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 and employment. 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. Therefore, credit affects the distribution of pay and employment both within and between firms.

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1 Introduction

Credit supply is traditionally viewed as a macroeconomic stabilization lever of the monetary policy toolkit. Yet the possibility that credit supply can also have distributional consequences is relevant not just due to concerns about inequality. Rather, the effectiveness of monetary policy itself depends on the anatomy of its impact at the micro level. At the same time, there is scant empirical evidence on the individual-level effects of macroeconomic adjustments of credit supply in the context of labor market outcomes like employment and wages. This is perhaps surprising in light of labor income making up the lion's share of total income for the average person.[1]

In this paper, we empirically quantify the distributional effects of credit supply in the labor market. To this end, we build a novel dataset, which links worker employment histories to firm financials and banking relationships in Germany. We exploit as a persistent credit supply shock banks' lending reaction to the introduction of negative monetary policy rates by the European Central Bank (ECB). We show that firms in preexisting relationships with banks that were more exposed to negative rates see 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 this relative reduction in credit supply on workers within and between firms. Within firms, initially lower-paid workers are more likely to leave employment, while higher-paid workers see relative wage declines. Between firms, wages fall by more at initially higher-paying employers. Altogether, we find that a monetary policy-induced reduction in credit supply leads to lower wage inequality within and between firms.

To guide our empirical investigation, we develop a simple equilibrium model with both search and credit frictions. Both frictions are necessary to understand the distributional consequences of credit supply: absent search frictions workers all receive identical pay across firms, while absent credit frictions there is no role for credit supply to affect 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 a credit tightening causes wages to decline by more among high-skill workers and more so 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

[1]For example, estimates based on income tax data by the Internal Revenue Service (2020) in the U.S. show that 83 percent of all tax returns and 67 percent of tax filing amounts pertain to salary and wage income.

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 that resulted in lending cuts to their preexisting borrowers (Heider et al., 2019; Eggertsson et al., 2023; Ulate, 2021). Our empirical strategy exploits heterogeneity in preexisting bank-firm relationships around the introduction of negative rates as a source of variation in firm-level credit supply. A crucial question is how such a credit supply shock was passed through to labor markets, and whether there were differential effects across subgroups of workers and firms.

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, and not just 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 substitute for the drop in credit by switching to other banks or to alternative sources of debt financing, with more affected firms reducing their leverage and building cash reserves.

In the second step, we study the effects of this monetary policy-induced credit supply shock on worker-level labor market outcomes. A reduction in credit leads to lower *average* firm-level wages and employment. 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 around 1.3 percent, and an increase in the probability of leaving employment of around 0.2 percentage points. These estimates control for state time trends and worker-firm match heterogeneity, which would be confounded with changes in worker composition absent individual-level microdata.

These average effects mask important heterogeneity across workers *within* firms. To shed light on this, we estimate individual wage equations with controls for worker-firm match-specific and time-varying firm pay components. 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-quintile wages of around 0.8 percent relative to workers in the bottom quintile. Consequently, within-firm wage inequality decreases. At the same time, the probability of leaving employment increases by around 0.2 percentage points per standard deviation of exposure among workers in the bottom relative to the top quintile.

We then extend our analysis to firm-level aggregates net of changes in worker composition. In doing so, we explicitly take into account 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 less within-firm wage inequality than those in preexisting relationships with low-deposit banks following the introduction of negative rates. This holds also 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 escalate the discrepancy between the top and bottom echelons within firms, by bringing in data on compensation for executive board members. Doing so, we confirm the inequality-reducing treatment effect. Importantly, and in line 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 in variable compensation. In addition, we confirm at the firm level that firms in preexisting relationships with high-deposit banks reduce employment, and especially of nonmanagerial staff. As the latter tend to hail 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 state time trends. We find that among firms equally exposed to the credit supply shock, wages decline by more at initially higher-paying firms. Over four years, wages at top-ranked firms fall by 11 percent relative to bottom-ranked firms, while the probability of leaving employment is 2 percentage points higher at bottom-ranked compared to top-ranked firms. Consequently, between-firm wage inequality decreases.

In summary, we show that monetary policy-induced credit supply affects the distribution of pay and employment within and between firms. Therefore, our findings suggest that credit supply, firms, and earnings inequality are interlinked, in line with the predictions of an equilibrium model of frictional labor markets with credit constraints.

Related literature. We contribute to an emerging literature on the distributional consequences of credit and monetary policy. 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 pass-through of other firm-level shocks. Previous work has studied the pass-through of shocks to revenue productivity (Guiso et al., 2005; Fagereng et al., 2018; Garin and Silvério, 2018; Bagger et al., 2020; Friedrich et al., 2021; Lamadon et al., 2021; Chan et al., 2021), innovation (Van Reenen, 1996; Kline et al., 2019; Aghion et al., 2019; Kogan et al., 2021, 2022), cash (Howell and Brown, 2020), and taxes (Arulampalam et al., 2012; Suárez Serrato and Zidar, 2016; Fuest et al., 2018).

A unique aspect of our work is that by studying the link between credit supply and wage inequality, it puts a spotlight on the role of firm heterogeneity in the labor market. Studying firms' responses to credit is important because firms have been shown to be the natural unit of analysis when it comes to wage setting and employment decisions (Card et al., 2013; Song et al., 2019). Related work by Michelacci and Quadrini (2009) and Guiso et al. (2013) shows that credit matters for new hires' average starting wages and subsequent wage growth. 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.

The credit channel has been the focus of a tradition of work on the *aggregate* effects of monetary policy (Bernanke and Gertler, 1995). We study the credit channel 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 monetary policy to banks to firms to workers—in Germany. This allows us to shed light on distributional consequences of credit at the worker level that would remain hidden in more aggregate data due to changes in the underlying worker composition.

