The Macroeconomics of Establishment-Level Employment Dynamics

by

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To José and Debra for their never ending support.

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ABSTRACT

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This dissertation comprises four chapters examining the macroeconomics of establishment-level employment dynamics. “Wage Rigidity and Employment Outcomes: Evidence from Administrative Data,” examines the relationship between wage rigidity and employment outcomes using employer-employee data. The estimates suggest wage rigidity prevents 24.5 percent of wage cuts that would have occurred in absence of wage rigidity. An establishment with the average level of wage rigidity is predicted to have a 0.7 percentage point increase in the layoff rate, a 1.8 and 1.3 percentage point reduction in the quit and hire rates.

“Wage Rigidity and Employment Outcomes: Theory” derives the analytics of a structural model, showing that wage rigidity works through two channels. First, a cost to cutting wages deters establishments from reducing the current wages, increasing layoffs and reducing hires to lower the wage bill in the face of an adverse shock. Second, forward-looking establishments reduce the cost of wage cuts in future periods through dampening wage increases in the current period. Counterfactual policy simulations suggest that inflation can “grease the wheels” of the labor market by facilitating real wage cuts when nominal wage cuts are costly to achieve.

“Wage Rigidity and Employment Outcomes: The Role of Establishment Size and Age” shows that, on average, large establishments have more wage rigidity than small establishments and old establishments have more wage rigidity than young establishments. Estimates show, however, that wage rigidity and employment has a stronger relationship in younger establishments compared to older establishments.

“Bank Balance Sheet Shocks and the Effects of the Financial Crisis” uses employer-employee data from Germany to examine the impact of exogenous shocks to bank capital. German regional banks’ trading losses from U.S. mortgage-backed securities cause a deep economic contraction in the banks exclusive geographic domain. Annual loan and output growth decline by 20 and 0.3 percentage points, respectively, and the annual unemployment rate rises by 1.4 percentage points in affected states compared to unaffected states, on average. The effect is stronger for privately-held than for publicly-listed firms. Private firms in affected states reduce net hiring by 24 percentage points and cut investment by one-half, relative to publicly-listed firms.

CHAPTER I

Introduction

A perennial debate in economics concerns the extent to which difficulties in reducing nominal wages affect employment outcomes. While there exists a large literature documenting the existence of wage rigidity in several modern, industrial economies, quantifying an empirical relationship between wage rigidity and employment outcomes has proven difficult. Further, no research to date has been able to quantify the cost to an establishment for reducing a worker's wage in the presence of wage rigidity.

A more recent debate, with the onset of the Great Recession, concerns the extent to which shocks to the supply of bank credit affect worker-level employment outcomes in firms exposed to those banking shocks. As the literature is realtively young, there exists little empirical evidence quantifying the relationship between bank shocks and establishment-level and firm-level employment, and, due to data limitations, no empirical evidence of the impact on worker outcomes. Both debates are of central importance to macroeconomics and the formation of macroeconomic policy, yet the extent and signfinicance of both topics are subjects of considerable uncertainty.

The research of this dissertation uses individual-level, employer-employee linked, administrative data to quantify the relationship between microeconomic employment adjustment and variables of key interest to macroeconomic policy: wage rigidity and financing constraints. A significant advantage of this research over previous work is that it uses high quality, adminstrative, individual worker data with establishment identifiers and a rich set of covariates from a modern industrial economy to quantify these relationships at the level where the employment decisions are made: the establishment. The main findings in the research on wage rigidity and employment outcomes are that there is a significant, empirical relationship between wage rigidity and layoffs, quits, and hires and that only moderate costs to cutting nominal wages are necessary to generate these movements in employment. The main findings in the research on banking shocks and employment outcomes are that, in response to a local banking shock, establishments reduce their levels of employment and that privately held establishments reduce their levels of employment by even more than establishments that are part of a publicly traded company. Below, I discuss the results from each chapter in more detail.

Chapter 2, "Wage Rigidity and Employment Outcomes: Evidence from Administrative Data" examines the empirical relationship between downward nominal wage rigidity and employment outcomes using linked employer-employee data and has four novel contributions. First, this chapter develops a theoretical model of an establishment's decision making problem in the presence of downard nominal wage rigidity, modeling wage rigidity through a downward nominal wage adjustment cost function with a fixed menu cost, a linear adjustment cost, and a quadratic adjustment cost. The model generates three directly testable predictions: establishment's with greater measured wage rigidity should exhibit a higher layoff rate and lower quit and hire rates.

Second, this chapter introduces a unique approach to measuring wage rigidity at the establishment level using worker-level year-over-year nominal wage changes and provides evidence from monte carlo simulations that the proposed measure is unbiased in the small worker sample sizes often found in the establishments of the sample. Using the proposed measure of wage rigidity, this chapter provides evidence on the degree of wage rigidity and the dispersion of wage rigidity present in the establishments in the sample. Establishment-level estimates suggest that wage rigidity prevents 24 percent of counterfactual wage cuts with an estimated standard deviation of 27 percent across all establishments, indicating significant variation in wage rigidity across establishments.

Third, chapter 2 estimates the empirical relationship between wage rigidity and layoffs, quits, and hires at the establishment-level and confirms the predictions of the theoretical model. An establishment with the sample average level of measured wage rigidity is predicted to have a 0.7 percentage point increase in the layoff rate, a 1.8 percentage point reduction in the quit rate, and a 1.3 percentage point decrease in the hire rate relative to an establishment with no measured wage rigidity. Wage rigidity interacts with movements in establishment revenue in economically meaningful ways, amplifying its relationship with employment outcomes. An establishment with the sample average level of measured wage rigidity is predicted to have a 1.5 percentage point increase in the layoff rate, a 3.5 percentage point decrease in the quit rate, and a 2.6 percentage point decrease in the hire rate in response to a one standard deviation movement in revenue growth.

Finally, this chapter uses the empirical results to estimate the structural parameters in the downward nominal wage adjustment cost function. Estimation of the parameters in the wage cutting cost function implies that wage rigidity reduces profits by approximately 3.3 percent at the average establishment, with the average establishment facing a per-worker menu cost of 774 euros annually of cutting nominal wages. These estimates are the first in the literature to quantify the costs an establishment faces to cut its workers' nominal wages and suggests that even moderate costs of cutting nominal wages generate meaningful effects on employment outcomes.

Chapter 3, "Wage Rigidity and Employment Outcomes: Theory", provides two contributions. First, it derives the simple analytics of the establishment decision making model to establish the intuition of how wage rigidity effects establishment choices. Second, it uses the estimated model to provide a series of counterfactual inflation policy simulations to determine whether inflation "greases the wheels" of the labor market.

The analytics of the theoretical model shows that wage rigidity works through two channels in the model. The first channel is immediately clear: a cost to cutting wages deters establishments from reducing the current wage of its workers, leading the establishment to pursue employment adjustment through increased layoffs and reduced hiring to lower the wage bill in the face of an adverse shock. Further, not reducing the wage in response to a negative shock leads to a lower quit rate at the establishment, further exacerbating the need for layoffs and reduced hiring to adjust employment. The second channel through which wage rigidity works in the model is that establishments are forward-looking and will reduce the cost of wage rigidity in future periods through dampening wage increases in response to positive shocks in the current period.

The series of counterfactual policy simulations provide establishment and worker outcomes under three inflation regimes: a low level of 0 percent inflation, a moderate level of 1.3 percent inflation (the average inflation rate over the sample period), and a high level of 5 percent inflation. Increasing the inflation rate in the model from 1.3 percent to the counterfactual 5 percent mitigates the effects of downward nominal wage rigidity, as the average establishment's profit loss falls from 3.3 percent in the baseline case to 1.7 percent in the case of 5 percent inflation. Further, the layoff rate attributed to wage rigidity falls from 1.1 percent in the baseline case to 0.7 percent in the 5 percent inflation case, and the quit and hire rates increase from 12.2 percent to 12.7 percent and 9.5 percent to 9.7 percent, respectively. Thus, the counterfactual simulations provide some evidence that inflation may "grease the wheels' of the labor market.

Chapter 4, “Wage Rigidity and Employment: The Role of Establishment Size and Age”, examines whether the relationship between establishment-level wage rigidity and layoffs, quits, and hires varies meaningfully across large and small establishments and young and old estbalishments. The results show that large establishments, on average, exhibit slightly more wage rigidity than smaller establishments and that older establishments exhibit significantly more wage rigidity, on average, than the younger establishments.

Interestingly, while the younger establishments exhibit lower estimated wage rigidity than the older establishments, the younger establishments display a much stronger relationship between wage rigidity and layoffs and hires. The average young establishment exhibits an increase in its layoff rate by 4.7 percentage points and a redcution in its hire rate by 7.1 percentage points relative to a young establishment with no wage rigidity, whereas the average old establishment see no significant increase in its layoff rate and only a 3.4 percentage point reduction in its hire rate. The young and old establishmets experience relatively similar experiences in quits, with 1.6 and 1.7 percentage point reductions, respectively.

A significant difference also exists between large and small establishments. Wage rigidity increases layoffs at the average large establishment by 1 percentage point and decreases quits and hires by 2.5 and 2.8 percentage points, respectively, whereas wage rigidity and layoffs have no statisitcal relationship in small establishments and decreases quits and hires by 1.2 and 2.5 percentage points respectively.

Chapter 5, “Bank Balance Sheet Shocks and the Effects of the Financial Crisis”, studies the impact of an exogenous shock to the supply of bank loan credit on employment using a unique, comprehensive, employer-employee dataset from German social security records. The geographically contained nature of the German banking system makes this dataset particularly well suited to study the worker impact of an exogenous contraction in the local supply of bank credit. German savings banks are required, by law, to only loan to and take deposits from customers within each bank's geographically specified region, and local bank customers–both businesses and private households–face significant costs to switching banks. Further, the German regional savings banks' trading losses from U.S. mortgage-backed securities provide the opportunity to study the impact of an internationally imported shock to bank capital completely unrelated to local demand.

The German savings banks' trading losses from U.S. mortgage-backed securities cause a deep economic contraction in the banks exclusive geographic domain. Using detailed, annual balance sheet data for the universe of German banks, this chapter shows that loan growth declines by 20 percentage points, on average over the four crisis years, in regions with exposed banks. The reduction in loan credit translates into meaningful macroeconomic outcomes, with regional output declining by 0.3 percentage points and the unemployment rate rising by 1.4 percentage points, on average, in affected regions compared to unaffected regions during each of the four crisis years.

Chapter 5 then explores a possible mechanism through which bank shocks transmit through the economy by examining whether establishment's belonging to privately-held firms that are more dependent on local bank loans to finance their business activities are differentially and adversely affected compared to establishments belonging to publicly-listed and traded firms that can access international and domestic equity markets for financing needs. The results show that the effect is stronger for privately-held firms than for publicly-listed firms. Establishments in privately-held firms in affected states reduce net hiring by 24 percentage points relative to establishments in publicly-listed firms.

CHAPTER II

Wage Rigidity and Employment Outcomes: Evidence from Administrative Data

> You say, "We know from repeated experience that the money price of labour never falls till many workmen have been for some time out of work." I know no such thing; and, if wages were previously high, I can see no reason whatever why they should not fall before many labourers are thrown out of work. All general reasoning, I apprehend, is in favour of my view of this question, for why should some agree to go without any wages while others were most liberally rewarded?
>
> *Letter of David Ricardo to Thomas Malthus, 1821*

2.1 Introduction[1]

A perennial debate in economics concerns the extent to which difficulty reducing nominal wages affects employment outcomes. This paper uses a novel dataset to estimate the extent of wage rigidity at a sample of West German establishments. It then examines the relationship between establishment-level wage rigidity and employment outcomes, specifically layoff, quit, and hire rates. The results are consistent with

[1]This chapter is joint work the Gabriel Ehrlich (Congressional Budget Office).

the predictions of a theoretical model of establishment decision-making in the face of downward nominal wage rigidity (simply "wage rigidity" hereafter). Establishments with more rigid wages exhibit higher layoff rates and lower quit and hire rates.

The paper introduces a measure of wage rigidity suitable for establishment-level analysis. The data are particularly useful for this task, as they contain total compensation histories for every worker at each of the sampled establishments. Furthermore, the compensation histories are taken from administrative data and should be free of measurement error, though have their own set of measurement issues. There are three major advantages to this approach to estimating wage rigidity. First, it uses cross-sectional and time variation in the position of the wage change distribution to identify wage rigidity, rather than relying solely on cross-sectional variation within each period. Second, in contrast to typical histogram location approaches to estimating wage rigidity in the literature, this wage rigidity estimator uses wage changes both above and below the median to estimate the counterfactual distribution. Third, it performs well regardless of whether the median wage change is above or below zero, a situation that can be problematic for estimators that rely only on cross-sectional variation in the wage change distribution within a period.

The estimates suggest that wage rigidity prevents 24.5 percent of wage cuts at the average establishment, with a standard deviation of 23.5 percent across establishments. Establishments in the construction and transportation supersectors display the least wage rigidity, with average levels of 3.2 percent and 11.5 percent of wage cuts prevented, respectively. Establishments in the public administration supersector display the most wage rigidity, with an average level of 45.9 percent of wage cuts prevented.

The paper establishes a clear empirical relationship between wage rigidity and employment outcomes. Because the data allow for the observation of employment flows at the individual level, including into and out of unemployment, layoffs, quits, and hires may be imputed with minimal assumptions. An establishment with the sample-average level of wage rigidity is predicted to have a 0.7 percentage point higher layoff rate, a 1.8 percentage point lower quit rate, and a 1.3 percentage point lower hire rate than an establishment with no wage rigidity.[2] Wage rigidity amplifies layoffs at establishments with shrinking revenues and dampens hires at establishments with gorwing revenues. Given a one standard deviation decrease in revenue growth, an establishment with average wage rigidity is predicted to increase its layoff rate by 1.5 percentage points more than an establishment with no wage rigidity. Given a one standard deviation increase in revenue growth, an establishment with average wage rigidity is predicted to increase its hire rate by 2.6 percentage points less than an establishment with no wage rigidity.

A significant advatange of this paper is that the administrative, individual-level wage data is a measure of total compensation that includes base salary, bonuses, and other forms of compensation. Previous studies focus on wage rigidity in base pay only due to data limitations. One can imagine, however, that establishments may avoid the implications of wage rigidity in base pay on employment outcomes through altering bonuses. Thus, a complete examination of the relationship between wage rigidity and employment outcomes should include a measure of total compensation as this paper does.

Using the empirical results to estimate the structural model via indirect inference suggests that wage rigidity reduces the average establishment's profits by 3.3 percent. The estimates suggest that the average establishment faces a per-worker menu cost of 774 euros annually of cutting nominal wages.

Several previous studies have documented the existence of wage rigidity in microeconomic datasets. Prominent examples using U.S. data include Card and Hyslop (1997), and Kahn (1997). Kahn (1997) estimates that wage earners experience nom-

[2]For comparison, the sample average layoff rate is 4.5 percent, the sample average quit rate is 9.2 percent, and the sample average hire rate is 17.2 percent.

inal wage reductions 47 percent less often than they would in the absence of wage rigidity. Daly and Hobijn (2013) show that the proportion of workers reporting a zero nominal wage change in the United States increased in the recent recession, from 12 percent in 2006 to 16 percent in 2011. Lebow, Saks, and Wilson (1999) estimate that, even accounting for benefits such as cash bonuses and health insurance, wage rigidity prevents 30 percent of reductions in total nominal compensation that would otherwise occur. Using European data, Knoppik and Beissinger (2009) conclude that wage rigidity prevents 37 percent of counterfactual wage cuts in the Euro area, and 28 percent of wage cuts in Germany specifically. Dickens et al. (2007) examine evidence in the United States and 15 European countries, and find that the fraction of workers covered by wage rigidity is 28 percent on average, ranging from 4 percent in Ireland to 58 percent in Portugal.

It has been difficult, though, to establish a link between wage rigidity and employment outcomes. Card and Hyslop (1997) find that "...nominal rigidities have a small effect on the aggregate economy...," while Altonji and Devereux (2000) report, "Our estimates of the effect of nominal wage rigidity on layoffs and promotions ... are too imprecise for us to draw any conclusions." Daly and Hobijn (2013) find that their model of nominal wage rigidity generates wage dynamics that are consistent with recent U.S. data, although their use of the Current Population Survey prevents them from studying the micro-level relationship between wage rigidity and employment outcomes. Akerlof, Dickens, and Perry (1996) find that wage rigidity makes a statistically insignificant difference in macroeconomic time series estimates of a Phillips Curve equation in the postwar period. Lebow et al. (1999) estimate that the non-accelerating inflation rate of unemployment is positively correlated with inflation, contrary to what would be predicted by an important role for nominal wage rigidity. They describe the apparent contradiction between the evidence on the extent of wage rigidity and the lack of evidence that it affects employment outcomes as a “micro-macro puzzle”. An exception to this pattern is Kaur (2012), who finds strong causal effects of wage rigidity on employment levels in informal agricultural labor markets in India.

Two possible solutions to the micro-macro puzzle have been proposed. Barro (1977) argues that in a long-term employment relationship, the wage at a particular point in time is less important than the path of wages over the life of the relationship. Therefore, apparently rigid wages may reflect optimal long-term contracting rather than difficulties in wage adjustment, and may not have meaningful implications for employment outcomes. Elsby (2009) notes that forward-looking, wage-setting firms will compress wage increases in the presence of wage rigidity. Smaller wage increases in good times reduce the need for wage cuts in the face of an adverse shock.

The paper proceeds as follows: Section 2 presents a model of establishment decision-making in the presence of wage rigidity and derives predictions for the effects of wage rigidity on layoffs, quits, and hires. Section 3 provides an overview of the data set and basic descriptive statistics. Section 4 introduces a method of measuring wage rigidity at the establishment level and describes the distribution of wage rigidity across establishments in the sample. Section 5 estimates the empirical relationship between wage rigidity and layoffs, quits, and hires. Section 6 uses these results to estimate the theoretical model by indirect inference, quantifies the costs of wage rigidity to establishments, and conducts a series of counterfactual policy simulations under various levels of the inflation rate. Section 7 concludes.

2.2 Model of Establishment Decision Making with Wage Rigidity

This section examines the dynamic wage and employment policies of a single establishment with heterogeneous worker types facing an imperfectly competitive labor market.[3] The establishment's goal is to maximize its discounted stream of expected future profits. The establishment experiences shocks to its marginal revenue product of labor and faces costs of adjusting its stock of labor and wage rate.

2.2.1 Establishment Environment

The establishment has infinite life and uses one input to production, labor, of which there are J distinct types. The establishment maximizes its discounted stream of expected per period profits, where per period profits are given as:

$$\Pi = \sum_{j=1}^{J} \Big(a_j n_j^{\alpha} - w_j n_j - c_h(h_j, n_{j,-1}) h_j - c_\ell \ell_j - g(w_j, w_{j,-1}) n_j \Big)$$

where n_j is the stock of type j labor used in production, α governs returns to scale, and w_j is the wage rate for type j labor. h_j and ℓ_j are the number of type j employees the establishment hires and lays off, respectively. $c_h(\cdot)$ is a per employee hiring cost function and c_ℓ is the cost per layoff. a_j is a stochastic process that shifts the marginal revenue product of labor.[4] a_j is the product of an establishment-wide productivity level z and a type j productivity level u_j.

Downward nominal wage rigidity enters the model through the wage adjustment cost function, $g\left(w_j, w_{j,-1}\right)$, which is specified in per-employee terms as a polynomial in nominal wage reductions:

$$\begin{aligned} g\left(w_j, w_{j,-1}\right) &= \lambda_0 \mathbb{1}_{(1+\pi) w_j < w_{j,-1}} + \lambda_1 \left(w_{j,-1} - (1+\pi) w_j\right) \mathbb{1}_{(1+\pi) w_j < w_{j,-1}} \\ &\quad + \lambda_2 \left(w_{j,-1} - (1+\pi) w_j\right)^2 \mathbb{1}_{(1+\pi) w_j < w_{j,-1}} \end{aligned} \tag{2.1}$$

λ_0 represents a fixed menu cost of cutting wages, while λ_1 and λ_2 represent linear and

[3] The analysis refers to an establishment rather than a firm to be consistent with the data set, which provides establishment identifiers rather than firm identifiers. Theoretically, however, the analysis would apply equally as well to a firm's problem.

[4] a_j may be conceptualized either as type j's level of labor productivity or as the level of its output price; the remainder of the paper refers to a_j as productivity for concreteness' sake.

quadratic costs of wage cuts, respectively. π represents the deterministic rate of price inflation. Both w_j and $w_{j,-1}$ are specified in real terms, but the establishment bears costs only when it cuts nominal wages. The nominal wage cut from the previous period to the present period is last period's real wage, $w_{j,-1}$, less this period's real wage, w_j, times the increase in the price level $1+\pi$, when this difference is negative, and zero otherwise. Thus, the cost of wage adjustment, $g(\cdot)$, is positive when nominal wages are cut and zero otherwise. The cost of cutting nominal wages gives rise to downward nominal wage rigidity in the model. This is the only place that nominal variables enter the model. Otherwise, the establishment cares exclusively about real payoffs, and all variables above are specified in real terms.

The model is agnostic regarding the precise mechanism generating wage rigidity in the economy. Multiple sources of wage rigidity have been proposed in the literature. Bewley (1999) emphasizes that wage cuts may reduce morale, thereby lowering worker productivity. Similarly, Elsby (2009) and Kaur (2012) both model wage rigidity as arising from reductions in morale associated with wage cuts. Hall and Milgrom (2008) and Christiano, Eichenbaum, and Trabandt (2013) emphasize the role of the bargaining process between employers and workers in insulating wages from business cycle conditions. The sources of wage rigidity remain a topic of discussion in the literature. However, this paper focuses on the consequences of wage rigidity rather than its sources.

The establishment's stock of type j labor evolves according to the equation:

$$n_j = n_{j,-1} - \delta(w_j)n_{j,-1} + h_j - \ell_j$$

where $\delta(w_j)$ is the quit rate of type j labor and $h_j, \ell_j \geq 0$. The establishment faces an imperfectly competitive labor market for each type of labor. The quit rate of type

j labor is given by the function

$$\delta\left(w_j\right) = \bar{\delta}\left(\frac{w_j}{\overline{w}}\right)^{-\gamma}, \quad \gamma > 0 \tag{2.2}$$

where $\bar{\delta}$ is the economy-wide average quit rate. The quit rate is decreasing in the wage rate, w_j. γ governs the degree of competition in the labor market: as γ increases, the quit rate becomes more sensitive to wages. In the limit as γ approaches infinity, the labor market becomes perfectly competitive.

The establishment faces a cost per hire given by the function

$$c_h(h_j, n_{j,-1}) = \phi_1\left(\frac{h_j}{n_{j,-1}}\right) + \phi_2\left(\frac{h_j}{n_{j,-1}}\right)^2 \tag{2.3}$$

The quadratic hiring cost function allows for increasing or decreasing returns to scale in the hire rate. Most studies of hiring costs indicate that they are subject to decreasing returns to scale, for instance Shapiro (1986), Blatter, Muehlmann, and Schenker (2008) and Muehlmann and Pfeifer (2013).

2.2.2 Solution to the Establishment's Problem

Because the establishment's profit function is a linear summation of the individual type j profit functions, the dynamic optimization problem can be written separately for each type of labor. For each labor type j, the establishment chooses the wage rate, level of hires, and layoffs to solve the following dynamic optimization problem:

$$\begin{aligned} V_j\left(z, u_j, w_{j,-1}, n_{j,-1}\right) = \max_{w_j, h_j, l_j} & \; a_j n_j^{\alpha} - w_j n_j - c_h(h_j, n_{j,-1})h_j - c_\ell \ell_j \\ & -g(w_j, w_{j,-1})n_j + \beta E\left[V_j\left(z', u_j', w_j, n_j\right)\right] \end{aligned} \tag{2.4}$$

subject to

$$\ln a_j = \ln z + \ln u_j \tag{2.5}$$

$$\ln z = (1-\psi_z)\ln \bar{z} + \psi \ln z_{-1} + \varepsilon_z,\ \varepsilon_z \sim N\left(0, \sigma_z^2\right) \tag{2.6}$$

$$\ln u_j = (1-\psi_u)\ln \bar{u} + \psi \ln u_{j,-1} + \varepsilon_{u_j},\ \varepsilon_{u_j} \sim N\left(0, \sigma_{u_j}^2\right) \tag{2.7}$$

$$n_j = (1-\delta(w_j))\, n_{j,-1} + h_j - \ell_j \tag{2.8}$$

$$h_j, \ell_j \geq 0 \tag{2.9}$$

The Bellman equation has 4 state variables: establishment-level productivity, z, labor type j-specific productivity u_j, last period's type j wage rate, $w_{j,-1}$, and last period's type j labor stock, $n_{j,-1}$. As specified in equations 2.6 and 2.7, both productivity levels evolve according to a mean reverting, $AR(1)$ process. The errors ε_z and ε_{u_j} are assumed to be independent.

The model solution uses standard value function iteration techniques to find the establishment's value and policy functions. The method of Tauchen (1986) approximates the autoregressive process for the productivity levels z and u_j. The model estimation strategy uses the empirical results from section 5.4 of the paper. Thus, a description of the model estimation is deferred until section 2.6.

2.2.3 Establishment Policy Functions and Simulations

Figure 2.1 displays the establishment's policy functions in the case of perfectly flexible wages, using the parameters decribed in section 2.6 but setting the cost of wage cut parameters λ_0, λ_1, and λ_2, to zero.[5] Panel A illustrates the establishment's wage policy function. The establishment pays lower wages when last period's employment level is higher. With a higher previous level of employment, the establishment can tolerate a higher quit rate while retaining enough employees to meet its desired

[5]The panels show the policy functions for a medium productivity level. The policy functions depend quantitatively on the productivity level but are qualitatively similar.

employment level, reducing the incentive to pay high wages. Panel B illustrates the establishment's employment policy function. As expected, the establishment's desired employment level increases with productivity and the last period's level of employment, but does not vary with the last period's wage. Panel C illustrates the establishment's quit rate policy function, which the establishment controls deterministically by setting the wage rate. The quit rate policy function varies inversely with the wage policy function. It is increasing in last period's employment level but decreasing with productivity but does not vary with the previous wage. Panel D illustrates the establishment's layoff policy function. In the absence of wage rigidity, the establishment never finds it optimal to lay off employees: it is always more profitable to lower wages, thus inducing quits, when employment is greater than desired. A key feature of figure 2.1 is that when wages are perfectly flexible, all policy functions are independent of the previous period's wage.

Figure 2.2 illustrates the establishment's policy functions with wage rigidity.[6] Panel A shows the establishment's wage policy function. The flat portion of the policy function towards the front of the figure is the area where wage rigidity is not binding because the optimal new wage is above the previous period's wage. The upward sloping portion is the area where wage rigidity binds and the establishment sets the current period's nominal wage equal to the previous period's nominal wage.

Panel B shows that the establishment's employment policy continues to be increasing in the previous period's employment level. However, the optimal employment level now depends on the previous period's wage. When wage rigidity binds, the resulting higher wages discourage the establishment from employing as many workers.

Panel C illustrates that the establishment's quit rate policy function remains inversely related to the wage policy function, consistent with equation (2.2). As before, the quit rate increases in last period's employment level. However, with rigid wages,

[6] Specifically, the cost of wage cut parameters λ_0, λ_1, and λ_2, are set to their estimated values as described in section 2.6.

the quit rate decreases with the previous period's wage in the area where wage rigidity is binding. The higher wages in this area of the state space induce fewer quits.

Panel D shows that, in contrast to the policy functions in figure 2.1, an establishment with rigid wages sometimes finds it optimal to lay off workers. Layoffs are never optimal in areas where wage rigidity does not bind, but where wage rigidity binds, the establishment responds by laying off workers. In those cases, the establishment finds it more profitable to lay off workers than to pay the cost of cutting nominal wages. This result is consistent with the intuition that wage rigidity leads to layoffs that would not occur with perfectly flexible wages.

Figure 2.3 displays a simulated wage change histogram in the case of no wage rigidity. As expected, the histogram is widely dispersed around the median and roughly symmetrical. Wage cuts are as prevalent as would be be expected given a symmetrical wage change distribution. Figure 2.4 displays a simulated wage change histogram in the case of rigid wages.[7] The distribution of wage changes is notably compressed realtive to the case of flexible wages and clearly asymmetrical. The portion corresponding to wage cuts is visibly compressed relative to the portion corresponding to wage increases.

Figure 2.5 presents results from simulating the model holding all parameters fixed except the wage rigidity parameters. The horizontal axis indexes the level of wage rigidity in the simulations.[8] The simulation uses 2,000 periods and drops the first 400 periods to allow for burn-in. Panel A shows the estimated level of wage rigidity using the estimator described in section 2.4.1. This is the same estimator applied to the actual data in section 5.4. Panel A shows that estimated wage rigidity increases with the actual level of wage rigidity in the model. Panel B shows the average layoff rate, which increases with wage rigidity. Panels C and D illustrate the average quit

[7]The λ parameters are again set to their estimated values from section 2.6.

[8]Specifically, the estimated λ parameters are each multiplied by the scale factors shown on the axis.

and hire rates, respectively, which both decrease with wage rigidity. Wage rigidity reduces the quit rate by occasionally "holding up" wages above their flexible level, thereby reducing worker turnover. The slower pace of worker turnover reduces the establishment's need to hire new workers as well. Forward-looking establishments also realize that if they hire workers in good times, they may have to pay the costs associated with wage rigidity, either from cutting nominal wages or from laying off workers, in response to future negative shocks.[9]

Therefore, the model predicts that establishments with more rigid wages should exhibit:

1. Higher layoff rates;

2. Lower quit rates; and

3. Lower hire rates.

Section 5.4 tests these three empirical predictions.

2.3 Data Description

2.3.1 Overview of Dataset

The paper employs administrative and survey data from the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB). The main analysis uses the Linked Employer-Employee Data of Integrated Labor Market Biographies (LIAB), matched with the annual IAB Establishment Panel Survey. The LIAB includes 5,293 West German establishments that participated in the annual IAB establishment survey each year either from 1999

[9]Theoretically, it is possible that the hire rate will increase with wage rigidity. However, this situation does not arise when realistic parameter values are used in the model.

through 2001 or from 2000 through 2002, and follows each such establishment every year of its existence from 1997 through 2003.[10]

The LIAB also provides complete labor market biographies for each employee liable to social security who was employed at a sampled, surveyed establishment at any point between 1997 and 2003. The data set follows these workers' entire employment, unemployment, and wage histories from 1993 through 2007, even if the workers move to an establishment outside the sample. The LIAB also provides the exact dates that an employment spell begins and ends for an employee at a given establishment.

The administrative nature of the individual worker data is an important advantage for studying wage rigidity. Establishments provide the individual worker wage data to the FDZ by law, and are subject to penalty for misreporting. Thus, the wage data for each individual should theoretically be without measurement error. Establishment identifiers and full employment samples for the surveyed establishments allow for the accurate calculation of the wage change distribution for each establishment.

Reported wages are the average daily compensation over the employment spell and include base salary and any bonuses, fringe benefits, or other monetary compensation received throughout the spell or year. Thus, the wage reported in the data corresponds more closely to a measure of total compensation than to a base wage rate. This more inclusive wage concept is a significant advantage for studying the relationship between employment adjustment and wage rigidity in light of Lebow et al.'s (1999) finding that establishments are partially able to circumvent wage rigidity by adjusting ancillary compensation.

The employment biographies provide information such as the start and end dates of each employment spell and the reason for each employment notification (e.g. end of or break in employment, required annual notification, etc.). Therefore, labor flows

[10]The East German establishments in the sample were excluded from the analysis.

such as layoffs, quits, and hires may be imputed with minimal assumptions.

Additionally, the LIAB provides an extensive set of employment-related characteristics such as the type of employment spell, professional and occupational status, and white-collar versus blue collar. The worker biographies also include detailed individual characteristics, such as gender, birth year, nationality, education, and vocational training. Finally, the annual IAB Establishment Panel Survey that is linked to the LIAB provides a rich set of establishment characteristics, including information on an establishment's revenue or business volume, and the presence or absence of a work council or wage bargaining agreement.

A disadvantage is that the dataset does not contain employee-level data on hours worked; therefore, a reduction in hours may appear as a wage cut using the wage measure in the data, the daily average wage rate. The data do distinguish between part-time workers working less than half of full-time, those working more than half of full-time, and full-time workers. The wage change distributions include only workers whose hours status does not change between periods to minimize the potential for measurement error. To the extent that this error still exists, it is likely to make wages appear less downwardly rigid than in the absence of hours variation.

Another disadvantage of the dataset is that reported compensation is top-censored at the contribution limit for the German social security system. Top-censoring affects roughly 7 percent of workers in the sample; the analysis excludes these workers from the sample for the purpose of estimating wage rigidity, but not for the purpose of calculating employment flows.[11]

The analysis also uses the Establishment History Panel (BHP) as an additional dataset. The BHP includes industry classification codes and state- and distrcit-level location identifiers for each establishment. In additon, the BHP contains an

[11] The exclusion is necessary because workers with earnings above the contribution limit are all assigned the same top-coded wage in a given year. Therefore, these workers' wage changes would not reflect their actual earnings but instead the change in the yearly contribution limit.

extension file with information on establishment births, deaths, and reclassifications. Supplementary data in this extension allows for the identification of establishment closures that are likely to be spin-offs or takeovers as opposed to true closures.

The final dataset used in the paper is the Sample of Integrated Labor Market Biographies (SIAB). The SIAB provides complete labor market biographies for a 2 percent random sample of all employees liable to social security. However, the SIAB does not provide worker biographies for all workers at a sampled establishment as in the LIAB, nor is it linked to the Establishment Survey Panel. Therefore, the paper focuses on the LIAB for the main analysis. However, because the SIAB is a representative sample of the German workforce, the dataset provides an opportunity to examine aggregate labor market statistics in section 2.3.3.

2.3.2 Descriptive Statistics

The analysis restricts the sample to the years 1997 through 2003, the period for which the data includes worker biographies for all workers at the sampled establishments. The analysis includes workers ages 20 through 60. The main unit of observation is the establishment-year. An establishment-year is excluded if the establishment has less than 50 employees or 10 valid wage changes in the year. An establishment is excluded altogether if it does not meet the criteria above for at least three years. Additionally, the analysis requires data on establishment revenues in both the current and previous years in order to calculate the establishment's change in revenue. These restrictions leave 2,250 establishments for the analysis.

Layoff, quit, and hire rates are measured as fractions of the establishment's total workforce as of December 31st of the preceding year. Following a convention for distinguishing involuntary layoffs and voluntary quits in the worker biographies similar to that of Blein and Rudolph (1989) and Haas (2000), a layoff is defined as an interruption between employment spells that results in the employee flowing into unemployment, as indicated by receipt of unemployment assistance, before the beginning of another employment spell. Conversely, a quit is defined as an employment interruption that does not contain an unemployment spell and results in an employee flowing into another job without receipt of unemployment assistance. The beginning of a new employment spell is classified as a hire if the employee's immediately preceding spell was either unemployment or employment at another establishment.[12] In the data, there are many instances of a spell reported as ending, but after which the worker resumes employment at the same establishment nearly immediately without collecting unemployment assistance. These occurences are classified as neither quits nor hires if the break between spells is less than 28 days. A separation is classified as neither a layoff nor a quit if the worker's biography contains neither a subsequent employment spell nor subsequent receipt of unemployment assistance (for instance, if the worker dies).[13] This situation arises in less than one percent of separations.

Table 2.1 shows the descriptive statistics of the layoff, quit, and hire rates for the sample of establishments from 1997 through 2003. The average annual layoff rate over the period is 4.5 percent with a standard deviation of 9.2 percent across establishment-years. The average annual quit rate over the period is 9.2 percent with a standard deviation of 14.7 percent. The average annual hire rate is 17.2 percent with a standard deviation of 32.7 percent. The average establishment employs 549 workers, versus 219 workers for the median establishment. The average nominal wage

[12] A fourth possibility for employment adjustment is that of a "spin", which can take the form of either an inflow or an outflow. Spin employment flows are those that involve employment movements either between establishments within a firm or a merger or acquisition of two establishments from different firms. An example of an employment movement between establishments covered under the former description is that of an establishment closure where a large proportion of employees from the closed establishment moves directly to another establishment within the same firm. The FDZ provides an extension file on establishment births, deaths, and reclassifications that allows for the identification of spin employment flows. Because the study focuses on the relationship between wage rigidity and the traditional employment flows, spin flows are excluded from the analysis.

[13] The establishment-level anlysis considers the period 1997 through 2003, but the worker biographies span the period 1993 to 2007, so most worker biographies extend beyond the end of the analysis period.

is 102.90 euros per day, with a standard deviation of 66.21 euros per day.[14]

The empirical strategy described in section 5.4 uses changes in establishment revenues to proxy for shifts in the marginal revenue product of labor. Each year, the survey asks each establishment to provide its total business volume (or sales) in the preceeding fiscal year (i.e. from January 1 through December 31).[15] The average establishment-level revenue growth in the sample is 3.5 percent per year with a standard deviation of 20.6 percent.

2.3.3 Aggregate Wage Change Distributions

The wage data from the SIAB provides a representative overview of wage changes for job stayers during the period 1997 through 2003. Figure 2.6 shows the annual aggregate nominal wage change distributions for this period. The plot labeled 2000 represents the distribution of wage changes from 1999 to 2000, et cetera.

Four conclusions are visually evident from observing the nominal wage change histograms and are confirmed through simple tabulations. First, the aggregate nominal wage change distributions exhibit a clear spike at the histogram bin containing a nominal wage change of zero (or the "zero bin" for short). The proportion of nominal wage changes in the zero bin ranges from 11.32 percent to 15.45 percent, with an average of 12.35 percent. Second, a nominal wage change of zero is the most common nominal wage change over the sample period. Third, while nominal wage cuts certainly occur, they are less frequent when compared to nominally zero and nominally positive wage changes. Further, it appears as if a part of the nominally negative portion of the wage change distribution is "missing" when compared to its nominally positive counterpart. From 1997 through 2003, the fraction of workers receiving a

[14] For the purposes of calculating these descriptive statistics, wages were imputed for top-coded earners using a procedure provided by the FDZ.

[15] While the sample only covers establishments with full employment biographies from 1997 through 2003, the survey spans from 1993 through 2008. Thus, the 2004 survey records the establishment's business volume from 2003, the 2003 survey records business volume from 2002, etc.

nominal wage cut ranges from 14.88 percent to 21.32 percent, with an average of 18.46 percent.

