012**	-0.012***	-0.006***	-0.005*
	(0.005)	(0.005)	(0.005)	(0.003)	(0.002)	(0.003)
Worker FE	Y	Y	N	Y	Y	N
Firm FE	Y	N	N	Y	N	N
Worker-firm FE	N	N	Y	N	N	Y
Year FE	Y	N	N	Y	N	N
Firm-year FE	N	Y	Y	N	Y	Y
N	62,389,766	61,920,221	60,379,179	60,280,277	59,856,219	58,536,104

Notes: The sample consists of full-time employees i at German corporations j in year t from 2010 to 2017. The dependent variable in the columns 1–3 is the natural logarithm of the wage of individual i at firm j in year t. The dependent variable in the columns 4–6 is an indicator variable that equals 1 if individual i is unemployed in year $t+1$ and 0 otherwise. $\textit{Deposit ratio}_j \in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. $\textit{After}(2014)_t$ is a dummy variable for the years 2014–2017. *Bottom* 33% (*Middle* 33%) $\textit{within firm}_i$ is an indicator variable for whether worker i's wage is in the bottom tercile (middle tercile) of the wage distribution of the firm where i was employed in the last available year during the preperiod from 2010 to 2013. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

Firm-Level Aggregation. At the worker level, we find that initially higher-paid workers see relative wage declines, while initially lower-paid workers are more likely to become unemployed. As a consequence, within-firm wage inequality decreases. Our worker-level analysis held worker composition constant via worker or worker-firm-match fixed effects. We now turn to firm-level outcomes that explicitly account for compositional changes arising from hiring and separations. We construct measures of within-firm wage inequality for all firms in each year and estimate:

$$y_{jt} = \beta \mathit{Deposit\ ratio}_j \times \mathit{After}(2014)_t + \psi_j + \zeta_{k(j)t} + \varepsilon_{jt}, \tag{13}$$

where y_{jt} is a measure of within-firm pay inequality at firm j in year t, ψ_j denotes firm fixed effects, and $\zeta_{k(j)t}$ are industry-year, state-year, or industry-state-year fixed effects.

Table 7 reports the results. Columns 1 and 2 take the log P90-P10 wage percentile ratio as the dependent variable and include all firms, yielding a coefficient of -0.011 (standard error of 0.006)—a modest reduction in within-firm wage inequality among more-affected firms, consistent with our worker-level finding of relative wage declines at higher pay ranks (Table 6). The result is qualitatively robust to using the Gini coefficient as an alternative measure of within-firm wage inequality in columns 3 and 4.

Motivated by evidence that larger, publicly listed firms exhibit greater within-firm wage inequality (Mueller et al., 2017), columns 5 and 6 reestimate the baseline specification on public firms only, where the reduction in within-firm inequality is even more pronounced. This subsample also overlaps with our DealScan syndicated loans data (Tables 3 and 4) and includes firms that borrow from non-euro-area banks unaffected by euro-area monetary policy. Columns 7 and 8 exploit this for a falsification test by adding an interaction between $\mathit{After}(2014)_t$ and $\mathit{Non\text{-}euro\ deposit\ ratio}_j \in [0, 1]$, the mean deposit ratio across non-euro-area lead arrangers (and other banks outside negative-rate currency areas) from which firm j received a syndicated loan from 2010 to 2013. Reassuringly, the placebo coefficient is close to zero and statistically insignificant.

While rich in many dimensions, the IAB's linked employer-employee data do not allow us to measure top-wage inequality as the data are winsorized at the Social Security contribution threshold, which falls around the 90$^{\text{th}}$ percentile of the earnings distribution in our sample. This winsorizing may be particularly relevant for the pay structure at public firms, which tend to offer high variable compensation to their top management (Bertrand and Schoar, 2003; Gabaix and Landier, 2008). A plausible way for firms to reduce pay at the top is by adjusting variable compensation.

Table 7: Firm-Level Effects of Monetary Policy-Induced Credit Supply on Within-Firm Inequality

	ln(P90/P10)	ln(P90/P10)	Gini	Gini	ln(P90/P10)	ln(P90/P10)	ln(P90/P10)	ln(P90/P10)
Sample	All	All	All	All	Public firms	Public firms	Public firms	Public firms
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Deposit ratio × After(2014)	-0.011**	-0.011*	-0.009***	-0.009***	-0.305**	-0.336**	-0.432***	-0.376**
	(0.006)	(0.006)	(0.001)	(0.001)	(0.138)	(0.164)	(0.155)	(0.173)
Non-euro deposit ratio × After(2014)							0.052	0.131
							(0.126)	(0.154)
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y
Industry-year FE	Y	N	Y	N	Y	N	Y	N
State-year FE	Y	N	Y	N	Y	N	Y	N
Industry-state-year FE	N	Y	N	Y	N	Y	N	Y
N	2,765,907	2,765,907	2,311,806	2,311,806	1,315	1,151	1,133	1,004

Notes: The unit of observation is the firm-year level jt. In columns 1–4, the sample consists of all German corporations j in year t from 2010 to 2017. In columns 5–8, the sample is limited to all publicly listed German corporations j that are active in the syndicated loans market in year t from 2010 to 2017. The dependent variable in columns 3–4 is the Gini coefficient based on firm j's wage distribution in year t. In all remaining columns, the dependent variable is the delta log of the wage at the 90^{th} versus 10^{th} percentile of firm j's wage distribution in year t. *Deposit ratio*$_j \in [0, 1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. *Non-euro deposit ratio*$_j \in [0, 1]$ is the average deposits-to-assets ratio, measured in 2013, across all non-euro area banks (and other banks not based in negative-rate currency areas) from which firm j received syndicated loans from 2010 to 2013. *After*(2014)$_t$ is an indicator variable for the years 2014–2017. Industry-year fixed effects are based on the one-digit industry code of firm j's industry. State-year fixed effects are based on the modal location (state) of firm j's establishments. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

To test this adjustment mechanism, we use compensation data for executive and nonexecutive board members of DAX-listed firms from BoardEx. Although large firms with capital market access tend to be more sheltered from credit supply shocks (Chodorow-Reich, 2014), we still find an effect on larger German firms that are active in the syndicated loans market. In columns 1, 3, and 5 of Table A.3 in Appendix A, we provide small-sample evidence that a negative credit supply shock is associated with a reduction of top-to-bottom wage inequality within the listed firms. Focusing on executive board members, column 1 shows a point estimate that is large and negative but noisily estimated and barely significant at the 10 percent level. Splitting the executive board's total pay into salary and bonus, we find a significant negative reduction in bonus (column 5) but not in salary (column 3). This suggests that firms take into account the availability of credit and associated future growth prospects when reducing top-earners' variable compensation due to tighter financial constraints.

In Germany, some company board positions are allocated to worker representatives and other nonexecutives (Jäger et al., 2021). For nonexecutive board members, who typically do not receive substantial variable compensation, we find no significant response in their relative pay (columns 2, 4, and 6). In line with our rationale, fewer of them receive any variable pay to start with, as seen by the number of observations in column 5 versus column 6.