Numerous policymakers have expressed interest in the *distributional* effects of monetary policy (Bullard, 2014; Yellen, 2014; Bernanke, 2015; Lagarde, 2020). Existing research by Doepke and Schneider (2006), Coibion et al. (2017), Wong (2019), Holm et al. (2021), Andersen et al. (2021), and Amberg et al. (2021) has linked monetary policy to *household* balance sheets and inequality. In contrast, our paper identifies monetary policy-induced *firm* credit as a source of wage inequality. This complements studies of the effects of credit on employment (Chodorow-Reich, 2014; Jiménez et al., 2017; Berton et al., 2018; Benmelech et al., 2019; Caggese et al., 2019; Barbosa et al., 2020; Bergman et al., 2021; Coglianese et al., 2021).[2] Relatedly, Broer et al. (2021) quantify the effects of monetary policy shocks on workers' job-finding and separation rates. While we confirm significant employment responses to credit, the effect of credit on wages matters for the vast majority of workers who remain employed. Consequently, it matters to what extent different workers are heterogeneously

[2]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), firm dynamics (Acabbi et al., 2020), organizational structure (Sforza, 2020), and aggregate output (Herreño, 2020).

affected by credit supply. Parallel work has been interested in the response of wages to firm-level credit shocks (Fonseca and Van Doornik, 2020; Adamopoulou et al., 2020; Arabzadeh et al., 2020; Jasova et al., 2021). A distinguishing feature of our analysis is its focus on the distributional effects of monetary policy-induced credit supply on pay within and between firms.

2 Equilibrium Model

There exists no benchmark model of how credit affects worker and firm heterogeneity in pay and employment. Therefore, the purpose of the following model is to provide a conceptual framework that links credit supply to the distribution of wages and employment within and between firms. To this end, we model multi-worker firms subject to frictions in both credit and labor markets. Credit frictions imply that firms finance working capital, including their wage bill and recruiting costs, subject to idiosyncratic borrowing constraints (Neumeyer and Perri, 2005; Jermann and Quadrini, 2012). We model the bank lending channel of monetary policy as affecting firms' idiosyncratic credit constraints, while—for simplicity—abstracting from other channels through which monetary policy affects the real economy. Labor market frictions imply that identical workers receive different pay across employers (Card et al., 2013; Song et al., 2019). Unlike existing models of match heterogeneity subject to credit and labor market frictions (Wasmer and Weil, 2004; Kehoe et al., 2019, 2020), we allow for multi-worker firms as in the seminal Burdett and Mortensen (1998) framework. We tractably extend this framework to include worker heterogeneity in skills and firm heterogeneity in credit constraints.

2.1 Workers

Workers are infinitely lived, risk neutral, and discount time at rate ρ. They differ ex ante in skill $a \in \{a_L, a_H\}$. We assume $0 < a_L < a_H$ and refer to worker types 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 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 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

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

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

Firms differ ex ante in their productivity $p > 0$ and credit limit $\xi > 0$, with $j = (p, \xi) \overset{c}{\sim} \Gamma(j)$.

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, defined as the sum of their wage bill $\sum_a w_a l_a(w_a, v_a)$ and recruiting costs $\sum_a c_a(v_a)$. Given interest rate $r > 0$, firms face idiosyncratic credit limits 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

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

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

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 $\psi_j > 0$ for constrained firms. Firms are more credit constrained if, all else equal, they have higher productivity p_j, which leads to higher labor demand, or a lower credit limit ξ_j.

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 $\psi_j > 0$. Firms facing a tighter credit constraint, as measured by a higher ψ_j, 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 characterizes the effect of credit supply on the distribution of wages and employment within and between firms across steady states of the economy.

Proposition 1 (Effects of credit on distribution of wages and employment)**.** *Suppose that credit constraints bind across firms. For any firm j, a decrease in the idiosyncratic credit limit ξ_j leads to*

(i) lower firm-level wages for identical workers,

(ii) lower firm-level employment,

(iii) lower within-firm wage inequality through a relatively greater reduction in wages among initially high-paid workers, and

(iv) lower between-firm wage inequality through a relatively greater reduction in wages at initially high-paying firms.

Proof. See Appendix B.2. □

Intuitively, Proposition 1 states that a tighter credit constraint lowers a firm's effective productivity, leading to a reduction in relative wages of high-paid workers. Thus, our simple equilibrium model has sharp predictions for the effect of credit supply on the distribution of wages and employment within and between firms. 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 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

[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 existing workers being fired or adjust slowly over time by new hires being reduced following worker separations.

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). However, without a dominant dimension of heterogeneity across banks that gives rise to differential lending responses, it is difficult to empirically disentangle fluctuations in credit supply due to changes in positive monetary policy rates from other factors, such as credit demand (Nakamura and Steinsson, 2018).