Finally, the aggregate nominal wage change distributions exhibit a significant "fall-off" in density from the zero bin to the nominally negative bin immediately to the left of zero. For example, in the year 2003, the zero bin contains 15.45 percent of all wage changes compared to only 4.84 percent in the bin immediately to the left, a fall-off of 10.61 percentage points. Throughout the sample period, the fall-off in density from the zero bin to the bin immediately to the left ranges from 6.17 to 10.61 percentage points and averages 7.90 percentage points. For comparison, the next largest average fall-off between any two histogram bins is 2.91 percentage points and only eight bins exhibit an average fall-off of more than one percentage point. This evidence suggests the existence of downward nominal wage rigidity in the aggregate German economy. The paper now turns to measuring the degree and extent of wage rigidity across German establishments.

2.4 Estimating Wage Rigidity

2.4.1 Methodology

Previous studies have proposed several methods of measuring downward nominal wage rigidity. However, those studies have measured wage rigidity at the aggregate level, whereas this study measures wage rigidity at the establishment level. The small size of many of the establishments in the sample poses a problem for these approaches in the context of this paper. The approach in this paper takes elements from Card and Hyslop (1997) and Kahn (1997), modified for the context of much smaller samples. Figures 2.7 through 2.10 illustrate the approach.

For each establishment i, let m_{it} represent the median wage change from time $t-1$ to time t expressed in percentage points. Then measure the proportion of wage changes in each year in one percentage point wide bins. Let $prop_{ijt}$, for $j \in \{-10, -9, -8, ..., -1, 1, ..., 8, 9, 10\}$, denote the size of the bin that is between j and $j+1$ percentage points away from the median wage change for that year. Figure 2.7 illustrates these measurements for a simulated establishment.[16] For example, $j = -1$ represents the bin between the median wage change and the median wage change minus one percent in each year, whereas $j = 1$ represents the bin between the median wage change and the median wage change plus one percent. In year 1, the median wage change is 2.9 percent, and bin $j = -1$ contains wage changes between 1.9 percent and 2.9 percent, whereas bin $j = 1$ contains wage changes between 2.9 percent and 3.9 percent. The median wage change always separates bins $j = -1$ and $j = 1$. However, the bin j containing nominal zero will vary with the median wage change over time. For instance, in year 1 the nominal wage change of zero lies in bin $j = -3$, which includes wage changes between -0.1 percent and 0.9 percent; in year 2, the median nominal wage change is 6.8 percent, and the nominal wage change of zero lies in bin $j = -7$. The analysis excludes all bins more than 10 percentage points from the median each year.

For each establishment, estimate the regression

$$prop_{ijt} = \delta_0 + \delta_1|j| + \delta_2 j^2 + \epsilon_{it},\ j + m_{it} > 0,\ \forall t \tag{2.10}$$

The regression in equation (2.10) restricts the sample to bins that reflect nominal wage increases only, and excludes the bin containing wage changes of nominal zero, as illustrated in figure 2.8. A data point in this regression is the proportion of wage changes in bin j, in year t, at establishment i. The regression pools the data across years within an establishment. Thus, for the simulated establishment in figure 2.8, the regression in (2.10) contains 74 data points, as there are a total of 74 nominally

[16]The data user agreement with the FDZ prohibits displaying the wage change histograms for a single establishment, necessitating the use of a simulated establishment for the illustration.

positive wage change bins across all six years. $|j|$ and j^2 represent the linear and quadratic distances from the median wage change, respectively. Therefore, equation (2.10) expresses the nominally positive portion of the wage change distribution as a quadratic function of the distance from the median wage change each year.

Next, this estimated function is used to predict what the nominally negative portion of the wage change distribution would be in the absence of wage rigidity. The estimated coefficients from equation (2.10) are used to predict the values $\widehat{prop}_{ijt}$ for the bins that contain negative wage changes, again excluding the bin that contains wage changes of nominal zero.[17] For example, in year 1 of figure 2.9, proportions are predicted for bins $j = -4$ through $j = -10$.

These predicted values are used to estimate the regression

$$prop_{ijt} = \gamma_i \times \widehat{prop}_{ijt} + u_{ijt},\ \ j + m_{it} + 1 < 0,\ \ \forall t \tag{2.11}$$

Equation (2.11) regresses the observed proportion of wage changes in bins corresponding to nominal wage cuts on the proportions that would be predicted from the regression in equation (2.10). Figure 2.10 illustrates how the regression operates. The dark bars in the nominally negative portion of the distribution represent the observed proportion of wage changes in these bins, while the light bars represent the predicted proportions of wage changes in these bins. $\hat{\gamma}_i$ represents the fraction of predicted wage cuts observed in the data. The measure of establishment-level wage rigidity is the proportion of counterfactual wage cuts that are "missing" from the data and is calculated as

$$\widehat{wr}_i = 1 - \widehat{\gamma}_i \tag{2.12}$$

Therefore, the wage rigidity estimate in equation (2.12) is a time-invariant charac-

[17] In cases where $prop_{ijt}$ would be predicted to be negative, $\widehat{prop}_{ijt}$ is set to zero.

teristic of the establishment. $\widehat{wr}_i$ has the natural interpretation that a value of 0.25 implies that 25 percent of counterfactual nominal wage cuts at establishment i were prevented by downward nominal wage rigidity over the sample period.[18]

There are three major advantages to this approach to estimating wage rigidity. First, it uses cross-sectional and time variation in the position of the wage change distribution to identify wage rigidity, rather than relying solely on cross-sectional variation within each period. Second, in contrast to typical histogram location approaches to estimating wage rigidity in the literature, this wage rigidity estimator uses wage changes both above and below the median to estimate the counterfactual distribution. Third, it performs well regardless of whether the median wage change is above or below zero, a situation that can be problematic for estimators that rely only on cross-sectional variation in the wage change distribution within a period. This situation arises in 8.02 percent of the establishment years in the sample.

This approach implicitly assumes that an establishment's counterfactual wage change distribution is symmetrical and has a constant variance across years. Card and Hyslop (1997) argue that, "...symmetry is a natural starting point for building a counterfactual distribution. ...if the individual wage determination process is stationary, then symmetry holds." It is also worth noting that the aggregate German wage change distributions shown in appendix B in section 2.9 appear to be roughly symmetrical around the median in the high inflation years of the late 1970s and early 1980s. When inflation is high, a smaller proportion of the wage change distribution is pushed against nominal zero compared to periods of low inflation. Thus, the shape of the wage change distribution in high inflation periods is likely to be indicative of the shape of the counterfactual distribution that would prevail in the absence of downward nominal rigidity.

[18] Nothing in this procedure prevents $\widehat{wr}_i$ from being negative. A value for $\widehat{wr}_i$ of -0.25 would imply that there are 25 percent more wage cuts in the data than would be predicted by the distribution of nominally positive wage changes.

A potential drawback of this approach is that it also implicitly assumes the nominally positive portion of the wage change distribution is unaffected by wage rigidity in order to predict the nominally negative portion. As emphasized by Elsby (2009), theory suggests that wage rigidity should affect the nominally positive portion of the wage change distribution as well as the nominally negative portion. Specifically, wage increases should be compressed in the presence of wage rigidity. This compression is evident in simulations of the theoretical model presented in section 2.2, as well. Monte Carlo simulations of the estimator presented here suggest that it is unbiased both with and without compression in the wage change distribution. Intuitively, this is because the estimator only attempts to estimate the fraction of counterfactual wage cuts prevented by wage rigidity, and not their magnitudes.

The Monte Carlo simulations suggest that there is some sampling error associated with the estimator. This samplng error will lead to attenuation bias in the estimates of the association between wage rigidity and employment outcomes presented in section 5.4. Therefore, the estimates of these associations are likely to underestimate the strength of the true associations. Please see appendix A in section 2.8 for a discussion of the Monte Carlo simulations.

2.4.2 The Distribution of Wage Rigidity in West Germany

Table 2.2 shows the mean, median, and standard devation of the distribution of wage rigidity estimates for individual establishments within the sample. The average establishment-level measure of wage rigidity is 24.5 percent, implying that wage rigidity prevents 24.5 percent of counterfactual wage cuts at the average establishment. The standard deviation of the estimates is 23.5 percent and the median is 21.8 percent. Thus, there is both a notable degree of estimated wage rigidity among establishments and significant variation across establishments.

Table 2.2 also shows the mean, median, and standard devation of the distribution of wage rigidity estimates within each of the ten supersectors of the economy to provide context as to where wage rigidity is present. The mean and median levels of wage rigidity vary widely across supersectors, with little difference between the mean and median within supersectors. The variation within supersectors, as measured by the standard deviation across establishments, ranges from 19 percent to 34 percent. Among supersectors, public administration exhibits the highest degree of wage rigidity, with an average of 45.9 percent of wage cuts prevented by wage rigidity across establishments. Administration, finance, and energy/water also display large amounts of wage rigidity, with an average of 35.5, 31.1, and 28.9 percent of wage cuts prevented, respectively. Construction, transportation, and mining/manufacturing exhibit the smallest degree of average wage rigidity, with 3.2 percent, 11.5 percent, and 12.8 percent of nominal wage cuts prevented, respectively.

2.5 Wage Rigidity and Employment Adjustment

2.5.1 Empirical Approach

The predictions from the theoretical model in section 2.2.3 imply empirical regressions of the form:

$$y_{it} = \beta_0 + \beta_1 wr_i + X_{it}'\Upsilon + \epsilon_{it} \tag{2.13}$$

where the unit of observation is an establishment-year. y_{it} represents an employment adjustment variable of interest: the layoff rate, the quit rate, or the hire rate. wr_i represents the estimated percentage of wage cuts prevented by downward nominal wage rigidity, as discussed in section 2.4.1. X_{it} represents a vector of control variables, including a dummy for the presence of a work council, the median year-over-year percentage wage change, a set of year, state, and sector fixed effects, dummies for establishment size groups, the fraction of the workforce that is female, and controls for workforce educational attainment and occupation. Estimates of equation (2.13) are presented in column 1 of tables 2.3, 2.4, and 2.5, for layoffs, quits, and hires, respectively. Figures 2.8, 2.9, and 2.10 illustrate the economic interpretation of these estimates.

It is natural to examine whether the association between wage rigidity and employment adjustment varies according to the economic shocks an establishment faces. In the theoretical model presented in section 2.2, layoffs are a response to negative shocks to the marginal revenue product of labor, while hires are a response to positive shocks to the marginal revenue product of labor.[19] Although the data do not permit explicit observation of marginal revenue product of labor shocks, data on revenue growth is likely to be informative about such shocks. Assuming changes in revenue growth reflect primarily shifts in the marginal revenue product of labor suggests the following additional specification for examining the relationship between wage rigidity and employment outcomes:

$$\begin{aligned} y_{it} &= \beta_0 + \beta_1 wr_i + \beta_2 posrev_{it} + \beta_3 negrev_{it} \\ &\quad + \beta_4 \left(wr_i \times posrev_{it}\right) + \beta_5 \left(wr_i \times negrev_{it}\right) + X_{it}'\Upsilon + \epsilon_{it} \end{aligned} \tag{2.14}$$

The variables $posrev_{it}$ and $negrev_{it}$ denote the year-over-year percentage change in revenue; $posrev_{it}$ is set to zero when this change is negative, while $negrev_{it}$ is set to zero when this change is positive. Specifying the change in revenue this way allows the estimation of a linear spline function over revenue growth, permitting disparate associations between revenue growth and employment adjustment depending on whether revenue growth is positive or negative.[20] The variables $(wr_i \times posrev_{it})$

[19] These responses can persist over time after the initial shock. Whether positive and negative shocks should have disparate relationships with quits is ambiguous in the model.

[20] The specification of revenue growth as a linear spline function with a kink at zero is similar to the specification of Holzer and Montgomery (1993), who also interpret changes in sales growth as reflecting primarily shifts in demand.

and $(wr_i \times negrev_{it})$ are interactions between estimated establishment wage rigidity and revenue growth, capturing possible interactions between wage rigidity and changes in revenue.

Tables 2.3, 2.4, and 2.5 present three sets of results for each outcome variable. Column (1) presents estimates of equation (2.13). Column (2) presents estimates of equation (2.14) without the wage rigidity-revenue growth interaction terms. Column (3) presents estimates of equation (2.14) including wage rigidity-revenue growth interaction terms. All regressions are weighted by establishment-year employment. Regression weights are likely to be appropriate for two reasons: first, because the estimate of wage rigidity is likely to be less noisy at larger establishments; and second, because larger establishments employ a much larger percentage of the total workforce, and thus their behavior has a larger effect on the aggregate economy.

2.5.2 Layoffs

Table 2.3 shows results from the layoff regressions. In column (1), the estimated coefficient on the wage rigidity variable is 2.9 percent and statistically significant. The sign is consistent with the predictions presented in section 2.2.3: establishments with higher degrees of wage rigidity exhibit higher layoff rates. Adding positive and negative revenue growth as regressors in column (2), the coefficient on estimated wage rigidity is again 2.9 percent and statistically significant. In column (3), which adds interactions between wage rigidity and revenue growth, the coefficient on the level of estimated wage rigidity is 2.5 percent and statistically significant. The coefficients on the uninteracted revenue growth terms are statistically insignificant, as is the coefficient on the interaction between wage rigidity and positive revenue growth. However, as suggested by the model, the coefficient on the interactions between wage rigidity and negative revenue growth is negative 19.3 percent and highly significant. Because $negrev_{it}$ enters the regression as a weakly negative number, the negative coefficient on the interaction implies that an establishment with more rigid wages exhibits more layoffs when its revenue declines than an establishment with less rigid wages.

Figure 2.11 shows the economic interpretation of these coefficients. The horizontal axis measures the estimated level of wage rigidity for an establishment, and the vertical axis measures the predicted increase in the layoff rate for an establishment relative to the case of no wage rigidity. The solid line in figure 2.11 presents the economic interpretation of the layoff regression results from column (1) of table 2.3 with no revenue variables or interaction terms. An establishment with the sample average level of wage rigidity of 24.5 percent is predicted to have a 0.7 percentage point higher layoff rate than an establishment with no wage rigidity. This difference corresponds to a 15.9 percent increase over the 4.5 percent sample average layoff rate. An establishment with estimated wage rigidity of 48 percent, one standard deviation above the sample average, is predicted to have a 1.4 percentage point higher layoff rate than an establishment with no wage rigidity, an increase of 30.9 percent relative to the average.

The dashed line in figure 2.11 presents the economic interpretation of the layoff regression results from column (3) of table 2.3, which include the revenue variables and the interaction of the revenue variables with wage rigidity. Specifically, the line shows an establishment's predicted increase in the layoff rate relative to the case of no wage rigidity given a negative, one standard deviation (-20.6 percent) movement in revenue growth. Including the revenue interaction terms in the empirical model significantly amplifies the positive relationship between estimated wage rigidity and layoffs. An establishment with the sample average level of wage rigidity and a one standard deviation negative movement in revenue growth is predicted to have a 1.5 percentage point higher layoff rate relative to an establishment with no wage rigidity, corresponding to a 34.5 percent increase. Further, an establishment with estimated wage rigidity one standard deviation above the sample average is predicted to have a 3.0 percentage point higher layoff rate, corresponding to a 66.7 percent increase in the layoff rate relative to the sample average.

2.5.3 Quits

Table 2.4 shows results from the quit regressions. In column (1), the estimated coefficient on the wage rigidity variable is -7.2 percent and statistically significant. This sign is consistent with the predictions from the theoretical model, as establishments with higher degrees of wage rigidity exhibit lower quit rates. In column (2), the coefficient on estimated wage rigidity is also -7.2 percent and statistically significant. In column (3), the coefficient on estimated wage rigidity is negative 5.2 percent and statistically significant. The coefficients on the interactions between wage rigidity and revenue growth imply that firms with more wage rigidity are predicted to experience fewer quits in response to an increase or decrease in revenue, although the coefficient on the negative revenue movement interaction is not statistically significant.

Figure 2.12 shows the economic interpretation of the coefficients in the quit rate regressions. The horizontal axis measures the estimated level of wage rigidity for an establishment, and the vertical axis measures the predicted decrease in the quit rate for an establishment relative to the case of no wage rigidity. The solid line presents the economic interpretation of the quit rate regression results from column (1) of table 2.4 with no revenue variables or interaction terms. An establishment with the sample average level of wage rigidity is predicted to have a 1.8 percentage point lower quit rate than an establishment with no wage rigidity, which corresponds to a 19.3 percent decrease relative to the 9.2 percent sample average quit rate. An establishment with one standard deviation higher than average wage rigidity is predicted to have a 3.5 percentage point lower quit rate relative to the case with no wage rigidity, which corresponds to a 38.0 percent decrease.

The dashed line in figure 2.12 presents the economic interpretation of the quit rate regression results from column (3) of table 2.4, which include the revenue variables and the interaction of the revenue variables with wage rigidity. The line shows an establishment's predicted decrease in the quit rate relative to the case of no wage rigidity given a one standard deviation negative movement in revenue growth. However, the specification including the revenue interaction terms is not statistically distinguishable from the specification that does not include the revenue interaction terms, and the difference is not economically meaningful.

2.5.4 Hires

Table 2.5 shows results from the hire regressions. In column (1), the estimated coefficient on the wage rigidity variable is -5.3 percent and statistically significant. This coefficient implies that establishments with greater wage rigidity exhbit lower hire rates, as predicted by the theoretical model. In column (2), the coefficient on estimated wage rigidity is -5.9 percent and statistically significant. In column (3), the coefficient on measured wage rigidity is -5.3 percent and statistically significant. The coefficient on the interaction between wage rigidity and positive revenue growth has the expected negative sign and is statistically significant. Establishments with more wage rigidity are predicted to engage in fewer hires when revenue increases.

Figure 2.13 shows the economic interpretation of the coefficients in the hire rate regressions. The horizontal axis of figure 2.13 measures the estimated level of wage rigidity for an establishment, and the vertical axis measures the predicted decrease in the hire rate for an establishment relative to the case of no wage rigidity. The solid line in figure 2.13 presents the economic interpretation of the hire regression results from column (1) of table 2.5 with no revenue variables or interaction terms. An establishment with the sample average level of wage rigidity is predicted to have a 1.3 percentage point lower hire rate than an establishment with no wage rigidity, which corresponds to a 7.5 percent decrease relative to the 17.2 percent sample average hire rate. An establishment with one standard deviation higher than average wage rigidity is predicted to have a 2.5 percentage point decrease in the hire rate relative to the case with no wage rigidity, which corresponds to a 14.8 percent decrease in the hire rate relative to the sample average.

The dashed line in figure 2.13 presents the economic interpretation of the hire regression results from column (3) of table 2.5, which include the revenue variables and the interaction of the revenue variables with wage rigidity. The line shows an establishment's predicted decrease in the hire rate relative to the case of no wage rigidity given a positive one standard deviation movement in revenue growth. Including the revenue interaction terms in the empirical model significantly amplifies the negative relationship between estimated establishment level wage rigidity and the hire rate. An establishment with the sample average level of wage rigidity and a one standard deviation positive movement in revenue growth is predicted to have a 2.6 percentage point lower hire rate than an establishment with no wage rigidity, which corresponds to a 14.8 percent decrease in the hire rate relative to the sample average. Further, an establishment with one standard deviation higher than average wage rigidity is predicted to have a 5.0 percentage point decrease in the hire rate, corresponding to a 29.1 percent decrease in the hire rate relative to the sample average.

2.6 Model Estimation

This paper employs a combination of methods to choose the parameters of the theoretical model described in section 2.2. The parameters β, π, α, $\overline{\delta}$, $\overline{w}$, γ, $\ln \bar{z}$, ψ_z, σ_z^2, $\overline{u}$, λ_1, and c_ℓ are calibrated externally or estimated directly from the LIAB microdata used throughout the paper. The parameters λ_0, λ_2, ψ_u, σ_u^2, ϕ_1, and ϕ_2 are estimated via indirect inference to match a set of simulated moments to their empirical counterparts in the LIAB microdata. The model period is taken to be one year.

2.6.1 Calibrated Parameters

Table 2.6 shows the values of the calibrated parameters. The inflation rate, π, of 1.33 percent is the average rate of consumer price inflation in Germany over the period 1997 through 2003 from the World Bank's World Development Indicators. The establishment discount rate β is calibrated from the World Bank's WDI tables to match the average German real lending interest rate for the period 1997-2002. The nominal lending interest rate is defined as the bank rate that meets the short- and medium-term financing needs of the private sector, and averages 9.5 percent for the period.[21] The real interest rate is calculated as the nominal rate minus the average inflation rate of 1.33 percent. β is then calibrated as $\frac{1}{1+0.095-0.0133}$, or 0.924. The parameter c_ℓ is taken to match German redundancy costs from the World Bank's Cost of Doing Business project, taking the average of the cost for workers with 1 year of tenure and workers with 5 years of tenure. To expedite the estimation, the parameter λ_1 is set to zero. The parameter $\bar{u}$ is normalized to 1. The average quit rate $\overline{\delta}$ and the average nominal daily wage $\overline{w}$ are taken directly from the microdata sample. $\overline{\delta}$ is the average quit rate across establishment-years, 9.2 percent. The average daily wage is 102.90 euros.

The returns-to-scale parameter α is chosen to match labor's average share of value added across all establishment-years in the microdata. For each establishment-year, the establishment's total wage bill is calculated from the worker biographies. The establishment's value added is calculated as total revenues minus intermediate inputs and external costs.[22] The theoretical model in section 2.2 abstracts from intermediate

[21]See http://data.worldbank.org/indicator/FR.INR.LEND/countries?page=2 for more detail. The rate is not available for 2003.

[22]Each year, the establishment survey panel includes a question regarding the share of revenue attributable to external costs. For instance, in the 2002 survey the question read:

> What share of sales was attributed to intermediate inputs and external costs in 2001, i.e. all raw materials and supplies purchased from other businesses or institutions, merchandise, wage work, external services, rents and other costs (e.g. advertising and agency expenses, travel costs, commissions, royalties, postal charges, insurance premiums, testing costs, consultancy fees, bank charges, contributions to chambers of

inputs, so there is no distinction in the model between revenue and value added. In the microdata, it is necessary to adjust for intermediate inputs to calculate labor's share of value added accurately. Labor's share of value added per establishment-year is simply the establishment's total wage bill divided by revenue less external costs. Averaging labor's share of value added across establishment-years yields an estimate of 0.65 for the parameter α.

The parameter γ is the elasticity of the quit rate with respect to wages in equation (2.2) from section 2.2.1. This equation is difficult to estimate directly due to its non-linearity in the wage, but a first-order Taylor's expansion yields the following linear approximation around the average wage, $\overline{w}$:

$$\delta(w) - \overline{\delta} \;\approx\; -\overline{\delta}\gamma\left(\frac{w - \overline{w}}{\overline{w}}\right) \tag{2.15}$$

Equation (2.15) expresses the deviation of the establishment-year quit rate from the average quit rate as a decreasing function of the percentage deviation of the establishment-year wage from the economy average wage.[23] Taking equation (2.15) to the data requires accounting for worker and establishment heterogeneity that is not present in the theoretical model.[24] A Mincer regression of individual log wages on worker and establishment observable characteristics allows for the removal of observable heterogeneity.[25] Thus, the residual from this regression provides a "cleansed" measure of the deviation of individual log wages from the market average. Averaging these residuals at the establishment-year level provides a log approximation to the

trade and commerce and professional associations)?

[23]Although the quit rate function in equation 2.2 is linear in logs, estimating the equation in logs is not feasible because quits are zero in some establishment-years.

[24]Neglecting to account for heterogeneity may yield biased inference if wages are correlated with other determinants of the quit rate. For example, if non-wage amenities such as pleasantness of the job are reflected in compensating wage differentials, a naive estimate of γ that does not account for heterogeneity will be biased toward zero.

[25]The covariates included in the Mincer regression are a set occupation dummies, a set of education dummies, gender, nationality, age and age squared, a set of year fixed effects, federal state, and a set of sector dummies.

term $\left(\frac{w-\overline{w}}{\overline{w}}\right)$ in equation (2.15).

To estimate equation (2.15), establishment-year quit rates minus the average quit rate were regressed on the average Mincer residuals, and a set of establishment and year fixed effects. The inclusion of establishment fixed effects identifies γ off of time series variation in wages within establishments, rather than cross-sectional variation in wages across establishments, which helps to account for the possible trade-off between wages and amenities. The estimated coefficient in regression equation (2.15) is -0.53, which corresponds to $-\overline{\delta}\gamma$. Dividing the estimated coefficient by $-\overline{\delta}$ yields an estimated γ of 5.75.

The average establishment-wide level of productivity, $\ln \bar{z}$, and the persistence, ψ_z, and variance, σ_z^2, of establishment-wide productivity shocks are estimated from the LIAB microdata using information on value added, VA_{it}, and the returns-to-scale parameter α calculated above. Using the measures of value added and the estimate of α described above, $\ln(VA_{it}) - \alpha \ln(n_{it})$ is regressed on a set of worker covariates, year fixed effects, and establishment fixed effects. The regression results allow for the calculation of annual establishment-level average productivity, p_{it}. Assuming the law of large numbers holds, the average level of u_{ijt} in the sample will be 1 each year, implying $z_{it} = p_{it}$. Regressing $\hat{p}_{it}$ on $\hat{p}_{i,t-1}$ and a set of establishment fixed effects yields estimates for the persistence of the establishment-level productivity process, ψ_z, and the variance of establishment-level productivity shocks, σ_z^2.

2.6.2 Estimated Parameters

The wage cut cost function parameters λ_0 and λ_2, the persistence and variance of shocks to worker-type productivity ψ_u and σ_u^2, and the hiring cost function parameters ϕ_1 and ϕ_2 are estimated through indirect inference to match a set of simulated moments from the theoretical model to their empirical counterparts in the data sample. For a given guess of these parameters, the establishment's optimal policy functions are computed and a series of wage change distributions and employment outcomes are simulated using a set of random shocks. Applying the same method of measuring establishment wage rigidity used in the empirical results to the simulated wage change distributions yields an estimate of wage rigidity for the establishment in the simulated data, $\hat{wr}^s$. For each guess of the wage rigidity parameters, the same procedure is implemented nine times, by multiplying the wage rigidity parameters by a multiplicative factor increasing linearly from zero to two, and the simulated measure of wage rigidity is calculated along with the simulated layoff, quit, and hire rates for each of the nine simulations. Regressing the simulated layoff, quit, and hire rates on $\hat{wr}^s$, as in equation 2.13 from section 2.5.1, yields the regression coefficients on wage rigidity, $\hat{\beta}^s = \begin{bmatrix} \hat{\beta}_\ell^{\,s} & \hat{\beta}_q^{\,s} & \hat{\beta}_h^{\,s} \end{bmatrix}'$, from the simulated data. Therefore, the indirect inference approach provides four natural targets to match from the data: $\hat{\beta}_\ell$, $\hat{\beta}_q$, $\hat{\beta}_h$, and $\hat{wr}$. Identifying the remaining model parameters requires additional moments from the data. These additional target moments are the predicted increase in the layoff rate associated with wage rigidity from the regression in column (1) of table 2.3,[26] the average hire rate in the data sample, the average wage level, the standard deviation of wage changes for job stayers, and the magnitude of the average negative wage change in the data sample.[27]

While there is not a one-to-one correspondence between the target moments and the estimated parameters, λ_0 and λ_2 are identified primarily by the average level of measured wage rigidity $\hat{wr}$ and the magnitude of the average negative wage change in the sample. The empirically estimated level of wage rigidity reflects the total cost of wage rigidity the establishment faces. Given the total cost of wage rigidity, the average negative wage change identifies how much of the cost stems from the menu cost wage

[26]In other words, 0.029 times the average measured level of wage rigidity, 0.245, or 0.7 percent. This quantity is used as a target rather than the sample average layoff rate because the only source of layoffs in the model is wage rigidity. Presumably, there are additional causes of layoffs in the real economy.

[27]All of the simulated moments except for $\hat{\beta}^s$ are taken from the single simulation in which the multiplicative factor on the wage rigidity parameters equals one.

adjustment term, λ_0, and how much from the quadratic wage adjustment cost, λ_2. If the quadratic adjustment cost predominates, the establishment will be willing to cut wages by only a small amount and the magnitude of the average negative wage change will be small. If, however, the menu cost predominates, the establishment will exhibit few small negative wage changes and the negative wage changes the establishment does make will be large on average.

The worker-type productivity shock parameters ψ_u and σ_u^2 are identified mainly by the predicted increase in the establishment layoff rate associated with wage rigidity and the standard deviation of the average wage change. The hiring cost parameters ϕ_1 and ϕ_2 are identified mainly by the average hiring rate and average wage. Inclusion of the wage rigidity regression coefficients β_ℓ, β_q, and β_h, serves two purposes. First, the regression coefficients provide additional discipline on the estimated model parameters by ensuring that the relationships between wage rigidity and employment outcomes hold in the estimated model. Second, the regression coefficients help to correct for any possible model misspecification, in the sense that the regression coefficients are generated using the same procedure in the model as in the data.

With nine target moments from the data and six parameters to estimate, the model is over identified. The parameters are estimated by minimizing the sum of the squared percent deviations of the simulated moments from their empirical counterparts.[28] Table 2.7 shows the empirical and simulated moments using the estimated parameters. The model matches most target moments reasonably well. It noticeably underpredicts the establishment hire rate and the magnitude of the regression coefficient on wage rigidity in the hire regression.

[28]More formally, let θ be a vector of the six structural parameters to be estimated and μ be a vector of the nine target moments. Let $\hat{\mu}^s(\theta)$ be the corresponding simulated moments for any guess of the parameters θ. Then the estimated structural parameters are

$$\hat{\theta} = \arg\min_{\theta}[\hat{\mu}^s(\theta) - \mu]' W^{-1}[\hat{\mu}^s(\theta) - \mu]$$

where W is an nine by nine diagonal weighting matrix with the squared target moments as its entries.

Table 2.8 shows the estimated parameter values along with standard errors. The persistence and variance of worker-type productivity shocks are 0.767 and 0.507, respectively, implying that shocks to individual worker types are larger and more persistent than shocks to overall establishment productivity. The estimated hiring cost parameters, $\hat{\phi}_1$ and $\hat{\phi}_2$, require context to be useful. Applying the estimated values to the hiring cost function in equation 2.3 at the simulated average hire rate of 9.5 percent yields a per-employee hiring cost of 8.9 weeks of total compensation, broadly consistent with Muehlman and Pfeifer's (2013) estimate of more than 8 weeks of wages for skilled German workers. The positive estimate for ϕ_2 implies that there are diseconomies of scale in hiring, also consistent with Muehlman and Pfeifer (2013).[29]

The estimated wage cut cost parameters, $\hat{\lambda}_0$ and $\hat{\lambda}_2$, imply that the average cost of a simulated wage cut is 1,238 euros. In the model simulations, wage rigidity reduces profits by 3.3 percent at the average establishment. Approximately 63 percent of the cost stems from the fixed cost λ_0 and approximately 37 percent stems from the quadratic adjustment cost λ_2.

2.7 Conclusion

This paper explores the relationship between downward nominal wage rigidity and employment outcomes theoretically and empirically using German administrative data. A novel contribution of the paper is the use of linked establishment-employee data to measure wage rigidity and employment adjustment at the establishment level. Establishment-level wage rigidity estimates suggest a substantial amount of downward nominal wage rigidity in Germany, with an average of 24.5 percent of counterfactual wage cuts prevented by wage rigidity. The paper introduces a theoretical model of an establishment's wage and employment decisions in the face real resource cost

[29] At the sample average hire rate of 17.2 percent, implied hiring costs are 18.8 weeks of total compensation.

for cutting nominal wages. The model predicts that more rigid wages should be associated with a higher layoff rate and lower quit and hire rates.

The empirical analysis is consistent with the predictions of the theoretical model. An establishment with the sample average level of measured wage rigidity is predicted to have a 0.7 percentage point higher layoff rate, a 1.8 percentage point lower quit rate, and a 1.3 percentage point lower hire rate than an establishment with no measured wage rigidity. The relationship between wage rigidity and employment outcomes is generally amplified by movements in establishment revenue. An establishment with the sample average level of wage rigidity and a one standard deviation decrease in revenue growth is predicted to have a 1.5 percentage point higher layoff rate relative to an establishment with no wage rigidity. An establishment with the sample average level of wage rigidity and a one standard deviation increase in revenue growth is predicted to have a 2.6 percentage point lower hire rate than an establishment with no wage rigidity.

Using the empirical results to estimate the structural model via indirect inference suggests that wage rigidity reduces the average establishment's profits by 3.3 percent. The estimates suggest that the average establishment faces a per-worker menu cost of 774 euros annually of cutting nominal wages.

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Figure 2.1: Establishment Policy Functions with No Wage Rigidity