We also consider the effects of the negative credit supply shock on firm-level employment. The key difference between this analysis and our previous worker-level analysis is that we now account for both new hires and separations. Table 8 presents the results from estimating specification (13) for different employment counts. All specifications in this table control for industry-year and state-year or industry-state-year fixed effects. Columns 1 and 2 show that firms more exposed to negative rates see a significant reduction in overall employment, in line with Part (ii) of Proposition 1 of our theoretical model. The coefficients range from -0.009 to -0.011, suggesting that a one standard deviation increase in firm-level exposure is associated with an up to $0.153 \times 0.011 = 0.2$ percent reduction in total employment.

Consistent with our worker-level effects in columns 4 to 6 of Table 6, columns 3 and 4 of Table 8 show that more-exposed firms see a significant reduction in their share of nonmanagerial workers, who are less likely to be at the top of the wage distribution. Finally, columns 5 and 6 show that the negative credit supply shock is also associated with a reduction in part-time work, suggesting that those workers are more likely to leave employment or be asked to work extra hours.

Table 8: Firm-Level Effects of Monetary Policy-Induced Credit Supply on Employment

	ln(no. of all employees)		Share nonmanagerial		Share part-time	
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio × After(2014)	-0.011**	-0.009*	-0.006***	-0.005***	-0.007***	-0.007***
	(0.005)	(0.005)	(0.001)	(0.001)	(0.001)	(0.001)
Firm FE	Y	Y	Y	Y	Y	Y
Industry-year FE	Y	N	Y	N	Y	N
State-year FE	Y	N	Y	N	Y	N
Industry-state-year FE	N	Y	N	Y	N	Y
N	2,797,072	2,797,072	2,797,072	2,797,072	2,797,072	2,797,072

Notes: The sample consists of all German corporations j in year t from 2010 to 2017. The dependent variable in columns 1–2 is the natural logarithm of the total number of employees at firm j in year t. The dependent variable in columns 3–4 is the ratio of nonmanagerial staff to all employees at firm j in year t, ranging from 0 to 1. The dependent variable in columns 5–6 is the ratio of part-time staff to all employees at firm j in year t, ranging from 0 to 1. $Deposit\ ratio_j \in [0, 1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. $After(2014)_t$ is an indicator variable for the years 2014–2017. Industry-year fixed effects are based on the one-digit industry code of firm j's industry. State-year fixed effects are based on the modal location (state) of firm j's establishments. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

Between-Firm Heterogeneity. While we have shown that the credit supply shock from negative rates led to lower wages on average, we now address the extent to which firms adjusted wages differently. To explore this, we estimate variants of specification (10) of our empirical strategy, which adds an interaction term indicating a firm's mean wage rank.

Table 9 presents the results. We always include worker-firm fixed effects so as to focus on incumbent workers. Column 1, which also includes year fixed effects, shows that initially higher-paying firms respond to the negative credit supply shock by reducing relative wages. This continues to hold true in column 2 after replacing year fixed effects with more granular industry-year and state-year fixed effects of the respective firms, or even more granular industry-state-year fixed effects in column 3.

Columns 4 to 6 test for differential unemployment effects across firm pay ranks. To this end, we replace the dependent variable with an indicator for whether a worker will be unemployed the next year. We obtain a negative estimate of the interaction coefficient of interest, which becomes statistically significant after replacing year fixed effects with more granular time-varying fixed effects.

Table 9: Effects of Monetary Policy-Induced Credit Supply on Wages and Employment, by Firms' Pay Rank

		ln(wage)			Unemployed next year $\in \{0,1\}$	
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio × After(2014) × Firm pay rank	-0.135***	-0.067**	-0.057**	-0.002	-0.046***	-0.051***
	(0.032)	(0.028)	(0.026)	(0.020)	(0.017)	(0.017)
Deposit ratio × After(2014)	0.059***	0.030	0.023	-0.030**	-0.009	-0.004
	(0.019)	(0.019)	(0.018)	(0.012)	(0.010)	(0.010)
After(2014) × Firm pay rank	0.173***	0.122***	0.115***	-0.087***	-0.066***	-0.062***
	(0.023)	(0.021)	(0.020)	(0.015)	(0.012)	(0.012)
Worker-firm FE	Y	Y	Y	Y	Y	Y
Year FE	Y	N	N	Y	N	N
Industry-year FE	N	Y	N	N	Y	N
State-year FE	N	Y	N	N	Y	N
Industry-state-year FE	N	N	Y	N	N	Y
N	67,435,825	67,418,017	67,418,017	58,958,314	58,942,622	58,942,622

Notes: The sample consists of full-time employees i at German corporations j in year t from 2010 to 2017. The dependent variable in columns 1–3 is the natural logarithm of the wage of individual i at firm j in year t. The dependent variable in columns 4–6 is an indicator variable that equals 1 if individual i is unemployed in year $t+1$ and 0 otherwise. *Deposit ratio*$_j \in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. *After*(2014)$_t$ is a dummy variable for the years 2014–2017. *Firm pay rank*$_j$ is the rank (from $0 =$ lowest to $1 =$ highest) of firm j in terms of its average pay in 2013. Industry-year fixed effects are based on the one-digit industry code of firm j's industry. State-year fixed effects are based on the modal location (state) of firm j's establishments. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

In terms of economic magnitudes, a one standard deviation increase in firms' exposure (*Deposit ratio$_j$*) is associated with an at least ($0.153 \times 0.057 =$) 0.9 percent (column 3) larger wage decline at the top-ranked relative to the bottom-ranked firm. In contrast, the same one-standard-deviation increase in exposure is associated with a ($0.153 \times 0.051 =$) 0.8 percentage point higher probability of leaving employment at the bottom-ranked relative to the top-ranked firm (column 6).

These findings complement our results on lower within-firm inequality in Table 6, where firm-year fixed effects absorb between-firm differences in the development of the average wage, Table 7, and Table A.3 in Appendix A in that initially higher-paying firms are more likely to have a larger portion of their payroll accrue to managerial and other staff with variable compensation. This, in turn, enables them to respond to a tightening of the credit supply by reducing their pay by a relatively larger amount.

As argued above, the mechanical effect of winsorization runs *counter* to our finding that between-firm inequality declines due to the credit supply shock. This is because, all else equal, initially higher-paying firms have a larger share of their workers in the winsorized range, which mechanically dampens the measured wage response at those firms. This is the exact opposite of what we find, namely, a greater reduction in wages at initially higher-paying firms.