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., 2023; Ulate, 2021). The second channel is due to the effect of negative rates on banks' net worth or equity 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 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 $\textit{Deposit ratio}_j$ de-

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

note 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 $\mathit{After}(2014)_t$ denote a dummy variable for the years 2014–2017. In a difference-in-differences setting, we define as our credit supply shock

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

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 in the labor market, 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 effects. While the credit supply shock is at the firm-year level, we study individual wages and employment 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:

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

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, $\mathit{Deposit\ ratio}_j$ 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 anytime from 2010 to 2013, $\mathit{After}(2014)_t$ is a dummy variable for the years 2014–2017, and θ_{ij} and $\zeta_{s(j)t}$ denote, respectively, worker-firm and state-year fixed effects corresponding to state $s(j)$ that firm j is located in. 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 response of y_{ijt} 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 state-year fixed effects, we absorb 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:

$$\begin{aligned}
y_{ijt} = {} & \beta_1 \mathit{Deposit\ ratio}_j \times \mathit{After}(2014)_t \times \mathit{Bottom}\ 20\%\ \mathit{within\ firm}_i \\
& + \beta_2 \mathit{Deposit\ ratio}_j \times \mathit{After}(2014)_t \times \mathit{Middle}\ 60\%\ \mathit{within\ firm}_i \\
& + \beta_3 \mathit{Deposit\ ratio}_j \times \mathit{Bottom}\ 20\%\ \mathit{within\ firm}_i + \beta_4 \mathit{Deposit\ ratio}_j \times \mathit{Middle}\ 60\%\ \mathit{within\ firm}_i \\
& + \beta_5 \mathit{After}(2014)_t \times \mathit{Bottom}\ 20\%\ \mathit{within\ firm}_i + \beta_6 \mathit{After}(2014)_t \times \mathit{Middle}\ 60\%\ \mathit{within\ firm}_i \\
& + \theta_{ij} + \eta_{jt} + \varepsilon_{ijt},
\end{aligned} \tag{9}$$

where *Bottom* 20% *within firm*$_i$ (*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 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.

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 20 percent and middle 60 percent of the wage distribution relative to workers in the top 20 percent, 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 the set of controls in (8), 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:

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

where *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 $\zeta_{s(j)t}$ denote, respectively, worker-firm and state-year fixed effects corresponding to state $s(j)$ that firm j is located in. 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.

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 financials data that comprise private and public firms' balance sheet information. 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 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 data at the firm level.

Board compensation (BoardEx). We supplement the IAB 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 collected by *Creditreform* (Huber, 2018). These data identify private and public 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.[5] Two-thirds of all firms indicate to be in a relationship with a single bank. Only 9% of all German firms report to be in a relationship with more than two banks.

[5] 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.

Table 1: Summary Statistics

Variable	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)	37,294	18,541	8,317	35,249	70,949	72,130,131
Unemployed next year	0.096	0.294	0	0	1	66,250,135
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	47.793	752.164	1.500	11	135	2,810,558
No. of nonmanagerial employees	44.812	692.138	1	10	127	2,810,558
No. of part-time employees	16.442	277.181	0	3	44	2,810,558

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. 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.

Wages are defined as the mean (log) daily earnings of full-time employees as reported in the IAB linked employer-employee data.[6] 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.

We use data from years 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 keep the main job j, which we define 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, *Deposit ratio*$_j$.

Our final sample covers approximately 36 percent of all full-time workers in Germany, which

[6] We separately study part-time versus full-time employment shares as an outcome in our firm-level analysis.

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

Variable	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 employees	27.634	497.686	1.5	9	78	88,899
Average annualized wage (euros)	27,530	11,290	11,728	26,118	48,483	88,899
Proportion female	0.252	0.320	0.000	0.111	1	88,899
Proportion foreigner	0.070	0.183	0.000	0.000	0.500	88,899
Proportion university	0.110	0.236	0.000	0.000	0.700	88,899
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 employees	74.729	990.003	1	12	219	87,150
Average annualized wage (euros)	33,116	13,989	12,642	31,490	58,499	87,150
Proportion female	0.297	0.317	0.000	0.200	1	87,150
Proportion foreigner	0.080	0.185	0.000	0.000	0.500	87,150
Proportion university	0.191	0.287	0.000	0.035	1	87,150
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 *Deposit ratio*$_j$, which is the average deposit ratio, measured in 2013, across all (typically German) banks that firm j reports to be in a banking relationship with anytime from 2010 to 2013.

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 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 47.8 full-time employees, the firm size distribution is positively skewed and fat-tailed. The mean number of nonmanagerial employees that work full-time is 44.8, while the mean number of part-time employees is 16.4.

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 workforce, 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 a mean of 27.6 employees, compared to 74.7 employees at firms in relationships with low-deposit banks. Mean pay at firms in relationships with high-deposit banks is 27,530 euros, a bit less than that at firms in relationships with low-deposit banks, which is 33,116 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 permanent as well as time-varying (unobserved) employer differences, subsuming any firm-specific trends.

5 Results

5.1 Effect of Negative Interest Rates on Credit Supply

We start out by estimating the extent to which German firms in relationships with high-deposit, rather than low-deposit, banks see 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' credit access. We focus on banks that act as lead arrangers in the syndication process. Lead arrangers are those members of a syndicate that are 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 be-

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
Variable	(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 the first three columns, the dependent variable is an indicator for any nonzero share of firm j's loans held by bank k in t. In the last three columns, 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.

tween 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. (2023), 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, *Deposit ratio*$_k$ is bank k's deposits-to-assets ratio, measured in 2013, *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 4: Impact of Negative Policy Rates on Lending to German Firms

Variable	Any loan share $\in \{0,1\}$ (1)	(2)	$\ln(1+\text{total loan volume})$ (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 first two columns, the dependent variable is an indicator for any nonzero share of firm j's loans held by bank k in t. In the last two columns, 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 (typically German) banks that firm j reports to be in a banking relationship with anytime 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.

Table 3 presents the results of estimating (11). In columns 1–2, the dependent variable is an indicator for any non-zero share of firm j's syndicated loans held by bank k in t. In the first column, we find that high-deposit banks reduce their credit supply after the introduction of negative rates. Using these estimates, a one standard deviation increase in bank-level deposit ratios implies a lower 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., their reliance on deposits) 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 on negative rates to most of their (typically household) depositors (Heider et al., 2019; Eggertsson et al., 2023; Heider et al., 2021).[7]

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

[7] In this sense, our analysis complements related work on the reversal interest rate (Brunnermeier and Koby, 2018).