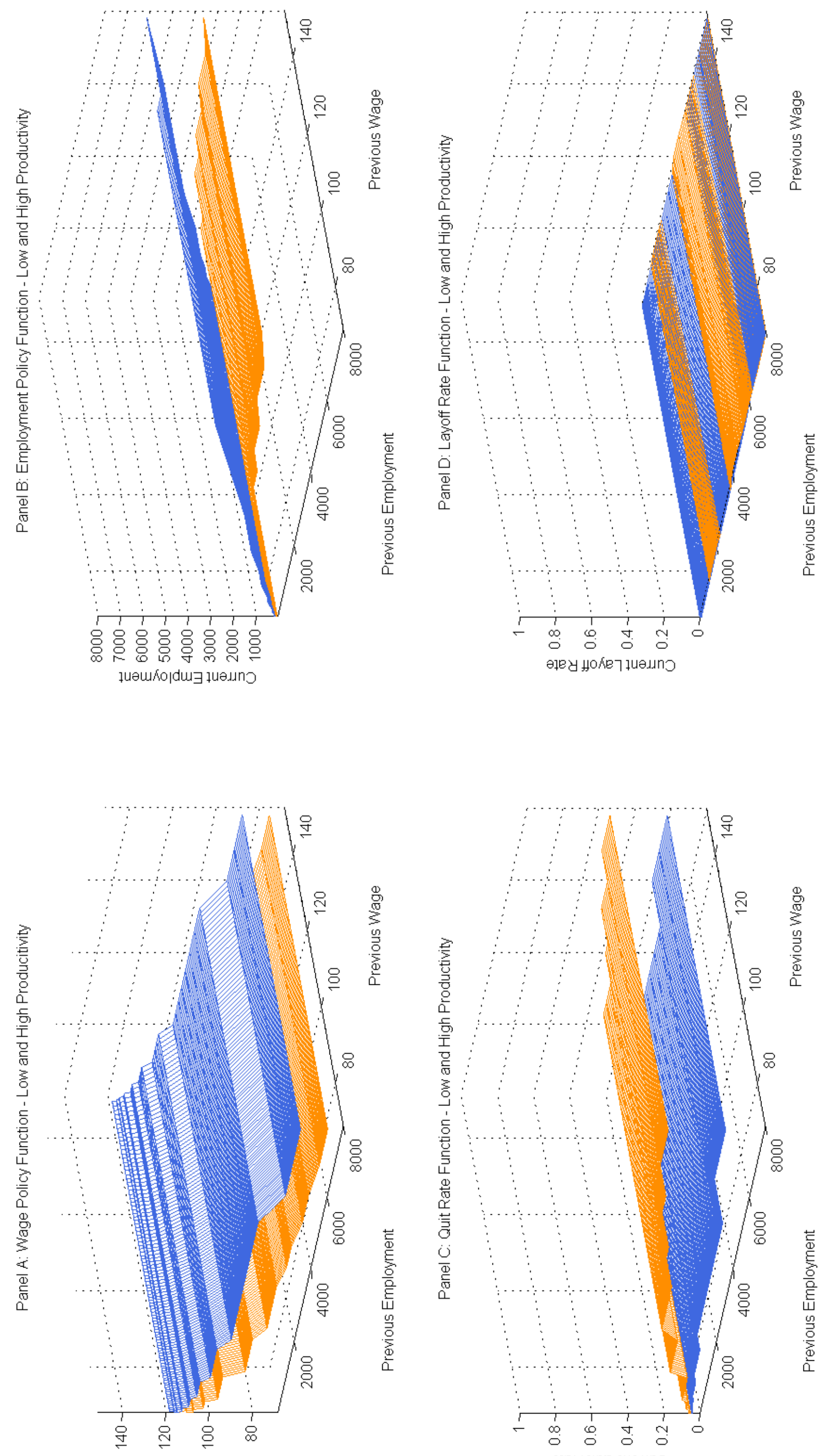


Figure 2.2: Establishment Policy Functions with Wage Rigidity

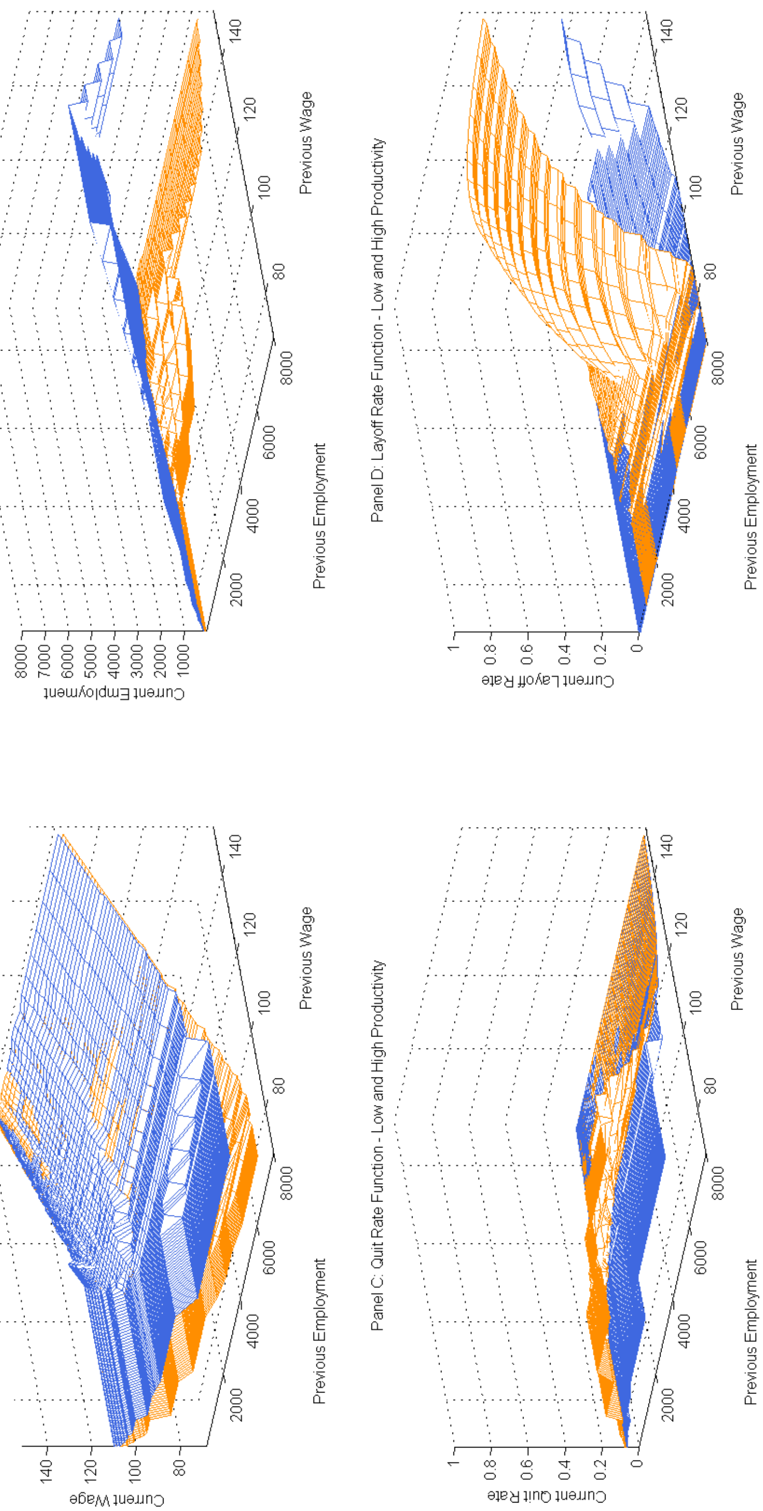


Figure 2.3: Wage Change Distributions with No Wage Rigidity

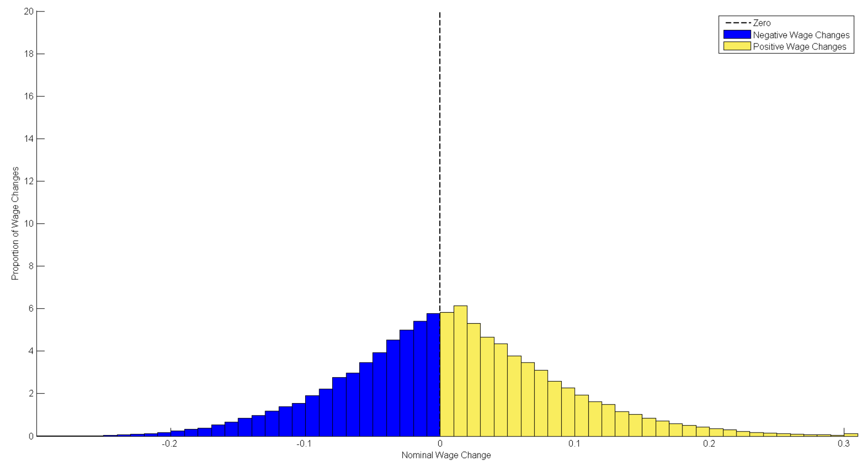


Figure 2.4: Wage Change Distributions with Wage Rigidity

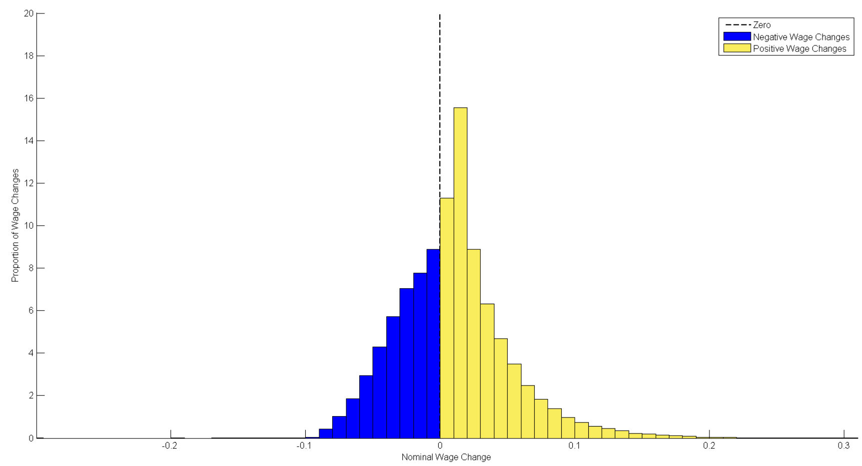


Figure 2.5: Simulated Moments with Different Levels of Wage Rigidity

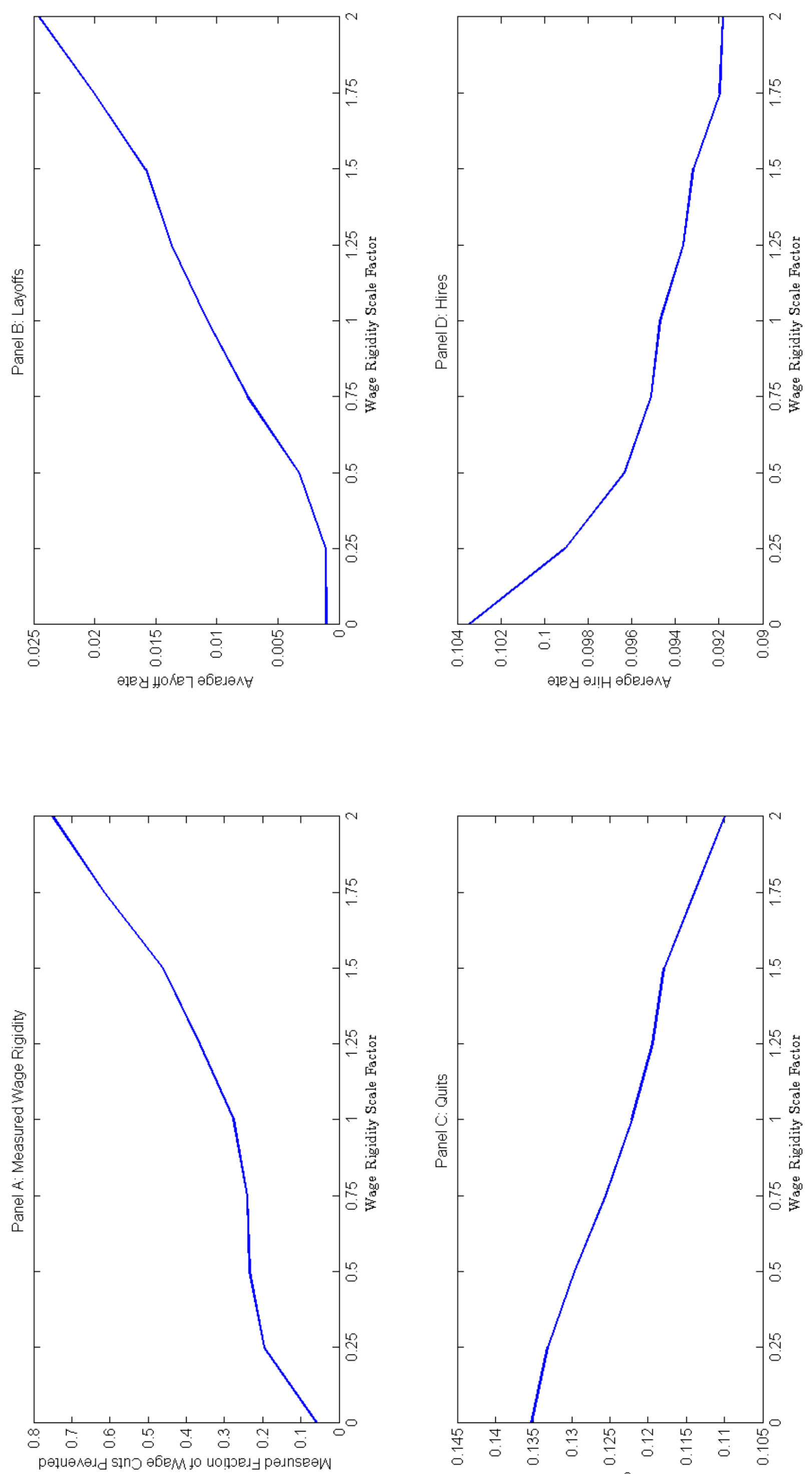


Figure 2.6: Aggregate Wage Change Distributions 1997 to 2003

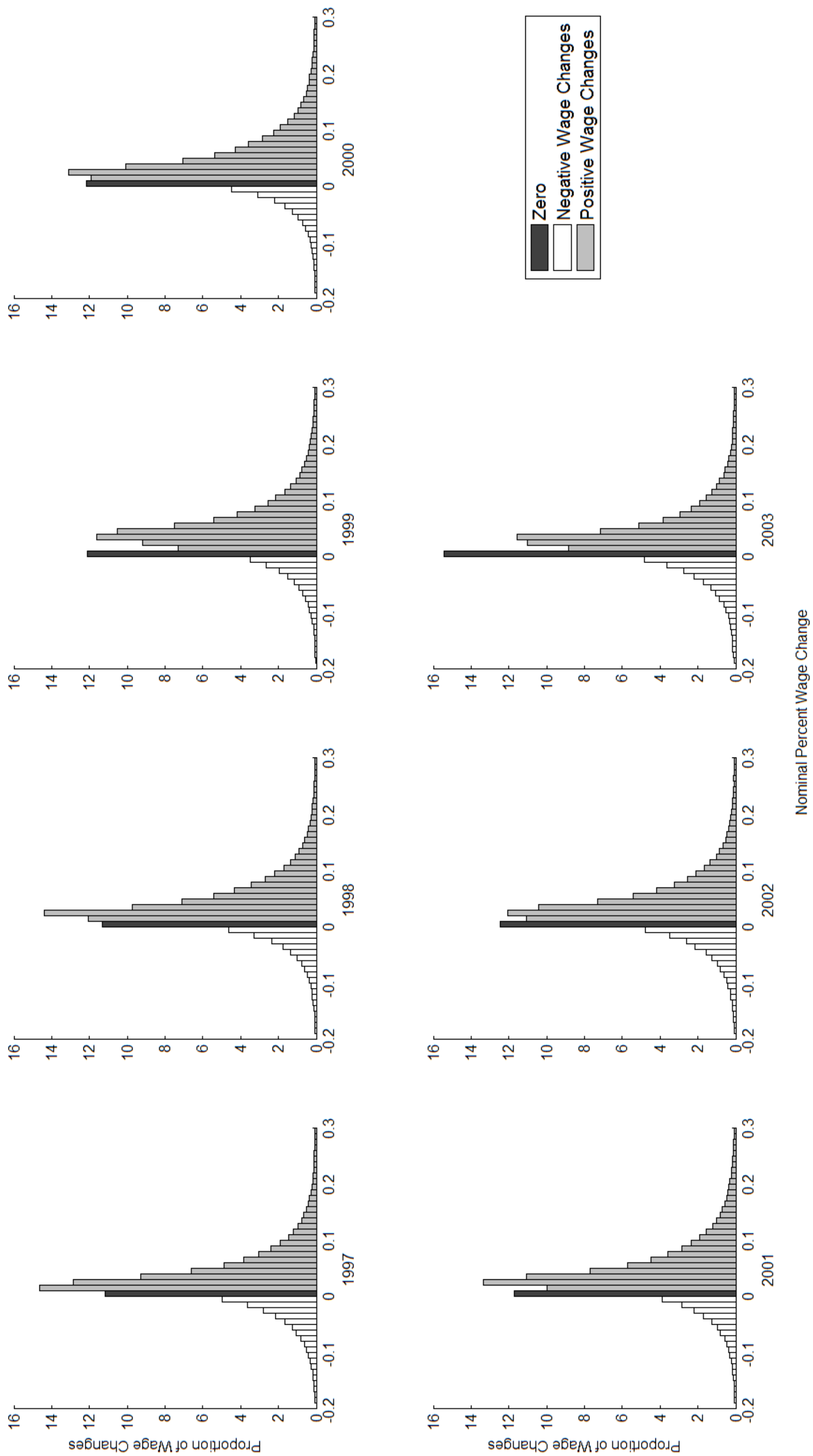


Figure 2.7: Illustration of Wage Rigidity Estimator

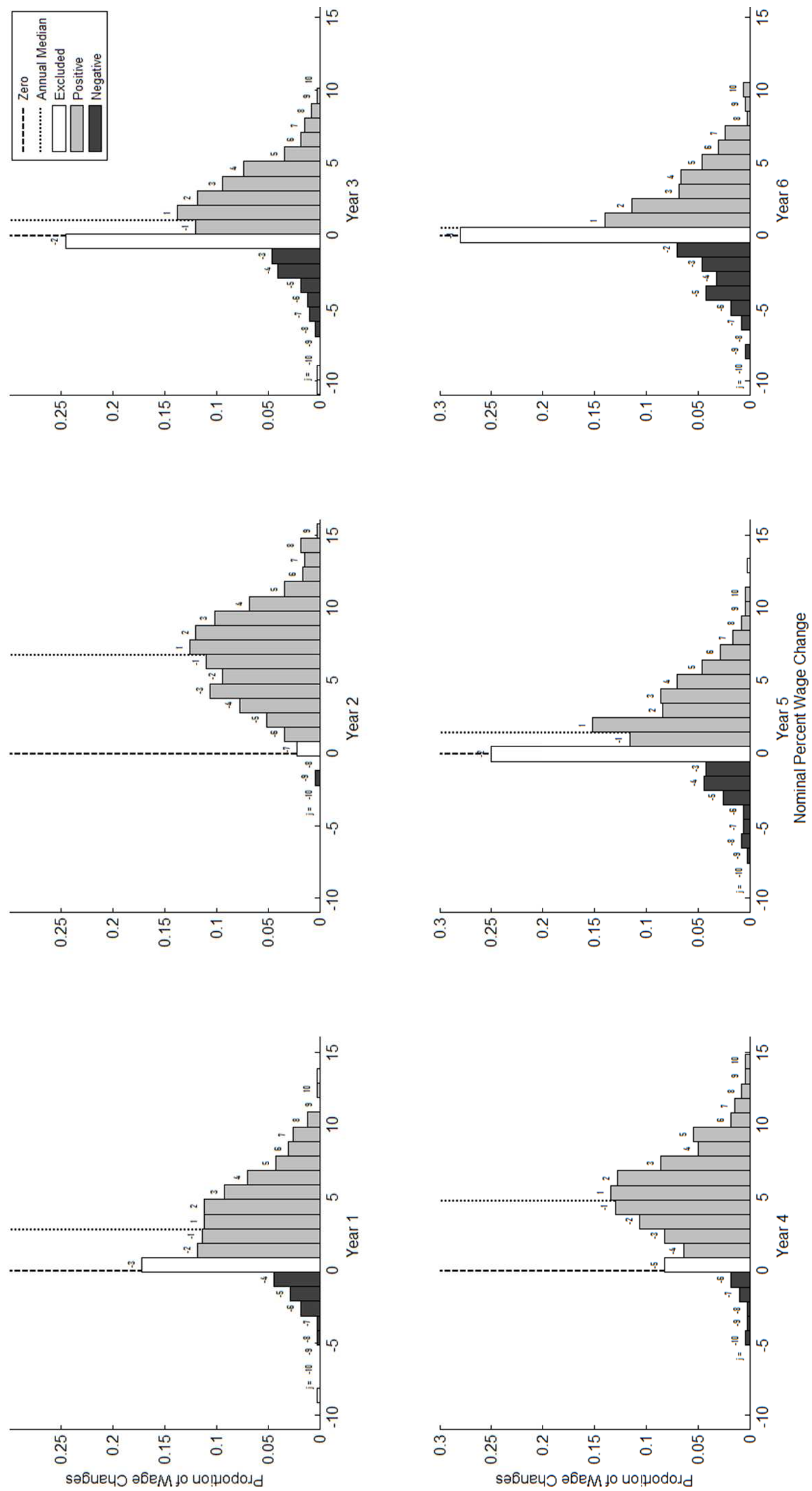


Figure 2.8: Illustration of Wage Rigidity Estimator: Estimating Counterfactual Distribution

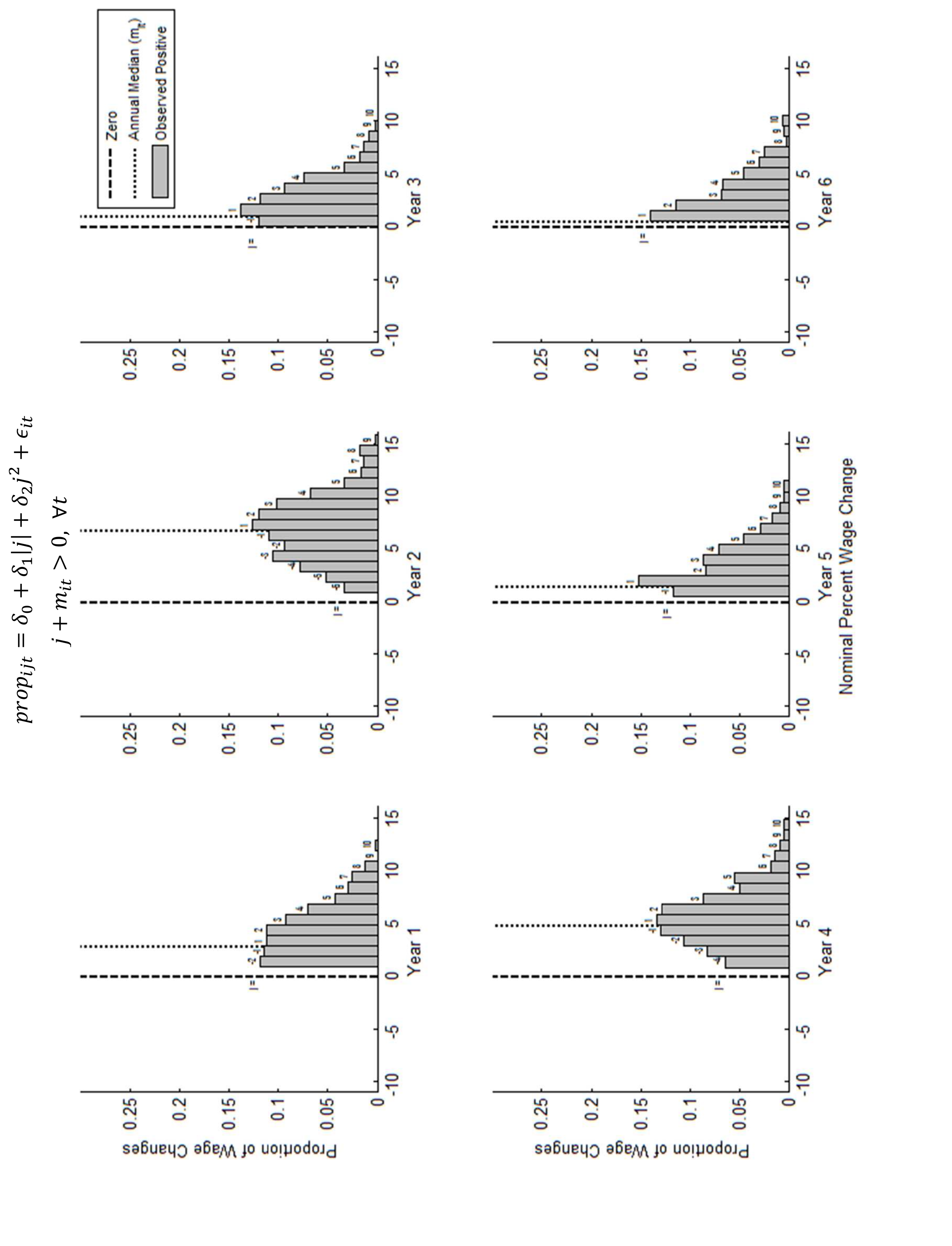


Figure 2.9: Illustration of Wage Rigidity Estimator: Counterfactual Negative Wage Changes

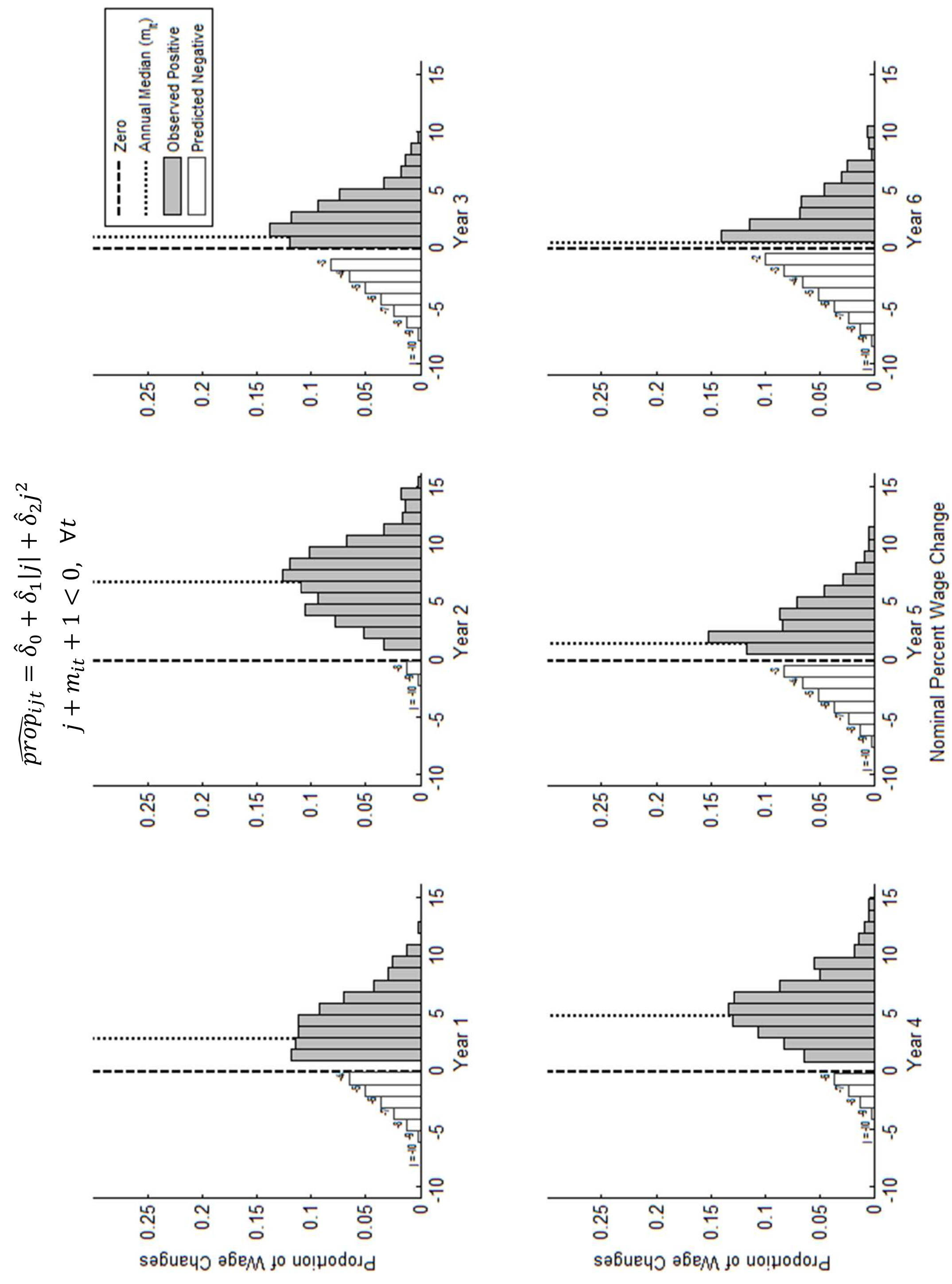


Figure 2.10: Illustration of Wage Rigidity Estimator: Estimating Missing Wage Cuts

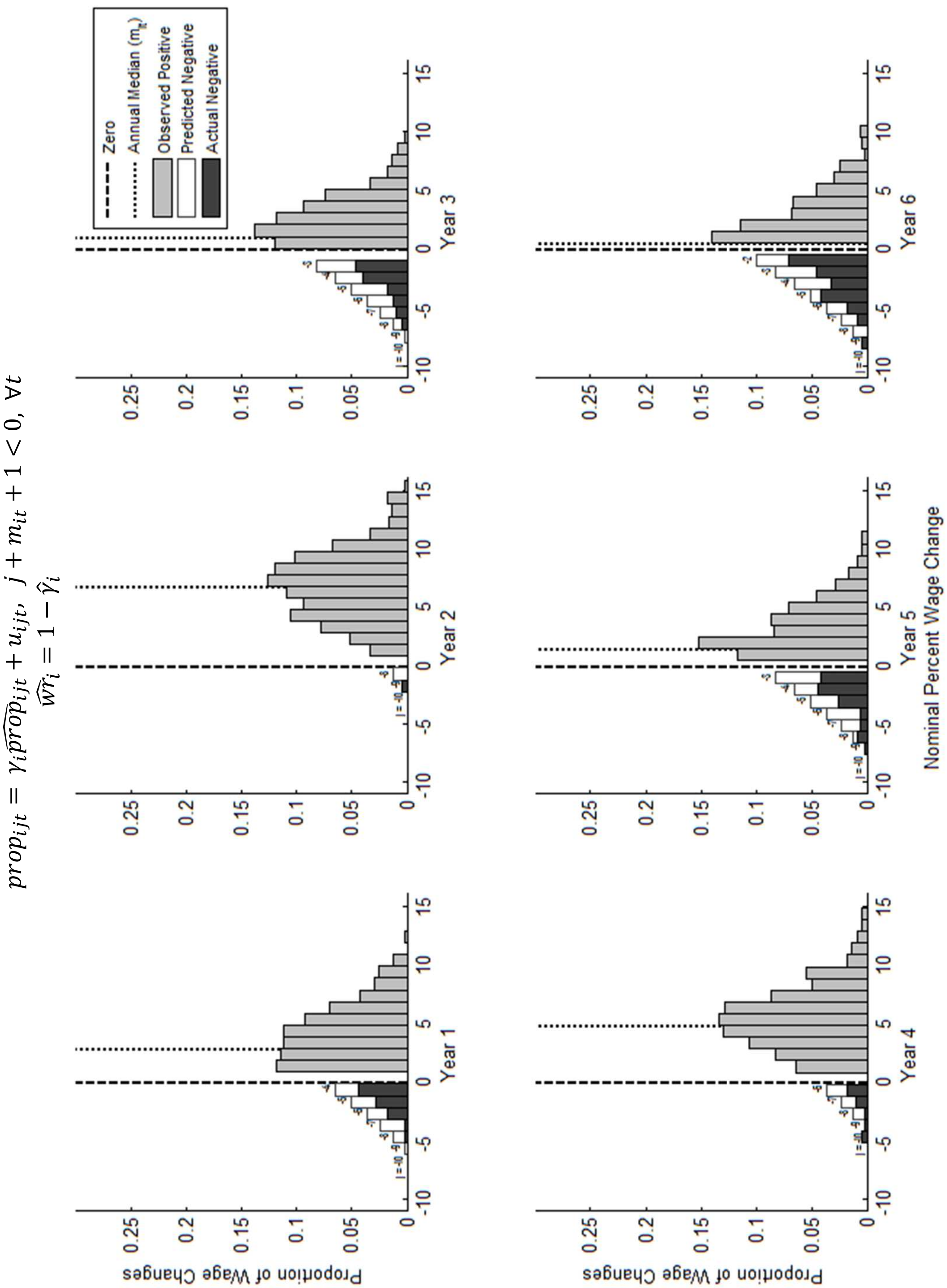


Figure 2.11: Layoff Regressions – Economic Significance

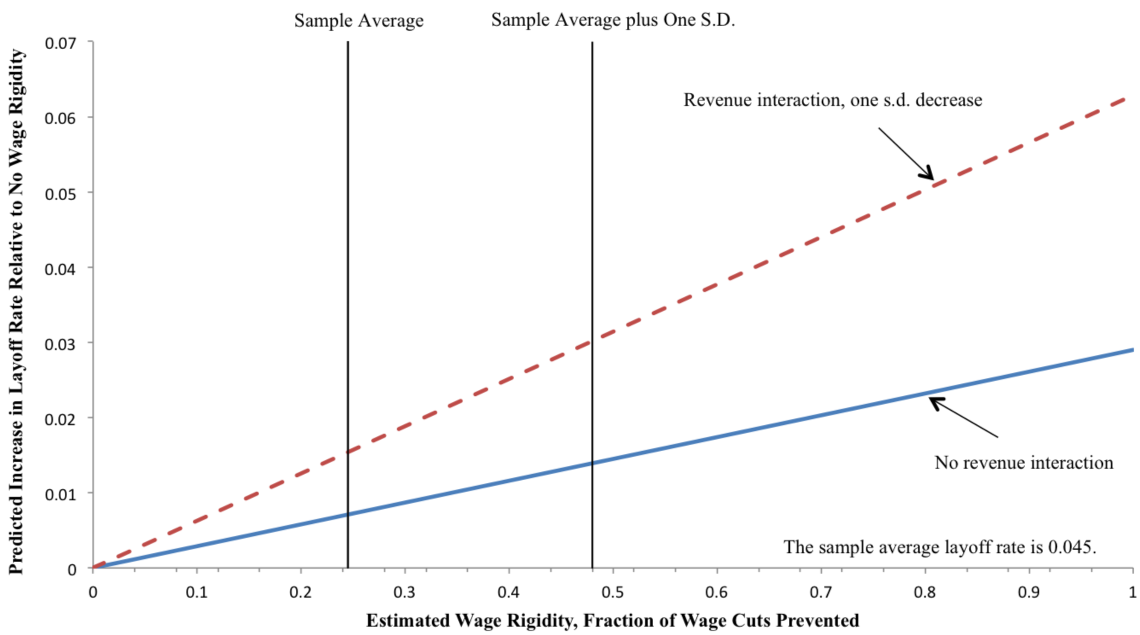


Figure 2.12: Quits Regressions – Economic Significance

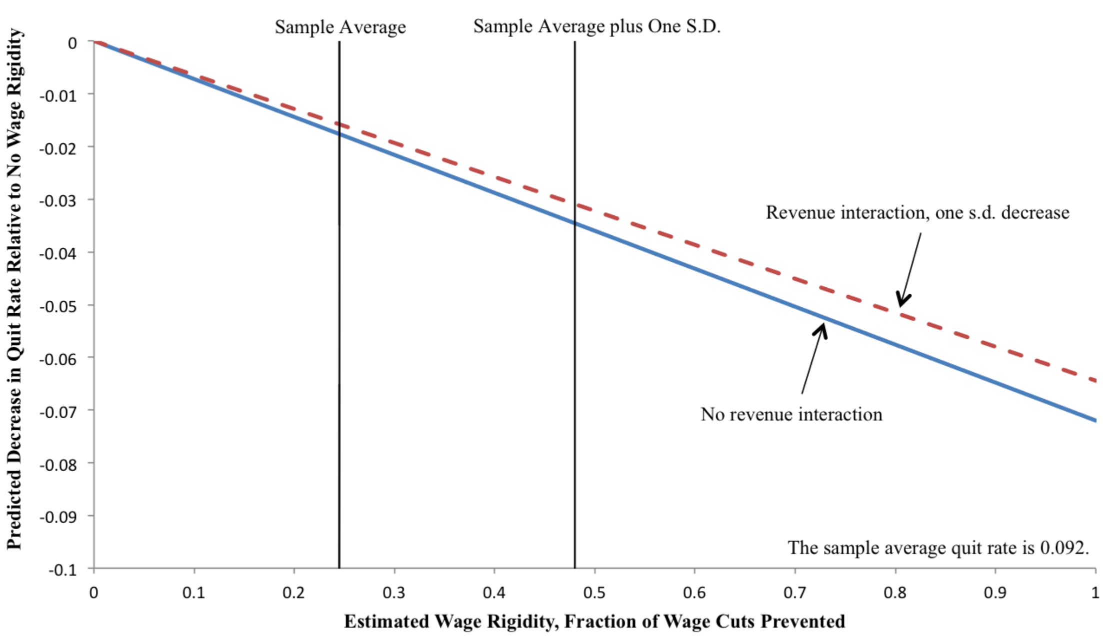


Figure 2.13: Hires Regressions – Economic Significance

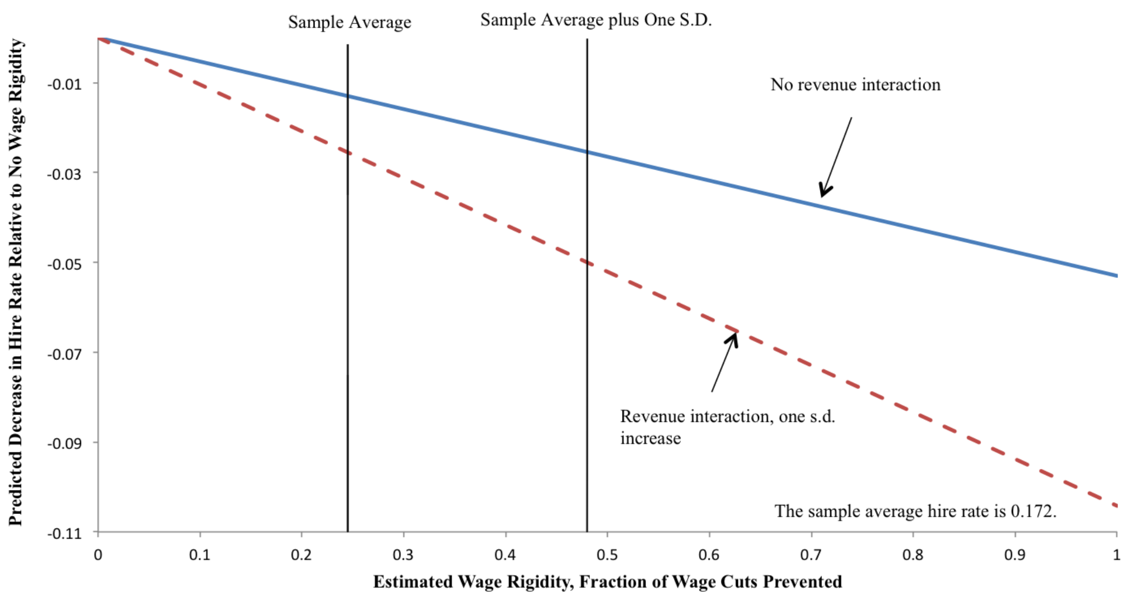


Table 2.1: Establishment-Level Descriptive Statistics

Number of Establishments	2,250
Sample Size, Establishment-Years	9,230
Mean Layoff Rate	0.045
	(0.092)
Mean Quit Rate	0.092
	(0.147)
Mean Hire Rate	0.172
	(0.327)
Mean Employees per Establishment	549
	(1,211)
Median Employees per Establishment	219
Mean Daily Wage, Level	102.899
	(66.207)
Median Daily Wage, Level	90.729
Average Revenue Growth	0.035
	(0.206)

Standard deviations in parentheses where applicable.

Table 2.2: Establishment-Level Wage Rigidity Estimates

	Estimated Wage Rigidity		
	Mean	Median	Standard Deviation
All Establishments	0.245	0.218	0.235
Supersector:			
Agriculture	0.132	0.094	0.320
Mining/Manufacturing	0.128	0.104	0.217
Energy/Water	0.289	0.268	0.283
Construction	0.032	0.034	0.190
Trade/Foodservice	0.212	0.173	0.284
Transportation	0.115	0.104	0.234
Finance	0.311	0.303	0.222
Real Estate	0.216	0.151	0.340
Public Administration	0.459	0.490	0.270
Administration	0.355	0.398	0.306

Wage rigidity is estimated as discussed in section 4. The wage rigidity estimator is a fixed characterisitic of the establishment and estimates which fraction of nominal wage cuts were prevented due to downard nominal wage rigidity.

Table 2.3: Wage Rigidity and Layoffs – Regression Results

Dependent Variable	Layoff Rate as a Fraction of Establishment Workforce		
	(1)	(2)	(3)
Wage Rigidity	0.029	0.029	0.025
	(0.005)	(0.005)	(0.006)
Positive Revenue Growth		0.005	0.012
		(0.005)	(0.007)
Negative Revenue Growth		-0.058	-0.015
		(0.010)	(0.014)
Wage Rigidity x Positive Revenue Growth			-0.038
			(0.024)
Wage Rigidity x Negative Revenue Growth			-0.193
			(0.044)
Work Council	-0.058	-0.052	-0.058
	(0.004)	(0.003)	(0.004)
R-squared	0.339	0.342	0.344
N	9,230	9,230	9,230

Standard errors in parentheses. Unit of observation is establishment-year. Layoffs are defined as a percentage of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated as described in section 4 and is fixed by establishment over sample period. Each regression includes a set of etablishment characteristics, individual charactersistics, and year dummies as controls. Establishment characteristics include a set of controls for the median year-over-year wage change, occupational mix and dummies for sector, federal state, and establishment size. Individual characteristics include controls for gender and workers' education. Positive revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is positive and zero otherwise. Negative revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is negative and zero otherwise. All regressions require establishments to have at least 20 employees in a given year, are weighted by the square root of the number of employees, and cover the period 1997 to 2003.

Table 2.4: Wage Rigidity and Quits – Regression Results

Dependent Variable	Quit Rate as a Fraction of Establishment Workforce		
	(1)	(2)	(3)
Wage Rigidity	-0.072	-0.072	-0.056
	(0.008)	(0.008)	(0.009)
Positive Revenue Growth		0.037	0.079
		(0.008)	(0.012)
Negative Revenue Growth		-0.215	-0.225
		(0.015)	(0.022)
Wage Rigidity x Positive Revenue Growth			-0.192
			(0.038)
Wage Rigidity x Negative Revenue Growth			0.043
			(0.069)
Work Council	-0.047	-0.047	-0.047
	(0.006)	(0.006)	(0.006)
R-squared	0.341	0.355	0.357
N	9,230	9,230	9,230

Standard errors in parentheses. Unit of observation is establishment-year. Quits are defined as a percentage of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated as described in section 4 and is fixed by establishment over sample period. Each regression includes a set of etablishment characteristics, individual charactersistics, and year dummies as controls. Establishment characteristics include a set of controls for the median year-over-year wage change, occupational mix and dummies for sector, federal state, and establishment size. Individual characteristics include controls for gender and workers' education. Positive revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is positive and zero otherwise. Negative revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is negative and zero otherwise. All regressions require establishments to have at least 20 employees in a given year, are weighted by the square root of the number of employees, and cover the period 1997 to 2003.

Table 2.5: Wage Rigidity and Hires – Regression Results

Dependent Variable	Hire Rate as a Fraction of Establishment Workforce		
	(1)	(2)	(3)
Wage Rigidity	-0.053	-0.059	-0.053
	(0.015)	(0.014)	(0.015)
Positive Revenue Growth		0.048	0.103
		(0.014)	(0.019)
Negative Revenue Growth		0.025	0.106
		(0.025)	(0.036)
Wage Rigidity x Positive Revenue Growth			-0.261
			(0.064)
Wage Rigidity x Negative Revenue Growth			-0.362
			(0.116)
Work Council	-0.115	-0.115	-0.115
	(0.010)	(0.010)	(0.010)
R-squared	0.639	0.637	0.639
N	9,230	9,230	9,230

Standard errors in parentheses. Unit of observation is establishment-year. Hires are defined as a percentage of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated as described in section 4 and is fixed by establishment over sample period. Each regression includes a set of etablishment characteristics, individual charactersistics, and year dummies as controls. Establishment characteristics include a set of controls for the median year-over-year wage change, occupational mix and dummies for sector, federal state, and establishment size. Individual characteristics include controls for gender and workers' education. Positive revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is positive and zero otherwise. Negative revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is negative and zero otherwise. All regressions require establishments to have at least 20 employees in a given year, are weighted by the square root of the number of employees, and cover the period 1997 to 2003.