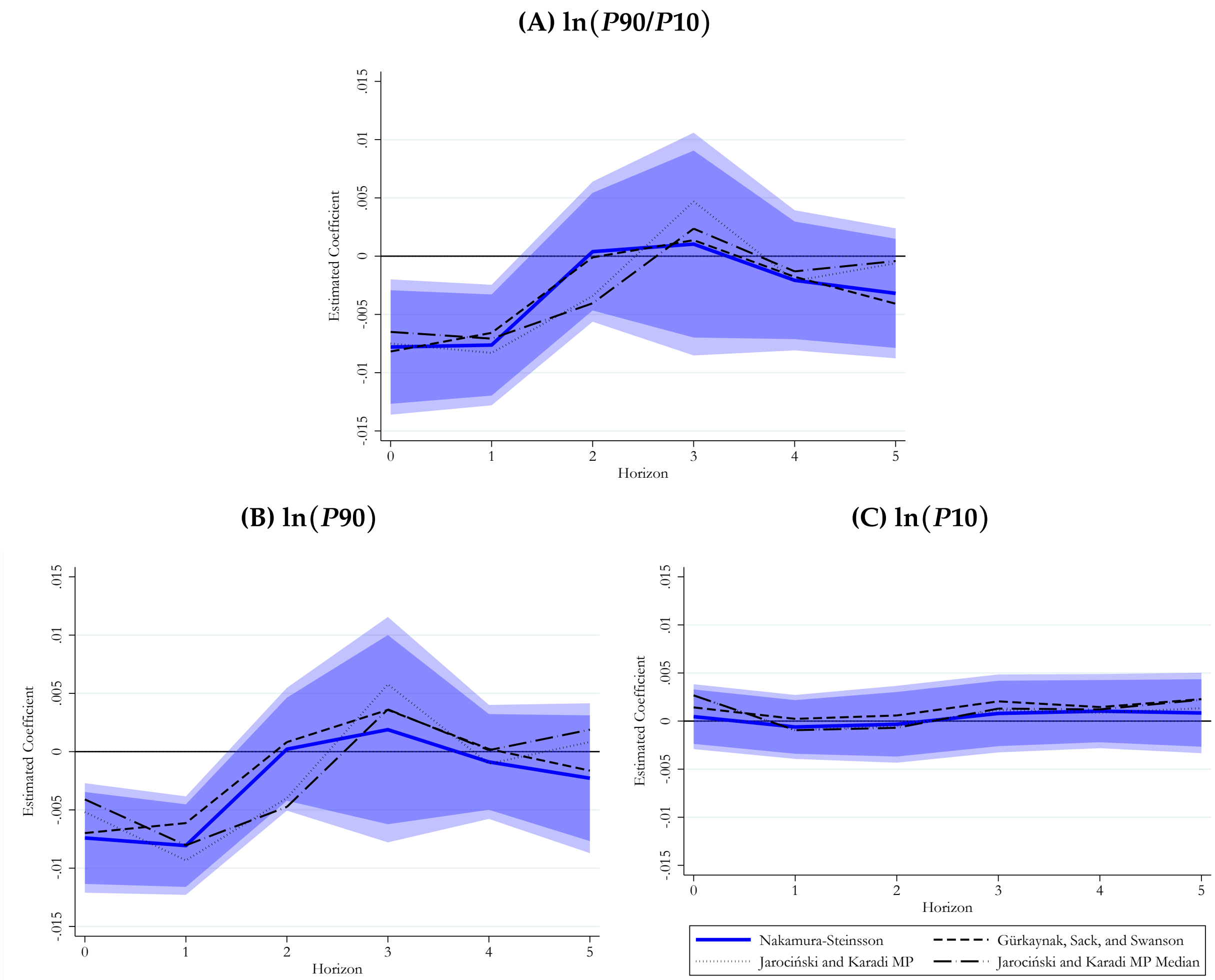
To summarize, we find that initially higher-paying firms implement relative wage cuts while retaining a relatively larger share of their workforce. As a consequence and in line with Part (iv) of Proposition 1 of our theoretical model, between-firm wage inequality decreases.

6 External Validity

Thus far, we have studied firms with differentially binding credit constraints as a function of their relationship banks' exposure to the introduction of negative monetary policy rates. In our setting, negative monetary policy rates led to a relative contraction in credit supply to firms in relationships with high-deposit banks, which translated into distributional effects for workers due to firms' varying ability to adjust workers' pay throughout the wage distribution. Our focus on monetary policy and wage inequality in Germany raises a question about the external validity of our findings—Germany being a large and highly bank-dependent economy in the euro area, while other countries like the US have more market-based financial systems.

To address this, we examine whether the same forces we highlight shape the US wage distribution's response to monetary policy. To this end, we use contractionary monetary policy shocks

Figure 3: Effects of Contractionary Monetary Policy and Wage Floors on Inequality in the US



Notes: The figure shows results from local projections, obtained by regressing h-step-ahead values of wage percentiles on monetary policy shocks interacted with a state-level indicator for having an above-median minimum wage in place, as described in Appendix C. The figure plots the estimated coefficients on this interaction term across the horizon, measured in years since impact. The dependent variables are the natural logarithm of the state-level P90/P10 wage ratio (Panel A), the 90$^{\text{th}}$ percentile (Panel B), and the 10$^{\text{th}}$ percentile (Panel C). The monetary policy shocks are collapsed to annual frequency and standardized to mean zero and standard deviation one. We use the shocks of Nakamura and Steinsson (2018), monetary policy path shocks of Gürkaynak et al. (2005), and the monetary shocks obtained from simple sign restriction and median rotation in Jarociński and Karadi (2020). The plot shows point estimates alongside 90% and 95% confidence bands, obtained from standard errors clustered at the state level. Confidence intervals are based on the estimated standard errors for the shock series of Nakamura and Steinsson (2018).

(based on the shock series by Nakamura and Steinsson, 2018; Gürkaynak et al., 2005; Jarociński and Karadi, 2020) in the US, interacted with differences in the level—and, thus, differences in the relative bindingness—of the minimum wage across states as a proxy for downward wage rigidity that makes it differentially difficult for firms to adjust pay at different points in the wage distribution. We present our results in the form of local projections of contractionary monetary policy shocks (using different shock series) on the log P90/P10 wage percentile ratio in a given state s and year t, relegating details on our estimation to Appendix C.

Figure 3 shows that contractionary monetary policy shocks are associated with a reduction in inequality in high as opposed to low minimum wage states. The effect persists for one additional year and is driven by the top 10% of the wage distribution, while the bottom 10% of the distribution shows barely any movement. This parallels our German findings in Table 6, which we have shown to aggregate up to the firm level in Table 7.

Finally, we add another dimension of heterogeneity across states by considering more versus less bank-dependent states, reflecting the importance of bank credit for our mechanism in Germany. Appendix Figure C.1 shows that the decline in inequality in high versus low minimum wage states in response to a contractionary monetary policy shock is particularly pronounced in more bank-dependent states. This lends support to the idea that our mechanism operates through bank credit pass-through to firms, rather than through any direct, non-bank channel of monetary policy. Credit supply affects inequality through similar channels across different financial systems when two conditions exist: downward wage rigidity for lower-paid workers and some degree of firm dependence on bank credit.