[8] Our results are robust to using the mean deposit ratio over a preperiod that concludes before 2012, likely reflecting

that high-deposit banks start reducing their credit supply 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 around the introduction of negative rates in June 2014. This reduces the likelihood of other bank-level events, including other monetary policy decisions, interfering 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 by the natural logarithm of one 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 information on each lead bank's share from DealScan to compute each lead bank's total loan amount granted to a firm in a given time period.[9]

Our results indicate that high-deposit banks reduce their credit supply 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. (2023). In the next step, we establish that firms in existing relationships with affected high-deposit banks do not only receive less credit but also cannot perfectly substitute for the drop in credit by switching to other banks. For this purpose, we extend our balanced panel so as to include all lending relationships, including those with banks outside of the euro area, from 2010 to 2017. Furthermore, we replace $\textit{Deposit ratio}_k$ in (11) by $\textit{Deposit ratio}_j$, 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.

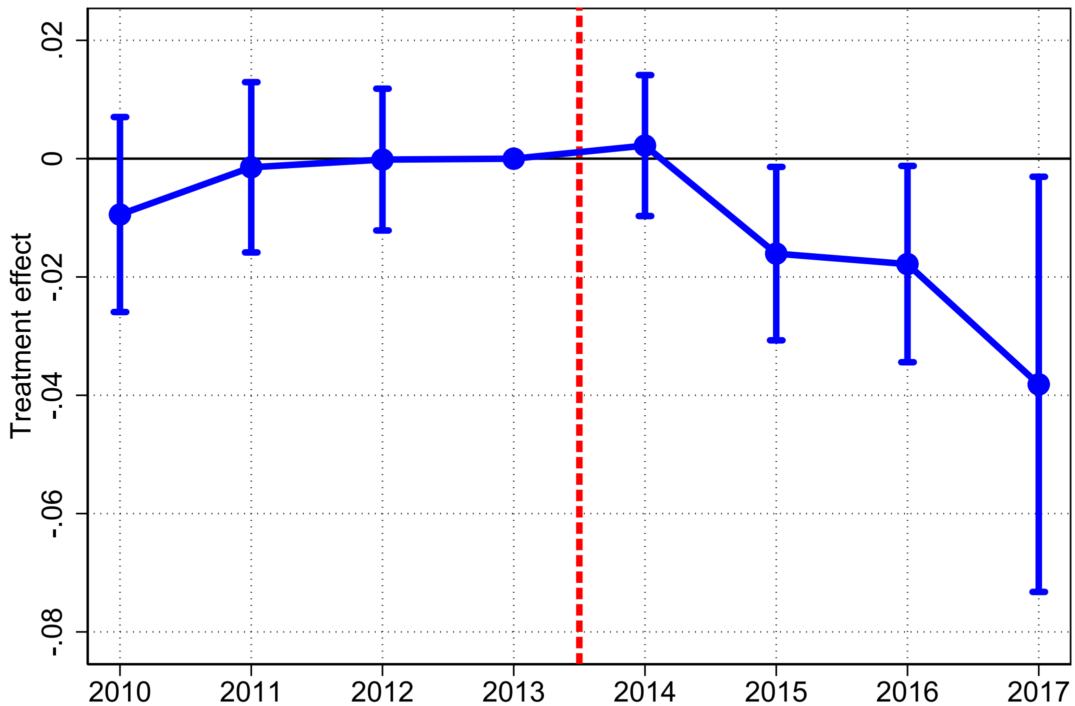
The results in Table 4 show a significant reduction in credit for firms in relationships with high-deposit banks, regardless of the bank from which they obtain credit after the introduction of negative rates in 2014. In the first column, we include only bank-firm and time fixed effects, and find that a one standard deviation increase in $\textit{Deposit ratio}_j$ is associated with a $0.127 \times 0.084 = 1.1$ percentage points lower likelihood of attaining any loan.

This estimate becomes even larger in the second column, 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 off firms in relationships with the

the stability of firms' relationship banks in our sample.

[9]Whenever available, we use loan shares as reported in DealScan. Otherwise, similar to 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 in the syndicate.

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



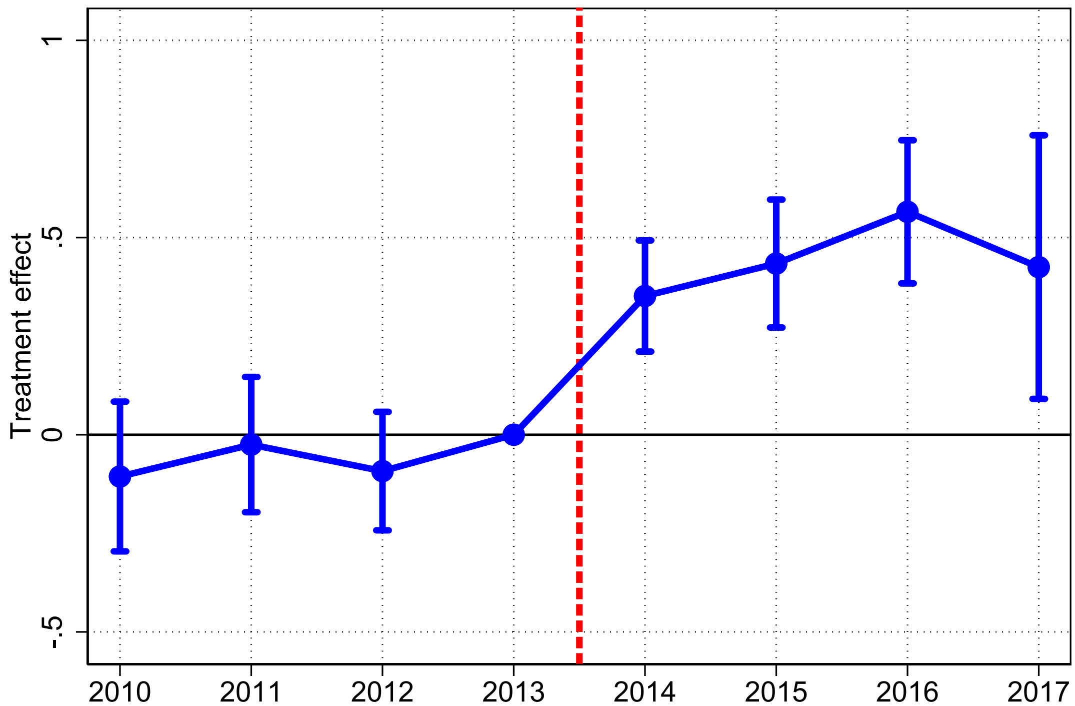
Notes: This figure plots estimates of β_τ, alongside 90 percent confidence bands, over time (each year represents the respective year-end) based on the event study specification in (12), using as dependent variable firm j's leverage ratio, estimated on the sample of German firms in the administrative linked employer-employee data merged with Orbis from 2010 to 2017.