Table 2.6: Calibrated Parameter Values

Parameter	Description	Value	Source
α	Returns to Scale in Production	0.650	Labor's Share of Value Added
$\overline{\delta}$	Average Quit Rate	0.092	Average Sample Quit Rate
$\overline{w}$	Average Daily Wage	103	Average Sample Wage
γ	Wage Elasticity of Quit Rate	5.75	Auxiliary Regression
$\ln \bar{z}$	Average Establishment-Wide Productivity	4.03x10^5	Auxiliary Regression
ψ_z	Persistence of Establishment-Wide Productivity	0.096	Auxiliary Regression
σ_z^2	Variance of Establishment-Wide Productivity Shock	0.325	Auxiliary Regression
c_l	Firing Cost	4,695	Redundancy Costs
$\bar{u}$	Worker Type Productivity	1	Normalization
π	Deterministic Inflation Rate	0.013	Average Inflation
β	Establishment Discount Rate	0.924	Real Interest Rate
λ_1	Linear Cost of Downward Wage Adjustment	0	Choice

The parameters for returns to scale in production, the average quit rate, the average daily wage, the wage elasticity of the quit rate, average establishment-wide productivity, the persistence of establishment-wide productivity, and the variance of the establishment-wide productivity shock are estimated through a series of empirical regressions and sample averages directly from the data, described in sections 2.3 and 2.6. The persistence of worker type productivity, the variance of the worker type productivity shock, the hiring cost parameters, and the nominal wage adjustment cost parameters are estimated by indirect inference as described in section 2.6. Firing costs are set to match German redundancy costs from the World Bank's Doing Business project, taking the average of the cost for workers with 1 year of tenure and workers with 5 years of tenure. The deterministic inflation rate is the average consumer price inflation rate from the World Bank's World Development Indicators. The real interest rate is the "lending interest rate" from the World Bank's World Development Indicators minus the calibrated inflation rate.

Table 2.7: Empirical and Simulated Moments

Moment	Sample Value	Simulated Value
Layoff Regression Coefficient on Wage Rigidity	0.029	0.037
Quit Regression Coefficient on Wage Rigidity	-0.072	-0.038
Hire Regression Coefficient on Wage Rigidity	-0.053	-0.014
Average Wage Rate	102.90	99.23
Standard Deviation of Percentage Wage Change	0.074	0.041
Measured Level of Wage Rigidity	0.245	0.285
Predicted Increase in Layoff Rate Associated with Wage Rigidity	0.007	0.011
Average Hire Rate	0.172	0.095
Average Negative Wage Change	-0.042	-0.028

The coefficients on wage rigidity in the layoff, quit, and hire regressions are from column 1 of tables 3, 4, and 5, respectively. Measured level of wage rigidity is mean wage rigidity for all establishments from table 2, calculated as described in section 4.

Table 2.8: Estimated Parameter Values

Parameter	Description	Value
ψ_u	Persistence of Worker Type Productivity	0.767
		(0.087)
σ_u^2	Variance of Worker Type Productivity Shock	0.507
		(0.133)
ϕ_1	Linear Hiring Cost	$5.29\text{x}10^4$
		$(2.26\text{x}10^4)$
ϕ_2	Quadratic Hiring Cost	$1.35\text{x}10^5$
		$(6.48\text{x}10^4)$
λ_0	Menu Cost of Downward Wage Adjustment, Euros	774
		(149)
λ_2	Quadratic Cost of Downward Wage Adjustment, Euros	$2.88\text{x}10^{-4}$
		$(1.74\text{x}10^{-5})$

Parameters are estimated by the indirect inference procedure described in section 2.6. Standard errors are reported in parentheses.

Appendix

2.8 Appendix A: Monte Carlo Simulations of Wage Rigidity Estimator

This section tests the performance of the estimator of wage rigidity proposed in section 3 using Monte Carlo simulations. We simulate wage change distributions for 500 establishments facing different levels of wage rigidity and generate the number of years' worth of wage changes observed in the sample for each establishment as a random integer uniformly distributed between three and seven. Next, we generate the number of employees per establishment as a random integer uniformly distributed between 15 and 500; the number of employees is fixed over the simulation period. For each establishment we generate the proportion of nominal wage cuts that will be prevented by downward nominal wage rigidity as a random variable uniformly distributed over the interval [0, 1]: $wr_i \sim U[0, 1]$.

To simulate counterfactual nominally flexible wage change distributions for each establishment in each year, begin by drawing the mean of the establishment-year wage change distribution from a normal distribution with a mean of four percent and a standard deviation of four percent: $\mu_{it} \sim N(.04, .04^2)$. We draw the standard deviation of the counterfactual wage change distribution from a uniform distribution over the interval [0 .05]: $\sigma_{it} \sim U[0, .05]$. We then draw the counterfactual flexible wage changes for each year from the normal distribution $\Delta \ln w_{ijt}^{cf} \sim N(\mu_{it}, \sigma_{it}^2)$, where $\Delta \ln w_{ijt}^{cf}$ is the counterfactual flexible log wage change of individual j at establishment i from year $t-1$ to year t.

Wage rigidity is introduced by replacing proportion wr_i of the counterfactual negative wage changes with positive wage changes that are distributed $N(.001, .005)$ to allow for prevented wage cuts to result in wage changes that are not exactly equal to zero. Wage cuts are chosen to replace randomly: there is no tendency for smaller wage cuts to be more likely to be prevented, for example. Finally, compression in the wage change distribution in the face of wage rigidity is introduced by multiplying counterfactual wage changes by a compression factor of $1 - 0.5wr_i$. That is, wage changes at an establishment with no wage rigidity will not be affected by wage compression, while wage changes at an establishment with 100 percent wage rigidity will be compressed by 50 percent. Introducing wage compression does not substantially change the simulation results.

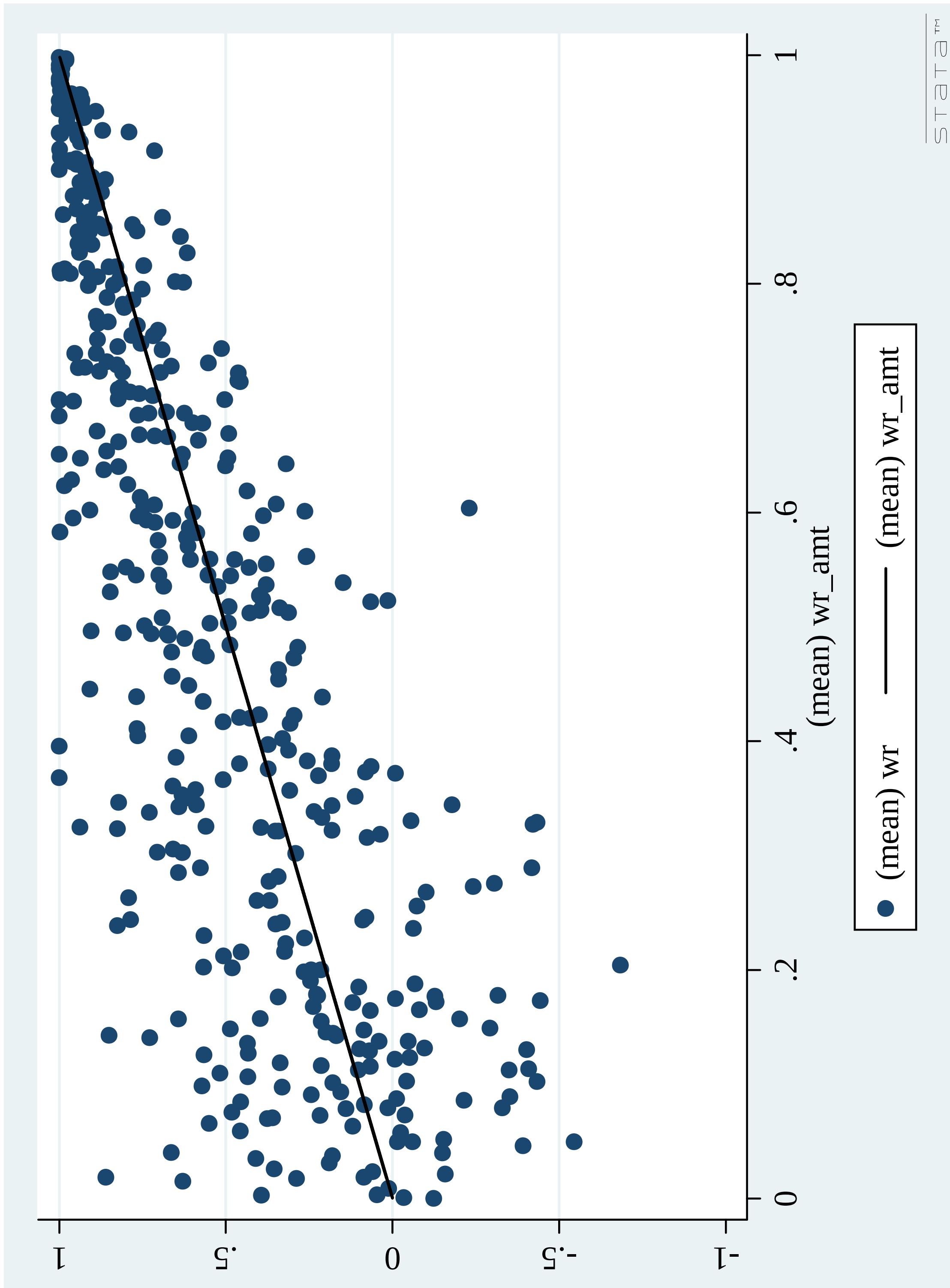
The simulations use a reduced form method of simulating the wage change distributions in order to test whether the estimator of $\widehat{wr}_i$ provides unbiased estimates of the true wr_i in a setting in which a constant fraction of counterfactual wage cuts are prevented by wage rigidity. In contrast, there is not a direct correspondence in the theoretical model presented in sections 5 and 7 between the cost of wage adjustment parameters λ_0, λ_1, and λ_2 and a fixed proportion of counterfactual wage cuts prevented.

Figure 2.14 displays the estimated and actual proportions of counterfactual wage cuts prevented by wage rigidity in these simulations. A regression of the form

$$\widehat{wr}_i = \alpha + \beta wr_i + u_i \tag{2.16}$$

gives an estimate for $\hat{\alpha}$ of 0.036 with a standard error of 0.024 and an estimate for $\hat{\beta}$ of 0.982 with a standard error of 0.041. Therefore, $\hat{\alpha}$ and $\hat{\beta}$ are not statistically distinguishable from 0 and 1, respectively. We interpret these results as suggesting that this estimator of wage rigidity is likely to be unbiased in this context. The standard error of the regression is 0.24, nearly equal in magnitude to the standard deviation of the true amount of wage rigidity, which is 0.29. As discussed in the main text, this noise is likely to cause attenuation bias in the estimates of the relationship between wage rigidity and employment outcomes, meaning the true associations are likely to be larger than estimated in this paper.

Figure 2.14: Monte Carlo Simulations of Wage Rigidity Estimates



2.9 Appendix B: Aggregate German Wage Change Distributions, 1976-2005

This section displays the aggregate German wage change distributions from 1976 to 2005. The data is taken from the Sample of Integrated Labor Market Biographies (SIAB) described in section 5.3. The dataset contains a 2 percent random sample of workers liable to social security in West Germany during the sample period. The histograms display nominal percent wage changes for job stayers. The sample includes workers whose earnings are top-censored at the social security contribution limit; these workers are excluded from the histograms as their true wage is not known.

Figure 2.15: Aggregate Wage Change Distributions 1976 to 1990

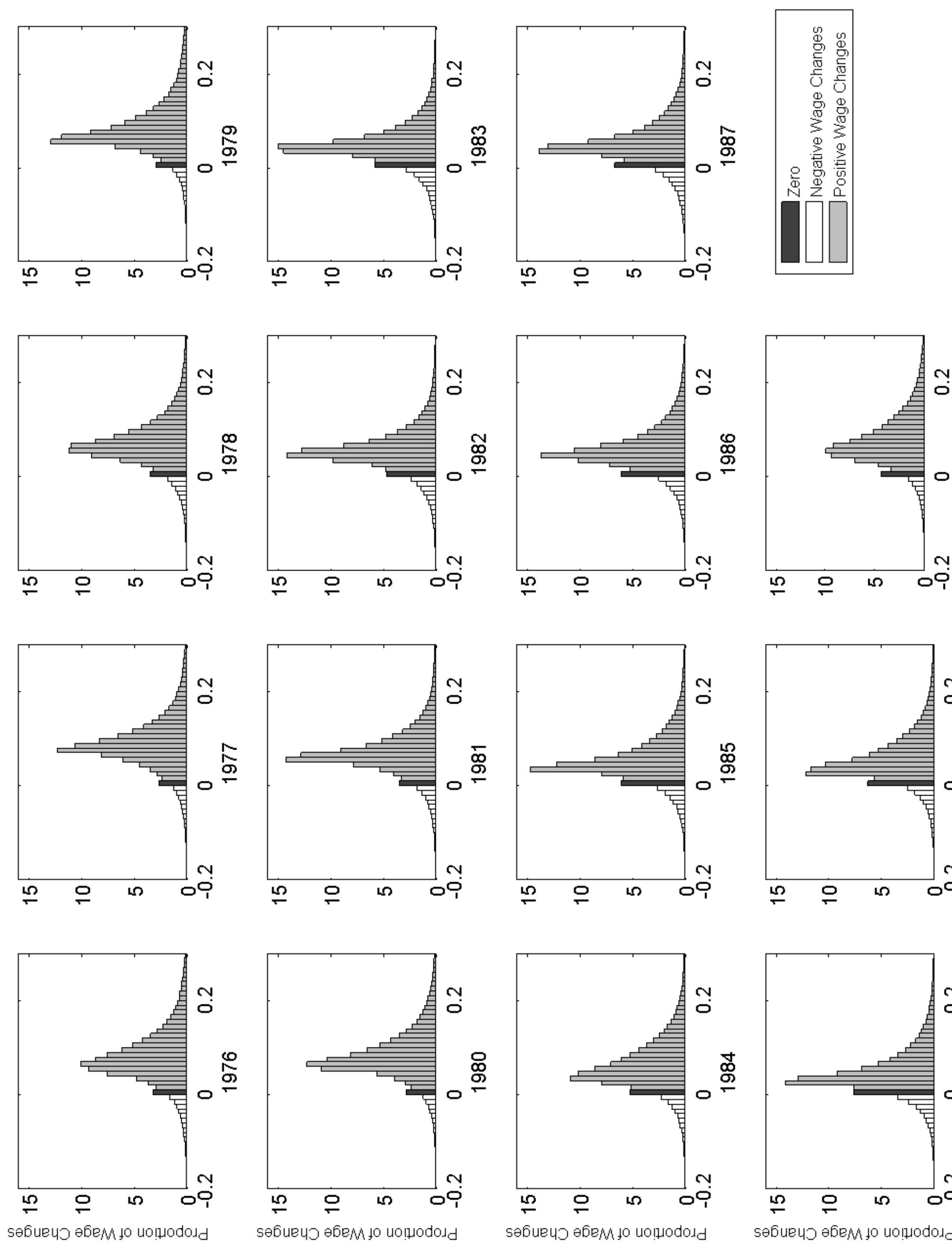
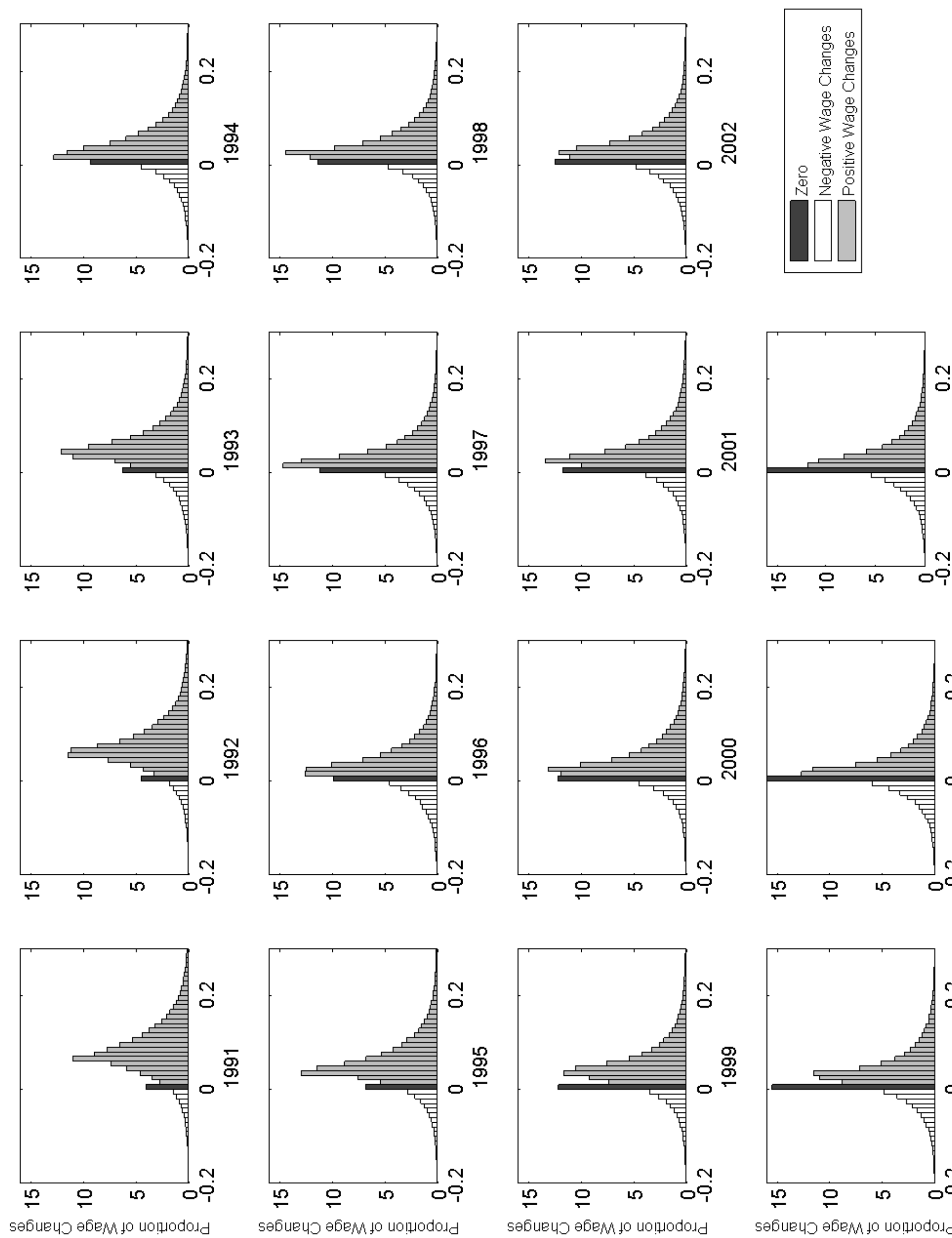


Figure 2.16: Aggregate Wage Change Distributions 1991 to 2005



CHAPTER III

Wage Rigidity and Employment Outcomes: Theory

3.1 Introduction

This paper explores the theoretical relationship between downward nominal wage rigidity and employment outcomes at the establishment decision making level. This paper provides the simple analytics of the model from chapter 2 that novelly models wage rigidity through a downward nominal wage adjustment cost function.

The analytics of the model show that as the cost of cutting nominal wages increases for an establishment, the establishment reduces its wage bill in response to a negative shock through laying off workers and reducing hires. Further, the model shows that downward nominal wage rigidity keeps the wages of its marginal workers above their marginal product and thus above the wage the marginal workers would receive in a fully flexible environment (and the wage the marginal worker would receive in the open market), reducing the rate of quits at an estabishment. The reduction in the quit rate thereby requires an even larger increasing in layoffs and reduction in hires by the establishment to optimally reduce its wage bill.

This paper then calibrates the model and performs a series of counterfactual policy simulations to study the effects of varying degrees of inflation on establishment profits and layoffs, quits, and hires in the presence of wage rigidity. The counterfactual policy simulation implies that inflation mitigates the costs of wage rigidity, as increasing the inflation rate reduces the profit reduction associated with wage rigidity from a 3.3 percent loss at the sample average annual inflation rate of 1.3 percent, to a 1.7 percent loss at an annual inflation rate of 5 percent. Further, the counterfactual policy simulation show that inflation can “grease the wheels” of the labor market, as higher levels of inflation facilitate real wage cuts even when nominal wage cuts are costly to achieve.

Akerlof, Dickens, and Perry (1996) find that wage rigidity makes a statistically insignificant difference in macroeconomic time series estimates of a Phillips Curve equation in the postwar period. Lebow et al. (1999) estimate that the non-accelerating inflation rate of unemployment is positively correlated with inflation, contrary to what would be predicted by an important role for nominal wage rigidity.

Barro (1977) argues that in a long-term employment relationship, the wage at a particular point in time is less important than the path of wages over the life of the relationship. Therefore, apparently rigid wages may reflect optimal long-term contracting rather than difficulties in wage adjustment, and may not have meaningful implications for employment outcomes. Elsby (2009) notes that forward-looking, wage-setting firms will compress wage increases in the presence of wage rigidity. Smaller wage increases in good times reduce the need for wage cuts in the face of an adverse shock.

The model in this paper incorporates the Elsby (2009) wage-compression effect, as it examines the optimal dynamic wage and employment decisions of an establishment that faces a real resource cost of cutting nominal wages. When this cost is large enough, the establishment will not cut wages in response to a negative shock to the marginal revenue product of labor, but will lay off workers instead. However, the effect of wage rigidity is not limited to the layoff margin of employment adjustment.

When wage rigidity prevents workers' wages from being cut, the workers will be less likely to quit. Prospective difficulties in cutting wages in the future also reduce forward-looking establishments' incentive to hire workers in the present. The model predicts that wage rigidity has meaningful effects on short-run employment outcomes, consistent with the empirical results.

Fehr and Goette (2005) use firm-level data from two Swiss firms (a large firm and a medium-sized firm) from the 1990s (a period of low inflation in Switzerland) to show that nominal wage rigidity is robust in a low inflation environment, constituting a considerable obstacle to real wage adjustments, and that the lack of nominal wage cuts due to wage rigidity correlate strongly to unemployment. This paper contributes to this portion of the literature by using a large sampe of establishments with individual-level wage data to quantify the relationship between the effects of wage rigidity and inflation.

Finally, this paper indirectly relates to a literature studying the sources of downward nominal wage rigidity. Capbell III and Kamlani (1997) survey 184 firms to investigte the sources of wage rigidity and find reducing turnover through quits, implicit contracts, and preventing reductions in effort associated with wage cuts as the primary reasons firms are hesitant to cut nominal wages. Bewley (1999) surveys over 300 executives and business leaders asking why firms and establishments appear to be reluctant to reduces its workers nominal wages and shows that a fear in the reduction of worker morale associated with wage cuts would reduce worker effort and thereby reduce worker productivity. While the theoretical model remains agnostic as to the sources of wage rigidity, a reduction in worker morale and worker effort, implicit contracts, and costs associated with worker turnover are consistent with the notion that cutting nominal wages is costly to the firm and establishment. This paper contributes by examining whether those costs can be mitigated through increases in inflation.

The paper proceeds as follows. Section 3.2 derives the analytics of the establishment decision making problem in the presence of wage rigidity and explores the model's theoretical underpinnings. Section 3.3 performs a series of counterfactual poilicy simulations under various levels of the inflation rate to study whether inflation "greases the wheels" of the labor market. Section 5.6 concludes.

3.2 Simple Analytics of the Establishment Decision Making Model with Wage Rigidity

This section provides the simple analytics of the partial equilibrium, establishment decision making model with wage rigidity proposed in section 2 of chapter 2. The model contains J heterogeneous worker types within a representative establishment. For simplicity, however, the analytical results of this section will suppress the J worker types and solve the establishment's problem as if workers are homogenous. Since each of the worker types, j, solve identical dynamic optimization problems, the analytical results presented here will be mostly identical[1] to the dynamic optimization problem solved by each worker type, j in chapter 2.

The infinitely-lived, representative establishment uses one input of production, labor, and maximizes the discounted stream of expected per period profits,

$$\Pi \;=\; pn^{\alpha} - wn - c_h(h, n_{-1})h - c_{\ell}\ell - g(w, w{-}1)n$$

where n is the stock of labor used in production, α governs returns to scale in production, w is the wage rate labor, h and ℓ are the number of employees the establishment hires and lays off, and $c_h(\cdot)$ and c_{ℓ} are costs of hiring and layoffs of labor, respectively.

[1]The only difference between the optimzation problem solved in chapter 2 and that solved in this chapter comes in the form of the prodctivity process. With heterogeneous workers types, each worker type, j, has a productivity process, u_j, that interacts with an establishment-level productivity process, z, to give overall worker productivity, p_j. Since all workers are homogenous in this section, overall worker productivity will be p across all workers. Accordingly, this section only models a productivity process, p, and makes no mention to processes z and u. This simplification does not affect the results.

$g(\cdot)$ is the cost of wage adjustment for labor, while p shifts the marginal revenue product of labor. p may be conceptualized as either the level of labor productivity or the level of its output price; for concreteness, this paper refers to p as productivity. The establishment is assumed to be concerned exclusively with real payoffs, and all variables above are specified in real terms. The rate of price inflation enters the model through the cost of wage adjustments function as described below.

The establishment chooses the current wage rate, w, level of hires, h, and level of layoffs, ℓ, to solve the following dynamic optimization problem:

$$\begin{aligned} V\left(p, w_{-1}, n_{-1}\right) &= \max_{w,h,l} pn^{\alpha} - wn - c_h(h, n_{-1})h - c_{\ell}\ell \\ &\quad -g(w, w_{-1})n + \beta E\left[V\left(p', w, n\right)\right] \end{aligned} \tag{3.1}$$

subject to

$$\ln p = (1 - \psi_p) \ln \bar{p} + \psi \ln p_{-1} + \varepsilon_p,\ \varepsilon_p \sim N\left(0, \sigma_p^2\right) \tag{3.2}$$

$$n = \left(1 - \delta\left(w\right)\right) n_{-1} + h - \mathbb{1}^{+}\ell \tag{3.3}$$

$$\begin{aligned} g\left(w, w_{-1}\right) &= \lambda_0 \mathbb{1}_{(1+\pi)w<w_{-1}} + \lambda_1 \left(w_{-1} - (1+\pi)w\right) \mathbb{1}_{(1+\pi)w<w_{-1}} \\ &\quad +\lambda_2 \left(w_{-1} - (1+\pi)w\right)^2 \mathbb{1}_{(1+\pi)w<w_{-1}} \end{aligned} \tag{3.4}$$

where (3.2) is the establishment's productivity process, which evolves according to a mean-reverting, $AR\left(1\right)$ process with error term, ε_p, (3.3) is the labor stock equation of motion, and (3.4) is the cost function of cutting wages facing the establishment and is how downward wage rigidity enters the model.

The paramterized functional form of the quit rate of labor is

$$\delta\left(w\right) = \bar{\delta}\left(\frac{w}{\overline{w}}\right)^{-\gamma}, \quad \gamma > 0$$

where $\bar{\delta}$ is the economy-wide average quit rate, and γ governs the degree of competition in the labor market: as γ increases, the quit rate becomes more sensitive to wages. The establishment also faces quadratic costs of hiring labor, given by

$$c_h(h, n_{-1}) = \phi_1 \left(\frac{h}{n_{-1}} \right) + \phi_2 \left(\frac{h}{n_{-1}} \right)^2$$

which allows for increasing or decreasing returns to scale.

The analytics of the estbalishment decision making model under wage rigidity on around the wage cutting cost function, $g(\cdot)$, in (3.4). Note that (3.4) only takes a non-zero value when the establishment cuts the nominal wage from the previous period (i.e. when $(1+\pi)w < w_{-1}$) and zero everywhere else. Therefore, $g(\cdot)$ is not a smooth function with a kink at the zero wage change and not differentiable when $(1+\pi)w = w_{-1}$. It follows that the derivative of $g(\cdot)$ with respect to the current period wage rate, w, takes the following piecewise form:

$$\frac{\partial g(w, w_{-1})}{\partial w} = \begin{cases} 0 & \text{if } (1+\pi)w > w_{-1} \\ \text{undefined} & \text{if } (1+\pi)w = w_{-1} \\ -\Big(1+\pi\Big)\Big[\lambda_1 + 2\lambda_2\big(w_{-1} - (1+\pi)\,w\big)\Big] & \text{if } (1+\pi)w < w_{-1} \end{cases} \tag{3.5}$$

The intuition from (3.5) is straightforward: If the establishment raises the workers' wages, then no cost of wage adjustment is incurred. If the establishment cuts the workers' wages, then the cost of wage adjustment increases linearly in λ_1 and increases quadratically in λ_2.

Similarly crucial to the analytics is how the labor stock equation of motion, (3.3), changes as the current period wage rate, w, changes. Differentiating (3.3) with resprect to w yields

$$\begin{aligned}
\frac{\partial n(w,h)}{\partial w} &= -\frac{\partial \delta(w)}{\partial w} n_{-1} \\
&= -\left[(-\gamma)\,\bar{\delta}\left(\frac{w}{\bar{w}}\right)^{-\gamma-1}\left(\frac{1}{\bar{w}}\right)\right] n_{-1} \\
&= \left(\frac{\gamma}{\bar{w}}\right)\bar{\delta}\left(\frac{w}{\bar{w}}\right)^{-\gamma}\left(\frac{\bar{w}}{w}\right) n_{-1} \\
\frac{\partial n(w,h)}{\partial w} &= \left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1} > 0 \qquad (3.6)
\end{aligned}$$

That is, the evolution of the establishment's labor stock depends positively on the wage rate. As the current period's wage rate increases, the number of quits in the current period (as a fraction of the end of the previous period's employment, n_{-1}) falls, and the establishment retains a larger portion of its workforce. The term, $\left(\frac{\gamma}{w}\right)$, in (3.6) captures the effect that the establishment's competitiveness in the labor market has on the evolution of the labor stock: as the establishment becomes more competitive, it will retain more of its workers.

Since the model is a discrete choice problem, the establishment will never find it optimal to both hire workers and lay off workers in the same period.[2] Therefore, the analysis of this section is partitioned into five distinct cases: Case 1 presents the case of no layoffs and wage increases ($\ell = 0,\ g(\cdot) = 0$). Case 2 presents the case with no layoffs and wage cuts ($\ell = 0,\ g(\cdot) > 0$). Case 3 presents the case of no layoffs and wages unchanged ($\ell = 0,\ g(\cdot) = 0,\ \frac{\partial g(\cdot)}{\partial w}$ is undefined). Case 4 presents the case of layoffs and wages unchanged (($\ell > 0,\ h = 0,\ g(\cdot) = 0,\ \frac{\partial g(\cdot)}{\partial w}$ is undefined). Case 5 presents the case of layoffs and wage cuts ($\ell > 0,\ h = 0,\ g(\cdot) > 0$).[3]

[2] The model presented in chapter 2 achieves both hires and layoffs in the same period through the heterogeneous J worker types. Distinct worker types experience hires and layoffs, but a single group worker type, j, will never experince both hires and layoffs in a given period.

[3] There is not a sixth case where an establishment both lays off workers and increases the wage. For an establishment that is laying off workers, wages must be less than or equal to w_{-1}, as there is no need to keep wages high to recruit and keep workers while the establishment actively reduces its workforce. Therefore, an establishment laying off workers must either keep wages unchanged or reduce wages in the model.

3.2.1 Case 1: No Layoffs, Nominal Wage Increases ($\ell = 0$, $g(w, w_{-1}) = 0$)

In the case of no layoffs ($\ell = 0$) and wage increases, (3.1) becomes

$$V(p, w_{-1}, n_{-1}) = \max_{w,h} pn^{\alpha} - wn - c_h(h, n_{-1})h + \beta E[V(p', w, n)] \tag{3.7}$$

The establishment chooses the current wage rate, w, and the level of hires, h, to solve its dynamic optimization problem. Further, $\ell = 0$ implies that the labor stock equation of motion constraint, (3.3), becomes

$$n = (1 - \delta(w))\, n_{-1} + h \tag{3.8}$$

so that the current period labor stock, n, is simply the number of workers retained from the previous period plus the number of workers hired.

Solving the establishment's dynamic optimization problem yields the following first order condition with respect to the wage rate, w:

$$\begin{aligned}
\frac{\partial V}{\partial w} \quad : \quad & \alpha p n^{\alpha-1}\frac{\partial n}{\partial w} - n - w\frac{\partial n}{\partial w} + \beta E\left[\frac{\partial V'}{\partial w}\right] = 0 \\
\implies \quad & \alpha p n^{\alpha-1}\left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1} + \left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1} \times \cdots \\
& \quad \beta E\left[\alpha p' n'^{\alpha-1}\Big(1 - \delta(w')\Big) + \phi_1\left(\frac{h'}{n}\right)^2 + 2\phi_2\left(\frac{h'}{n}\right)^3\right] = \cdots \\
& n + w\left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1} + \cdots \\
& \quad \left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1}\beta E\left[\Big(w' + g(w', w)\Big)\Big(1 - \delta(w')\Big)\right] + \cdots \\
& \quad \beta E\left[n'\frac{\partial g(w', w)}{\partial w}\right]
\end{aligned} \tag{3.9}$$

where

$$g\left(w', w\right) \begin{cases} 0 & \text{if} \quad (1+\pi)w' \geq w \\ \lambda_0 + \lambda_1\Big(w + (1+\pi)\, w'\Big) + \lambda_2\Big(w - (1+\pi)\, w'\Big)^2 & \text{if} \quad (1+\pi)w' < w \end{cases} \tag{3.10}$$

and

$$\frac{\partial g\left(w', w\right)}{\partial w} \begin{cases} 0 & \text{if} \quad (1+\pi)w' > w \\ \text{undefined} & \text{if} \quad (1+\pi)w' = w \\ \Big[\lambda_1 + 2\lambda_2\big(w - (1+\pi)\, w'\big)\Big] > 0 & \text{if} \quad (1+\pi)w' < w \end{cases} \tag{3.11}$$

The first order condition in (3.9) implies that the establishment sets the wage so that the marginal benefit of the wage change equals the marginal cost. The left-hand side of (3.9) shows the marginal benefit of a wage increase to the establishment. The first term on the left-hand side of (3.9) is the extra output associated with retaining more workers from a wage increase induced lower quit rate. The second term on the left-hand side of (3.9) is the continuation value of the extra output from those workers in future periods. An increase in the wage lowers this period's quit rate, and the establishment carries those workers into the future period. Those workers produce output in the future period if the future wage rate, w', does not induce those workers to quit in the future period. The third term on the left-hand side of (3.9) is the continuation value of the current period's wage increase on future hiring costs. Increasing wages today increases the establishment's stock of labor, n, through reducing quits and carries that stock into the future period. A larger stock of labor, n, carried into the future period reduces the future period's hiring costs, $c\left(h', n\right)$, as hiring costs are decreasing in the stock of labor.

The right-hand side of (3.9) shows the marginal cost of a wage increase to the establishment. The first two terms on the right-hand side of (3.9) show the establishment's total cost, this period, of the wage increase, as the establishment now pays all workers a higher wage. The third term on the right-hand side of (3.9) is the continuation cost of increasing wages today in the form of future wage costs. Increasing wages in the current period reduces the current period quit rate and increases the stock of labor, n, the establishment brings into the future period. When the wage increases in the current period, fewer workers quit and the establishment's stock of labor, n, carried into the future period increases. Of those workers who do not quit in the future given the future wage, w', the establishment pays them w' and must pay $g(w', w)$ from (3.10) if the establishment cuts the nominal wage from the current period to the future period. If the nominal wage change from the current to the future period is zero or positve, then $g(w', w)$ is zero as shown in (3.10), and the establishment only pays the wage rate, w', to those workers in the future.

The fourth and final term on the right-hand side of (3.9) is the continuation cost of increasing wages today in the form of future costs to cutting wages that increasing the wage in the current period imposes on the establishment in the future. The per worker cost of cutting wages in the future period, $g(w', w)$, is weakly increasing the the current period's wage, w, as shown in (3.11). If the establishment finds itself in a future productivity state that requires a wage cut, the cost of cutting the wage in the future period will be greater if the establishment increases the wage today than if the establishment did not increase the wage today. This term represents the Elsby (2009) effect of wage rigidity: forward-looking establishments will dampen wage increases in the current period knowing that the establishment will have to pay a future cost of cutting wages in response to a future adverse shock.

The first order condition with respect to hires, h, yields

$$
\begin{aligned}
\frac{\partial V}{\partial h} \quad : \quad & \alpha p n^{\alpha-1}\frac{\partial n}{\partial h} - w\frac{\partial n}{\partial h} - \cdots \\
& \left[c_h\left(h, n_{-1}\right) + \frac{\partial c_h\left(h, n_{-1}\right)}{\partial h}h\right] + \beta E\left[\frac{\partial V'}{\partial h}\right] = 0 \\
\Longrightarrow \quad & \alpha p n^{\alpha-1} + \cdots \\
& \beta E\left[\alpha p' n'^{\alpha-1}\left[1-\delta\left(w'\right)\right] + \phi_1\left(\frac{h'}{n}\right)^2 + 2\phi_2\left(\frac{h'}{n}\right)^3\right] = \cdots \\
& w + \left[2\phi_1\left(\frac{h}{n_{-1}}\right) + 3\phi_2\left(\frac{h}{n_{-1}}\right)^2\right] + \cdots \\
& \beta E\left[\left[w' + g\left(w', w\right)\right]\left[1-\delta\left(w'\right)\right]\right]
\end{aligned}
\tag{3.12}
$$

where $\frac{\partial n}{\partial h} = 1$. The first order condition in (3.12) implies that the establishment will recruit workers to the point where the marginal benefit of the hires equals the marginal cost. The left-hand side of (3.12) shows the marginal benefit of the hires. The first term is the extra output produced by the hired workers in the current period. The second term is the continuation value of the hires through both the expected increase in output the hired workers that do not quit produce in the future period and the expected reduction in the future period's total hiring costs associated with an increase in n. The right-hand side of (3.12) shows the marginal costs of the hires. The first term on the right-hand side of (3.12) is the cost the establshment incurs in the form of wages to employ the hires, whereas the second term captures the hiring costs incurred by the establishment to hire the new workers. The third term on the right-hand side of (3.12) is the continuation cost of hires in the form of expected future wages. If the workers hired in the current do not quit in the future period, the establishment must pay the workers wage w', and, if the establishment cuts the workers wage in the future period, the establishment must pay the nominal wage adjustment cost as described in (3.10).