7 Conclusion

We empirically study the effects of credit supply on the distribution of wages and employment within and between firms. A neoclassical view of labor markets without frictions suggests that identical workers should earn the same pay across employers, and that credit supply to firms should not matter for the determination of employment or wages. In contrast, we develop a simple equilibrium model in which credit and search frictions interact to yield predictions about the elasticities of wages and employment to credit supply, both within and between firms. We test our model's predictions using a novel dataset that spans the complete credit chain—from banks to firms to workers—in Germany. We find that lower credit supply reduces firm-level average wages and employment, while having differential effects across workers within and between firms, in line with our theoretical predictions.

Our analysis suggests several directions for future work. First, the relative adjustments of wages versus employment we find are consistent with existing studies of incumbent workers' wage rigidity (Schoefer, 2021). It would be interesting to further explore the wage rigidity channel by combining variation in credit supply like ours with direct measures of wage floors across sectors, occupations, and time. Second, while our analysis focuses on workers' wages and employment, a natural extension could explore other margins of adjustment to credit supply shocks, including firms' investment choices and workers' job-search choices. Third and finally, the effects of credit in our study are estimated over a relatively short time window since June 2014. Understanding the medium- and long-term effects of credit supply through the channels highlighted in our work deserves further attention.

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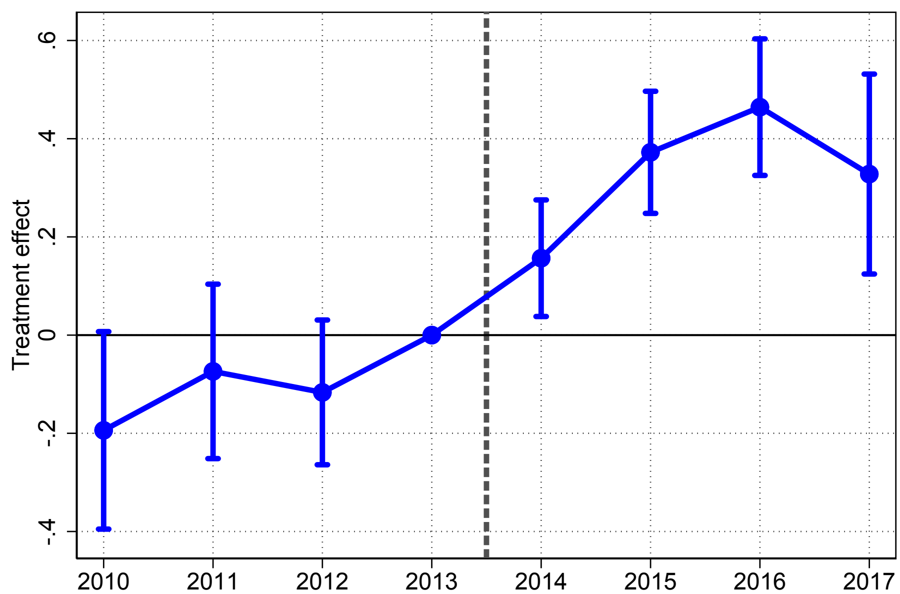
Appendix for
"Credit Supply, Firms, and Earnings Inequality"

Christian Moser   Farzad Saidi   Benjamin Wirth   Stefanie Wolter

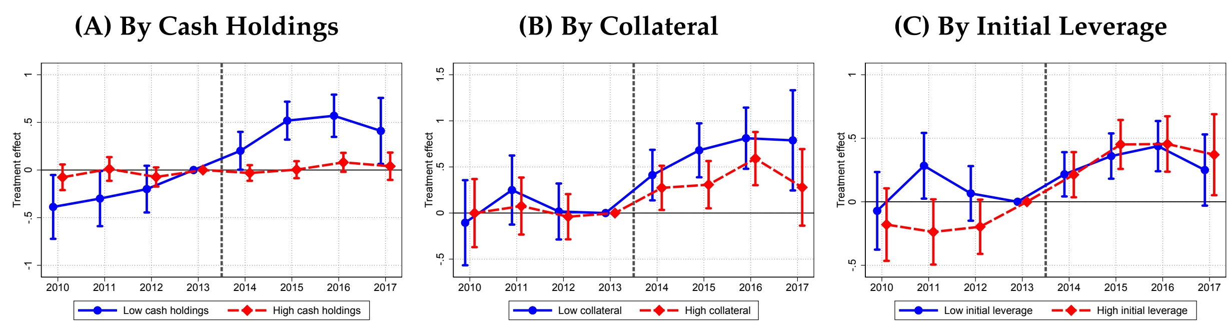
A Data Appendix

Figure A.1: Impact of Negative Policy Rates on Firms' Cash Holdings



Notes: This figure plots estimates of β_τ, alongside 95 percent confidence bands, over time (each year represents the respective year-end) based on the event study specification in (12), using as dependent variable the natural logarithm of firm j's holdings of cash and cash equivalents, estimated on the sample of German firms in the administrative linked employer-employee data merged with Orbis from 2010 to 2017.

Figure A.2: Heterogeneity in the Impact of Negative Policy Rates on Firms' Cash Holdings



Notes: This figure plots estimates of β_τ, alongside 95 percent confidence bands, over time (each year represents the respective year-end) based on the event study specification in (12), using as dependent variable the natural logarithm of firm j's holdings of cash and cash equivalents, estimated on the sample of German firms in the administrative linked employer-employee data merged with Orbis from 2010 to 2017. The specification is separately estimated by, and the results are shown for, subgroups below and above median cash holdings (Panel A), collateral defined as the capital intensity or capital per worker (Panel B), and initial leverage (Panel C).

Table A.1: Effects of Monetary Policy-Induced Credit Supply on Wages and Employment, by Workers' Within-Firm Pay Rank: Across Quintiles

	ln(wage)			Unemployed next year $\in \{0,1\}$		
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio × After(2014) × Bottom 20% within firm	0.028	0.063***	0.050***	0.030***	0.020***	0.031***
	(0.018)	(0.019)	(0.017)	(0.007)	(0.006)	(0.008)
Deposit ratio × After(2014) × Middle 60% within firm	-0.018***	-0.013*	-0.015**	0.029***	0.025***	0.029***
	(0.007)	(0.007)	(0.007)	(0.006)	(0.004)	(0.004)
Deposit ratio × After(2014)	-0.009			-0.022**		
	(0.008)			(0.010)		
Deposit ratio × Bottom 20% within firm	-0.137***	-0.138***		0.003	0.016***	
	(0.021)	(0.018)		(0.006)	(0.006)	
Deposit ratio × Middle 60% within firm	-0.117***	-0.109***		0.005	0.013***	
	(0.015)	(0.014)		(0.005)	(0.004)	
After(2014) × Bottom 20% within firm	0.148***	0.135***	0.069***	0.016***	0.025***	0.047***
	(0.012)	(0.013)	(0.011)	(0.005)	(0.005)	(0.006)
After(2014) × Middle 60% within firm	0.009*	0.006	-0.011**	-0.014***	-0.006**	-0.004
	(0.005)	(0.005)	(0.005)	(0.004)	(0.003)	(0.003)
Worker FE	Y	Y	N	Y	Y	N
Firm FE	Y	N	N	Y	N	N
Worker-firm FE	N	N	Y	N	N	Y
Year FE	Y	N	N	Y	N	N
Firm-year FE	N	Y	Y	N	Y	Y
N	60,936,322	60,465,690	58,934,938	58,938,217	58,513,100	57,202,523