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 by the actual loan amounts granted by lead banks through syndication in columns 3–4.

These findings imply that firms in relationships with high-deposit banks, on average, receive less credit from *any* bank, including those outside of 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, the latter cannot readily compensate for a loss in 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:

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

Figure 2: Impact of Negative Policy Rates on Firms' Cash Position



Notes: This figure plots estimates of β_τ, alongside 90 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 total cash and cash equivalents, estimated on the sample of German firms in the administrative linked employer-employee data merged with Amadeus from 2010 to 2017.

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

Figure 1 plots estimates of β_τ from equation (12) alongside 90 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 (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 10 percent level starting with the first full year of negative rates in 2015. As a result, firms experience what corresponds to a

Table 5: Effects of Monetary Policy-Induced Credit Supply on Wages and Employment

		ln(wage)			Unemployed next year $\in \{0,1\}$	
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio $\times$ After(2014)	-0.019**	-0.077***	-0.083***	0.007**	0.011***	0.013***
	(0.009)	(0.010)	(0.009)	(0.003)	(0.004)	(0.004)
Worker FE	Y	N	N	Y	N	N
Firm FE	Y	N	N	Y	N	N
Worker-firm FE	N	Y	Y	N	Y	Y
Year FE	Y	Y	N	Y	Y	N
State-year FE	N	N	Y	N	N	Y
N	70,137,681	67,731,621	67,722,380	65,253,153	63,505,552	63,495,556

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 (typically German) banks that firm j reports to be in a banking relationship with anytime from 2010 to 2013. *After*(2014)$_t$ is a dummy variable for the years 2014–2017. 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.

tightening of their idiosyncratic credit constraint ξ_j in our theoretical model in Section 2.

Supporting our interpretation of the introduction of negative rates as a credit supply shock, Figure 2 shows that firms in relationships with high-deposit banks hoard significantly more cash following the introduction of negative rates. This is in line with an increased self-financing motive (Almeida et al., 2004). What is more, across both outcomes, firms in relationships with banks of different deposit reliance exhibit parallel trends during the preperiod, which supports our identification assumption for the effects of credit supply.

5.2 Effects on the Distribution of Wages and Employment

So far, we have established that firms in relationships with high-deposit banks receive less credit, both within preexisting relationships and also from other banks and external debt financing sources. Our next goal is to estimate the effects of this credit supply shock on the distribution of wages and employment in our worker-level data. In doing so, we pursue the hypothesis that credit supply has distributional effects across workers and firms in the labor market.

Mean effects. We first look at the effects of credit supply on mean wages and unemployment. Table 5 shows results from estimating variants of equation (8) using either worker i's log wage at firm j or her employment status the following year 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) and higher unemployment risk (column 4).

Columns 2 and 5 show that the effects on wages and unemployment become stronger when including worker-firm match effects. Here, the coefficient of interest, β, is estimated off workers that were either employed at the same firm before and after 2014, or leave employment after 2014. Based on these estimates, 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 and a $0.153 \times 0.011 = 0.2$ percentage points increase in the probability of becoming unemployed. These estimates become somewhat larger when we include state-year fixed effects in columns 3 and 6, which additionally controls for time-varying unobserved regional heterogeneity.

In summary, these findings are in line with parts (i)–(ii) of Proposition 1 of our theoretical model.

Within-firm heterogeneity. These estimated mean effects on wages and employment may mask important heterogeneity across worker groups within firms. 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.

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

We find that individuals that used to earn a wage in the bottom quintile of their respective firms' wage distributions see their wages grow more at more exposed firms after the introduction of negative rates than the top quintile (i.e., the omitted category). This result remains robust after adding worker-firm fixed effects in column 3, and is consistent with part (iii) of Proposition 1 of our theoretical model. A one standard deviation increase in firms' exposure as captured by