3.2.2 Case 2: No Layoffs, Nominal Wage Cuts ($\ell = 0$, $g(w, w_{-1}) > 0$)

When the establishment cuts wages but does not lay off workers, it incurs a cost $g(w, w_{-1})$. Thus, the value function takes the form

$$\begin{aligned} V(p, w_{-1}, n_{-1}) &= \max_{w,h} pn^{\alpha} - wn - c_h(h, n_{-1})h \\ &\quad - g(w, w_{-1})\, n + \beta E\left[V(p', w, n)\right] \end{aligned} \tag{3.13}$$

The establishment chooses the current wage rate, w, and the level of hires, h, to solve its dynamic optimization problem. Since $\ell = 0$ as in section 3.2.1, the labor stock equation of motion remains as stated in (3.8).

Solving the establishment's dynamic optimization problem in the case of wage cuts yields the following first order condition with respect to the wage rate, w:

$$\begin{aligned} \frac{\partial V}{\partial w} \quad : \quad & \alpha p n^{\alpha-1}\frac{\partial n}{\partial w} - n - w\frac{\partial n}{\partial w} - \cdots \\ & \left(\frac{\partial g(w, w_{-1})}{\partial w} n + g(w, w_{-1})\frac{\partial n}{\partial w}\right) + \beta E\left[\frac{\partial V'}{\partial w}\right] = 0 \\ \Longrightarrow \quad & \alpha p n^{\alpha-1}\left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1} + \cdots \\ & (1+\pi)\Big[\lambda_1 + 2\lambda_2\left(w_{-1} - (1+\pi)w\right)\Big] n + \cdots \\ & \beta E\left[\alpha p' n'^{\alpha-1}\Big(1 - \delta(w')\Big) + \phi_1\left(\frac{h'}{n}\right)^2 + 2\phi_2\left(\frac{h'}{n}\right)^3\right] = \cdots \\ & n + \Big[w + g(w, w_{-1})\Big]\left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1} + \cdots \\ & \left(\frac{\gamma}{w}\right)\delta(w)\, n_{-1}\beta E\left[\Big(w' + g(w', w)\Big)\Big(1 - \delta(w')\Big)\right] + \cdots \\ & \beta E\left[n'\frac{\partial g(w', w)}{\partial w}\right] \end{aligned} \tag{3.14}$$

The intuition for (3.14) is similar to (3.9) in that the establishment will cut wages up until the point where the marginal cost of cutting wages equals the marginal benefit. The left-hand side of (3.14) shows the marginal cost to the establishment of cutting wages. The first term on the left-hand side of (3.14) represents the loss of output for the establishment as a result of a wage cut induced increase in quits. The second term on the left-hand side of (3.14) is the marginal cost the establishment pays in order to execute the nominal wage cut for all workers, n. The third and final term on the left-hand side of (3.14) represents the continuation cost of the wage cut in the form of expected loss in future output and expected future hiring costs. As wage induced quits increase this period, the establishment enters the next period with fewer workers, all else equal. If the establishment experiences a high productivity state in the future, it will have to hire more workers to meet the labor needs of the high prouctivity state than if workers never quit. The additional hires lead to additional hiring costs.

The right-hand side of (3.14) shows the marginal benefit to the establishment of cutting wages. The first term on the right-hand side, n, is the number of workers in the current period that take a wage cut and thus represents the decrease in the establishment's wage bill from all employed workers. The second term on the right-hand side represents the decrease in the establishment's wage bill as a result of the wage cut induced quits. For the workers that quit, the establishment no longer has to pay the wage rate, w, nor does the estbalishment have to pay the cost associated with cutting those workers' wages. The third term on the right-hand side of (3.14) is the continuation value to the establishment for cutting wages this period in the form of expected future wage savings from this periods wage induced quits. The final term on the right-hand side is the continuation value of the nominal wage cut in terms of paying the nominal wage adjustment cost in the future, where $\frac{\partial g(w',w)}{\partial w}$ is defined in (3.11). If the establishment cuts wages this period, it reduces the wage at which next period's nominal wage adjustment cost function binds, and thus makes it less likely that the establishment pays the nominal wage adjustment cost in the future.

The first order condition with respect to hires, h, in the case of wage cuts yields

$$\begin{aligned}
\frac{\partial V}{\partial h} \quad &: \quad \alpha p n^{\alpha-1}\frac{\partial n}{\partial h} - w\frac{\partial n}{\partial h} - \left[c_h\left(h, n_{-1}\right) + \frac{\partial c_h\left(h, n_{-1}\right)}{\partial h}h\right] - \cdots \\
& \qquad g\left(w, w_{-1}\right)\frac{\partial n}{\partial h} + \beta E\left[\frac{\partial V'}{\partial h}\right] = 0 \\
\Longrightarrow \quad & \alpha p n^{\alpha-1} + \cdots \\
& \qquad \beta E\left[\alpha p' n'^{\alpha-1}\left[1 - \delta\left(w'\right)\right] + \phi_1\left(\frac{h'}{n}\right)^2 + 2\phi_2\left(\frac{h'}{n}\right)^3\right] = \cdots \\
& w + g\left(w, w_{-1}\right) + \left[2\phi_1\left(\frac{h}{n_{-1}}\right) + 3\phi_2\left(\frac{h}{n_{-1}}\right)^2\right] + \cdots \\
& \qquad \beta E\left[\left[w' + g\left(w', w\right)\right]\left[1 - \delta\left(w'\right)\right]\right] \qquad (3.15)
\end{aligned}$$

The left-hand side of (3.15) is the marginal benefit to the establishment for hiring, and the intuition of the marginal benefit is the same as discussed with (3.12). The right-hand side of (3.15) is the marginal cost to the establishment for hiring workers. The intuition for the marginal cost of hiring is again the same as discussed with (3.12) with one exception: the marginal cost of hiring now contains an additional term, $g\left(\cdot\right)$, for the cost of cutting the hired workers' wages.

3.2.3 Case 3: No Layoffs, Nominal Wages Unchanged ($\ell = 0$, $g\left(w, w_{-1}\right) = 0$)

In the case of no layoffs and wages unchanged ($\left(1+\pi\right)w = w_{-1}$), the value function is the same as (3.7) in case 1. However, a wage first order condition does not exist when nominal wages are unchanged. The function, $g\left(w, w_{-1}\right)$, is not smooth at w_{-1} and, thus, not differentiable. Therefore, the profit function is not differentiable in w at w_{-1}. The hires first order condition is unchanged from (3.12) in case 1, and the intuition remains the same.

3.2.4 Case 4: Layoffs, Nominal Wages Unchanged ($\ell > 0$, $h = 0$, $g(w, w_{-1}) = 0$)

In the case of layoffs ($\ell > 0$) and no wage cuts from w_{-1}, it must be the case that $(1+\pi)w = w_{-1}$ (nominal wages unchanged) and $h = 0$. When laying off workers, the establishment has no need to keep wages high to recruit workers and keep employees. Therefore, (3.1) becomes

$$V(p, w_{-1}, n_{-1}) = \max_{w,h} pn^{\alpha} - wn - c_{\ell}\ell + \beta E\left[V(p', w, n)\right] \tag{3.16}$$

where $g(w, w_{-1}) = 0$ since nominal wages are unchanged from w_{-1}. The establishment chooses the current wage rate, w, and the level of layoffs, ℓ, to solve its dynamic optimization problem. Further, $\ell > 0$ and $h = 0$ imply that the labor stock equation of motion constraint, (3.3), becomes

$$n = (1 - \delta(w))\, n_{-1} - \ell \tag{3.17}$$

so that the current period labor stock, n, is simply the number of workers retained from the previous period less the number of workers the establishment lays off.

As in case 3, wages are unchanged ($(1+\pi)w = w_{-1}$), $g(w, w_{-1}) = 0$, and $\frac{\partial g(w,w_{-1})}{\partial w}$ is undefinined at w_{-1}. Therefore, the profit function is not differentiable at w_{-1} and a wage first order condition does not exist.

The first order condition with respect to layoffs, ℓ, yields

$$\begin{aligned} \frac{\partial V}{\partial \ell} \quad : \quad & \alpha p n^{\alpha-1}\frac{\partial n}{\partial \ell} - w\frac{\partial n}{\partial \ell} - c_{\ell} + E\left[\frac{\partial V'}{\partial \ell}\right] = 0 \\ \Longrightarrow \quad & w + \beta E\left[\Big(w' + g(w', w)\Big)\Big(1 - \delta(w')\Big)\right] = \alpha p n^{\alpha-1} + c_{\ell} + \cdots \\ & \beta E\left[\alpha p' n'^{\alpha-1}\Big(1 - \delta(w')\Big) + \phi_1\left(\frac{h'}{n}\right)^2 + 2\phi_2\left(\frac{h'}{n}\right)^3\right] \end{aligned} \tag{3.18}$$

where $\frac{\partial n}{\partial l} = -1$.

The first order condition in (3.18) implies that the establishment will lay off workers up until the point where the marginal marginal benefit of the layoffs equals the marginal cost. The left-hand side of (3.18) shows the marginal benefit to the establishment of laying off workers. The first term on the left-hand side of (3.18) represents the wage savings for the establishment as a result of not paying wages to laid off workers. The second term on the left-hand side of (3.18) represents the continuation value of the layoffs in the form of expected future wage savings: the establishment will not have to pay the wage rate in the future period, w', or the nominal wage adjustment cost (if the establishment cuts the nominal wage in the future period) $g\left(w', w\right)$, for the workers laid off in the current period that otherwise would not have quit.

The right-hand side of (3.18) shows the marginal cost to the establishment of laying off workers. The first term on the right-hand side of (3.18) represents the loss of output for the establishment as a result of laying off workers. The second term on the right-hand side represents the firing cost the establishment must pay for each layoff. The third term on the right-hand side of (3.18) is the continuation cost to the establishment for laying off workers this period in the form of expected future lost output and increased expected future hiring costs. As layoffs increase this period, the establishment enters the next period with fewer workers, all else equal. If the establishment experiences a high productivity state in the future, it will have to hire more workers to meet the labor needs of the high prouctivity state than if th establishment never laid off workers. The additional hires lead to additional hiring costs.

3.2.5 Case 5: Layoffs, Nominal Wage Cuts ($h = 0$, $g\left(w, w_{-1}\right) > 0$)

In the case of layoffs ($\ell > 0$) and nominal wage cuts from, it must be the case that $\left(1 + \pi\right) w < w_{-1}$ and $h = 0$. When laying off workers, the establishment has no need to keep wages high to recruit workers and keep employees. Therefore, (3.1) becomes

$$V\left(p, w_{-1}, n_{-1}\right) = \max_{w,h} pn^{\alpha} - wn - c_{\ell}\ell - g\left(w, w_{-1}\right)n + \beta E\left[V\left(p', w, n\right)\right] \quad (3.19)$$

where $g\left(w, w_{-1}\right) > 0$ since the establishment cuts nominal wages from w_{-1}. The establishment chooses the current wage rate, w, and the level of layoffs, ℓ, to solve its dynamic optimization problem. Further, as in case 4 from section A.4, $\ell > 0$ and $h = 0$ imply that the labor stock equation of motion constraint remains as is in (3.17), so that the current period labor stock, n, is simply the number of workers retained from the previous period less the number of workers the establishment lays off.

Solving the establishment's dynamic optimization problem in the case of wage cuts yields the same first order condition with respect to the wage rate, w, as in (3.14) from case 2 in section 3.2.2.

The first order condition with respect to layoffs, ℓ, yields

$$\begin{aligned} \frac{\partial V}{\partial \ell} \quad : \quad & \alpha p n^{\alpha-1}\frac{\partial n}{\partial \ell} - w\frac{\partial n}{\partial \ell} - c_{\ell} - g\left(w, w_{-1}\right)\frac{\partial n}{\partial \ell} + E\left[\frac{\partial V'}{\partial \ell}\right] = 0 \\ \implies \quad & w + g\left(w, w_{-1}\right) + \beta E\left[\left(w' + g\left(w', w\right)\right)\left(1 - \delta\left(w'\right)\right)\right] = \alpha p n^{\alpha-1} + \cdots \\ & c_{\ell} + \beta E\left[\alpha p' n'^{\alpha-1}\left(1 - \delta\left(w'\right)\right) + \phi_1\left(\frac{h'}{n}\right)^2 + 2\phi_2\left(\frac{h'}{n}\right)^3\right] \quad (3.20) \end{aligned}$$

where $\frac{\partial n}{\partial l} = -1$.

The intuition for (3.20) is the same as (3.18) in section 3.2.4 with the exception that the marginal benefit of a layoff now includes a term for the nominal wage adjustment cost funciton, $g\left(w, w_{-1}\right)$. This term captures a key tradeoff for the establishment, as the establishment avoids paying the nominal wage adjustment cost by instead laying off the worker.

3.3 Counterfactual Policy Simulation

The theoretical model can be used to conduct counterfactual policy simulations that examine the effects of changes in the structural parameters on economic outcomes. This section examines the effects of alternative deterministic inflation rates, holding all other model parameters constant, including the cost of nominal wage cut parameters λ_0, λ_1, and λ_2. Specifically, the simulations study an alternative low inflation rate of 0 percent and an alternative high inflation rate of 5 percent relative to the sample average, moderate inflation rate of 1.3 percent.

The model consists of 18 parameters. This chapter uses the calibration and estimation procedure discussed in section 6 of chapter 2 to pick the parameter values. Table 3.1 shows the parameter values used in the counterfactual inflation policy simulations below.

Figure 3.1 displays the simulated wage change distribution from the model with zero percent inflation. The asymmetry of this histogram is quite pronounced, and there is a large spike at the wage change bin containing a nominal wage change of zero. Furthermore, the distribution is highly compressed, displaying the forward-looking effects of wage rigidity as put for by Elsby (2009).

Figure 3.2 displays the simulated wage change distribution from the model with sample average, moderate inflation rate of 1.3 percent. The median nominal wage change shifts to the right, consistent with the higher inflation rate relative to the case in figure 3.1. The distribution of wage changes becomes slightly more dispersed, though the compression associated with the Elsby (2009) effect is still starkly noticeable.

Figure 3.3 displays the simulated wage change distribution from the model with 5 percent inflation. The median nominal wage change again shifts noticeably to the right, consistent with the increase in inflation to 5 percent. The distribution also becomes more symmetrical about the median and more dispersed, which is more reminiscent of a wage distribution with no wage rigidity, than of figures 1 and 2. This change occurs because the higher inflation rate facilitates real wage cuts compared to the lower inflation cases. This effect is visually evident in the histogram, as all nominal wage changes less than 5 percent are real wage cuts.

Table 3.2 shows the results of these simulations for profits and employment outcomes. Profits are lowest in the case of zero percent inflation, with a 4.1 percent reduction compared to the perfectly flexible wage model. In the five percent inflation case, profits are 1.7 percent lower than in the perfectly flexible wage model. The different inflation rates also yield different employment outcomes. The simulated layoff rate declines from 1.2 percent in the case with 0 percent inflation to 0.7 percent in the case with 5 percent annual inflation.[4] The simulated quit rate rises from 12.1 percent to 12.7 percent when inflation increases from 0 to 5 percent, and the simulated hire rate rises from 9.4 percent to 9.7 percent. All of these movements are consistent with the idea that higher inflation mitigates the effects of wage rigidity, thus providing some evidence in support of the idea that inflation may "grease the wheels" of the labor market. However, the model presented here does not address the welfare costs of inflation, so a benefit-cost analysis of different inflation rates is beyond the scope of this paper.

3.4 Conclusion

This paper explores the theoretical relationship between downward nominal wage rigidity and employment outcomes at the establishment decision making level. This paper provides the simple analytics of the novel model from chapter 2 that models wage rigidity through a downward nominal wage adjustment cost function presented in chapter 2. The model incorporates the Elsby (2009) effect that forward-looking

[4]It is worth recalling that wage rigidity is the only source of layoffs in the model, which is why the simulated rates are smaller than the sample average layoff rate.

firms incorporate future costs to cutting nominal wages by dampening wage increases. However, the analytics of this model and its estimation shows that, even when incorporating the forward-looking behavior of firms, downward nominal wage rigidity still leads to meaningful movements in layoffs, quits, and hires and reduces profits.

Using the calibrated and estimated model from chapter 2, this paper performs a series of counterfactual policy simulations to study the effects of varying degrees of inflation on establishment profits and layoffs, quits, and hires in the presence of wage rigidity. Increasing inflation from a moderate level of 1.3 percent (the average inflation rate over the sample period) to a high level of 5 percent leads to an increase in establishment profits, a decrease layoffs, and an increase in quits and hires. Further, the counterfactual policy simulation suggests that inflation can “grease the wheels” of the labor market by facilitating real wage cuts even when nominal wage cuts are costly to achieve.

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Figure 3.1: Wage Change Distributions with Wage Rigidity and 0 Percent Inflation

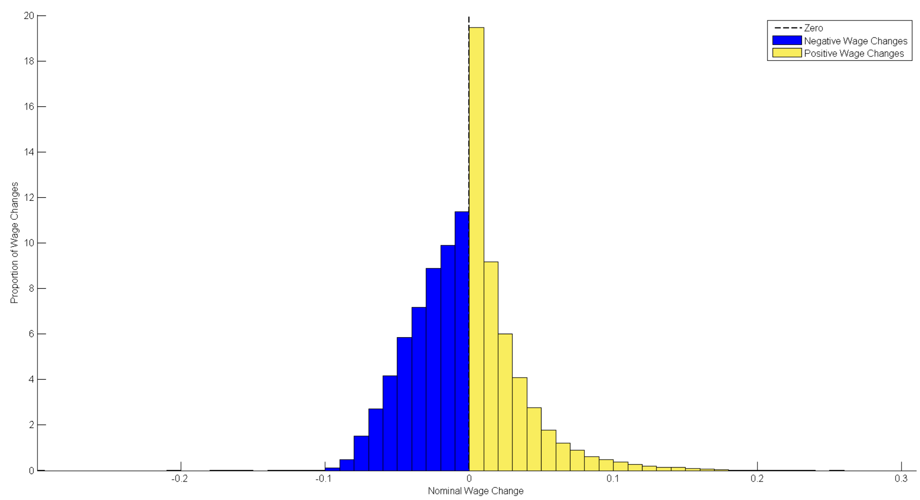

Figure 3.2: Wage Change Distributions with Wage Rigidity and 1.3 Percent Inflation

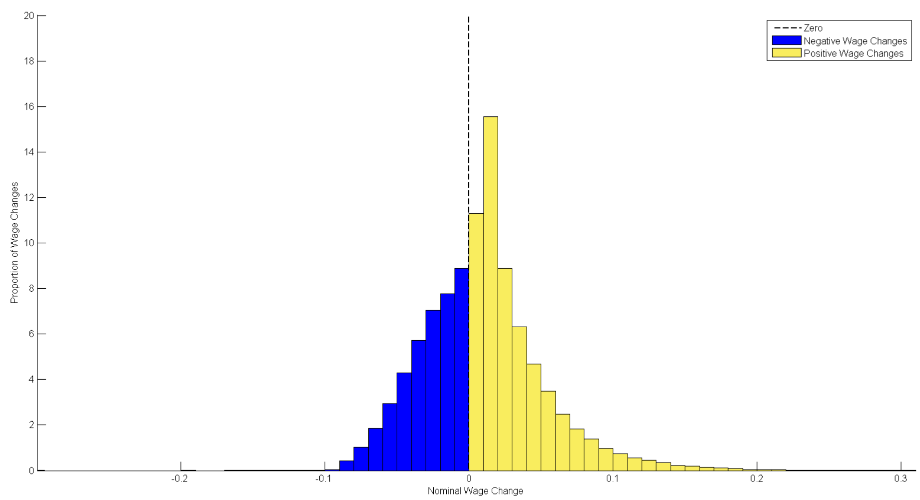

Figure 3.3: Wage Change Distributions with Wage Rigidity and 5 Percent Inflation

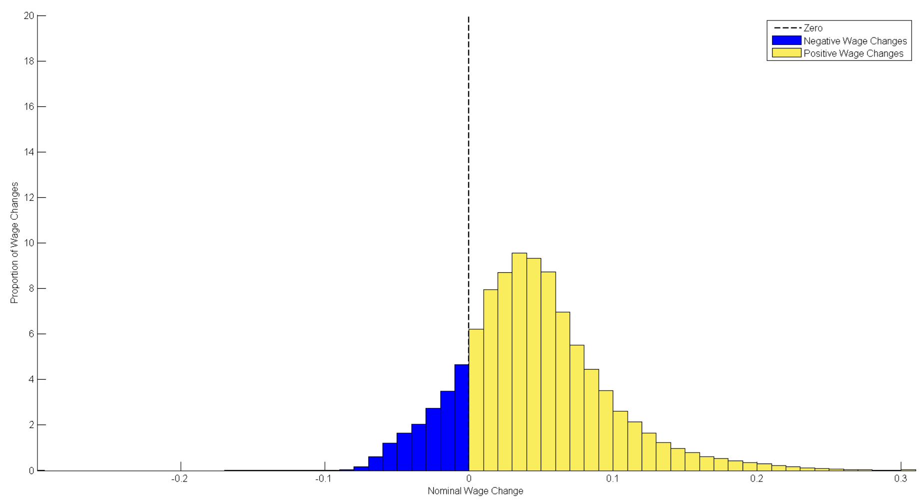

Table 3.1: Calibrated Parameter Values

Parameter	Description	Value	Source
α	Returns to Scale in Production	0.650	Labor's Share of Value Added
$\overline{\delta}$	Average Quit Rate	0.092	Average Sample Quit Rate
$\overline{w}$	Average Daily Wage	103	Average Sample Wage
γ	Wage Elasticity of Quit Rate	5.75	Auxiliary Regression
$\ln \bar{z}$	Average Establishment-Wide Productivity	4.03x10^5	Auxiliary Regression
ψ_z	Persistence of Establishment-Wide Productivity	0.096	Auxiliary Regression
σ_z^2	Variance of Establishment-Wide Productivity Shock	0.325	Auxiliary Regression
c_l	Firing Cost	4,695	Redundancy Costs
$\bar{u}$	Worker Type Productivity	1	Normalization
π	Deterministic Inflation Rate	0.013	Average Inflation
β	Establishment Discount Rate	0.924	Real Interest Rate
λ_1	Linear Cost of Downward Wage Adjustment	0	Choice
ψ_u	Persistence of Worker Type Productivity	0.767	Estimation, Chapter2
σ_u^2	Variance of Worker Type Productivity Shock	0.507	Estimation, Chapter2
ϕ_1	Linear Hiring Cost	5.29x10^4	Estimation, Chapter2
ϕ_2	Quadratic Hiring Cost	1.35x10^5	Estimation, Chapter2
λ_0	Menu Cost of Downward Wage Adjustment	774	Estimation, Chapter2
λ_2	Quadratic Cost of Downward Wage Adjustment	2.88x10^{-4}	Estimation, Chapter2

For details on parameter calibration and estimation, please see chapter 2, section 6. The parameters for returns to scale in production, the average quit rate, the average daily wage, the wage elasticity of the quit rate, average establishment-wide productivity, the persistence of establishment-wide productivity, and the variance of the establishment-wide productivity shock are estimated through a series of empirical regressions and sample averages directly from the data, described in chapter 2, section 6. The persistence of worker type productivity, the variance of the worker type productivity shock, the hiring cost parameters, and the nominal wage adjustment cost parameters are estimated by indirect inference as described in chapter 2, section 6. Firing costs are set to match German redundancy costs from the World Bank's Doing Business project, taking the average of the cost for workers with 1 year of tenure and workers with 5 years of tenure. The deterministic inflation rate is the average consumer price inflation rate from the World Bank's World Development Indicators. The real interest rate is the "lending interest rate" from the World Bank's World Development Indicators minus the calibrated inflation rate.

Table 3.2: Counterfactual Simulated Outcomes with Alternative Inflation Rates

	Annual Inflation Rate		
Outcome	0.00	0.013 (Baseline)	0.05
Profit Reduction	0.041	0.033	0.017
Layoff Rate	0.012	0.011	0.007
Quit Rate	0.121	0.122	0.127
Hire Rate	0.094	0.095	0.097

Profit reduction is relative to case with baseline parameters except for λ_0, λ_1, and λ_2 set to zero.

CHAPTER IV

Wage Rigidity and Employment Outcomes: The Role of Establishment Size and Age

4.1 Introduction

This paper uses a novel, individual-level, administrative dataset with worker-level wages and employer-employee matched identifiers to examine the differential relationship between establishment-level wage rigidity and the employment outcomes of layoffs, quits, and hires by establishment size and age.

The dataset is well suited to both examine which sectors, occupations, and education levels exhibit greater degrees of wage rigidity and to test whether there exists a differential, establishment-level relationship between measured wage rigidity and employment based on establishment size and age. In particular, the data contain total compensation histories for every worker in the sample. Since the compensation histories are taken from administrative data, they should theoretically be free of measurement error, a significant advantage of previous studies of wage rigidity that rely on survey responses. Furthermore, in addition to providing worker wage, the indivual data contains detailed information on the worker's industry, occupation and job status, and education level throughout each worker's entire employment history.

Since individuals in the dataset are randomly sampled, the study provides estimates of the degree of wage rigidity across supersectors, occupation classes, and education levels that are representative of the whole economy. The wage rigidity estimates show that supersectors that are more skill intensive, such as finance and public administration, exhibit greater degrees of wage rigidity that less skill intensive supersectors such as mining, trade, foodservice, and transportation. Similarly, the occupation and education level wage rigidity estimates show that more skilled occupations and more educated workers exhibit more measured wage rigidity than their less skilled and less educated counterparts.

Next, the paper makes use of the establishment-level identifiers, the dataset's sampling of all employees in an establishment, and data on each establishment's birth and death years to study whether the wage rigidity and employment outcomes relationship established in chapter 2 differs across estabishment size and age, where establishments are categorized into large and small establishments relative to the median establishment's level of employment and young and old establishments, with an establishment being labeled as young if it is no more than 5 years old and labeled as old if it is more than 5 years old. Splitting the sample by establishment size shows that large establishments, with an average of 25.5 percent of counterfactual wage cuts prevented by wage rigidity, exhibit more wage rigidity than smaller estabishments, with an average of 22.3 percent fo wage cuts prevented due to wage rigidity, and old establishments, with 24.5 percent of wage cuts prevented, exhibit more wage rigidity than younger establishments, with 16.1 percent of wage cuts pervented.

Finally, the empirical analysis shows that the establishment-level relationship between wage rigidity and employment outcomes is stronger for large establishments compared to small establishments and for young establishments compared to young. Wage rigidity increases layoffs at the average large establishment by 1 percentage point and decreases quits and hires by 2.5 and 2.8 percentage points, respectively, whereas wage rigidity and layoffs have no statisitcal relationship in small establishments and decreases quits and hires by 1.2 and 2.5 percentage points respectively. For the young establishments, wage rigidity increases layoffs by 4.7 percentage points and decreases quits and hires by 1.6 and 7.1 percentage points, respectively, on average, wherease the average old establishment exhibits no significant increase in layoffs and a 1.7 and 3.4 percentage point reduction in quits and hires, respectively.

This paper fits into a small but recently developing literature using large, administrative dataset to study the role of establishment and firm age play in employment dynamics. Davis and Haltiwanger (1992) use a large, administrative, establishment-level dataset on manufacturing firms to document a significant degree of heterogeneity of gross job creation and job destruction rates, with size and age playing a meaningful role. They find that job reallocation rates of older, larger, and multi-unit plants, in particular display pronounced countercyclical patterns of variation. Haltiwanger et al. (1999) use administrative firm level data to show that older and larger firms are more productive than their smaller and younger counterparts.

Haltiwanger et al. (2013) examine job creation in the lens of of small, large, young, and old firms and find that, once controlling for firm age, there is no clear relationship between firm size and growth. Fort et al. (2013) study the of firm sensitivity to the business cycle by size and age and find that small and young firms are more sensitive to the business cycle than larger and older firms. This paper adds to the existing literature through the lens of downward nominal wage rigidity, documenting that larger and older establishments exhibit greater measure wage rigidity than smaller and younger establishments, but that there is a greater relationship between wage rigidity and the employment outcomes of layoffs, quits, and hires in larger and younger establishmenets than smaller and older establishments.

The paper proceeds as follows. Section 5.3 describes the data used throughout the paper and presents the summary statistics by establishment size and age. Section 4.3 presents the estimates of wage rigidity across supersectors, occupations, and education levels, and presents the establishment-level distibutions of wage rigidity by establishment size and age. Section 5.4 presents the establishment-level wage rigidity regressions by establishment size and age for layoffs, quits, and hires. Section 5.6 concludes.

4.2 Data and Summary Statistics

The paper employs administrative and survey data from the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB). The main analysis uses the Linked Employer Employee Data of Integrated Labor Market Biographies (LIAB), matched with the annual IAB Establishment Panel Survey. The LIAB includes 5,293 West German establishments that participated in the annual IAB establishment survey each year either from 1999 through 2001 or from 2000 through 2002, and follows each such establishment every year of its existence from 1997 through 2003.

The LIAB also provides complete labor market biographies for each employee liable to social security who was employed at a sampled, surveyed establishment at any point between 1997 and 2003. The data set follows these workers entire employment, unemployment, and wage histories from 1993 through 2007, even if the workers move to an establishment outside the sample. The LIAB also provides the exact dates that an employment spell begins and ends for an employee at a given establishment.

The second dataset used in the paper is the Sample of Integrated Labor Market Biographies (SIAB). The SIAB provides complete labor market biographies for a 2 percent random sample of all employees liable to social security. However, the SIAB does not provide worker biographies for all workers at a sampled establishment as in the LIAB. Therefore, the paper uses the SIAB for the estiamtes of wage rigidity by supersector, occupation, and education and the LIAB establishment-level analysis.

Table 4.1 shows descriptive statistics of the full sample in column 1, splitting the sample between large and small establishments in columns 2 and 3, and splitting the sample between young and old establishments in columns 4 and 5. This paper follows the convention of Fort et al. (2013) and categorizes an establishment as young if it is no more than 5 years old and old if it is more than five years old. Fort et al. (2013) also categorize esablishments that are less than 20 employees as small. However, this paper only has establishments with at least 50 employess and thus simply uses the sample median to distinguish between small and large. Therefore, it is more true to consider the below analysis as a comparison between large- and medium-sized establishments; however, the study uses the small/large terminology for simplicity. There are 4,608 establishment-years in the large establishments sample, 4,622 establishment-years in the small establishments sample, 680 establishment-years in the young sample, and 8,550 establishments in the old sample.

Yearly establishment-level layoff, quit, and hire rates are measured as fractions of the establishment's total workforce as of December 31 of the preceding year. The full-sample average layoff rate is 6.8 percent, the full-sample average quit rate is 11.1 percent, and the full-sample average hire rate is 21.8 percent with standard deviations of 12.0 percent, 16.1 percent, and 33.5 percent, respectively.

Columns 2 and 3 show the layoff, quit, and hire rate averages for the large and small establishments, respectively. It is clear from the descriptive statistics that the small estbalishments have a higher degree of worker turnover than the lager establishments. The average small establishment lays off 7.9 percent of its workforce, hires 24.2 percent of its workforce, and has 12.3 percent of its workforce quit in the average year, whereas the average large establishment lays off 5.7 percent of its workforce, hires 19.4 perrcent of its workforce, and has 10.0 percent of its workforce quit in the average year. However, there is very little difference in the dispersion of layoff, quit, and hire rates between small establishments and large establishments.

Columns 4 and 5 show the layoff, quit, and hire rate averages for the young and old establishments, respectively. The young establisments have much larger degree of turnover compared to the older establishments. Young establishments, on average, have more than a 3 percentage point higher annual layoff rate relative to the old estbalishments at 9.7 percent per year. Quits are more than 5 percentage points higher per year, on average, for young estbalishments compared to old establishments at 15.9 percent per year, whereas hires are nearly 12.5 percentage points higher in young establishments veruses their older counterparts at 32.3 percent per year. Similarly, dispersion in the layoff, quit, and hire rates, as measured by the standard deviation, is also much larger for young establishments compared to old establishments. The significantly larger hire rate for young establishments compared to the old estbalishments is indicative of the younger establishments growing faster in size than the older, more established establishments.

Next, table 4.1 shows the average size an establishment in the full sample, the large establishment sample, the small establisment sample, the young establishment sample, and the old establishment sample. The mean size of an establishment in the full sample is 549 employees. There is considerable dispersion in size among establishments in the full sample, as the standard deviation of establishment size is 1,211 employees.

Splitting the establishments into large and small in columns 2 and 3 shows that the mean size of an estbalishment in the large sample is 987 employees with a standard deviation of 1,598, whereas the mean size of an establishment in the small sample is 113 employees with a standard deviation of 48. The median number of employees in the full sample is 219, as discussed above. The median number of employees at a large establishment is 539, whereas the median number of employees at a large establishment is 102.

Columns 4 and 5 split the establishments into young and old and show that the mean size of an estbalishment in the young sample is 377 employees with a standard deviation of 673, whereas the mean size of an establishment in the old sample is 563 employees with a standard deviation of 1,243. The median number of employees at a young establishment is 176, whereas the median number of employees at an old establishment is 222. The young establishents are noticeably smaller than the old establishments. However, the descriptive statstics show that there is a signifncant overlap in the distribution of establishment size by employees between the young and old establishments.

The mean, nominal daily wage for an employee in the full-sample is 102.90 euros with a standard deviation of 66.21 and a median of 90.73 euros per day. Breaking the sample into large and small establishments, the mean, nominal daily wage for an employee in the large establishment sample is 104.86 euros with a standard deviation of 67.39 and a median of 92.19 euros per day, whereas the mean, nomnal daily wage for an employee in the small establishment sample is 85.42 euros with a standard deviation of 49.80 and a median of 77.53 euros per day. Clearly, the larger establishents pay there emplyoees more than the smaller establishments and also have a much larger degree of dispersion in pay across its employees.

Splitting the sample into young and old establishments, the mean, nominal daily wage for an employee in a young establishment is 102.32 euros with a standard deviation of 65.26 and a median of 89.72 euros per day, whereas the mean, nomnal daily wage for an employee in the old establishment sample is 101.78 euros with a standard deviation of 65.66 and a median of 90.20 euros per day. The nominal wage statistics are nearly identical aross young and old establishments.

Finally, table 4.1 shows a breakdown of revenue growth amonth establishments in the full sample and large, small, young, and old estbalishments. In the full sample, the average establishment has an annual revenue growth rate of 3.5 percent with a standard deviation 20.6 percent. Large establishmnets, however, have average, annual revenue growth of 4.1 percent with a standard deviation of 19.6 percent compared to average, annual revenue growth of 2.9 percent for smaller esatblishments with a standard deviation 21.5 percent. Therefore, smaller establishments in the sample have lower average, annual revenue growth but, at the same time, more volatile revenue growth when compared to large establishments.

The difference in annual average revenue growth is larger between the young and old establishments than the small and large. The average young establishment saw its revenue grow by 6.7 percent per year, whereas the average old establishment saw its revenue grow by 3.3 percent per year. Younger establishments also had more dispersion in revenue growth, with a standard devaition of 23.6 percent, comapred to the larger establishments, with a stanadard deviation of 20.4 percent.

4.3 Distribution of Wage Rigidity

To provide some context as to "where" and "how much" wage rigidity is present in the sample, this section provides estimates of the prevalence of wage rigidity in each supersector, broad occupation class, and education level before turning to the establishment-level analysis. At the establishment-level, wage rigidity is dileneated between small and large establishments and young and old to provide context as to which types of estabishments, if any, have more rigid wages. The extent of wage rigidity is estimated using the methodology described in section 2.4, but taking these larger groups as the unit of analysis rather than the establishment.

4.3.1 Wage Rigidty by Supersector, Occupation, and Education

This section makes use of the 2 percent random sample of German workers provided by the SIAB to provide a representative picture of wage rigidity by supersector, occupation, and education. For a wage change to qualify as valid by supersector, occupation, and education categories, the employee must be classified in the same establishment and in either the same sector, occupation class, or education level for two consecutive years, with the employee's hour status unchanged between those years.

Table 4.2 shows wage rigidity estimates for each supersector, occupational class, and education level in the sample. Among supersectors, public administration exhibits the highest degree of wage rigidity, with an estimated 46.6 percent of wage cuts prevented by wage rigidity over the sample period. Finance and Administration also display large amounts of wage rigidty, with 28.8 and 26.4 percent of wage cuts prevented, respectively. Construction and mining/manufacturing exhibit exhibit the least wage rigidity over the sample period, with -3.2 percent and 3.4 percent of negative wage cuts prevented due to wage rigidty. That is, the estimates suggest that there are more wage cuts in the construction sector than one would expect from the nominally positive portion of the wage change distribution, although this estimate is statistically indistinguishable from zero.

Occupational classes are classified according to the Blossfeld grouping, a common, broad classification of occupations used in Germany. The middle portion of Table 4.2 shows that the highest degrees of wage rigidity are concentrated among the semi-professional, skilled commercial and administrative, and engineer occupations, with 35.2, 34.5, and 32.9 percent of wage cuts prevented due to wage rigidity, respectively, over the sample. On the other end of the spectrum, the lowest degrees of wage rigidity are concentrated among simple manual, skilled manual, and simple service occupations with -0.6, 2.8, and 6.7 percent of wage cuts prevented, respectively. One oddity that emerges from the table is that the wage rigidity estimate for agricultural occupatons is 31.2 percent, compared to an estimated 9.7 percent in the agricultural supersector. However, the agricultural supersector contains a large fraction of workers who are not employed in agricultural occupations. Many of these workers are employed in simple and skilled manual occupations that display low degrees of wage rigidity. Interestingly, skilled occupations consistently exhibit more wage rigidity than their less skilled counterparts.

Similarly, the bottom portion of table 4.2 shows that more educated groups generally exhibit higher degrees of wage rigidity. Workers whose highest educational attaintment is a secondary school leaving certificate with no vocation are estimated to have only 6.3 percent of counterfactual wage cuts prevented by wage rigidity. Graduates of a University of High Science are estimated to have 41.3 percent of counterfactual wage cuts prevented. Workers who receive vocational training in both the secondary and upper secondary levels of education exhibit roughly 7 percent more rigid wages than their non-vocational counterparts.

Consistently, though, workers who receive more education are, on average, less likely to experiece wage cuts in the sample. One possible explanation for this observation is that more educated workers may build up higher degrees of firm/establishment specific human capital. The theoretical model introduced and discussed in chapters 2 and 3 show that cutting wages for workers would induce quits. Losing workers with high reserves of human capital is costly to employers in terms of training costs for new workers. These costs would prohibit the reduction of the wages for workers with high levels of accumulated, firm/establishment specific human capital.