Notes: The sample consists of full-time employees i at German corporations j in year t from 2010 to 2017. The dependent variable in the first three columns is the natural logarithm of the wage of individual i at firm j in year t. The dependent variable in the last three columns is an indicator variable for whether individual i is unemployed in year $t+1$. *Deposit ratio*$_j \in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. *After*(2014)$_t$ is a dummy variable for the years 2014–2017. *Bottom* 20% (*Middle* 60%) *within firm*$_i$ is an indicator variable for whether worker i's wage is in the bottom 20 percent (middle 60 percent) of the wage distribution of the firm where i was employed in the last available year during the preperiod from 2010 to 2013. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

Table A.2: Effects of Monetary Policy-Induced Credit Supply on Wages and Employment, by Workers' Within-Firm Pay Rank: Below vs. Above Median

	ln(wage)			Unemployed next year $\in \{0,1\}$		
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio × After(2014) × Bottom 50% within firm	0.027***	0.029***	0.021**	0.028***	0.025***	0.033***
	(0.008)	(0.008)	(0.008)	(0.004)	(0.003)	(0.004)
Deposit ratio × After(2014)	-0.028***			-0.006		
	(0.010)			(0.009)		
Deposit ratio × Bottom 50% within firm	-0.107***	-0.104***		0.008**	0.013***	
	(0.014)	(0.014)		(0.004)	(0.004)	
After(2014) × Bottom 50% within firm	0.036***	0.036***	0.007	0.004	0.006**	0.013***
	(0.005)	(0.005)	(0.005)	(0.003)	(0.002)	(0.003)
Worker FE	Y	Y	N	Y	Y	N
Firm FE	Y	N	N	Y	N	N
Worker-firm FE	N	N	Y	N	N	Y
Year FE	Y	N	N	Y	N	N
Firm-year FE	N	Y	Y	N	Y	Y
N	60,182,633	59,711,416	58,211,678	58,174,328	57,748,797	56,465,121

Notes: The sample consists of full-time employees i at German corporations j in year t from 2010 to 2017. The dependent variable in the first three columns is the natural logarithm of the wage of individual i at firm j in year t. The dependent variable in the last three columns is an indicator variable for whether individual i is unemployed in year $t+1$. *Deposit ratio*$_j \in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. *After*(2014)$_t$ is a dummy variable for the years 2014–2017. *Bottom* 50% *within firm*$_i$ is an indicator variable for whether worker i's wage is in the bottom half of the wage distribution of the firm where i was employed in the last available year during the preperiod from 2010 to 2013. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

Table A.3: Effects of Monetary Policy-Induced Credit Supply Shock on Within-Firm Inequality: Board Members of Publicly Listed Firms

	ln(p50 board total/p5)		ln(p50 board salary/p5)		ln(p50 board bonus/p5)	
Board members	Executive	Nonexecutive	Executive	Nonexecutive	Executive	Nonexecutive
	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio × After(2014)	-0.877*	-0.311	-0.696	0.097	-0.888*	-0.295
	(0.485)	(0.548)	(0.456)	(0.577)	(0.532)	(1.450)
Firm FE	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y
N	266	266	266	266	263	105

Notes: The unit of observation is the firm-year level jt. The sample consists of DAX-listed German corporations j in year t from 2010 to 2016 for which we have board-compensation data from BoardEx. The dependent variable in columns 1 and 2 is the delta log of the median total compensation of executive and nonexecutive board members, respectively, at firm j in year t versus the annualized wage at the 5^{th} percentile of firm j's wage distribution in year t. The dependent variable in columns 3 and 4 is the delta log of the median salary of executive and nonexecutive board members, respectively, at firm j in year t versus the annualized wage at the 5^{th} percentile of firm j's wage distribution in year t. The dependent variable in columns 5 and 6 is the delta log of the median bonus (conditional on being nonzero) of executive and nonexecutive board members, respectively, at firm j in year t versus the annualized wage at the 5^{th} percentile of firm j's wage distribution in year t. $\textit{Deposit ratio}_j \in [0, 1]$ is the average deposits-to-assets ratio, measured in 2013, across all banks that firm j reports to be in a banking relationship with from 2010 to 2013. $\textit{After}(2014)_t$ is an indicator variable for the years 2014–2016. Robust standard errors (clustered at the firm level) are in parentheses. ***, **, and * denote significance at the 1 percent, 5 percent, and 10 percent level, respectively.

B Model Appendix

B.1 Equilibrium Definition

Definition 1. *A stationary search equilibrium is a set of worker value functions $\{S_a, W_a\}_a$ and policy functions $\{\phi_a\}_a$; a firm value function Π and policy functions $\{w_a, v_a\}_a$; wage offer distributions $\{F_a(w)\}_a$; measures of unemployed workers $\{u_a\}_a$, aggregate job searchers $\{U_a\}_a$, aggregate vacancies $\{V_a\}_a$, and labor market tightnesses $\{\theta_a\}_a$; job offer arrival rates $\{\lambda_a^u, \lambda_a^e\}_a$; and firm sizes $\{l_a\}_a$ such that for all a:*

- *Given $F_a(w)$ and $\{\lambda_a^u, \lambda_a^e\}$, the value functions S_a and W_a satisfy equations (1) and (2);*
- *Unemployed workers' job acceptance policy follows a threshold rule with reservation wage*

$$\phi_a = b_a + (\lambda_a^u - \lambda_a^e) \int\limits_{w' \geq \phi_a} \frac{1 - F_a(w')}{\rho + \delta_a + \lambda_a^e [1 - F_a(w')]} dw', \quad \forall a,$$

  and employed workers with wage w accept any job w' such that $w' > w$;
- *Given $l_a(\cdot)$, firms' value function Π and optimal policy functions $\{w_a, v_a\}$ are consistent with the problem in equations (3)–(4);*
- *Measures of unemployed workers are given by*

$$u_a = \frac{\delta_a}{\delta_a + \lambda_a^u}, \quad \forall a,$$

  aggregate job searchers are given by

$$U_a = \mu_a [u_a + s_a^e(1 - u_a)], \quad \forall a,$$

  aggregate vacancies are given by

$$V_a = E \int\limits_j v_a(j)\, d\Gamma(j), \quad \forall a,$$

  and labor market tightness θ_a is given by

$$\theta_a = \frac{V_a}{U_a}, \quad \forall a.$$

- *Given θ_a, the job offer arrival rates satisfy*

$$\lambda_a^u = \chi_a \theta_a^\alpha, \qquad \lambda_a^e = s_a \lambda_a^u.$$

- *Given $F_a(w)$, $\{\lambda_a^u, \lambda_a^e\}_a$, and V_a, steady-state firm sizes satisfy*

$$l_a(w, v) = \left(\frac{1}{\delta_a + \lambda_a^e [1 - F_a(w)]} \right)^2 \frac{1}{V_a} \mu_a u_a \lambda_a^u (\delta_a + \lambda_a^e) v, \quad \forall a.$$