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

	ln(wage)			Unemployed next year $\in \{0,1\}$		
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio $\times$ After(2014) $\times$ Bottom 20% within firm	0.034*	0.069***	0.051***	0.009**	0.004	0.013***
	(0.018)	(0.019)	(0.017)	(0.004)	(0.004)	(0.004)
Deposit ratio $\times$ After(2014) $\times$ Middle 60% within firm	-0.017**	-0.012*	-0.014**	0.018***	0.016***	0.019***
	(0.007)	(0.007)	(0.007)	(0.002)	(0.002)	(0.002)
Deposit ratio $\times$ After(2014)	-0.008			-0.008**		
	(0.007)			(0.003)		
Deposit ratio $\times$ Bottom 20% within firm	-0.136***	-0.142***		0.004	0.009**	
	(0.021)	(0.018)		(0.004)	(0.004)	
Deposit ratio $\times$ Middle 60% within firm	-0.112***	-0.106***		0.001	0.003	
	(0.015)	(0.013)		(0.003)	(0.003)	
After(2014) $\times$ Bottom 20% within firm	0.154***	0.141***	0.071***	0.029***	0.032***	0.050***
	(0.013)	(0.013)	(0.011)	(0.002)	(0.002)	(0.003)
After(2014) $\times$ Middle 60% within firm	0.010**	0.007	-0.011**	-0.005***	-0.001	0.000
	(0.004)	(0.005)	(0.005)	(0.002)	(0.001)	(0.002)
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	61,987,235	61,519,347	59,839,079	58,204,386	57,773,587	56,308,377

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 (typically German) banks that firm j reports to be in a banking relationship with anytime 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.

Deposit ratio$_j$ translates into a $0.153 \times 0.051 = 0.8$ percent relative reduction in wages of workers in the top quintile versus those in the bottom quintile of the within-firm wage distribution. Since the coefficient of interest for the wage regression is now estimated off workers who stay at the same employer before and after the introduction of negative rates, these results are driven by wage effects on incumbents rather than new hires.

In columns 4–6, we estimate specification (9) with the dependent variable replaced by an indicator for whether worker i is unemployed in year $t+1$. We find significant unemployment effects for workers in the middle 60 percent of the within-firm wage distribution across all three specifications. In column 4 and column 6, when including worker-firm match fixed effects, we find that all workers outside of the top quintile of the within-firm wage distribution face higher risk of being laid off following the negative credit supply shock. Quantitatively, the additional risk of leaving employment for workers below the top quintile of the within-firm wage distribution amounts to between $0.153 \times 0.013 = 0.2$ and $0.153 \times 0.019 = 0.3$ percentage points based on our preferred specification in column 6.

The empirical observation that wages are more rigid for lower-paid workers may partly reflect that Germany introduced a federal minimum wage 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 the tightening of their credit constraint.

As alluded to earlier, the German administrative data on earnings are winsorized around the 90^{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, as reflected in the fact that their winsorized share of worker-years is 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 plus differential firm-level wage growth due to firms' differential exposure to the credit supply shock. The aggregate wage growth component by itself pushes more employees at firms in relationships with 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, namely that within-firm inequality declines by more at firms in relationships with high-deposit banks. This suggests that our results are not driven by this statistical artifact.

Firm-level aggregation. In our worker-level analysis, 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. Throughout this analysis, we have been holding constant worker composition by including worker fixed effects or worker-firm match fixed effects. Of independent interest are outcomes aggregated to the firm level, which we now turn to. In doing so, we explicitly take account of changes in worker composition due to hiring and separations.

To this end, we construct measures of within-firm wage inequality for all firms in each year. We then estimate the following specification at the firm-year level:

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

where y_{jt} is a measure of within-firm pay inequality for firm j in year t, ψ_j denotes firm fixed effects, and $\zeta_{s(j)t}$ are state-year fixed effects corresponding to state $s(j)$ that firm j is located in.

Table 7 presents the results from estimating specification (13) for different inequality measures and different samples in our data. Columns 1–3 take as dependent variable y_{jt} the log P90-P10 wage percentile ratio. All three specifications include firm and state-year fixed effects, thereby controlling for time-invariant firm-specific and time-varying regional heterogeneity. Column 1, which includes all firms in our sample, indicates a modest reduction in within-firm wage inequality at more affected firms, with a coefficient estimate of -0.013 (standard error of 0.006). This is consistent with our worker-level finding of relative wage declines among higher pay ranks within firms, as in Table 6.

Motivated by evidence that larger, publicly listed firms may exhibit greater within-firm wage inequality (Mueller et al., 2017), we estimate the same regression specification separately for public firms in column 2. In doing so, we find that the reduction in within-firm inequality due to the negative credit shock is even more emphasized for firms in this small subsample.

One advantage of using this subsample is that it comprises firms that are large and covered also in our syndicated loans data from DealScan, which we have used in Tables 3 and 4. Those

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

	ln(P90/P10)	ln(P90/P10)	ln(P90/P10)	ln(P50 board total/P5)	ln(P50 board salary/P5)	ln(P50 board bonus/p5)
Sample	All	Public firms	Public firms	DAX firms	DAX firms	DAX firms
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio $\times$ After(2014)	-0.013**	-0.373**	-0.510***	-0.877*	-0.696	-0.888*
	(0.006)	(0.160)	(0.183)	(0.485)	(0.456)	(0.532)
Non-euro deposit ratio $\times$ After(2014)			-0.029			
			(0.117)			
Firm FE	Y	Y	Y	Y	Y	Y
State-year FE	Y	Y	Y	N	N	N
Year FE	N	N	N	Y	Y	Y
N	2,771,902	1,324	1,149	266	266	263

Notes: The unit of observation is the firm-year level jt. In column 1, the sample consists of all German corporations j in year t from 2010 to 2017. In columns 2 and 3, 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. In the last three columns, 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. In the first three 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. The dependent variable in column 4 is the delta log of the median total compensation, consisting of a salary and a potential bonus, of executive board members at firm j in year t versus the wage at the 5^{th} percentile of firm j's wage distribution in year t. The dependent variable in column 5 is the delta log of the median salary of executive board members at firm j in year t versus the wage at the 5^{th} percentile of firm j's wage distribution in year t. The dependent variable in column 6 is the delta log of the median bonus (conditional on being nonzero) of executive board members at firm j in year t versus the wage at the 5^{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 (typically German) banks that firm j reports to be in a banking relationship with anytime 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 anytime from 2010 to 2013. *After*(2014)$_t$ is an indicator variable for the years 2014–2017 in the first three columns (2014–2016 in all remaining columns). 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.