4.3.2 Wage Rigidity by Establishment Size and Age

The remained of the analysis employs the LIAB to study establishment-level wage rigidity and its relatinship to employment by establishment age and size. Table 4.3 shows the the distribution of estimated wage rigidity across establishments in the full sample and the sample split between large and small establishments. In the full sample mean of estimated wage rigidity across all establishments is 24.5 percent, suggesting that approximately 24.5 percent of nominal wage cuts were prevented due to wage rigidity at the average establishment. The standard deviation of the full sample, establishment-level wage rigidity estimates is 23.5 percent and the median is 21.8 percent.

Splitting the sample between large and small establishments, however, provides noticeable differences in the estimates when compared to both each other and the full sample estimates.The mean wage rigidity across the large establishments in the sample is 25.5 percent, suggesting that approximately 25.5 percent of nominal wage cuts were prevented due to wage rigidity at the average, large establishment. The standard deviation of estimated wage rigidity among the large establishments is 24.6 percent and the median is 22.1 percent. For the sample of small establishments, the mean of estimated wage rigidity is 22.3 percent, suggesting that approximately 22.3 percent of nominal wage cuts were prevented due to wage rigidity at the average, small establishment. The standard deviation of wage rigidity among small establishments is 29.0 percent and the median is 16.9 percent. Therefore, larger establishments, on average, have noticeably higher levels of estimated wage rigidity measured by both the mean and median when compared to smaller establishments, whereas the smaller establishments have more dispersion in their estimates of wage rigidity compared to larger establishments.

Differentiating establishment-level wage rigidity estimates into young and old categories provides an even more noticeable difference than the large versus small split provided. The mean estimated level of wage rigidity for young estbalishments in a given year is 16.1 percent of counterfactual wage cuts prevented due to wage rigidty compared to 24.5 percent for the older establishmnets, nearly an 8.5 percentage point difference. Similarly, the older establishments median level of estimated wage rigidity is 8 percentage points larger than the median of the younger establishments, with an estimated median of 20.2 percent for the old establishments compared to 10.2 percent for the young establishments. These estimates clearly suggest that older establishments have higher levels of wage rigidity relative to younger establishemements.

4.4 Empirical Results

This section presents the establishment-level wage rigidity regressions by establishment size and age. Section 4.4.1 shows the layoff regression results, section 4.4.2 shows the quits regression results, and section 4.4.3 shows the hires regression results.

4.4.1 Layoffs

Column 1 of table 4.4 shows the results for the full-sample layoffs regression, weighted by the square root of the number of employees per establishment-year. The coefficient on the level of estimated wage rigidity is 2.5 percent and statistically significant at the 99 percent confidence level. The coefficients on the uninteracted revenue growth terms are statistically insignificant, as is the coefficient on the interaction between wage rigidity and positive revenue growth. However, the coefficient on the interactions between wage rigidity and negative revenue growth is negative 19.3 percent and highly significant. Economically, a one standard deviation negative movement in revenue growth at an establishment with the sample average level of wage rigidity is predicted to have a 1.5 percentage point increase in the layoff rate, a 34.5 percent increase in the sample average layoff rate.

Columns 2 and 3 of table 4.4, present the unweighted regression results from the sample split by establishment size. The association between wage rigidity and layoffs differs meaningfully between small and large establishments. The coefficient on the level of wage rigidity for the large establishments in column 2 is 2.9 percent and highly significant, while the coefficient in column 3 is -1.0 percent and statistically insignificant. The coefficients on the wage rigidity-revenue growth interaction terms are not significant in either column, though the point estimate on the wage rigidity-negative revenue growth interaction term for the large establishments is just outside the 90 percent confidence level.

However, the coefficient on the wage-rigidity-revenue growth interaction term for the large establishments is economically meaningful. Figure 4.1 displays a graph illustrating the economic relationship between estimated wage rigidity and the predicted increase in the layoff rate associated with a given level of estimated wage rigidity relative to the case of no wage rigidity. The solid line depicts the relationship for the large establishments, whereas the dashed line depicts the relationship for the small establishments. Each line illustrates the predicted layoff increase-estimated wage rigidity relationship given a one standard devation decrease in revenue growth.

The estimates suggest that negative movement in revenue growth amplifies the positive relationship between wage rigidity and layoffs. The solid line in figure 4.1 shows that large establishments with the average level of estimated wage rigidity of 24.5 percent of wage cuts prevented would be predicted to increase layoffs by 1 percentage point in response to a one standard deviation decrease in revenue, similarly to the results for the full sample. The dashed line in figure 4.1 shows that small establishments with the average level of estimated wage rigidity of 22.3 percent of wage cuts prevented would be predicted to decrease layoffs by 0.3 percentage points in response to such a decrease in revenue, although this prediction is not statistically significant from zero.

Columns 4 and 5 of table 4.4, present the regression results from the sample split by establishment age. The association between wage rigidity and layoffs differs noticeably between young and old establishments. The coefficient on the level of wage rigidity for the young establishments in column 2 is 5.5 percent and highly significant, while the coefficient in column 5 for the old establishments is 0.7 percent and statistically insignificant. The coefficient on the wage rigidity-revenue growth interaction terms is significant and quite large in magnitude, with the expected, negative sign, for the negative revenue growth term for the young estbalishments with a point estiamte of -0.576, but insignificant for all other terms.

Figure 4.2 displays a graph illustrating the economic relationship between estimated wage rigidity and the predicted increase in the layoff rate associated with a given level of estimated wage rigidity relative to the case of no wage rigidity by establishment age. The solid line depicts the relationship for the young establishments, whereas the dashed line depicts the relationship for the old establishments. Each line illustrates the predicted layoff increase-estimated wage rigidity relationship given a one standard devation decrease in revenue growth at all points on the lines.

The estimates show that negative movement in revenue growth amplifies the positive relationship between wage rigidity and layoffs for the young establishments, as a one standard deviation decrease in revenue growth for a young establishment with the sample average level of wage rigidity implies a 4.7 percentage point increase in the layoff rate relative to a young establishment with no measured wage rigidity. The 4.7 percentage point increase in the layoff rate for the average young establishment corresponds to a 46.6 percent increase relative to the sample average 9.7 percent layoff rate. The dashed line for the old estabishments in figure 4.2 is statistically insignificant from zero.

4.4.2 Quits

Column 1 of table 4.5 shows the results for the full sample quits regression, weighted by the square root of the number of employees per establishment-year. The coefficient on estimated wage rigidity is negative 5.2 percent, suggesting that establishments with more wage rigidity are predicted to have fewer quits. The coefficients on the interactions between wage rigidity and revenue growth also have the expected signs. However, the difference between the magnitudes of the wage rigidity interactions with positive and negative reveune growth are stark, with the coefficient on the interaction between wage rigidity and positive revenue growth large and statistically significant while the coefficient on the interaction between wage rigidity and negative revenue growth is smaller and statistically insignificant. Economically, a one standard deviation movement in negative revenue growth at an establishment with the sample average level of wage rigidity is predicted to have 1.6 percentage point decrease in the quit rate, a 17.3 percent decrease in the sample average quit rate.

Columns 2 and 3 of table 4.5, present the results from the unweighted quit rate regressions for the samples split by establishment size. The association between wage rigidity and quits varies slightly between small and large establishments. The coefficient on the level of wage rigidity in column 2 is -5.5 percent compared to -4.4 percent in column 3; both are statistically signficant. The coefficients on the wage rigidity-negative revenue growth interaction terms are generally statistically signifcant and have the same, expected, positive signs in both columns, implying that a negative movement in revenue growth amplifies the negative relationship between wage rigidity and quits.The magnitude of the wage rigidity-revenue growth interaction terms are markedly different between small and large establishments, as the interaction terms are at least three times as large for the large establishments compared to the small establishments.

Figure 4.3 shows the economic significance of the quit regression results by establishment size from columns 2 and 3 of table 4.5. The solid line represents the predicted decrease in the quit rate for a given level of estimated wage rigidity for the large establishments, and the dashed line represents the relationship for the small establishments. Both lines assume a one standard deviation movement in revenue growth at all points on the line.

A large establishment with the sample average level of wage rigidity would be predicted to experience a quit rate 2.5 percentage points lower than a large establishment with no wage rigidity in response to a one standard deviation decrease in revenue. The corresponding decrease for small establishments is 1.2 percentage points. These differences translate into 25.1 percent and 9.6 percent reductions in the sample average quit rates for large and small establishments, respectively.

Columns 4 and 5 of table 4.6, present the results from the quit rate regressions for by establishment age. The relationship between wage rigidity and quits varies noticeably between young and old establishments. The coefficient on the level of wage rigidity for the young establishmentd is -9.3 percent compared to -4.8 percent for the old; both are statistically signficant at the 99 percent confidence level. The coefficients on the wage rigidity-negative revenue growth interaction terms are, however, statistically insignificant for both the young and old establishments.

Figure 4.4 shows the economic significance of the quit regression results by establishment age. The solid line represents the predicted decrease in the quit rate for a given level of estimated wage rigidity and a one standard deviation movement in revenue growth at all points on the line for the young establishments, and the dashed line represents the same relationship for the old establishments.

A young establishment with the sample average level of wage rigidity would be predicted to experience a quit rate 1.6 percentage points lower than a young establishment with no wage rigidity in response to a one standard deviation decrease in revenue. The corresponding decrease for old establishments is 1.7 percentage points. These differences translate into 10.1 percent and 15.7 percent reductions in the sample average quit rates for young and old establishments, respectively.

4.4.3 Hires

Column 1 of table 4.6 shows the results for the full sample hires regression, weighted by the square root of the number of employees per establishment-year. The coefficient on estimated wage rigidity is negative 5.3 percent. The coefficients on the interactions between wage rigidity and positive revenue growth have the expected negative sign, implying that establishments with both positive levels of estimated wage rigidity and positive movements in revenue growth hire less than establishments with no measured wage rigidity and positive movements in revenue growth. This follows from the establishments forward looking nature, as discussed in chapter 3's treatment of the theoretical model: establishments hiring workers in the current period know that, in response to future negative movements in revenue, they will face the costs of future wage cuts and future firing costs for workers hired today, reducing the incentive to hire in the current period. Economically, an establishment with the sample average level of wage rigidity and a one standard deviation positive movement in revenue growth is predicted to have a 2.6 percentage point lower hire rate than an establishment with no wage rigidity, which corresponds to a 14.8 percent decrease in the hire rate relative to the sample average.

Columns 2 and 3 of table 4.6 present the results from the hire rate regressions split by establishment size. The coefficient on the level of wage rigidity for the large establishments in column 2 is -4.8 percent compared to -3.3 percent in column 3 for the small establishments, though only the former is statistically significant. The coefficients on the wage rigidity-revenue growth interaction terms are statistically signifcant for positive revenue growth, have the expected sign, and are similar in magnitude. These coefficients have the unexpected sign for the negative revenue growth interaction term but are not statistically significant, although the magnitude of the negative revenue growth interaction term is much smaller for the small establishments in the sample compared to the large establishments.

Figure 4.5 shows the economic significance of the hire regression results by establishment size from columns 2 and 3 of table 4.6. The solid line represents the predicted decrease in the hire rate for a given level of estimated wage rigidity for the large establishments and given a one standard deviation movement in positive revenue growth, and the dashed line represents the relationship for the small establishments. A large establishment with the sample average level of wage rigidity is predicted to have a hire rate 2.8 percentage points lower than a large establishment with no wage rigidity in response to a one standard deviation increase in revenue. The corresponding decrease for small establishments is 2.5 percentage points. These differences translate into 16.2 percent and 14.5 percent reductions in the sample average hire rates for large and small establishments, respectively.

Columns 4 and 5 of table 4.6 present the results from the hire rate regressions by establishment age. The coefficients on the level of wage rigidity for the young and old stablishments are relatively similar, with point estiamtes of -9.2 percent for the young in column 4 and -8.3 percent for the old in column 5. However, the point estiamte on the level wage rigidity term is statistically insignificant for the young establishments. The coefficients on the wage rigidity-revenue growth interaction terms are statistically signifcant for positive revenue growth, have the expected sign, and are similar in magnitude for both the young and old establishments, though the point estimate on the positive revenue growth interation term is more than three times as large for the young establishments compared to the old establishments.

Figure 4.5 shows the economic significance of the hire regression results by establishment age. The solid line represents the predicted decrease in the hire rate for a given level of estimated wage rigidity and given a one standard deviation movement in positive revenue growth for the young establishments, and the dashed line represents the relationship for the old establishments. A young establishment with the sample average level of wage rigidity is predicted to have a hire rate 7.1 percentage points lower than a young establishment with no wage rigidity in response to a one standard deviation increase in revenue. The corresponding decrease for the old establishments is 3.4 percentage points. These differences translate into 22.0 percent and 16.3 percent reductions in the sample average hire rates for young and old establishments, respectively.

4.5 Conclusion

This paper explores the relationship between downward nominal wage rigidity and employment outcomes by establishment size and age. Supersector-, occupation-, and education-level wage rigidity estimates suggest that more highly skilled workers are less likely to receive wage cuts, as workers in supersectors typically employing high skilled workers, workers with more skill-intensive occupations, and workers with higher degrees of education display higher degrees of estimated downward nominal wage rigidity.

Establishment-level wage rigidity estimates suggest a substantial amount of downward nominal wage rigidity in Germany, with an average of 24.5 percent of counterfactual wage cuts prevented by wage rigidity for the full sample. Splitting the sample by establishment size shows noticeable variation in establishment-level wage rigidity between large and small establishments and young and old establishments. Large establishments, with an average of 25.5 percent of counterfactual wage cuts prevented by wage rigidity, exhibit more wage rigidity than smaller estabishments, with an average of 22.3 percent fo wage cuts prevented due to wage rigidity, and old establishments, 24.5 percent of wage cuts prevented, exhibit more wage rigidity than younger establishments, with 16.1 percent of wage cuts pervented.

The empirical analysis shows that the establishment-level relationship between wage rigidity and employment outcomes is stronger for large establishments compared to small establishments and for young establishments compared to young. Wage rigidity increases layoffs at the average large establishment by 1 percentage point and decreases quits and hires by 2.5 and 2.8 percentage points, respectively, whereas wage rigidity and layoffs have no statisitcal relationship in small establishments and decreases quits and hires by 1.2 and 2.5 percentage points respectively.

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Figure 4.1: Layoff Regressions by Establishment Size – Economic Significance

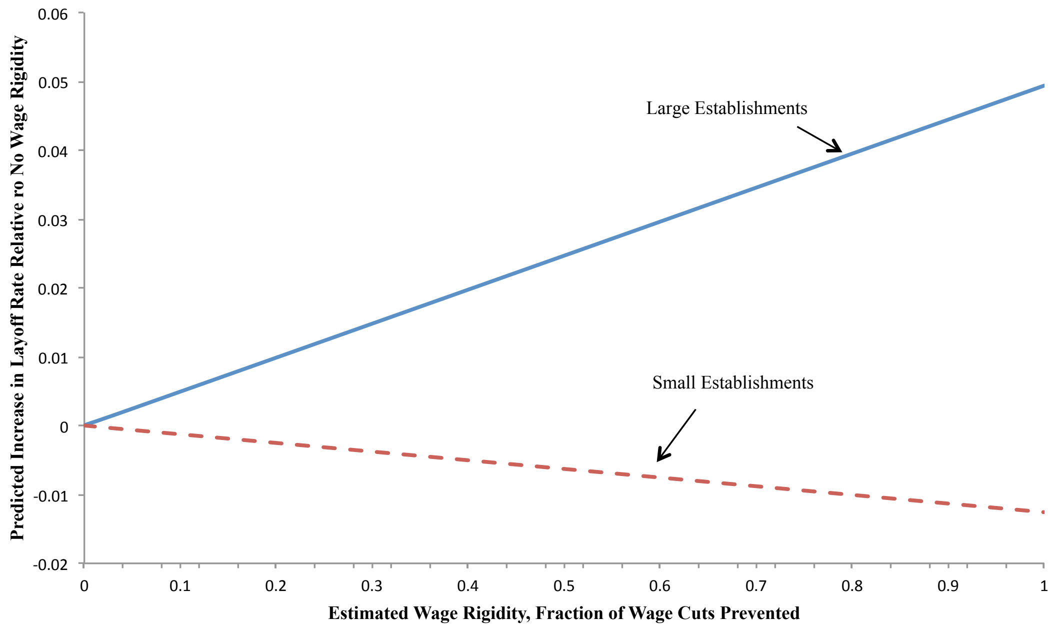

Figure 4.2: Layoff Regressions by Establishment Age – Economic Significance

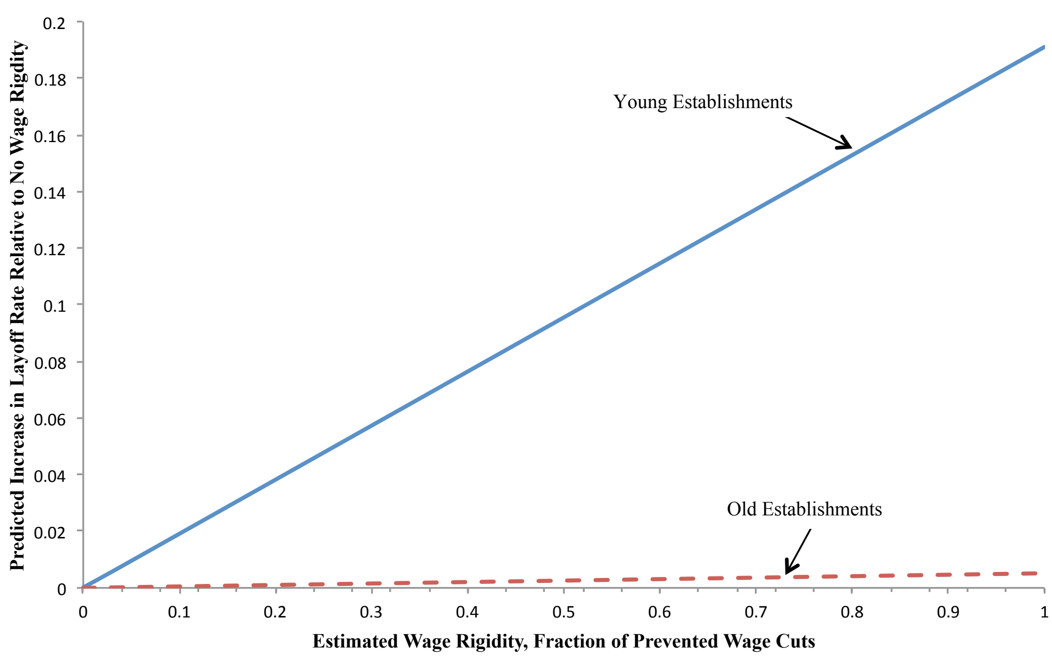

Figure 4.3: Quits Regressions by Establishment Size – Economic Significance

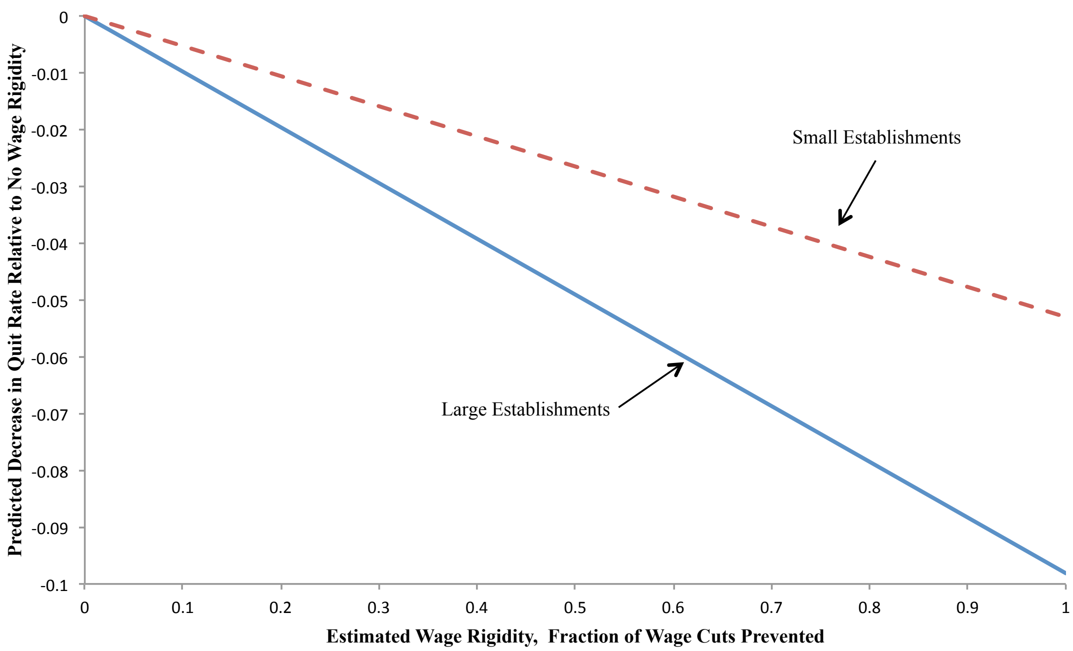

Figure 4.4: Quits Regressions by Establishment Age – Economic Significance

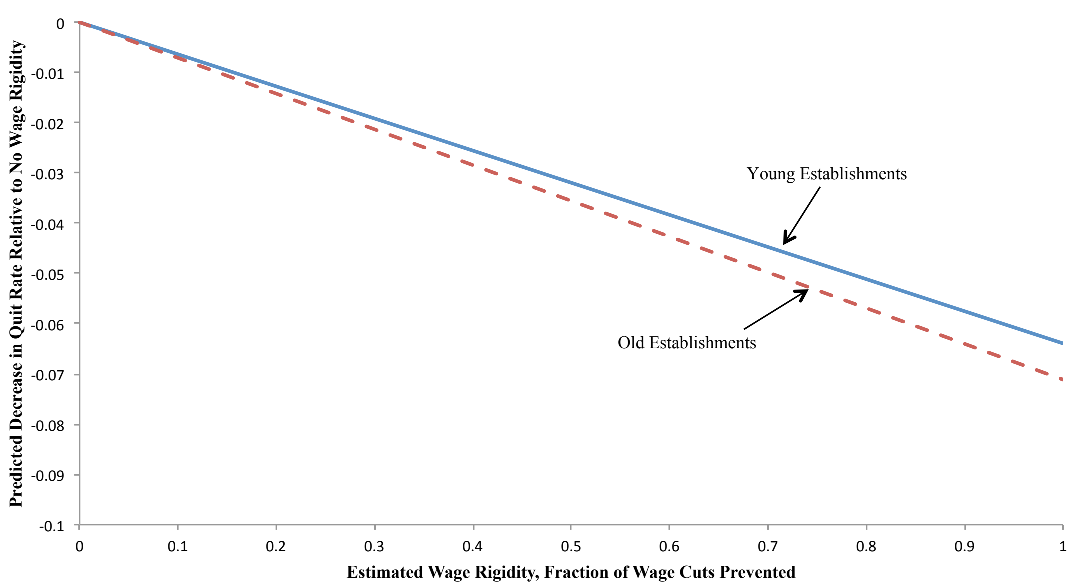

Figure 4.5: Hires Regressions by Establishment Size – Economic Significance

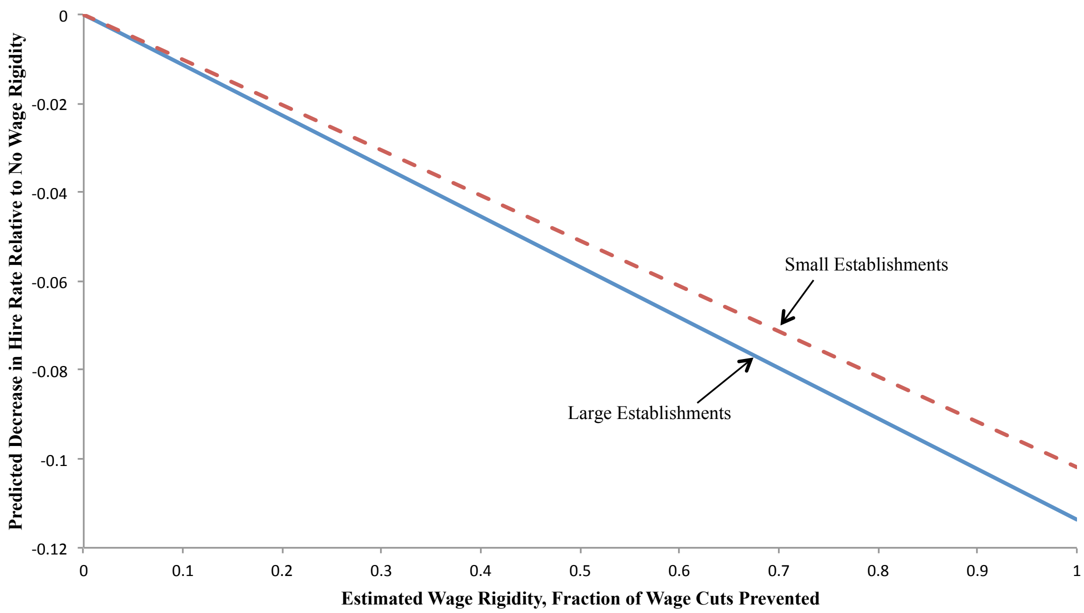

Figure 4.6: Hires Regressions by Establishment Age – Economic Significance

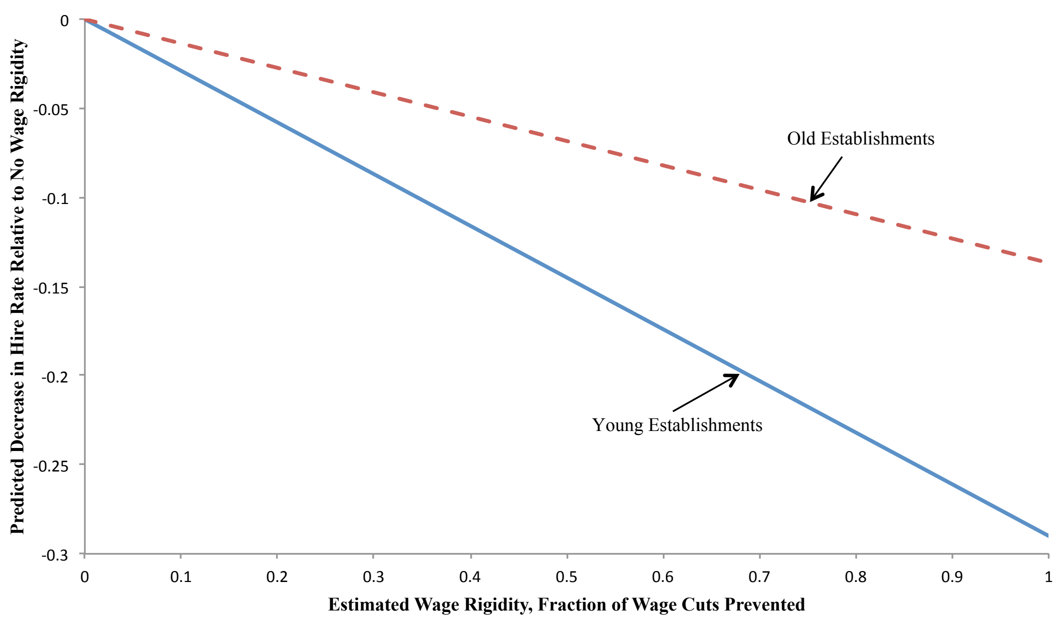

Table 4.1: Descriptive Statisitcs by Establishment Size and Age

	(1) All Establishments	(2) Min. 220 Employees	(3) Max. 219 Employees	(4) Young ($\leq$ 5 Years Old)	(5) Old ($>$ 5 Years Old)
Sample Size, Establishment-Years	9,230	4,608	4,622	680	8,550
Mean Layoff Rate	0.068	0.057	0.079	0.097	0.066
	(0.120)	(0.120)	(0.118)	(0.165)	(0.115)
Mean Quit Rate	0.111	0.100	0.123	0.159	0.108
	(0.161)	(0.158)	(0.164)	(0.240)	(0.153)
Mean Hire Rate	0.218	0.194	0.242	0.323	0.209
	(0.335)	(0.336)	(0.331)	(0.499)	(0.316)
Mean Employees per Establishment	549	987	113	377	563
	(1,211)	(1,598)	(48)	(673)	(1,243)
Median Employees per Establishment	219	539	102	176	222
Mean Daily Wage, Level	102.90	104.86	85.42	102.32	101.78
	(66.21)	(67.39)	(49.80)	(65.26)	(65.66)
Median Daily Wage, Level	90.73	92.19	77.53	89.72	90.20
Average Revenue Growth	0.035	0.041	0.029	0.067	0.033
	(0.206)	(0.196)	(0.215)	(0.236)	(0.204)

Standard deviations in parentheses where applicable.

Table 4.2: Aggregate Wage Rigidity Estimates by Select Classifications

Classification	Estimated Wage Rigidity
Supersector	
Agriculture	0.097
Mining/Manufacturing	0.034
Energy/Water	0.063
Construction	-0.032
Trade/Foodservice	0.058
Transportation	0.050
Finance	0.288
Real Estate	0.170
Public Administration	0.466
Administration	0.264
Blossfeld Occupational Classification	
Agricultural occupations	0.312
Simple manual occupations	-0.006
Skilled manual occupations	0.028
Technicians	0.256
Engineers	0.329
Simple service	0.067
Qualified service	0.191
Semi-professions	0.352
Professions	0.139
Simple commercial and administrative occupations	0.242
Skilled commercial and administrative occupations	0.345
Manager	0.201
Educational Attainment	
Secondary/Intermediate without Vocation	0.063
Secondary/Intermediate with Vocation	0.138
Upper Secondary without Vocation	0.308
Upper Secondary with Vocation	0.373
University of High Science	0.413
University Degree	0.306

Table 4.3: Estimated Wage Rigidity by Establishment Size and Age

	Estimated Wage Rigidity		
	Mean	Median	Standard Deviation
All Establishments	0.245	0.218	0.235
Establishment Size by Median			
Min. 220 Employees	0.255	0.221	0.246
Max. 219 Employees	0.223	0.169	0.290
Establishment Age			
Young ($\leq$ 5 years)	0.161	0.102	0.291
Old (>5 years)	0.245	0.202	0.267

Table 4.4: Wage Rigidity and Layoffs by Establishment Size and Age

Dependent Variable	Layoff Rate as a Fraction of Establishment Workforce				
	Full	Establishment Size		Establishment Age	
	Sample	Large	Small	Young	Old
	(1)	(2)	(3)	(4)	(5)
Wage Rigidity	0.025	0.029	-0.010	0.055	0.007
	(0.006)	(0.009)	(0.008)	(0.028)	(0.006)
Positive Revenue Growth	0.012	-0.001	0.022	0.016	0.007
	(0.007)	(0.011)	(0.013)	(0.031)	(0.007)
Negative Revenue Growth	-0.015	-0.026	-0.121	-0.013	-0.042
	(0.014)	(0.023)	(0.021)	(0.067)	(0.013)
Wage Rigidity x Positive Revenue Growth	-0.038	-0.014	0.019	-0.031	-0.040
	(0.024)	(0.036)	(0.036)	(0.093)	(0.023)
Wage Rigidity x Negative Revenue Growth	-0.193	-0.104	0.012	-0.576	0.009
	(0.044)	(0.066)	(0.058)	(0.177)	(0.041)
Work Council	-0.058	-0.065	-0.040	-0.075	-0.056
	(0.004)	(0.006)	(0.004)	(0.020)	(0.003)
Weights	No. of Employees	None	None	No. of Employees	No. of Employees
Sample Restriction	Min. 20 Employees	Min. 220 Employees	Max. 219 Employees	Young ($\leq$ 5 Years Old)	Old (> 5 Years Old)
R-squared	0.344	0.450	0.226	0.778	0.342
N	9,230	4,608	4,622	676	8549

Standard errors in parentheses. Unit of observation is establishment-year. Columns (2) and (3) divide the sample into large and small establishments, respectively, by the median establishment size. Columns (4) and (5) divide the sample into young that are 5 years old or less and old establishments that are more than 5 years old. Layoffs are defined as a percentage of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated as described in section 4 and is fixed by establishment over sample period. Each regression includes a set of etablishment characteristics, individual charactersistics, and year dummies as controls. Establishment characteristics include a set of controls for the median year-over-year wage change, occupational mix and dummies for sector, federal state, and establishment size. Individual characteristics include controls for gender and workers' education. Positive revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is positive and zero otherwise. Negative revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is negative and zero otherwise. All regressions require establishments to have at least 20 employees in a given year and cover the period 1997 to 2003.

Table 4.5: Wage Rigidity and Quits by Establishment Size and Age

Dependent Variable	Quit Rate as a Fraction of Establishment Workforce				
	Full	Establishment Size		Establishment Age	
	Sample	Large	Small	Young	Old
	(1)	(2)	(3)	(4)	(5)
Wage Rigidity	-0.056	-0.055	-0.044	-0.093	-0.048
	(0.009)	(0.012)	(0.011)	(0.044)	(0.010)
Positive Revenue Growth	0.079	0.083	0.065	0.091	0.068
	(0.012)	(0.015)	(0.017)	(0.048)	(0.012)
Negative Revenue Growth	-0.225	-0.130	-0.076	-0.045	-0.248
	(0.022)	(0.030)	(0.028)	(0.104)	(0.022)
Wage Rigidity x Positive Revenue Growth	-0.192	-0.220	-0.070	-0.257	-0.15
	(0.038)	(0.047)	(0.047)	(0.143)	(0.040)
Wage Rigidity x Negative Revenue Growth	0.043	0.220	0.042	-0.123	0.098
	(0.069)	(0.088)	(0.076)	(0.274)	(0.073)
Work Council	-0.047	-0.040	-0.044	-0.19	-0.039
	(0.006)	(0.008)	(0.005)	(0.030)	(0.006)
Weights	No. of Employees	None	None	No. of Employees	No. of Employees
Sample Restriction	Min. 20 Employees	Min. 220 Employees	Max. 219 Employees	Young (≤ 5 Years Old)	Old (> 5 Years Old)
R-squared	0.357	0.442	0.290	0.420	0.348
N	9,230	4,608	4,622	676	8549

Standard errors in parentheses. Unit of observation is establishment-year. Columns (2) and (3) divide the sample into large and small establishments, respectively, by the median establishment size. Columns (4) and (5) divide the sample into young that are 5 years old or less and old establishments that are more than 5 years old. Quits are defined as a percentage of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated as described in section 4 and is fixed by establishment over sample period. Each regression includes a set of etablishment characteristics, individual charactersistics, and year dummies as controls. Establishment characteristics include a set of controls for the median year-over-year wage change, occupational mix and dummies for sector, federal state, and establishment size. Individual characteristics include controls for gender and workers' education. Positive revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is positive and zero otherwise. Negative revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is negative and zero otherwise. All regressions require establishments to have at least 20 employees in a given year and cover the period 1997 to 2003.

Table 4.6: Wage Rigidity and Hires by Establishment Size and Age

Dependent Variable	Hire Rate as a Fraction of Establishment Workforce				
	Full	Establishment Size		Establishment Age	
	Sample	Large	Small	Young	Old
	(1)	(2)	(3)	(4)	(5)
Wage Rigidity	-0.053	-0.048	-0.033	-0.092	-0.083
	(0.015)	(0.021)	(0.021)	(0.093)	(0.024)
Positive Revenue Growth	0.103	0.157	0.235	0.275	0.139
	(0.019)	(0.026)	(0.035)	(0.100)	(0.030)
Negative Revenue Growth	0.106	0.119	0.001	0.150	0.145
	(0.036)	(0.051)	(0.055)	(0.219)	(0.055)
Wage Rigidity x Positive Revenue Growth	-0.261	-0.335	-0.321	-0.838	-0.264
	(0.064)	(0.080)	(0.094)	(0.301)	(0.099)
Wage Rigidity x Negative Revenue Growth	-0.362	-0.287	-0.088	-0.484	-0.329
	(0.116)	(0.148)	(0.152)	(0.577)	(0.179)
Work Council	-0.115	-0.112	-0.123	-0.306	-0.091
	(0.010)	(0.013)	(0.010)	(0.064)	(0.015)
Weights	No. of Employees	None	None	No. of Employees	No. of Employees
Sample Restriction	Min. 20 Employees	Min. 220 Employees	Max. 219 Employees	Young (≤ 5 Years Old)	Old (> 5 Years Old)
R-squared	0.639	0.647	0.310	0.497	0.213
N	9,230	4,608	4,622	676	8549

Standard errors in parentheses. Unit of observation is establishment-year. Columns (2) and (3) divide the sample into large and small establishments, respectively, by the median establishment size. Columns (4) and (5) divide the sample into young that are 5 years old or less and old establishments that are more than 5 years old. Hires are defined as a percentage of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated as described in section 4 and is fixed by establishment over sample period. Each regression includes a set of etablishment characteristics, individual charactersistics, and year dummies as controls. Establishment characteristics include a set of controls for the median year-over-year wage change, occupational mix and dummies for sector, federal state, and establishment size. Individual characteristics include controls for gender and workers' education. Positive revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is positive and zero otherwise. Negative revenue growth is defined as the year-over-year percentage change in revenues when revenue growth is negative and zero otherwise. All regressions require establishments to have at least 20 employees in a given year and cover the period 1997 to 2003.

CHAPTER V

Bank Balance Sheet Shocks and the Effects of the Financial Crisis

5.1 Introduction [1]

It is well known that the financial crisis had a profound effect not only on the financial sector, but also on real firms both in the U.S. and abroad. In contrast, because of data limitations, much less is known about the effect on individual employees employment outcomes, job changes within and across firms, wage dynamics, relocation, and other human costs as a result of the crisis and associated lending constraints of financial institutions. Yet, understanding such individual labor outcomes are key to understanding firm productivity in general and the welfare effects of the financial crisis in particular.

To provide a basis to address some of these questions, this paper uses a unique administrative data set from Germany that offers several key advantages compared to existing sources. First, the data set includes both privately-held firms, such as partnerships and limited liability companies, and publicly-listed firms. The dataset not only contains individual employees wages and job titles, but also each individuals

[1] This chapter is joint work the Daniela Hochfellner (FDZ-IAB), Martin Schmalz (University of Michigan, Ross School of Business), and Denis Sosyura (University of Michigan, Ross School of Business

complete employment history as well as geographical information of establishment locations and workers private residence. The dataset thus allows for the tracking of labor flows within firm, across firms and firm types, and geographic boundaries. In contrast, most of the existing literature has focused on the firm or establishment as the unit of observation and has been unable to follow workers from one job to another, and, as a result, the activity within firms has remained beyond scope of academic research.