- *The offer distribution satisfies $F_a(w) = \int_j v_a(j) \mathbf{1}[w_a(j) \leq w]\, d\Gamma(j) / V_a$.*

B.2 Proof of Proposition 1

As in equation (7) of the main text, we first reformulate the firm's problem by defining *effective productivity* as

$$\tilde{p} = p\frac{1+r}{1+(1+\psi)r}, \tag{14}$$

where ψ is the Lagrange multiplier on a firm's credit constraint. From here, the proof closely follows that in Morchio and Moser (2026), which we adapt to our setting, where credit constraints codetermine firms' effective productivity.

B.2.1 Part (i) of Proposition 1

Claim: For any firm j, a decrease in their idiosyncratic credit limit ξ_j leads to lower firm-level average wages for the same workers at firm j.

Proof. To prove this part, we proceed in two steps.

Step 1. In the first step, we prove monotonicity of w_a^* in the composite productivity $\tilde{p}$. We can rewrite the firm's FOCs as

$$[\partial w_a]: \quad 1 = (\tilde{p} - w_a)\frac{2\lambda_a^e f_a(w_a)}{\delta_a + \lambda_a^e(1 - F_a(w_a))} \tag{15}$$

$$[\partial v_a]: \quad c_a^{v,0}\frac{\partial \tilde{c}^v(v_a)}{\partial v_a} = T_a(\tilde{p} - w_a)\left(\frac{1}{\delta_a + \lambda_a^e(1 - F_a(w_a))}\right)^2, \tag{16}$$

where $T_a = \mu_a[u_a\lambda_a^u(\delta_a + \lambda_a^e)]/V_a$. Equation (15) already shows that the optimal wage w_a is independent of the cost of posting vacancies, proving the first statement. Now consider equation (16); because the term on the right-hand side is always positive for $\tilde{p} > \phi_a$, it follows that optimal vacancies $v_a^*(\tilde{p}, c_a^{v,0})$ are always strictly positive for any firm with productivity $p > \phi_a$.

We now show that the derivative of wages with respect to $\tilde{p}$ is always positive. Define $h_a(\tilde{p}) = F_a(w_a^*(\tilde{p}))$. Thus:

$$h_a(\tilde{p}) = \frac{\int_{\tilde{p}' \geq \phi_a}^{\tilde{p}} \overline{v}_a^*(\tilde{p})\gamma_a(\tilde{p})}{V_a}\, d\,\tilde{p}' \tag{17}$$

$$h_a'(\tilde{p}) = f_a(w_a^*(\tilde{p}))\, w_a^{*\prime}(\tilde{p}) \tag{18}$$

$$f_a(w_a^*(\tilde{p})) = h_a'(\tilde{p})/w_a^{*\prime}(\tilde{p}), \tag{19}$$

where $\overline{v}_a^*(\tilde{p}) = \int v_a^*(\tilde{p}, c')\gamma_a^c(c'|\tilde{p})\, dc'$ is the integral of optimal vacancies conditional on $\tilde{p}$ and $\gamma_a^c(c|\tilde{p})$ is the density of vacancy posting costs $c_a^{v,0}$ conditional on $\tilde{p}$, $\gamma_a(\tilde{p})$ is the marginal density of composite productivity $\tilde{p}$ and $\partial w_a^*(\tilde{p})/\partial \tilde{p} = w_a^{*\prime}(\tilde{p})$ is the derivative of equilibrium wage with respect to $\tilde{p}$. Thus, we can rewrite $h_a'(\tilde{p}) = \frac{\overline{v}_a^*(\tilde{p})}{V_a}\gamma(\tilde{p})$ by differentiating equation (17) using Leibniz's integral rule.

Using these identities, we can write $f_a(w_a^*(\tilde{p})) = \frac{\overline{v}_a^*(\tilde{p})}{V_a}\gamma_a(\tilde{p})\partial\tilde{p}/\partial w_a^*(\tilde{p})$. Thus, we can rewrite equation (15) as

$$\frac{\partial w_a^*(\tilde{p})}{\partial \tilde{p}} = (\tilde{p} - w_a^*) \frac{2\lambda_a^e}{\delta_a + \lambda_a^e(1 - h_a(\tilde{p}))} \frac{\overline{v}_a^*(\tilde{p})}{V_a} \gamma_a(\tilde{p}). \tag{20}$$

Because the right-hand side of this expression is always positive for $\tilde{p} > \phi_a$, it follows that $\partial w_a^*(\tilde{p})/\partial \tilde{p} > 0$, thus proving that equilibrium wage is increasing in $\tilde{p}$.

Step 2. Having demonstrated above that the wage policy function $w_a^*(\tilde{p})$ is increasing in effective productivity $\tilde{p}$, the result that wages are increasing in productivity p and decreasing (constant) in the Lagrange multiplier on the credit limit ψ for high-ability (low-ability) workers follows from the definition of $\tilde{p}$ in equation (14) above. As a result, a reduction in credit supply leaves the wages of low-skill workers constant while reducing the wages of high-skill workers, thus resulting in lower firm-level average wages for the same workers at a given firm. □

B.2.2 Part (ii) of Proposition 1

Claim: For any firm j, a decrease in their idiosyncratic credit limit ξ_j leads to lower firm-level employment at firm j.