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

Variable	ln(no. of all employees) (1)	ln(no. of nonmanagerial employees) (2)	Share nonmanagerial (3)	Share part-time (4)
Deposit ratio × After(2014)	-0.015***	-0.021***	-0.006***	-0.011***
	(0.005)	(0.005)	(0.001)	(0.001)
Firm FE	Y	Y	Y	Y
State-year FE	Y	Y	Y	Y
N	2,803,152	2,803,152	2,803,152	2,803,152

Notes: The unit of observation is the firm-year level jt. In the first four columns, the sample consists of all German corporations j in year t from 2010 to 2017. The dependent variable in column 1 is the natural logarithm of the total number of employees at firm j in year t. The dependent variable in column 2 is the natural logarithm of the number of nonmanagerial employees at firm j in year t. The dependent variable in column 3 is the ratio, between 0 and 1, of nonmanagerial staff over all employees at firm j in year t. The dependent variable in column 4 is the ratio, between 0 and 1, of part-time staff over all employees at firm j in year t. $\textit{Deposit ratio}_j \in [0,1]$ is the average deposits-to-assets ratio, measured in 2013, across all (typically German) banks that firm j reports to be in a banking relationship with anytime from 2010 to 2013. $\textit{After}(2014)_t$ is an indicator variable for the years 2014–2017. 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.

firms are likely to receive syndicated loans not only from German and other euro-area banks, but also from non-euro area banks whose supply of credit should not be affected by monetary policy in the euro area. This enables us to conduct a falsification test in column 3 by adding an interaction term between $After(2014)_t$ and $\textit{Non-euro deposit ratio}_j \in [0, 1]$, which is the mean deposit ratio across all non-euro area lead arrangers (and other banks not based in negative-rate currency areas) that firm j received a syndicated loan from in the preperiod from 2010 to 2013. Reassuringly, we find that the coefficient on the placebo term is close to zero and statistically insignificant.

While rich in many dimensions, the IAB linked employer-employee data do not allow us to measure top-wage inequality due to the data being winsorized at the social security contribution threshold, which falls around the 90^{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 of the distribution is by adjusting variable compensation.

To test for this adjustment mechanism, we use information on compensation for executive board members of 26 of the DAX-listed firms from BoardEx.[10] 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 loan market. In columns 4–6 of Table 7, we provide small-sample evidence that a negative credit supply shock is associated with a reduction of top-to-bottom wage inequality within said listed firms. Column 4 shows a point estimate that is large and negative but noisily estimated and barely significant at the 10 percent level. Splitting board pay further into salary and bonus pay, we find a significant negative reduction in bonus (column 6), but not in salary (column 5). This suggests that firms take into account the availability of credit, with associated future growth prospects, when reducing top-earners' variable compensation due to tighter financial constraints.

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 take into account 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 firm and state-year fixed effects. Column 1 shows that firms more exposed to negative rates see a

[10] Since in Germany some company board positions are allocated to worker representatives and other nonexecutives (Jäger et al., 2021), we drop these from our data. For nonexecutive board members, who typically do not receive substantial variable compensation, we find no significant response in their relative pay—see Table A.1 in Appendix A.

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

	ln(wage)			Unemployed next year $\in \{0,1\}$		
Variable	(1)	(2)	(3)	(4)	(5)	(6)
Deposit ratio $\times$ After(2014) $\times$ Firm pay rank	-0.050	-0.137***	-0.113***	-0.028***	-0.009	-0.020**
	(0.037)	(0.031)	(0.029)	(0.009)	(0.010)	(0.010)
Deposit ratio $\times$ After(2014)	-0.017	0.060***	0.045**	0.002	-0.017***	-0.010*
	(0.024)	(0.019)	(0.019)	(0.005)	(0.006)	(0.006)
After(2014) $\times$ Firm pay rank	-0.034	0.173***	0.177***	-0.033***	-0.065***	-0.066***
	(0.028)	(0.023)	(0.021)	(0.006)	(0.007)	(0.007)
Worker FE	Y	N	N	Y	N	N
Firm FE	Y	N	N	Y	N	N
Worker-firm FE	N	Y	Y	N	Y	Y
Year FE	Y	Y	N	Y	Y	N
State-year FE	N	N	Y	N	N	Y
N	69,627,349	67,372,241	67,363,297	64,700,521	63,076,967	63,067,608

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 (typically German) banks that firm j reports to be in a banking relationship with anytime 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. 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.

significant reduction in overall employment. We estimate a coefficient of -0.015 (standard error of 0.005), suggesting that a one standard deviation increase in firm-level exposure is associated with a $0.153 \times 0.015 = 0.2$ percent reduction in total employment.

Consistent with our mean worker-level effects in columns 4–6 of Table 6, column 2 shows that this effect is 40 percent larger for nonmanagerial employees, which tend to hail from the bottom of the wage distribution. Column 3 shows that, as a result, more exposed firms see a significant reduction in their share of nonmanagerial workers. Finally, column 4 shows that the negative credit supply shock is also associated with a reduction of part-time work, suggesting that those workers are more likely to leave employment or else are asked to work extra hours.