Second, institutional features of the German financial system allow for the use of exogenous shocks to the banking systems lending capacity that is sharply confined to geographical regions. Specifically, local savings banks and their state-level organization, which span nearly 40 percent of the German banking market, are strictly prohibited to conduct business outside their geographical domain in Germany. Some of these banks, however, speculated in U.S. mortgage-backed securities during the run-up to the financial crisis and lost billions when the market collapsed, creating shocks to local bank capital that are unrelated to regional economic activity. Combining the unique labor data with the sharp, geographically confined financial shocks to bank capital allow for the clean identification of the effect of these shocks on individual employment outcomes.

Third, the dataset contains financial information, both for the publicly-listed and privately-held firms, in addition tothe worker information. As a result, the dataset allows for the distinction between the impact of financial shocks on individual employment outcomes of privately-held versus publicly-listed firms, differentiate the effect on firms of different legal forms (e.g., private limited liability companies versus partnerships) and investigate the relationship of these labor outcomes with corporate financial policy. This is in contrast to the existing literature that has focused on small versus large firms as proxies for access to financing and has not typically had access to private firms financial information.

To motivate that the bank capital shocks are geographically contained, the paper first document the effects of the shocks on the regional economy as a whole before turning to individual-level analyses. Using detailed, bank-level balance sheets, the analysis shows that affected banks experienced a 20 percentage point lower bank loan growth rate after suffering their losses relative to banks in unaffected states. Next, the reduction in bank lending led to an output reduction by 0.3 percentage points and a 1.4 percentage point increase in the unemployment rate in each of the four crisis years, on average, in affected states compared to states whose savings banks were unaffected by the shock.

The analysis then turns to worker-level data and shows that the employment effect of these shocks particularly affected the privately-held partnerships and limited liability companies and not the publicly-listed firms. Establishments belonging to privately-held firms in affected states reduced net hiring by 24 percentage points and cut investment by one-half, relative to those establishments that belong to publicly-listed firms. This result is consistent with the notion that small firms are more bank-dependent than large publicly-listed corporations, as the privately-held firms are, on average, smaller than the publicly-listed firms. However, the dataset includes both small publicly-listed firms and large privately-held firms, and the results are robust to the inclusion of size controls, suggesting that the firm's legal form is the key distinction for the effect of a negative shock to loan credit and not simply size.

These findings provide insights for public policy, especially with respect to how to manage shocks to the financial sector and their effects on the real economy. First, if firms use financial shocks to make adjustments to their labor demand and aggregate and individual wages, as well as to the composition and hierarchical structure of the work force in an attempt to restore greater efficiency, the best policy might be to avoid mitigating the shock. In contrast, if the destruction of firm-specific human capital is the main effect on firms and individuals, the appropriate policy response may be to contain such exogenous shocks. The results clearly indicate that in such a case, the support should be geared towards the smaller, privately-held firms and the retraining of individual employees in such firms, as those are most affected by a shock to the banking system.

Second, the results reflect a tradeoff, known at least since Smith (1776), which the regulator faces when designing a financial system. On the one hand, the region-specific relationship banking system prevalent in the German economy may help alleviate problems of informational asymmetries between borrowers and lenders in normal times. Also, the geographic confinement of the banks domains mitigates spillovers across banks of a financial shock. On the other hand, the regional economy is left vulnerable to financial shocks that are magnified through their impact on the real economy in such a system. The costs associated with a region-specific banking system offers a rationale why the U.S. financial system transitioned from a regional to a national banking system.

The paper proceeds as follows. Section 5.2 describes the features of the German financial system that are necessary to explain the empirical design. Section 5.3 describes the data and provides summary statistics. Section 5.4 details the empirical strategy and discusses results: section 5.4.1 quanitifies the bank capital shock's impact on loan growth in the affected federal state level, section 5.4.2 esimates the impact of the local bank credit crisis on the federal state macroeconomies, and section 5.4.3 shows the disproportionate impact of the local bank credit crisis on establishments in privately-held firms comapred to their publicly-listed counterparts. Section 5.6 concludes.

5.2 German Banking System and Identification of Shocks

The central hypothesis of this paper is that an exogenous reduction in local credit supply leads adverse effects on local economic outcomes, with a key transmission mechansim being that local credit supply shocks adversely and differentially affect employment and investment in property, plant, and equipment at privately-held partnerships and limited liability companies compared to publicly-listed capital corporations, as the publicly-listed corporations can more readily access capital from both domestic and international equity markets. Therefore, identifying an exogenous shock unrelated to local conditions that leads to a reduction in the supply local bank credit is crucial for testing the paper's hypothesis. This section first discusses the institutional background of the German banking system that makes Germany a convenient setting to study the effects of an exogenous shock to bank credit and then identifies candidate banks that were exposed to U.S. subprime mortgages during the financial crisis, and as a result, saw significant losses in their assets.

5.2.1 German Bank Institutional Background

The German banking system is comprised of three types of banks: public banks, cooperative banks, and private banks. The private banks include commercial banks, investment banks, and private banking and asset management companies, among others. They have a significant international client base. The cooperative banks are built on a membership-share where each member has one vote, regardless of its capital share in the cooperative. The public banks consists of the local savings banks (Sparkassen) and Landesbanks, both of which are not-for-profit entities.[2] The Landesbanks are a group of regional state-owned banks that serve three main functions: first, the Landesbanks' business is predominantly wholesale banking services, and they are the head banking institutions (i.e. central banks and clearing houses) for the local savings banks in each Landesbank's particular region; second, the Landesbanks serve as the lending houses to their federal state(s) to finance infrastructure and social housing projects; third, the Landesbanks serve as commercial banks, primarily

[2]Surpluses are broadly committed to social issues, including the arts, sports, cultural development, and educational issues within the region.

providing funding to big companies (Moody's 2004, Hughes 2008). Each Landesbank is owned by the federal state (Bundesland) where it is located and the collection of local savings banks (directly or indirectly through regional associations) within its particular region. Landesbanks are governed by regional rather than federal law. Which savings banks own a particular Landesbank is solely determined by the region in which the local savings bank resides, and a savings bank is prohibited by law to own shares in any Landesbank outside of its region.

Table 5.1 shows that the private banks own nearly 54 percent of the total bank assets from 1997 through 2011 and comprise 16 percent of all banks, whereas the cooperative banks own 8.5 percent of total bank assets and comprise nearly 55 percent of all banks. Table 5.1 also illustrates the significant role the public banks play in the German banking system, making up nearly 38 percent of the total bank assets in the German economy over the sample period from 1997 through 2011 and comprise roughly 29 percent of all banks.

The German Savings Bank Association (Deutscher Sparkassen- und Giroverband) describes the structure and operatational precedures of the local savings banks to follow five constitutive elements: a public mandate, the regional principle, municipal trusteeship, its legal status and ownership, and its decentralized group nature (Deutscher Sparkassen- und Giroverband 2012). First, the public mandate requires the savings banks to practice a "sustainable business philosophy" focusing on the "appropriate and adequate provision of credit" to all customers, including private customers, regardless of personal income and current financial situation, and for a "sustainable committment to the development of local businesses." The public mandate is issued by law and is the foundation for the savings banks' primary mission: to serve and promote economic and business development within the region. Further, the public mandate allows the savings banks to extend credit to customers that otherwise would not receive credit from other financial institutions.

Second, the regional principle stipulates that the savings banks are authorized to operate only within their region and "that their loan activities should focus on that region" alone. The regional principle therefore prohibits savings banks from lending outside of their region and searching out or accepting customers in more economically viable regions. Further, the savings banks are required by mandate to turn down any loan requests from customers outside of their respective regions. Thus, the regional principle creates an environment of narrow banking by turning "local deposits into local loans." [3] While German savings banks are locally specialized, they are sectorally diversified (OECD 2014).

Third, the municipal trusteeship element of the savings banks attempts to ensure that locally anchored savings banks remain legally and financially independent. A municipality is "the responsible pubic body of a savings bank – but not its owner", where a municipality may refer to a city, town, district, or an "association of local authority for the purpose of jointly running a savings bank". However, a savings bank is not an asset of the municipality, the municipality owns no shares of the savings bank, and the municipality cannot sell the savings banks for revenue generation. A supervisory board, consisting of local council representatives and employees within the region, ensures that the savings banks satisfy their public mandate. However, the supervisory board has no direct impact on day-to-day operations of the savings banks.

Fourth, the legal status and ownership are structured to prevent savings banks from being taken over by private institutions with the primary goal to maximize profits. The savings banks "are incorporated as institutions under public law" that are "legally and financially independent" and have no owners. Their legal form aims to ensure that region's population is "adequately represented", as the supervisory board consists of city council officials and local citizens. However, mangement board

[3] See Kobayakawa and Nakamura (2000) for a survey of narrow banking.

runs the banks' day-to-day operations and consists of banking professionals–and not local politicians–to ensure impartial behavior.

A fifth feature of German savings banks is their decentralized network. Since individual savings banks are relatively small entities alone, the formation of savings bank groups allows for the diversification of risk. Savings banks may only form associations with other savings banks from close local regions (and within the same federal state), and members of these associations share the same common name, typically "Spakasse".

It is through these regional associations that the savings banks construct their ownership (either minority or majority ownership) in the Landesbanks. Federal law requires the local savings bank associations to legally support and maintain liability for the Landesbanks (Moody's 2004). Therefore, Landesbanks may rely on the savings banks for support, and shocks to the equity of a Landesbank directly transmit to the savings banks' balance sheets in its region, and further, local savings banks provide funds to the regional Landesbanks in response to a shock to the equity of the Landesbank.

In addition, there are significant costs to firms associated with switching from the local savings bank to a different bank. As a result, a shock to the capital of a Landesbank that is entirely unrelated to the local economy can have consequences for the provision of credit to local businesses, and properly identifying an exogenous shock to a Landesbank's capital potentially allows for the measure of the effect of an exogenous shock to the banking system on economic outcomes. Further, identifying an exogenous shock that affects only some of the Landesbanks rather than all would allow for the credible exploitation of regional variation in economic outcomes due to the contained, regional nature of the German banking system.

5.2.2 Identification of Bank Shocks

Figure 5.1 shows a the regional distribution of the German Landesbanks and the federal states each Landesbank covers as of 2007 and prior to the crisis. In 2007, 10 Landesbanks covered the 16 German federal states. Each Landesbank served as the central bank in typically one federal state, with the Helba and Nord/LB Landesbanks covering three federal states each.

The first signs of a Landesbank's exposure to the U.S. subprime crisis occured on August 17, 2007, when Sachsen LB, the Landesbank of the German federal state Saxony with total assets of 68 billion euros in 2007, was forced to take an emergency rescue loan in the amount of 17.3 billion euros from the German savings bank association, Sparkassen-Finanzgruppe, due to its exposure to U.S. asset backed securities (Simensen 2007). Sachsen LB's exposure to the U.S. subprime crisis stemmed from an off-balance sheet subsidiary, Ormund Quay, located in Dublin, Ireland. Ormund Quay borrowed significantly in short-term commercial paper and invested in long-term asset-backed securities, a transaction supported by a credit line from Sachsen LB.

As the U.S. subprime crisis unfolded, investors refused to refinance Ormund Quay's commercial paper debt, and Sachsen LB was unable to meet its pledged line of credit, necessitating the emergency credit bailout (Moody's 2008). At the time, *Spiegel Online* reported that Sachsen LB's losses due to direct involvement in subprime mortgages approached 500 million euros[4], whereas the German newspaper *Suddeutsche Zeitung* reported Sachsen LB had as much as 65 billion euros in five funds at Ormund Quay. State officials announced on August 26, 2007, that Sachsen LB would be sold to Landesbank Baden-Württemberg (LBBW), the central clearing house for the savings banks located in Baden-Württemberg and Rheinland-Palatinate, due to

[4] http://www.spiegel.de/international/business/debt-exposure-and-off-balance-sheet-loans-banks-in-germany-wobble-a-500833.html

the subprime losses. Sachsen LB no longer existed as a separate entity as of April 2008, at which point the local savings banks of Saxony transferred their holdings to LBBW. LBBW would now serve as the central clearing house for the savings banks in Baden-Württemberg, Rheinland-Palatinate, and Saxony.[5]

The second Landesbank to report losses due to exposure to the U.S. subprime crisis was HSH Nordbank, the central clearing house for the savings banks of the federal state Schleswig-Holstein and city-state Hamburg and with total assets of 174 billion euros. Though reporting strong profits for most of 2007, on August 23, 2007, HSH Nordbank said it had 1.8 billion euros invested in securities backed by U.S. subprime mortgages, primarily through its subsidiaries, Poseidon and Carrera, and HSH chief executive Hans Berger remarked, "We have a liquidity squeeze in the market, especially for lending between banks" (Kirchfield and Schmidt 2007). Berger stepped down in September 2008 as a result of the exposure to the U.S. crisis and subsequent writedowns and announced a plan to retructure its business and focus more on its core in Northern Germany going forward. HSH Nordbank had writedowns of 1.1 billion euros and a loss of 210 million euros in 2007 (Seuss and Kirchfield 2008).

Moody's downgraded HSH Nordbank's long term outlook in a Novermber 2008 report, citing its increased risk profile and stretched financial profile due to direct exposure to Lehman Brothers. Moody's also expected HSH to rely on strong support from the public banks going forward (Moody's 2008). In December 2008, HSH Nordbank was guaranteed notes of 30 billion euros from the German federal government's rescue fund. On February 24, 2009, HSH Nordbank announced a deal with the federal state Schleswig-Holstein and the city state of Hamburg to receive a capital injection of 3 billion euros and a state backed credit guarantee of 10 billion euros.

WestLB, the Landesbank of the federal state North Rhine-Westphalia and central clearing house for the state's savings banks, was the next Landesbank to announce

[5]The politicaly aftermath of the Sachsen LB emergency bailout and sale resulted in Georg Milbradt, the premier of Saxony, resigning from his position in April 2008.

losses due to exposure to the U.S. subprime crisis on August 27, 2007, saying it had nearly 1.25 billion euros of direct exposure to U.S. mortgage-backed securities (Clark 2007). WestLB was the third largest Landesbank at the time, with total assets of 285 billion euros in 2007. Its exposure stemmed from five subsidiaries which borrowed money by selling short-term commercial paper and investing those funds in securities backed by U.S. mortgages. In an attempt to limit the bank fallout from the exposure to U.S. mortgage assets, WestLB announced on December 3, 2007, it would guarantee full liquidity to its subsidiaries exposed to U.S. asset-backed securities, with each having the option of drawing as much as 25 billion euros (Dougherty 2007). However, in February 2008, WestLB received a 5 billion euro rescue package from the state of North Rhine-Westphalia and the two local savings banks associations (Rheinischer Sparkassen- und Girover- band and Westfalisch-LippischerSparkassen-undGirover- band). Losses of up to 2 million euros were to be absorbed by the owners of WestLB according to the respective ownership stakes in the bank, of which the two savings bank associations of North Rhine-Westpahlia owned more than 50 percent (Puri et al. 2011). On April 2, 2008, West LB reported a net loss of 1.6 billion euros for 2007, directly citing exposure to U.S. mortgage-backed securities.

In addition to the 5 billion euro rescue package in February 2008, WestLB announced the creation of a bad bank called Erste Abwicklungsanstalt (EAA), and 85 billion euros in toxic assets were transfeed from WestLB to EAA in November 2009. While referred to as a bad bank, EAA is not a bank as it does not hold a banking license nor attempts to generate new business. EAA was established as a specialty agency with the mission to wind up a financial portfolio of toxic assets responsibly and sustainably.[6] Therefore, EAA was viewed as serving a public function. On December 20, 2011, the European Commission approved a liquidiation plan for WestLB submitted by the German government. After June 30, 2012, WestLB stopped taking

[6]For more information on EAA's specific functions, please see https://www.aa1.de/en/about-us/faq/general-questions/.

new banking business (Lienemeyer and Magnus, 2011) and was disolved.[7]

Germany's second largest Landesbank with assets of 353 billion euros in 2007, BayernLB, the Landesbank of the fedral state Bavaria and the central clearing house for Bavaria's savings banks, was the fourth Landebank to report significant losses due to the U.S. subprime crisis. The state of Bavaria and the savings banks association, Sparkassenverband Bayern, each owned 50 percent of BayernLB in 2007. BayernLB announced on February 13, 2008, it would write down 1.9 billion euros with direct losses of 150 million euros due to U.S. subprime related investments in 2007 (Morajee 2008). BayernLB's chief executive, Werner Schmidt, resigned less than a week later over the losses (Morajee and Atkins 2008). By March 2008, BayernLB's writedowns reached 4.3 billion euros, with estimated losses at 6 billion euros. Of the estimated 6 billion in losses, Bayern LB would be responsible for 1.2 million, whereas the two owners of Bayern LB, the state of Bavaria and the savings bank association, Sparkassenverband Bayern, would be responsible for 2.4 million euros each (Reuter 2008). In April 2008, a *Spiegel Online* report brought BayernLB under heavy criticism, as it discovered the Landesbank knew about its U.S. subprime related losses in the second half of 2007, but did not reveal those losses to the public until February 2008.[8] Losses in the second half of 2007 would place the U.S. subprime crisis's impact on Bayern LB on a similar timeline to the impact on Sachsen LB and WestLB.

On October 21, 2008, BayernLB became the first bank to draw on support from the German federal government's 500 billion euro bailout fund, applying for 5.4 billion euros of the rescue funding. BayernLB also announced it faced an additional loss of up to 3 billion euros by the end of 2008 due to further exposure to the U.S. subprime crisis and the recent collapse of Lehman Brothers. The additional unexpected losses

[7] At this point, what remained of WestLB began to operate as Portigon Financial Services AG. EAA and the Landesbank of Hessian and Thuringian, Helba, handled the bankruptcy and carried on with the core functions of the former WestLB.

[8] See http://www.spiegel.de/wirtschaft/parteichef-am-pranger-bayernlb-krise-erschuettert-csu-huber-in-not-a-545159.html for more details.

prompted the resignation of the Bavaria's finance minister, Erwin Huber, the first politician to resign over Landesbank crisis.[9] In November 2012, BayernLB began repaying the aid received in 2008 with a payment of 350 million euros to the state of Bavaria. To complete the agreement for receiving the 2008 aid, BayernLB must repay the full 5.4 billion euros of rescure funding by 2019 and reduce its balance sheet to half its 2008 level (Suess 2012).

The fifth and final Landesbank to report losses directly attributed to exposure in the U.S. subprime crisis was Germany's largest Landesbank, Landesbank Baden-Württemberg (LBBW), with total assets in 2007 of 443 billion euros and an ownership structure of 40.5 percent by the State of Baden-Württemberg, 40.5 percent by the savings bank associations of Baden-Württemberg and Rhineland-Palatinate, and 19 percent by the City of Stuttgart (Moody's 2008). LBBW serves as the central clearing house for the savings banks of three German federal states: Baden-Württemberg, Rheinland-Palatinate, and Saxony.[10] While LBBW remained bullish on its operating business in early 2008, due to its strong market position in the core businesses of Baden-Württemberg and Rheinland Palatinate, LBBW annonced in November 2008 that it faced 800 million euros of writedowns and 1.1 billion euros of losses, citing direct exposure to U.S. subprime mortgage-backed securities (Luttmer and Simensen 2008). By the end of 2008, LBBW reported a loss of 2.1 billion euros.

In November 2008, the state of Baden-Württemberg, the city of Stuttgart, and the regional savings bank assocciations of Baden-Württemberg and Rheinland-Palatinate agreed to a 5 billion euro capital injection and a 12 billio euro lifeline to support LBBW. While a Moody's (2008) review of LBBW viewed the capital injection and

[9]See http://www.spiegel.de/international/germany/financial-crisis-aftermath-bavarian-finance-minister-quits-over-bank-losses-a-585739.html for more details.

[10]While always serving as the central bank for the saving banks of Baden-Württemberg, LBBW assumed complete central banking responsibilities for Saxony in April 2008, after SachsenLB failed due to its exposure to U.S. subprime asset-backed securities, and for Rheinland-Palatinate in July 2008 when Landesbank Rheinland-Palatinate was completely integrated into LBBW and LBBW assumed a 100 percent ownership share of Landesbank Rheinland-Palatinate.

LBBW's committment to reduce secondary market activities and related investments as a long-term positive, Moody's also expected this to be a slow process. LBBW did not return to profit until 2012.

Table 5.2 summarizes the identified Landesbanks exposed to the U.S. subprime crisis, the date each bank announced its first losses, the time period when each bank was expected to experience its first losses due to the crisis, and the resulting affected federal states and savings banks where each Landesbank served as the central bank. The approach reveals exposure to the U.S. subprime crisis by five Landesbanks–SachsenLB, HSH Nordbank, WestLB, BayernLB, and LBBW–that serve as the central bank for savings banks in seven of the sixteen German federal states–Saxony, Schleswig-Holstein, Hamburg, North Rhine-Westphalia, Bavaria, Baden-Württemberg, and Rheinland-Palatinate. SachsenLB, HSH Nordbank, and WestLB all announced first losses due to exposure to the U.S. subprime crisis within ten days of each other in August 2007. Further, while BayernLB did not annonce its first losses until February 2008, there is considerable evidence that BayernLB experienced its first losses in the third quarter of 2007, a similar time to that of SachsenLB, HSH Nordbank, and WestLB. Thus, the narrative suggests that four Landesbanks were in crisis due to exposure to the U.S. subprime crisis in the third quarter of 2007, affecting the savings banks in five German states–Saxony, Schleswig-Holstein, Hamburg, North Rhine-Westphalia, and Bavaria. The final exposed Landesbank, LBBW, went into crisis a year later, affecting the savings banks of Baden-Wurttemberg and Rheinland-Palatinate. Accordingly, 2007 marks the beginning of the crisis for the remainder of the study.

Figure 5.2 shows a map of the affected and unaffected German federal states. The map shows significant geographical dispersion in affected states, as there are affected states located in the north, east, south, and west. Further, each affected state borders an unaffected state, allowing for stark regional variation. .

5.3 Data and Summary Statistics

This section describes the data[11] used throughout the analysis and its descriptive statistics. The data comprise several unique features that allow for the observation of a large fraction of the German economy over a long period of time in a stratified linked-employer-employee sample. Section 5.3.1 discusses the the bank-level data used to quantify the bank shock, section 5.3.2 discusses the establishment- and worker-level data derived from German social security records, and section 5.3.3 discusses the regional macroeconomic variables used to quantify the state-level macro effects of the bank capital shocks.

5.3.1 Bank Data Overview and Summary Statistics

The paper uses Bankscope data provided by the Bureua van Dijk.[12]Bankscope provides the universe of annual German bank balance sheets over the sample period from 1997 through 2011. These balance sheets include detailed information on the banks' annual total assets, including annual total loan value, and the banks' annual liabilities and equity. Bankscope also provides each bank's registered name, street address, postcode, and specialty (e.g. savings bank, cooperative bank, investment bank, etc.), allowing for the identification of the savings banks that were exposed to the U.S. subprime crisis through ownership in their respective Landesbanks and specific German federal states[13] these local savings banks serve. As described in section 5.2.1, the local savings banks are required by mandate to only serve customers within its district; thus, the geographical, bank specialty, and asset and loan information in the Bankscope data set allows for the identification of the regional bank shocks that arose from the subprime crisis.

[11]Data access was provided through the ISR-FDZ at the University of Michigan, an U.S. on-site location of the Research Data Center of the Federal Employment Agency at the Institute for Employment Research.

[12]For further details, see www.bankscope.com.

[13]The German federal state is similar to the state in the United States.

Table 5.3 shows the summary statistics for the aggregate, federal state bank balance sheets in 2005 euros for the whole sample, 1997 through 2011, in the first column, and for the crisis period, 2007 through 2010, in the second column. [14] Comparing the values of total average assets and total average loans during the crisis period to that of the entire sample shows that the level value of assets and loans were, on average, larger during the crisis years than earlier in the sample. This observation is mostly attributed to trend growth for both variables over the sample. Comparing the average annual growth rates in total assets and total loans during the crisis period to that of the entire sample shows a clear, yet small contraction both asset and loan credit during the crisis. Average total asset growth is 3.11 percentage points lower during the crisis period compared to the whole sample, whereas loan growth is only 1.93 percentage points lower. The average annual growth rate and standard deviation for total loans are on par with those of total assets indicating that fluctuations in total loans drive fluctuations in total assets over the sample and crisis period.

Table 5.4 shows the summary statistics for the aggregate, federal state bank balance sheets in 2005 euros for only the crisis period, 2007 through 2010, divided into the federal states unaffected and affected by the Landesbank exposure to the U.S. subprime crisis. The first column shows the summary statistics for the 9 federal states that did not have its Landesbank exposed to the U.S. subprime crisis. Total asset growth was positive and relatively large in the unaffected states during the crisis period, averaging a growth rate of 4.72 percent with a standard deviation of 21.87 percent. Similarly, total loan growth was positive and relatively large in the unaffected states, averaging 4.62 percent with a standard deviation of 25.75 percent.

Conversely, the 7 affected states with Landesbanks exposed to the U.S. subprime crisis saw a significant contraction in both asset and loan growth during the crisis

[14] The category of total assets for a bank consists for total loans, other earning assets such as advances to banks, derivatives, and securities, and fixed assets, whereas total loans for a bank consists of mortgage loans, consumer and retail loans, and corporate and commercial loans.

period. Total asset growth averaged -10.48 percent in the affected states during the crisis, with a standard deviation of 23.49 percent, and total loan growth averaged -8.02 percent, with a standard deviation of 18.55 percent. Comparing to the unaffected states, the affected states experience a 15.20 percentage point lower growth rate in total assets and a 12.64 percentage point lower growth rate in total loans during the crisis period.

5.3.2 Establishment- and Worker-Level Data and Summary Statistics

The paper employs administrative and survey data from the Institute for Employment Research (IAB) of the German Federal Employment Agency (BA) at the Institute for Employment Research. The analysis sample is constructed from various data sources held at IAB to combine information on employers and employees. The establishment sample is based on the IAB Establishment Panel, which is a survey that is conducted annually by IAB since 1993 for West Germany and since 1996 for West and East Germany. The sample includes all West German establishments that completed a survey in at least one year from 1997 through 2011 and identified themselves as an establishment that is part of an unlisted, privately-held, partnership company or limited liability company, or as part of a listed, publicly-traded, capital corporation. An establishment in this sample always refers to a local unit of a firm–that is, a specific plant or building.

In addition, it is possible to merge in firm-level data for these establishments at the FDZ. These firm data come from Bureau vam Dijk's Orbis database and include balance sheets of all privately-held, limited liability and publicly-listed capital coporation firms operating in Germany from 1991 to 2014. The FDZ provides a crosswalk from establishment to firm in order to identify which establishments operate under one firm. Balance sheets on both privately-held and publicly-listed firms is a considerable advantage of this study, as comparable datasets, such as Compustat, typically contain firm balance sheet information only for publicly-listed firms.

For each of the sampled 15,392 establishments, the sample also includes complete worker-level histories for every worker who was employed and liable to social security at these establishments at least one day between 1997 and 2011. The worker-level labor market biographies provide detailed information on exact employment start and end dates and an extensive set of employment-related characteristics such as the type of employment, wages, professional and occupational status, and white-collar versus blue collar jobs plus detailed individual characteristics, such as gender, birth year, nationality, education, and vocational training. Additionally, the samples follow these workers' entire employment, unemployment, and wage histories from 1975 through 2010, even if the workers move to an establishment outside the sample. A detailed description of the variables included can be obtained from Dorner et. al (2010).

The resulting individual-level employment histories are complemented with administrative establishment data from the Establishment History Panel (BHP) (please see Spengler 2008 for detailed information on the BHP). The BHP includes industry classification codes and state- and district-level[15] location identifiers for each establishment. Regional identification of establishments is especially important for this paper, as the identification of those establishments exposed to the exogenous shock to bank credit relies on the regional variation inherent in the German banking system. In addition, the BHP contains an extension file with information on establishment births, deaths, and re-classifications. Supplementary data in this extension allows for the identification of establishment closures that are likely to be spin-offs or takeovers as opposed to true closures.

The paper also make use of the survey responses from the IAB Establishment Panel Survey. The survey is supported from the German Minister of Labor, and therefore shows a response rate of about 80 percent among the establishments that

[15]Districts in Germany are comparable to counties in the United States.

stay in the panel (Janik and Kohaut 2014). It provides information regarding an establishment's investment in plant, property, and equipment, including, but not limited to, the total euro value of investment spending at the establishment within the given calendar year for every year, the fraction of investment financed through methods such as cash flow, equity, bank loans, and government subsidies, and qualitative measures such as whether or not the establishment had difficulty acquiring bank loans within the given calendar year. Further, the Establishment Panel Survey devotes a special section of the 2010 survey to examine the impact the economic and financial crisis that started in 2007 had on establishment behavior during the crisis period. The survey questions help shed light on whether establishments in privately-held companies qualitatively had a differential and worse experience than establishments in publicly-traded companies as a result of the crisis, aiding the identification of the empirical work developed throughout the remainder of this study.

Table 5.5 shows the establishment-level descriptive statistics where the unit of analysis is the establishment-year. The employment statistics suggest there is substantial variation in establishment size across the data set and that some very large establishments are pulling up the arithmetic average. The average establishment-year net hire rate is -0.9 percent over the sample, where an establishment's net hire rate in a given year is defined as the total number of employee inflows minus the total number of employee outflows as a fraction of total employment as of January 1 of the given year. The median net hire rate is -0.2 percent, and the standard deviation is 27.0 percent, indicating substantial variation on net employment flows across establishment-years. The small and even slightly negative average establishment net hire rates are consistent with the macro-level employment growth statistics. Germany's working age population steadily declined over the sample period, and the 2001 through 2004 recession and 2007 through 2010 bank crisis period cover 8 out of the 15 years in the sample. Both contributed to muted average employment growth.

Worker inflow and outflow rates, however, are substantially larger over the sample period, indicating significant worker movement across establishments.

The mean value of investment per employee is 12,822 euros per establishment year over the sample, with a median level of investment per employee of 6,651 euros and a standard deviation of 55,053 euros. An investment per employee mean nearly twice that of its shows that some establishments in the data set had years with investments significantly larger than the other establishments in the dataset.

Table 5.6 shows a summary of the employment level, net hire rate, and investment per employee statistics broken down by establishments belonging to privately-held and publicly-listed firms. The statistics suggest clear distinctions between establishments in privately-held and publicly-listed firms. That the median employment level is much lower for publicly-listed with 278 employees than privately-held with 48 employees when both are compared to their respective means suggests that a small number of very large establishments are pulling up the mean much more for the publicly-listed than for the privately-held, though there are a number of quite large establishments in the privately-held firms. However, the summary statistics on employment levels indicate a distribution for both privately-held and publicly-listed that includes a significant amount of both small and large establishments. However, the average net hire rates over the entire sample paint more similar picture between establishments in privately-held firms and those in publicly-traded firms and are on par with the whole sample average presented in table 5.5.

Investment per employee differes noticeably between the privately-held and publicly-traded establishments. The mean and median investment per employee for the privately-held establishments is 10,829 and 5,068 euros respectively compared to 15,741 and 9,185 euros, respecitively, for the publily-listed. Clearly, are investment per employee is much larger for publicly-listed establishments than the pivately-held. The standard devation of investment per employee is also significantly larger for the publicly-listed at 76,359 euros compared to 33,314 for the privately-held.

Tables 5.7 and 5.8 show the establishments' responses to survey questions directly related to financing methods and the crisis. The survey responses suggest that the establishments belonging to privately-held firms depend more on bank loans than establishments in publicly-listed firms to finance their business activities and that privately-held had a more difficult time obtaining the bank loans during the bank credit crisis than the publicly-listed. Further, the privately-held in affected states with banks exposed to the U.S. subprime crisis reported being more strongly affected by the crisis than publicly-listed (regardless of whether the publicly-listed establishments were in affected or unaffected states) and their privately-held counterparts in unaffected states.

5.3.3 Macroeconomic Data and Summary Statistics

Annual federal state-level macro statistics are provided by the official German statistical agency and are used to quantify the regional macro economic impact of the bank shock. These statistics include the 2005 consumer price index and state-level output, household income, total employment levels, full-time and part-time employment levels, and unemployment rates. The 2005 CPI deflates all nominal, euro-valued variables used in the paper to real variables.

Table 5.9 shows the summary statistics for key federal state-level macroeconomic variables over the sample period, 1997 through 2011. The first column shows the summary statistics covering the entire sample, the second column shows the summary statistics over the 2001 through 2004 recession, and the third column shows the summary statistics spanning the bank loan crisis period from 2007 through 2010. The unit of observation is a given federal state in a given year, and all statistics shown in the table are weighted by the square root of the federal state's population in the given year to provide a better indication of overall German macroeconomic performance. All euro valued variables are expressed in real values of 2005 euros, and all growth rates are calculated as the percent change in the federal state from year $t-1$ to year t.

Average real output growth at the federal state-level over the entire 15 year sample is low, which can largely be explained through Germany experiencing a long recession from 2001 through 2004, with a slow recovery out of the recession after 2004, and a second recession during with the bank loan crisis period from 2007 through 2010. Output growth during the 2001 through 2004 recession was only 0.2 percent with a small standard deviation of 1.0 percent across observations, indicating a relatively uniform growth experience across federal states from 2001 through 2004. Comparing these numbers to the bank loan crisis period from 2007 through 2010, average annual output growth during the crisis was slightly higher at 0.4 percent. However, the standard deviation of output growth during the crisis period was 3.5 percent, noticeably larger than the standard deviation of output during the 2001 through 2004 recession, and indicating a much different growth experience across federal states during the crisis.

Average real household income growth was 0.5 percent over the entire sampl, with a standard deviation of 1.1 percent. Household income grew at a similar pace during the 2001-2004 recession with an average growth rate of 0.6 percent and a standard devition of 1.0 percent. Somewhat remarkably, though, average real household income grew at a strong pace during the bank loan crisis period when compared to both the entire sample and the 2001 through 2004 recession, with an average growth rate of 1.6 percent. The standard deviation during the crisis period was also larger, measuring at 1.6 percent, again indicating a more varied experience across federal state over the crisis period than over other periods during the sample. That household income experienced relatively strong average growth during the bank crisis period helps to support that the contraction in loan growth over the period is not demand driven.

Total employment growth over the entire sample period was 0.7 percent with a standard deviation of 1.6 percent, a low growth rate which is explained by both a steady decline in the overall and working populations in Germany and the extended downturns through the time period. During the 2001 through 2004 recession, average total employment growth was -0.7 percent with a standard deviation of 1.2 percent. Remarkably, though, average total employment grew steady pace over the loan crisis period from 2007 through 2010 at an average rate of 1.1 percent and a standard deviation of 1.1 percent.

Breaking employment growth into both full-time employment growth and part-time employment growth helps to explain some of the dynamics occurring in the total employment growth numbers. Full-time employment at the federal state-level contracted, on average, from 1997 through 2011, whereas part-time employment grew at a strong rate. During the 2001 through 2004 recession, full-time employment fell indicating a relatively uniform contraction in growth rates across federal states, whereas part-time employment grew. The loan crisis period from 2007 through 2010 saw a modest average growth rate in full-time employment albeit with strong variation in the growth rates across federal states, as some federal states saw full-time employment grow and some federal states saw full-time employment contract. However, most federal states experienced strong part-time employment growth during the crisis period.

The stark differences between full-time and part-time employment growth can be explained through a series of labor market policies instituted in Germany over the sample period. In 1999, Germany introduced the first of the Hartz Reforms of the labor market that was instituted in four waves through 2005. A second German labor market policy that helps explain the strong growth in part-time employment during the bank loan-crisis period is the institution of short-time work.[16] While short-time

[16]Short-time work is a program in which workers accept a reduction in work hours in exchange for government subsidies through employers to augment their salary to make up for wages lost due

work as a labor market policy has been available in Germany since the 1970s, it was only sparingly used until the the crisis, when Germany experience a large spike in the number of workers on short-time, peaking at nearly 1.5 million workers at the height of the crisis. The increase in short-time work has largely been viewed as a success in avoiding mass layoffs, reducing the rise in the unemployment rate by roughly half of what it would have been without the policy during the crisis (Brenke et al. 2011) and saving nearly 500,000 jobs (Balleer et al. 2013).

The unemployment rate over the entire sample averaged 9.4 percent, with a standard deviation of 4.5 percent. The large standard deviation in the unemployment rate stems from the noticeably larger unemployment rates in the East German federal states compared to the West German federal states. The average unemployment rate rose to 10.6 percent with a standard deviation of 5.1 percent during the 2001 through 2004 recession. However, the unemployment rate fell to 7.7 percent during the loan crisis period when compared to the whole sample and the 2001 through 2004 recession, with a markedly lower measured standard deviation of 3.1 percent. These trends during the crisis can again be attributed to active German labor market policies meant to combat mass layoffs and high unemployment rates during the crisis.

5.4 Regression Results

This section provides the empirical results quantifying the effect of the bank capital shock on regional bank loan credit, the impact of the bank loan credit contraction on regional macroeconomic outcomes, and the differential effect of the bank loand credit shock on establishments in privately-held firms compared to establishments belonging to publicly-listed firms.

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5.4.1 Bank Loan Credit During the Crisis

Section 5.2.2 documents that five Landesbanks, owned by the savings banks in seven federal states, were exposed to and experienced significant losses directly attributed to the U.S. subprime crisis starting in 2007. This leads to a key question of the analysis: did bank loan credit from savings banks siginicantly fall in the seven federal states with Landesbanks that experinced losses due to direct exposure to U.S. subprime asset-backed securities?

While the differences in average growth rates between the affected and unaffected states presented in table 5.4 are quite stark, isolating the effect and statisitcal significance of the crisis on the bank activity in the affected states requires a more formal analysis. The nature of the banking shock allows for the exploitation of both time-series and cross-sectional variation to isolate its impact on bank asset and loan growth, exploiting time variation through the pre-crisis and crisis periods and exploiting cross-sectional variation through the affected and unaffected states through each state's Landesbank exposure to the U.S. subprime crisis. Thus, to quantify the impact of the shock on asset and loan growth in the affected states during the crisis, the analysis relies on a the following difference-in-difference (diff-in-diff) specification:

$$\begin{aligned} BankActivity_{it} &= \beta_0 + \beta_1 Crisis_t + \beta_2 AffectedState_i \\ &\quad + \beta_3 (Crisis_t \times AffectedState_i) + W'\Upsilon + \epsilon_{it} \end{aligned} \tag{5.1}$$

where $BankActivity_{it}$ is either total asset growth or total loan growth in federal state i from period $t-1$ to t, $Crisis_t$ takes a value of 1 during the crisis period from 2007 through 2010 and 0 otherwise, $AffectedState_i$ takes the value of 1 if the federal state is one of the seven states to have its Landesbank exposed to the U.S. subprime crisis and zero otherwise, $(Crisis_t \times AffectedState_i)$ takes a value of 1 for an affected state during the crisis period and zero otherwise, and W is a vector of control variables including a dummy variable for whether a federal state is in East Germany and a dummy variable that takes a value of 1 for the 2001-2004 recession and 0 otherwise.