Proof. Expected profits per worker contacted by a firm is

$$\pi_a(\tilde{p}, w) = h_a(w) J_a(\tilde{p}, w),$$

where $h_a(w)$ is the acceptance probability and $J_a(\tilde{p}, w)$ is the value of employing a worker to a firm with composite productivity $\tilde{p}$ providing wage w. Let us define the relative speed of climbing the job ladder among workers of type a as $\kappa_a \equiv \lambda_a^e/\delta_a$. Under the assumption that firms maximize long-run profits, the value of employing a worker is simply

$$\begin{aligned} J_a(\tilde{p}, w) &= \frac{\tilde{p} - w}{\delta_a + \lambda_a^e(1 - F_a(w))} \\ &= \frac{(\tilde{p} - w) \,/\, (\delta_a)}{1 + \kappa_a^e \left(1 - F_a\left(w\right)\right)}, \end{aligned}$$

Given the cross-sectional wage distribution $G_a(w)$ for workers of type a, the acceptance probability for a firm offering w to workers of type a is

$$\begin{aligned} h_a(w) &= \frac{u_a + s_a^e \left(1 - u_a\right) G_a\left(w\right)}{u_a + s_a^e \left(1 - u_a\right)} \\ &= \frac{\delta_a + s_a^e \left(\lambda_a^u\right) G_a\left(w\right) \left(\delta_a + \lambda_a^u\right)}{\delta_a + s_a^e \left(\lambda_a^u\right) \left(\delta_a + \lambda_a^u\right)} \\ &= \frac{1 + s_a^e \kappa_a^u G_a\left(w\right) \left(1 + \kappa_a^u\right)}{1 + s_a^e \kappa_a^u \left(1 + \kappa_a^u\right)} \\ &= \frac{1 + s_a^e \kappa_a^u \left[\frac{F_a(w)}{1 + \kappa_a^e [1 - F_a(w)]}\right] \left(1 + \kappa_a^u\right)}{1 + s_a^e \kappa_a^u \left(1 + \kappa_a^u\right)} \\ &= \frac{1 + \kappa_a^e \left[1 - F_a\left(w\right)\right] + s_a^e \kappa_a^u F_a\left(w\right) \left(1 + \kappa_a^u\right) \left[1 + \kappa_a^e \left[1 - F_a\left(w\right)\right]\right]}{\left[1 + s_a^e \kappa_a^u \left(1 + \kappa_a^u\right)\right] \left[1 + \kappa_a^e \left[1 - F_a\left(w\right)\right]\right]}, \end{aligned}$$

where $\kappa_a^u = \lambda_a^u / \delta_a$. Combining expressions, the expected profits per contacted worker are

$$\pi(\tilde{p}, w) = h(w) J(\tilde{p}, w)$$
$$= \frac{\{1 + \kappa_a^e [1 - F_a(w)] + s_a^e \kappa_a^u F_a(w)(1 + \kappa_a^u)[1 + \kappa_a^e [1 F_a(w)]]\}(\tilde{p} - w)}{[1 + s_a^e \kappa_a^u (1 + \kappa_a^u)][1 + \kappa_a^e (1 - F_a(w))]^2 (\delta_a)}. \tag{21}$$

Given a vacancy filling rate q_a for workers of type a, the firm's problem becomes

$$\max_{w,v} \{\pi_a(\tilde{p}, w) v q_a - c_a(v)\}.$$

Therefore, the optimal wage and vacancy policy functions satisfy

$$w_a^*(\tilde{p}, \cdot) = \arg\max_w \pi_a(\tilde{p}, w)$$
$$\frac{\partial c_a(v^*(\tilde{p}, \cdot))}{\partial v} = \max_w \pi_a(\tilde{p}, w). \tag{22}$$

Since the vacancy cost function $c(\cdot)$ is convex, and $\pi(\tilde{p}, w)$ in equation (21) is strictly increasing in $\tilde{p}_j$, then it follows from an application of the envelope theorem to equation (22) that $v^*(\tilde{p}_j, \cdot)$ is strictly increasing in $\tilde{p}_j$ as well. Therefore, $v_a^*(\cdot)$ is strictly increasing in productivity p_j and strictly decreasing (constant) in the Lagrange multiplier on the credit constraint, ψ_j, for constrained (unconstrained) firms. □

B.2.3 Part (iii) of Proposition 1

Claim: For any firm j, a decrease in their idiosyncratic credit limit ξ_j leads to lower within-firm wage inequality, in the sense of a relatively greater reduction in wages among initially higher-paid workers at firm j.

Proof. The proof follows by combining the result in Part (i) of Proposition 1 with the fact that wages for low-ability workers are equal to the constant flow value of unemployment. Specifically, by Part (i), at constrained firms, wages of high-skill workers, w_{a_H}, are strictly increasing in ξ_j but wages of low-skill workers, w_{a_L}, are invariant to the credit limit ξ_j. Therefore, a reduction in the credit limit ξ_j that increases the Lagrange multiplier ψ_j strictly reduces the top-to-bottom wage difference,

$$\frac{\partial(w_{a_H} - w_{a_L})}{\partial \psi_j} = \frac{\partial w_{a_H}}{\partial \psi_j} < 0 \tag{23}$$

for credit constrained firms with $\psi_j > 0$. While equation (23) proves the result for one particular measure of within-firm wage inequality, an analogous result applies more generally due to the fact that

$$w_{a_L} = b_{a_L} \leq b_{a_H} < w_{a_H} \tag{24}$$

and

$$\frac{\partial w_{a_H}}{\partial \psi_j} < 0 = \frac{\partial w_{a_L}}{\partial \psi_j}. \tag{25}$$

□

B.2.4 Part (iv) of Proposition 1

Claim: For any firm j, a decrease in their idiosyncratic credit limit ξ_j leads to lower between-firm wage inequality, measured by a lower distance between the average wage at firm j and that at firm j' with rank $r(j') = 0$.

Proof. First, we have $\phi_{a_L} < \phi_{a_H}$ since the reservation wage ϕ_a satisfies

$$\phi_a = b_a + (\lambda_a^u - \lambda_a^e) \int_{w=\phi_a}^{\infty} \frac{1 - F_a(w)}{\rho + \delta_a + \lambda_a^e(1 - F_a(w))} dw, \tag{26}$$

combined with the fact that $b_{a_L} \leq b_{a_H}$ and $\lambda_{a_L}^e = 0 < \lambda_{a_H}^e$. Next, we have that the firm with the lowest composite productivity $\underline{\tilde{p}}$ pays exactly workers' reservation wages, $w_{a_L}(\underline{\tilde{p}}) = \phi_{a_L} = b_{a_L}$ and $w_{a_H}(\underline{\tilde{p}}) = \phi_{a_H} > b_{a_L}$. Note that the latter statement is independent of the bindingness of credit constraints. Finally, we have that any firm with higher composite productivity $\tilde{p}_j$ pay low-ability workers their reservation wage, $w_{a_L}(\tilde{p}_j) = \phi_{a_L}$, but high-ability workers some wage strictly above their reservation wage, $w_{a_H}(\tilde{p}_j) > \phi_{a_H}$.