Between-firm heterogeneity. While we have shown that the credit supply shock due to negative rates led to lower wages on average, we now address the extent to which different firms adjusted wages differentially. 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. Column 1, which includes worker, firm, and year fixed effects, shows that initially higher-paying firms respond to the negative credit supply shock with a relative reduction in wages, but the estimated coefficient falls short of being statistically significant at conventional levels. After including worker-firm fixed effects and therefore focusing on incumbent workers in column 2, the coefficient almost triples and becomes significant, suggesting that changes in worker composition are an important margin of adjustment. This continues to hold true in column 3 after replacing year fixed effects by more granular state-year fixed effects of the respective firms.

Columns 4–6 test for differential unemployment effects across firm pay ranks. To this end, we replace the dependent variable by an indicator for whether a worker will be unemployed the next year. Column 4 shows a negative and significant estimate of the interaction coefficient of -0.028 (standard error of 0.009). In our preferred specification with worker-firm and state-year fixed effects in column 6, the coefficient is still negative and statistically significant, but it is insignificant in column 5 when using year fixed effects instead of state-year fixed effects.

Our interpretation of these findings builds on our evidence in Table 7, 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 credit supply by decreasing their pay by relatively more.

As argued above, the mechanical effect of winsorization goes *against* our finding that between-firm inequality declines due to the credit supply shock. This is because initially higher-paying firms have, all else equal, 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 administer relative wage cuts while at the same time retaining weakly more 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 Conclusion

We empirically study the effects of monetary policy-induced credit supply on the distribution of wages and employment within and between firms. To this end, we build a novel dataset that spans the complete credit chain—from banks to firms to workers—in Germany. We find that firms in relationships with more deposit-reliant banks see a relative reduction in credit supply following the ECB's introduction of negative rates in June 2014. Lower credit in turn leads to a negative effect on firm-level wages and employment. Within firms, initially lower-paid workers are more likely to leave employment, while initially higher-paid workers receive lower relative wages. Between firms, wages decline by more at initially high-paying firms. In this way, credit affects the distribution of pay and employment within and between firms.

These findings are relevant in light of recent models emphasizing the importance of "unequal incidence" of aggregate fluctuations as a key determinant for monetary policy propagation (Kaplan et al., 2018). They can be rationalized through the lens of an equilibrium model of frictional labor markets with credit constraints. We present such a framework that allows us to link firm pay heterogeneity to frictions in the labor and credit markets. Consistent with the predictions from this framework, our empirical results point toward firm heterogeneity as a relevant transmission mechanism for the real effects of credit in the labor market.

Our analysis suggests several interesting directions for future work. First, while our analysis focuses exclusively on workers' wages and employment, a natural extension of our analysis could explore other margins of adjustment to credit supply, including firms' technology choices and workers' investment in human capital. Second, while we restrict attention to a particular episode of negative interest rates in Germany, it would be valuable to study other instances of conventional and unconventional monetary policy. Third and finally, the effects of credit in our study are estimated off a relatively short time window since June 2014. Understanding the medium- and long-term effects of monetary policy and credit supply through the channels highlighted in our work is deserving of further investigation.

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Online Appendix—Not for Publication

A Supplementary Tables

Table A.1: Effects of Monetary Policy-Induced Credit Supply Shock on Within-Firm Inequality: Nonexecutive Board Members

	ln(p50 board total/p5)	ln(p50 board salary/p5)	ln(p50 board bonus/p5)
Sample	DAX firms	DAX firms	DAX firms
Variable	(1)	(2)	(3)
Deposit ratio × After(2014)	-0.311	0.097	-0.295
	(0.548)	(0.577)	(1.450)
Firm FE	Y	Y	Y
Year FE	Y	Y	Y
N	266	266	105

Notes: The unit of observation is the firm-year level jt. In column 1, the sample consists of all German corporations j in year t from 2010 to 2017. 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 column 1 is the delta log of the median total compensation of nonexecutive board members at firm j in year t versus the annualized wage at the 5$^{\text{th}}$ percentile of firm j's wage distribution in year t. The dependent variable in column 2 is the delta log of the median salary of nonexecutive board members at firm j in year t versus the annualized wage at the 5$^{\text{th}}$ percentile of firm j's wage distribution in year t. The dependent variable in column 3 is the delta log of the median bonus (conditional on being nonzero) of nonexecutive board members at firm j in year t versus the annualized wage at the 5$^{\text{th}}$ 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 (typically German) banks that firm j reports to be in a banking relationship with anytime from 2010 to 2013. *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*

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

  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

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

  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*

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

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

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

- *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

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

where ψ is the Lagrange multiplier on a firm's credit constraint. From here, the proof follows closely that in Morchio and Moser (2020), which we adapt to our setting.

B.2.1 Part (a)

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

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

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

where $T_a = \mu_a[(u_a + s_a^G)\lambda_a^u(\delta_a + \lambda_a^G + \lambda_a^e)]/V_a$. 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.

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

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

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. That optimal wages w_a^* are strictly increasing in productivity p and strictly decreasing (constant) in the Lagrange multiplier on the credit limit ψ for workers of high (low) ability follows from the definition of $\tilde{p}$ in equation (14) above. $\square$

B.2.2 Part (b)

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. Under the assumption that firms maximize long-run profits, the value of employing a worker is simply

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

The acceptance probability for a firm offering w 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, expected profits per contacted worker are

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

Then the firm's problem becomes

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

Therefore, the optimal wage and vacancy policy functions satisfy

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

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

B.2.3 Part (c)

Proof. The proof follows by combining the result in part (a) of Proposition 1 with the fact that wages for low-ability workers are equal to the constant flow value of unemployment. Specifically, by part (a), 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 (d)

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 composive 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 the 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 (a) of Proposition 1. Therefore,

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

for workers of any ability level a, with strict inequality for workers with high ability $a = a_H$. $\square$

Appendix References

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