Now consider the impact of a decrease in their credit limit ξ_j for some firm j. At the lowest-paying firm, $\tilde{p}_j = \underline{\tilde{p}}$ and wages are invariant to the credit limit. At any higher-paying firm, $\tilde{p}_j > \underline{\tilde{p}}$ and wages of high-ability workers are strictly decreasing in the Lagrange multiplier on the credit constraint ψ_j, while wages of low-ability workers are invariant, by Part (i) of Proposition 1. Therefore,

$$\frac{\partial(w_a(\tilde{p}_j) - w_a(\underline{\tilde{p}}))}{\partial \psi_j} = \frac{\partial w_a(\tilde{p}_j)}{\partial \psi_j} \leq 0 \tag{27}$$

for workers of any ability level a, with strict inequality for workers with high ability $a = a_H$ at firms with rank $r(j) > 0$. Because the firm-level average wage is a weighted mean of high-skill workers (with $\partial w_a(\tilde{p}_j)/\partial \psi_j < 0$) and low-skill workers (with $\partial w_a(\tilde{p}_j)/\partial \psi_j = 0$), the firm-level average wage at any firm j with rank $r(j) > 0$ is decreasing relative to that at firm j' with rank $r(j') = 0$. □

B.2.5 Statement of Proposition 1 Regarding Binding Versus Slack Credit Constraints

Claim: The comparative statics for any firm j with a binding credit constraint ($\psi_j > 0$) involve

- strictly lower wage outcomes at interior firm ranks ($r(j) > 0$), as in (i), (iii), and (iv) above, and
- strictly lower employment, as in (ii) above.

Any firm j with a strictly slack credit constraint ($\psi_j = 0$) experiences no changes in wages or employment.

Proof. On the one hand, if the credit constraint of firm j is initially binding, then a reduction in the credit limit ξ_j will strictly increase the Lagrange multiplier $\psi_j > 0$. On the other hand, if the credit constraint of firm j is initially strictly slack, then a reduction in the credit limit ξ_j will leave the Lagrange multiplier constant at $\psi_j = 0$. The claim then follows directly from the results in (i)–(iv), given the dependence of effective productivity $\tilde{p}_j$ on the credit constraint's Lagrange multiplier ψ_j. Note that the lowest-ranked firm will not change the levels of pay for either worker type, which are pinned down through their outside options, but will reduce employment instead. □

C External Validity Appendix

For Figure 3 in the main text and Figure C.1 below, we estimate local projections by estimating the following series of regressions for state s in census region r and year t:

$$y_{st+h} = \beta^h MP\ Shock_t \times MW_{st} + \theta MW_{st} + \psi y_{st-1} + \omega \mathbf{X}_{st} + \gamma_s + \delta_{rt} + \epsilon_{st} \quad \forall h = \{0, ..., 5\}, \quad (28)$$

where

- y_{st} is the natural logarithm of state-level wage percentiles (P90 or P10) or the log P90/P10 ratio. The data are drawn from the ACS and cover 2001–2024. We clean the data using the routine from Bauluz et al. (2024), restricting the sample to private-sector full-time employees who earn at least 50 percent of annualized federal minimum wage, and we adjust for top-coding following the approach from Autor et al. (2008).

- $MPShock_t$ is a series of monetary policy shocks for the United States, collapsed to annual frequency and standardized. We focus on the shocks of Nakamura and Steinsson (2018); MP path shocks of Gürkaynak et al. (2005)—both obtained via `https://www.acostamiguel.com/research.html`; and the monetary shocks obtained from simple sign restriction and median rotation in Jarociński and Karadi (2020)—obtained via `https://github.com/marekjarocinski/jkshocks_update_fed_202401`. We multiply the shock series by -1, so that a "positive" shock corresponds to a contractionary monetary policy shock in our figures.

- MW_{st} is an indicator for a state s having an above-median state-level minimum wage in place in year t. The data are downloaded from FRED.

- $\mathbf{X}_{st}$ are state-level controls: the natural logarithms of disposable income and GDP per capita (both obtained via the BEA) as well as the unemployment rate (obtained via the BLS).

- γ_s and δ_{rt} are fixed effects at the state and census region-year level, respectively.

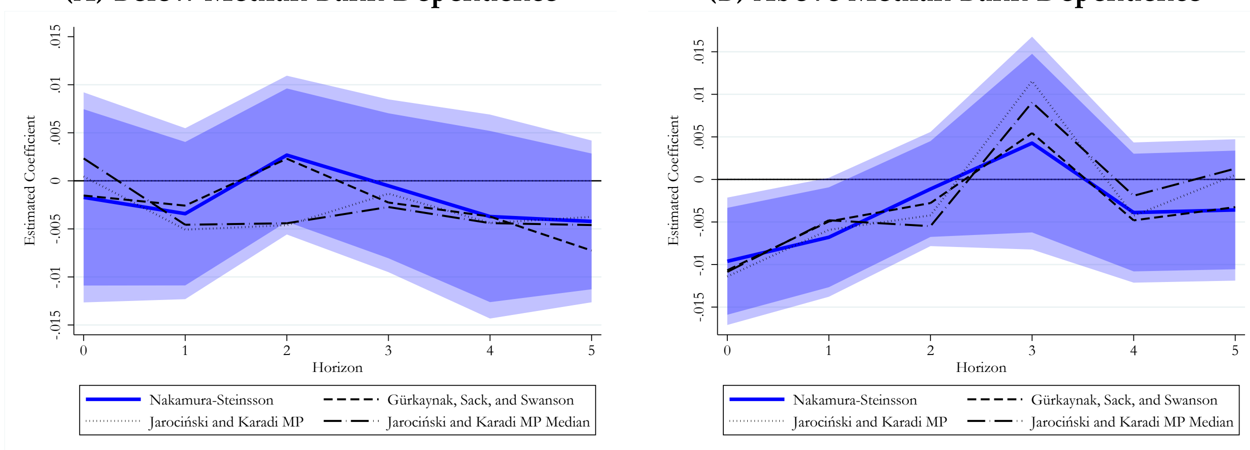
- Standard errors are clustered at the state level.

- All figures show the point estimates $\beta^h \, \forall \, h \in \{0, ..., 5\}$ alongside 95 percent and 90 percent confidence intervals.

In Figure C.1, we estimate (28) separately for states with values above and below the year-specific median of bank dependence across all states. Bank dependence is calculated by merging syndicated loans data from DealScan with Compustat and computing the share of total loan volume relative to total debt at the state-year level.

Figure C.1: Effects of Contractionary Monetary Policy and Wage Floors on Inequality in the US, by Bank Dependence

(A) Below Median Bank Dependence **(B) Above Median Bank Dependence**



Notes: The figure shows results from local projections, obtained by regressing *h*-step-ahead values of the natural logarithm of the state-level P90/P10 wage ratio on monetary policy shocks interacted with a state-level indicator for having an above-median minimum wage in place, as described in (28). The figure plots the estimated coefficients on this interaction term across the horizon, measured in years since impact, separately for states with values below (Panel A) and above (Panel B) the year-specific median of bank dependence across all states. Bank dependence is defined as the share of total syndicated loan volume over total debt at the state level. The monetary policy shocks are collapsed to annual frequency and standardized to mean zero and standard deviation one. We use the shocks of Nakamura and Steinsson (2018), monetary policy path shocks of Gürkaynak et al. (2005), and the monetary shocks obtained from simple sign restriction and median rotation in Jarociński and Karadi (2020). The plot shows point estimates alongside 90 percent and 95 percent confidence intervals, obtained from standard errors clustered at the state level. The confidence intervals are based on the estimated standard errors for the shock series of Nakamura and Steinsson (2018).

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