Validity of the diff-in-diff approach relies on satisfying the parallel trend assumption. The parallel trend assumption applied to the bank activity framework in equation (5.1) requires that total assets and total loans for both the unaffected and affected states follow the same trend in absence of the affected states' Landebanks exposure to the U.S. subprime crisis. Figure 5.3 shows the time series of total real loans (in 2005 euros) aggregated for all 9 of the unaffected states and the time series of total real loans aggregated for all 7 of the affected states, both indexed to 100 in 2004. The figure appears to tell a very clear story: the trends for total loans in both the unaffected and affected states are nearly identical until 2007, when there is a clear break. Beginning in 2007 (the start of the crisis period), the total loans time series for the affected states begins a clear downward trajectory while the total loans for the unaffected states continues to grow. Therefore, the behavior of the total loans time serieses for the unaffected and affected states strongly suggests that the parallel trend assumption applies.

Table 5.10 shows the results for the estimated equation (5.1), with column 1 showing the results where the federal state total bank asset growth rate is the dependent variable, and column 2 showing the results where the federal state total loan growth rate is the dependent variable. The key estimated coefficient of interest for both the asset and loan growth regressions is $\hat{\beta}_3$–the coefficient on the $(Crisis_t \times AffectedState_i)$ interaction variable–which isolates the impact that exposure to the U.S. subprime crisis had on the bank asset and loan growth rates in the affected states under the diff-in-diff. The estimated coefficient on the interaction term in the asset growth rate regression takes a value of -.245 with a standard error of .077, and the estimated coefficient on the interaction term in the loan growth rate regression takes a value of -.202 with a standard error of .076. Both coefficients are statistically significant at the 99 percent confidence-level.

The interpretation of the estimated coefficients on the interaction term for both regressions is as follows: Total bank asset growth was 24.5 percentage points lower in the 7 federal states with Landesbanks exposed to the U.S. subprime crisis during the crisis period, and total loan growth was 20.2 percentage points lower in affected states compared to the unaffected states during the crisis.[17] Both estimates, however, indicate a severe and marked contraction in bank activity in the affected states during the crisis.

While the regression results in table 5.10 provide clear evidence of a severe contraction in bank activity in the affected states during the crisis, the question still remains as to whether the contraction is due to a reduction in the supply of available bank credit in the affected states or the reduction in the demand for bank credit. The narrative in section 5.2.2 suggests the exposure to the U.S. subprime crisis led to a marked reduction in the affected states' bank balance sheets, and thus a marked reduction in the supply of credit. A formal analysis to disentangle the supply and demand effects, however, requires analyzing whether the number of loan applications fell in the affected states during the crisis or whether the loan acceptance rate decreased in the affected states during the crisis. Unfortunately, data on the number of loan applications and the acceptance rate of loan applications is not publicly available nor is it available in the restricted access datasets made available for this study. Puri et al. (2011) use individual bank loan application data from the Bundesbank spanning the time period from the third quarter of 2006 through the second quar-

[17]The difference between the estimated coefficients on the interaction term in the asset growth and loan growth regressions compared to the simple summary statistics from Table 5.4 comparing asset and loan growth in the affected and non affected states can be explained by the estimated coefficient, $\hat{\beta}_2$, on the affected state dummy. These estimated coefficients indicate that affected states have, on average, a 9.3 percentage point higher gorwth rate in assets and a 8.0 percentage point higher growth rate in loans over the sample period. Adding $\hat{\beta}_2$ and $\hat{\beta}_3$ in both the asset growth and loan growth regressions yields a difference of 15.2 and 12.2 percentage points, for each respective regression, when comparing the affected states versus the not affected states. These numbers are nearly identical to the differences calculated from the means in table 5.4.

ter of 2008 to study exactly whether the number of loan applications to the public savings banks decreased in the affected states during the crisis or whether the loan rejection rate increased. They find clear evidence that loan acceptance rates at the public savings bank fell by nearly 12.5 percentage points (from a loan acceptance rate of 97.34 percent in the third quarter of 2006 to 84.93 percent in the second quarter of 2008) in the affected states compared to no decrease in the unaffected states and that any contraction in the available credit can be attributed to a reduction in the supply of bank loans.

5.4.2 The Macroeconomic Impact of the Bank Loan Credit Shock

Section 5.4.1 provides stark evidence that bank loan credit fell drastically in the federal states with Landesbanks exposed to the U.S. subprime crisis, with loan growth rates contacting by 20 percentage points compared to loans in federal states without exposed Landesbanks. However, the question remains as to whether the bank credit crunch in the affected federal states had any impact on these states' real economy.

The following diffrence-in-difference empirical specification again exploits the time-series and cross-sectional variation in the bank shock to quantitatively examine the impact that a reduction in bank loan credit due to Landesbank exposure to the U.S. subprime crisis has on the macroeconomic variables at the federal state-level:

$$\begin{aligned} MacroVariable_{it} &= \delta_0 + \delta_1 Crisis_t + \delta_2 AffectedState_i \\ &\quad + \delta_3 (Crisis_t \times AffectedState_i) + X'\Psi + \eta_{it} \qquad (5.2) \end{aligned}$$

where $MacroVariable_{it}$ is a vector including output growth, total employment growth, full-time employment growth, part-time employment growth, and the unemployment rate in federal state i and period t; $Crisis_t$ takes a value of 1 during the bank loan crisis period of 2007 to 2010; $AffectedState_i$ takes a value of 1 if one of the 5 Landesbanks exposed to the U.S. subprime crisis serves as the central bank to the savings banks of that federal state; $(Crisis_t \times AffectedState_i)$ takes a value of 1 for an affected state during the crisis and 0 otherwise; and X is a vector of control variables including a dummy variable for whether a federal state is in East Germany; a dummy variable that takes a value of 1 for the 2001 through 2004 recession and 0 otherwise; a dummy variable that takes a value of 1 for the years 2005 and after to account for the implementation of the most significant stage of the Hartz reform; and the growth rate of the number of short-time workers in the economy to control for active German labor market policies.

The key coefficient of interest in the estimation of equation (5.2) is δ_3, the coefficient on the $(Crisis_t \times AffectedState_i)$ interaction term. The estimated coefficient, $\hat{\delta}_3$, quantifies the differential effect of the bank balance sheet shock on the macroeconomic variables in the affected states with exposed Landesbanks during the crisis relative to the unaffected states in which its Landesbanks were not exposed during the crisis. Therefore, the point estimate of $\hat{\delta}_3$ should isolate the impact of the negative shock to bank credit on the macroeconomic variable of choice and will be the focus of the analysis below.

Table 5.11, column 1 shows the results of the estimated equation (5.2) with real output growth serving as the dependent variable. The key estimated coefficient of interest, $\hat{\delta}_3$, on the $(Crisis_t \times AffectedState_i)$ is -.003, implying that affected states with Landesbanks exposed to the U.S. subprime crisis had a 0.3 percentage point lower output growth rate than unaffected states, on average, during the four year crisis period. However, the point estimate on the interaction coefficient is statistically insignificant at 90 percent confidence level. However, the point estimate on the $Crisis_t$ variable is -0.011 and statistically significant at the 95 percent confidence level implying that all states experienced an average decline in output growth of 1.1 percentage points, on average, over the four crisis years. This fall in output growth corresponds to approximately 2.5 times the sample average output growth rate.

Columns 2 through 4 of table 5.11 show the results of the estimated equation (5.2) with total employment growth rate, full-time employment growth rate, and part-time employment growth rate as the respective dependent variables. The point estimate of $\hat{\delta}_3$ isolating the impact of the bank balance sheet shock in the affected states during the crisis is negative in all three equations, implying that an exogenous contraction in loan growth due to exposure to the U.S. subprime crisis results in a contraction in employment growth. The estimated coefficient is statistically significant at the 95 percent confidence level for both the total employment growth and the full-time employment growth regressions, whereas the estimated coefficient is statistically insignificant for the part-time employment growth regression. Economically, though, the point estimates imply that the bank loan credit shock in an affected state during the crisis resulted in a 0.9 percentage point reduction in total employment growth, a 0.8 percentage point reduction in full-time employment growth, and a 0.9 percentage point reduction in part-time employment growth relative to unaffected states during the crisis. These employment growth reductions are equivalent to 128.5, 225.0, and 28.1 percent reductions in total employment growth, full-time employment growth, and part-time employment growth relative to each respective sample average.

Table 5.11, column 5 shows the estimated results of the unemployment rate regression. The point estimate on the $(Crisis_t \times AffectedState_i)$ variable isolating the isolating the differential effect of the bank credit shock on the unemployment rate in affected states relative to unaffected states is large in magnitude at 0.14 and statistically significant at the 90 percent confidence level. A positive point estimate on the interaction term suggests that the bank loan contraction during the crisis increased the unemployment rate in the affected states compared to the unaffected states by 1.4 percentage points, on average, during the four crisis years. A 1.4 percentage point increase in the unemployment rate corresponds to a 14.9 percent increase in the unemployment rate relative to the 9.4 percent sample average.

Overall, the macroeconomic regressions present clear evidence that the bank credit contraction resulting from the exposure of the 5 Landesbanks covering 7 federal states had an effect on the respective federal states' real economy. Employment growth (total employment, full-time, and part-time) all fell noticeably in the affected states relative to the unaffected states during the crisis, whereas the unemployment rate in the affected states rose significantly. Further, output growth contracted significantly in all states during the crisis period.

5.4.3 The Impact of the Bank Credit Shock: Privately-Held versus Publicly-Listed Companies

Section 5.4.1 showed that federal states with a Landesbank exposed to the U.S. subprime crisis saw a signifcant 20.2 percentage contraction in loan growth compared to federal states without an exposed Landesbank, and section 5.4.2 mapped the bank credit contraction into adverse real economic outcomes in output and employment for the affected federal states. This section explores a mechanism through which exogenous shocks to bank credit work their way through the economy: Shocks to bank credit differentially affect privately-held partnerships and limited liability firms compared to publicly-listed firms. A publicly-listed company has access to a wide array of methods to finance its business activity, including access to both domestic and international equity markets, cash from an international base of customers, and private bank credit from local banks, and more typically, commercial banks with an international reach. Privately-held companies, however, rely much more heavily on access to local bank credit and do not have access to the same equity markets and international commercial banks as publicly-listed companies. Therefore, when a bank shock that leads to a contraction in the supply of local credit hits the economy, publicly-listed companies can turn to equity makerts and international lenders to finance their business activities, whereas privately-held companies will be forced to reduce their business activities, thereby negtively impacting real economic outcomes. This section investigates the privately-held versus publicly-traded companies mechanism using matched employer-employee data linked with the annual establishment-level survey data from German Federal Employment Agency.

The survey results from section 5.3.2 lend credence to the central hypothesis of the paper that establishments belonging to privately-held firms are more dependent on bank loans to finance business activities than establishments belonging to publicly-listed firms and that an exogenous shock to bank capital, reducing the supply of loans, will differentially and adversely affect the net hiring and investment at establishments in privately-held firms relative to those in publicly-listed firms. This sections turns to the comprehensive, administrative, employer-employee matched dataset to directly test the hypothesis.

The following diff-in-diff empirical specification exploits both the cross-sectional and time-series variation of the bank capital shock identified in section 5.2.2 to estimate the differential effect bank exposure to the U.S. subprime has on establishment-level net hiring rates and levels of investment per employee for establishments in privately-held firms in affected states during the crisis:

$$\begin{aligned}
EstabDecision_{it} \quad = \quad & \alpha_0 + \alpha_1 Crisis_t + \alpha_2 AffectedState_i + \alpha_3 Private_{it} \\
& + \alpha_4 (Crisis_t \times AffectedState_i) \\
& + \alpha_5 (AffectedState_i \times Private_{it}) \\
& + \alpha_6 (Crisis_t \times Private_{it}) \\
& + \alpha_7 (Crisis_t \times AffectedState_i \times Private_{it}) \\
& + Y'\Omega + \xi_{it} \qquad (5.3)
\end{aligned}$$

where $EstabDecision_{it}$ is a vector including the net hiring rate and investment per employee, in 2005 euros, in establishment i and period t, $Crisis_t$ takes a value of 1 during the bank loan crisis period of 2007 from 2010 and 0 otherwise, $AffectedState_i$ takes a value of 1 if 1 of the 5 Landesbanks exposed to the U.S. subprime crisis serves as the central bank to the savings banks of that federal state and 0 otherwise, $Private_{it}$ takes a value of 1 if the esablishment belongs to a privately-held firm and 0 if the establishment belongs to a publicly-listed firm, $(Crisis_t \times AffectedState_i)$ takes a value of the 1 for the affected states during the crisis in 0 otherwise, $(AffectedState_i \times Private_{it})$ takes a value of 1 for privately-held establishments in affected states and 0 otherwise, $(Crisis_t \times Private_{it})$ takes a value of 1 for a privately-held establishment during the crisis and 0 otherwise, $(Crisis_t \times AffectedState_i \times Private_{it})$ takes a value of 1 during the crisis for establishments belonging to privately-held firms in the affected states and 0 otherwise, and Y is a vector of control variables including a dummy variable for 1999, the year of the first wave of Hartz Reforms, a dummy variable that takes a value of 1 starting in 2005 to account for the period after all Hartz reforms were implemented, a dummy variable for establishments that have a work council, a dummy for establishments that have collectively bargained wage agreements, a dummy variable for the 2001 through 2004 recession, controls for the establishment's occupational mix, the fraction of female employees, and the average age of the establishment's workforce.

The key coeffcient of interest in the estiamtion of equation (5.3) is α_7, the coefficient on the $(Crisis_t \times AffectedState_i \times Private_{it})$ interaction term. The estimated coefficient, $\hat{\alpha}_7$, quantifies the differential affect on net hiring rate and investment per employee on privately-held establishments compared to the publicly-listed in the affected states with exposed Landesbanks during the crisis.

Tables 5.12 and 5.13 show the results from the net hire rate and investment per employee regressions, respectively. Columns 1 and 2 differ based on the inclusion of the establishment-level employment measure dummies that indicate whether an establishment has instituted a policy of short-time work, reduced the offering of overtime hours, on average, for its employees, or reduced normal working hours for its employees within the given year. Column 1 does not include dummies for these establishment-level employment measures, whereas column 2 includes the employment measure dummies. However, the results are quantitatively similar regardless of whether the employment measures are included for both the net hire rate and investment per employee regressions. Therefore, the remaining analysis focuses on the results presented in column 2 of both tables.

For the net hire rate regression, the key estimated coefficient, $\hat{\alpha}_7$, on the triple interaction term, $(Crisis_t \times AffectedState_i \times Private_{it})$, is -0.239 and statisitically significant at the 99 percent confidence level, suggesting that establishments belonging to privately-held firms in affected state had a 23.9 percentage point lower hire rate during the crisis relative to establishments belonging to publicly-listed firms in affected states during the crisis. For comparison, the regression shows that net hire rates fell by roughly 3 percentage points during the 2001 through 2004 recession, a difference of nearly 8 times.

For the investment per employee regression, the estimated coefficient, $\hat{\alpha}_7$, on the triple interaction term, $(Crisis_t \times AffectedState_i \times Private_{it})$, is -6,021 and statistically significant at the 99 percent confidence level, suggesting that establishments belonging to privately-held firms in affected state had a differential reduction in investment per employee of 6,021 euros during the crisis relative to establishments belonging to publicly-listed firms in affected states during the crisis. Table 5.6 shows that the average level of investment per employee at establishments in privately-held firms is 10,829 euros over the sample, meaning that a 6,021 reduction in investment per employee corresponds to a 56 percent reduction relative to its sample average.

Overall, the establishment-level empirical results show a clear, differential, and adverse effect on net hiring at establishments belonging to privately-held firms compared to establishments belonging to publicly-listed firms in affeced states during the crisis. These results are consistent with the hypothesis that an exogenous shock to bank credit will affect more bank-dependent, privately-held firms compared to publicly-listed establishments that have access to equity markets that they can turn to during a bank loan credit crisis.

5.5 Related Literature

A large literature has argued that loan supply shocks affect real economic activity. Initial identification concerns were alleviated one study at a time (Ben S. Bernanke and Cara S. Lown, 1991; Diana Hancock and James A. Wilcox, 1992, 1997; Anil K Kashyap et al., 1993; Kashyap and Jeremy C. Stein, 1994a, b; Peek and Rosengren, 1995a, 1995b,1997, 2000, Calomiris 2003). All these studies are conducted at the level of the macroeconomy, economic region, or the firm.

The recent financial crisis has led to renewed interest in how bank shocks are transmitted through the financial system and what the real consequences are in terms of investment, employment, and economic activity. De Haas and Van Horen (2012) document that shocks from the U.S. financial crisis were transmitted internationally through inter-bank lending relationships. Aiyar (2012) documents the effect of such drying up of banks' international funding sources on the UK economy. Similarly, but in a domestic context, Ivashina and Scharfstein (2010) show that U.S. banks that had better access to deposit financing cut their lending less in response to the drying up of commercial paper markets in the U.S., indicating that the dry-up of banks' funding markets was a major determinant of their reduction in lending. In contrast, Puri et al. 2011 show the effect of a shock to bank capital (as opposed to a shock to the banks' funding liquidity) on lending behavior in Germany, using the same institutional setting as this paper. Whatever the mechanism leading to reduced bank lending, if firms were able to costlessly switch to healthier banks, such shocks would be of no consequence to real economic activity.

Chodorow-Reich (2014) shows that especially smaller firms suffer more from reduced bank lending and shed more labor as a result. The observation that small firms are especially vulnerable to loan supply shocks is reflected also in studies by Khwaja and Mian (2008) and Greenstone and Mas (2012).

There are two main distinctions of this study from the above ones. First, the unit of observation is the individual employee in this paper, whereas the finest level of granularity in existing work is the establishment. Second, this paper makes a substantial step in terms of identification. The identification in this study relies on multiple, geographically confined banking shocks, rather than a single shock implied by the 2007 financial crisis and the Lehman bankruptcy which may also have affected firms' and banks' expectations about future economic prospects. Also, the shocks employed here are entirely imported from a different economic system and thus more clearly exogenous than studies of the effects of U.S. financial system shocks on U.S. labor outcomes. These imported shocks are solely due to faulty corporate governance mechanisms in the Landesbanks and have nothing to do with local economic activity. Also, the funding of the savings banks is almost entirely local, which isolates the shock to capital from a dry-up of funding markets. Moreover, the shock used here applies only to sharply delineated geographies.[18]

More subtle distinctions from previous work on the real effect of banking shocks are that the dataset in this paper allows for the study of differential effects of banking shocks not only on small versus large firms but also but also for privately-held ver-

[18]The granularity of the banking shocks in the data and the resulting identification benefits is also a key distinction from the study by Duygan-Bump et al. (2010) and Jimenez (2014). While the identification in this study bases statements on bank supply reductions on bank balance sheet data, this paper checks the consistency of the identification with real firms' survey evidence on financial constraints. In contrast to Campello et al. (2010), the dataset used in ths paper allows for the distinction between affected and unaffected states within a country and between privately-held and publicly-listed firms. This granularity allows for a more clean differentiation of local bank loan supply shocks due to capital constraints from reductions in credit supply due to worries about future economic prospects of the economy.

sus publicly-listed firms while controlling for size. Consistent with previous results, small firms are more affected by bank shocks. This study adds, however, that a key mechanism is not necessarily size, but the legal form of the firm and the varying methods of financing business activities associated with those legal forms. Relatedly, the results speak to a literature on labor relations across different types of firms. On the one hand, Bach (2010), Bassanini et al 2012, Ellul and Pagano (2013), and Sraer and Thesmar (2007) show that family firms provide more employment and wage insurance than firms without family control[19]. The results of this paper indicate that the provision of wage and employment insurance is particularly vulnerable to funding shocks in private firms.[20]

The results are consistent with the existence of a "financial accelerator" Bernanke et al. (1996) in the sense that bank capital is an important state variable for the aggregate economy. The present paper also relates to Becker (2104) in the isolation of the effect of bank loan supply shocks to corporate policies.

5.6 Conclusion

This paper uses a comprehensive employer-employee dataset from German social security records to examine the impact of exogenous shocks to bank capital on firms' employment and investment. A narrative approach identifies 5 German re-

[19]See also Mueller and Phillipon (2006).

[20]Similar to Chava and Purnanandam (2011) and Ago (2012), the results indicate that firms that rely more on banks are more affected by larger shocks to their bank's capital than firms with access to other forms of financing. The key distinction is that the primary outcome variables of this study are labor outcome variables rather than firm value or investment, which the previous studies could not address due to data constraints. Also, the geographical heterogeneity across affected and unaffected states allows for even stronger causal claims about the correlations documented here. Compared to Cornett et al. (2011), the effect of banking shocks in Germany is propagated through the banking system because of the commitment of banks within a state to replenish each others' and the associated Landesbanks' capital, whereas the previous study focuses on banks in isolation. Second, this study link the banking shocks to real outcomes. Another subtle distinction is similar to Ashcroft (2005). This study makes clear that bank failures are not necessary to induce large economic contractions as a result of shocks to bank capital. None of the banks in the sample failed, yet the liquidity support from related banks led to a reduction in lending on their behalf.

gional Landesbanks covering 7 federal states with significant trading losses from U.S. mortgage-backed securities. The local savings banks in the affected states directly absorbed their respective Landesbanks' trading losses onto their balance sheets, causing a deep economic contraction in the banks' exclusive geographic domain. Loan growth and output growth decline by an average of 20 and 0.3 percentage points, respectively, and the unemployment rate rises by 1.4 percentage points in affected states, compared to unaffected states in each of the four crisis years. The effect is stronger for establishments belonging to privately-held, bank-dependent firms than for establishments in publicly-listed firms. Private firms in affected states reduce net hiring by 24 percentage points and cut investment by more one-half, relative to publicly listed firms.

Future extensions of this work will focus on identifying the individual-level costs placed on workers due to the shock to local bank capital. The rich, individual-level, administrative data allows for quantifying the impact of the crisis on affected workers' wages, displacement, and unemployment duration. Further, detailed job, industry, education, and geographical identifiers will allow for the study of whether affected workers moved to unaffected states to find new work, received further education or took part in job re-training programs in order to facilitate an occupational switch, or if the publicly-listed firms benefited from the adverse effect of the bank loan crisis on the privately-held firms through obtaining the privately-held's human capital.

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Figure 5.1: German Landesbanks as of 2007

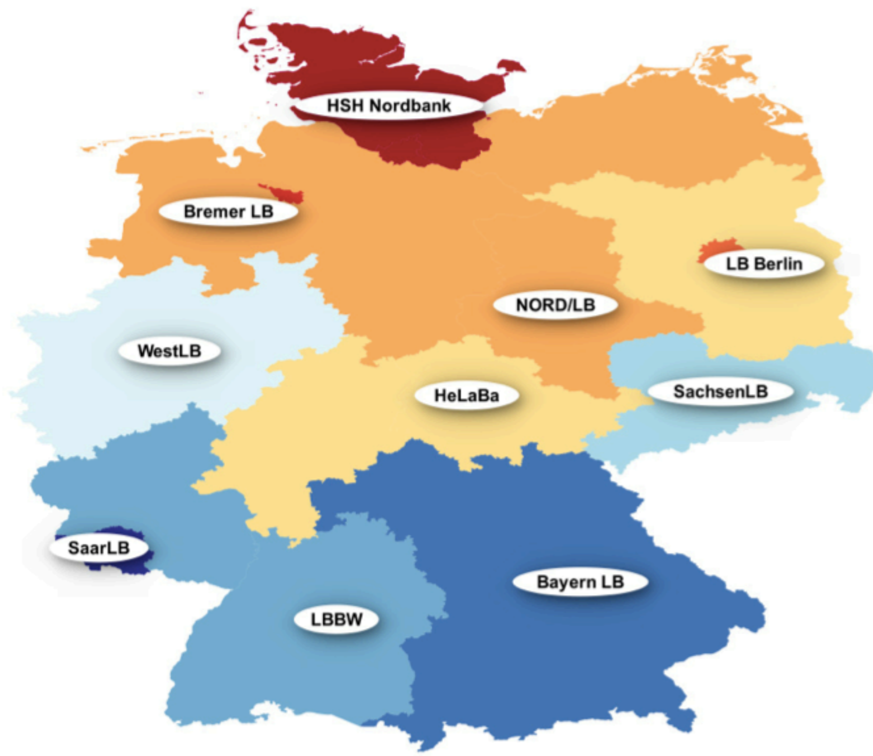

Figure 5.2: Affected States versus Unaffected States due to Landesbank Exposure

Figure 5.3: Total Real-Valued Bank Loans, Affected versus Unaffected State

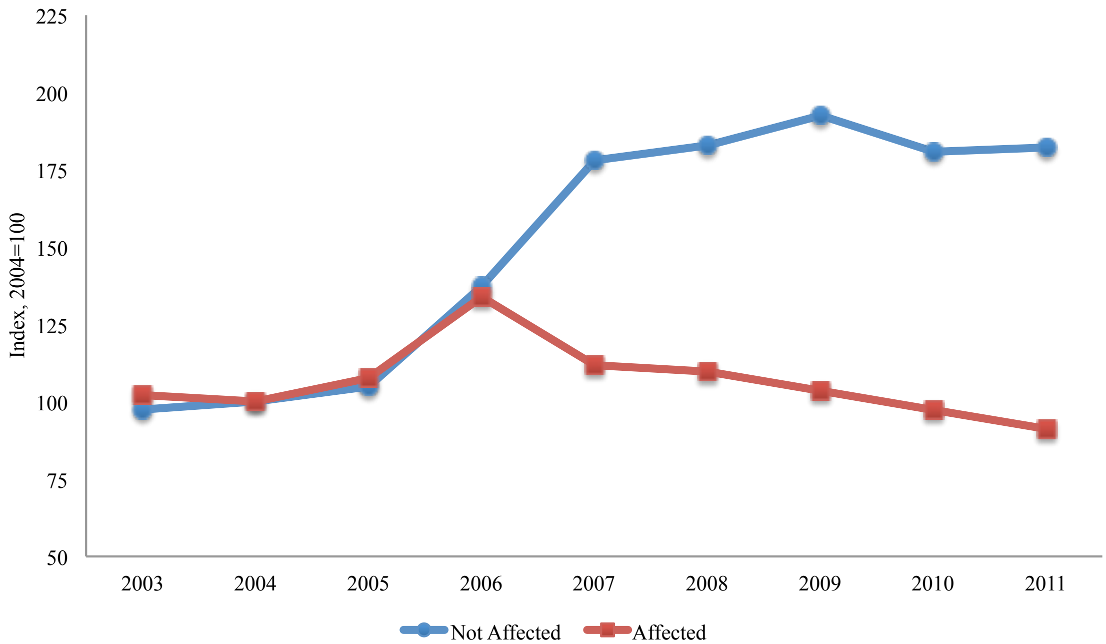

Table 5.1: German Bnak Specialization, 1997-2011

	Fraction of	
Specialization	Total Assets	Total Banks
Private Banks	53.8%	16.8%
Cooperative Banks	8.5	54.5
Public Savings Banks	37.7	28.7

All values expressed in percent terms. The fraction of total assets is calculated on an annual basis and averaged across the years in the indicated column. Total assets for each year is calculated by aggregating the assets of all German banks. The fraction of total banks is calculated on an annual basis and averaged across the years in the indicated column. Total banks for each year is calculated by counting all German banks operating in the economy. *Source:* Bankscope and authors' calculations.

Table 5.2: German Public Banks Exposed to the U.S. Subprime Crisis

Exposed Landesbank	Date of Announced First Losses	Quarter of Expected First Losses	Resulting Affected Federal States	Savings Bank Associations with Ownership Share in Exposed Landesbank
SachsenLB	August 17, 2007	2007 : 3	Saxony	Sachsen Finanzgruppe
HSH Nordbank	August 23, 2007	2007 : 3	Schleswig-Holstein Hamburg	Sparkassen- und Giroverband für Schleswig-Holstein
WestLB	August 27, 2007	2007 : 3	North Rhine-Westphalia	Rheinischer Sparkassen- und Giroverband Westfälisch-Lippischer Sparkassen- und Giroverband
BayernLB	February 13, 2008	2007 : 3	Bavaria	Sparkassenverband Bayern
LBBW	November 27, 2008	2008 : 4	Baden-Württemberg Rheinland-Palatinate Saxony	Sparkassenverband Baden-Württemberg Sparkassen- und Giroverband Rheinland-Pfalz

Table 5.3: German Federal State-Level Bank Summary Statistics

Time Period	1997-2011	2007-2010
Total Assets, Mean (Millions)	369,481 (463,428)	435,619 (542,012)
Asset Growth, Mean	1.18% (26.52)	-1.93% (23.66)
Total Loans, Mean (Millions)	156,084 (182,677)	190,994 (220,619)
Loan Growth, Mean	1.02% (25.91)	-0.91% (23.58)
N	240	64

Standard deviations in parentheses. Unit of observation is the federal state-year. Total assets and total loans are real variables expressed in 2005 euros.

Table 5.4: German Federal State-Level Bank Summary Statistics, 2007-2010

State	Not Affected	Affected
Total Assets, Mean, (Millions)	418,137 (638,121)	458,096 (395,850)
Asset Growth, Mean	4.72% (21.87)	-10.48% (23.49)
Total Loans, Mean, (Millions)	178,216 (256,415)	207,422 (166,860)
Loan Growth, Mean	4.62% (25.75)	-8.02% (18.55)
N	36	28

Standard deviations in parentheses. Unit of observation is the federal state-year. Total assets and total loans are real variables expressed in 2005 euros.

Table 5.5: Establishment Sample Summary Statistics

Key Variables	Mean	Median	Standard Deviation
Employment Level	283	55	1,257
Net Hiring Rate = (Inflows less Outflows)/Employment	-0.009	-0.002	0.270
Worker Inflow Rate	0.202	0.109	0.869
Worker Outflow Rate	0.211	0.114	0.833
Investment per Employeee	12,822	6,651	55,053
Establishment Counts	Total		
Total Establishments	15,392		
Partnership Establishments	51		
Limited Liability Establishments	13,870		
Publicly Traded/Capital Coporation Establishments	1,471		

Unit of observation for the key variables is the establishment-year. Employment flows are a fraction of the establishment's employment level on January 1 of the given year. Investment per employee is a real variable expressed in 2005 euros.

Table 5.6: Establishment Sample Summary Statistics by Legal Form

Emplyoment Level	Mean	Median	Standard Deviation
Privately Held -- Partnerships and Limited Liability	184	48	465
Publicly Traded/Capital Corporations	1,137	278	3,541
Net Hire Rate			
Privately Held -- Partnerships and Limited Liability	-0.008	0.000	0.238
Publicly Traded/Capital Coprorations	-0.010	-0.004	0.309
Investment per Employee			
Privately Held -- Partnerships and Limited Liability	10,829	5,068	33,314
Publicly Traded/Capital Corporations	15,741	9,185	76,359

Unit of observation is the establishment-year. Net hire rates are a fraction of the establishment's employment level on January 1 of the given year. Investment per employee is a real variable expressed in 2005 euros.

Table 5.7: Establishment Survey Questions and Responses

In retrospect, did the economic and financial crisis of the past two years affect your establishment/office? This questions refers to both negative and positive effects.

	2010
Full Sample	
Yes	57.12%
No	34.32
Don't Know	8.35
No Response	0.21
Affected State	
Yes	60.39%
No	31.96
Don't Know	7.65
No Response	0.00
Not Affected State	
Yes	54.87%
No	35.94
Don't Know	9.19
No Response	0.00

How strongly would you say your establishment was affected by the economic and financial crisis?

	2010
Full Sample	
Strong/Very Strong	44.63%
Moderate	36.44
Slight	18.93
Affected State & Private	
Strong/Very Strong	47.74%
Moderate	35.72
Slight	16.54
Affected State & Public	
Strong/Very Strong	44.83%
Moderate	32.76
Slight	22.41
Not Affected State & Private	
Strong/Very Strong	42.34%
Moderate	37.29
Slight	20.37
Not Affected State & Public	
Strong/ Very Strong	43.06%
Moderate	34.72
Slight	22.22

Unit of observation for each question response is the establishment in 2010.

Table 5.8: Establishemnt Survey Questions and Responses

What fraction of your overall financing was provided by the following methods?

(Note: The 2003 survey does not dilineate between a cash response and an equity response and are thud lumped together.)

	2003	2004	2007
Full Sample			
Cash Flow	78.01%	71.73%	73.98%
Equity		10.55	7.90
Bank Loans	19.22	14.45	14.98
Subsidies	2.74	3.20	3.12
Private			
Cash Flow	77.66%	72.07%	73.59%
Equity		9.70	7.79
Bank Loans	19.66	14.91	15.49
Subsidies	2.65	3.29	3.10
Public			
Cash Flow	81.30%	68.39%	78.70%
Equity		19.01	9.29
Bank Loans	15.13	9.93	8.68
Subsidies	3.57	2.40	3.33

Did you have difficulties acquiring a loan capital from private credit institutions?

	2008	2009	2010
Full Sample			
Yes	3.35 %	14.70 %	15.90 %
No	87.12	84.67	84.10
No Response	9.52	0.63	0.00
Private			
Yes	3.56%	15.41%	16.46%
No	86.95	84.02	83.54
No Response	9.49	0.57	0.00
Public			
Yes	10.79%	8.43%	8.89%
No	89.21	91.57	91.11
No Response	0.00	0.00	0.00

What fraction of your overall financing was provided by the following methods?: Unit of observation for each year is the estbalishment. The establishment provides the proportion of investments financed through cash, equity, private loans, and government subsidies, which add to 100% for each establishment, each year. The table reports the average fraction for each financing method across all establishments within a given year... *Did you have difficulties acquiring a loan capital from private credit institutions?* The unit of observation for each year is the establishment. The survery asks the establishment to qualitatively answer the question Yes/No/No response. The table reports the fraction of establishments reporting Yes/No/No response in each year.

Table 5.9: German Regional Macroeconomic Summary Statistics

Time Period	1997-2011	2001-2004	2007-2010
Output Growth Rate	0.007 (0.021)	0.002 (0.010)	0.004 (0.035)
Household Income Growth Rate	0.005 (0.011)	0.006 (0.010)	0.016 (0.016)
Employment Growth Rate, All Workers	0.007 (0.016)	-0.007 (0.012)	0.011 (0.011)
Employment Growth Rate, Full-Time	-0.004 (0.018)	-0.018 (0.013)	0.005 (0.013)
Employment Growth Rate, Part-Time	0.032 (0.030)	0.023 (0.023)	0.045 (0.014)
Unemployment Rate	0.094 (0.045)	0.106 (0.051)	0.077 (0.031)
N	240	64	64

Standard deviations in parentheses. Unit of observation is the federal state-year. Output and household income are real variables expressed in 2005 euros.

Table 5.10: Federal State Banks Asset and Loan Growth Rate Regressions

Dependent Variable	Asset Growth Rate	Loan Growth Rate
	(1)	(2)
Crisis	0.052 (0.060)	0.074 (0.059)
Affected State	0.093 (0.043)	0.080 (0.042)
Crisis x Affected State	-0.245 (0.077)	-0.202 (0.076)
2001-2004 Recession	0.035 (0.046)	0.035 (0.046)
East	0.001 (0.038)	0.010 (0.038)
R-squared	0.052	0.036
N	240	240

Standard errors in parentheses. Unit of observation is a federal state-year. Bank asset growth and bank loan growth are year-over-year growth rates of the euro-value of total real bank assets and total real bank loans, respectively, aggregated at the federal-state level, both expressed in terms of 2005 euros. Crisis takes a value of 1 during the years of the financial crisis, 2007-2010. The variable Affected State takes a value of 1 for the federal states exposed to the U.S. subprime crisis: Northrine Westfalia, Baden-Wurttemberg, Bavaria, and Saxony. The variable (Crisis x Affected State) is the interaction between the Crisis and Affected State variables and thus takes a value of 1 for the exposed states during the years of the financial crisis. The variable East takes a value of 1 if the federal state is located in East Germany.

Table 5.11: Regional Macroeconomic Regressions

Dependent Variable	Output Growth Rate	Employment Growth Rate, All Workers	Employment Growth Rate, Full-Time	Employment Growth Rate, Part-Time	Unemployment Rate
	(1)	(2)	(3)	(4)	(5)
Crisis	-0.011	0.012	0.010	0.018	-0.022
	(0.006)	(0.003)	(0.003)	(0.006)	(0.007)
Affected State	0.003	0.004	0.004	0.004	-0.025
	(0.003)	(0.002)	(0.002)	(0.004)	(0.004)
Crisis x Affected State	-0.003	-0.009	-0.008	-0.009	0.014
	(0.006)	(0.004)	(0.004)	(0.007)	(0.008)
Recession, 2001-2004	-0.009	-0.018	-0.018	-0.029	-0.006
	(0.008)	(0.005)	(0.005)	(0.009)	(0.010)
East	0.007	0.026	0.036	-0.044	-0.032
	(0.008)	(0.005)	(0.005)	(0.009)	(0.010)
R-squared	0.079	0.549	0.572	0.476	0.718
N	240	240	240	240	240

Standard errors in parentheses. Unit of observation is a federal state-year. Output and household income are real variables defined in terms of 2005 euros. Crisis takes a value of 1 during the years of the financial crisis, 2007-2010. The variable Affected State takes a value of 1 for the federal states exposed to the U.S. subprime crisis: Northrine Westfalia, Baden-Wurttemberg, Bavaria, and Saxony. The variable (Crisis x Affected State) is the interaction between the Crisis and Affected State variables and thus takes a value of 1 for the exposed states during the years of the financial crisis and zero otherwise. The variable East takes a value of 1 if the federal state is located in East Germany. All regressions include a set of labor market reform dummy variables. The labor market reform dummies consist of three variables: one variable takes a value of 1 in the year 1999 when the first wave of the Hartz reform lead to a large influx of part-time workers; a second variable takes a value of 1 after 1999 to account for a trend break due to the first Hartz reform; a third variable takes a value of 1 after 2004 to account for the final wave of the Hartz reform that changed the unemployment insurance system.

Table 5.12: Establishment-Level Net Hire Rate Regressions

Dependent Variable	Net Hire Rate (1)	Net Hire Rate (2)
Crisis	-0.170	-0.144
	(0.053)	(0.056)
Affected State	-0.143	-0.142
	(0.030)	(0.030)
Private	-0.098	-0.097
	(0.034)	(0.030)
Crisis x Affected State	0.117	0.117
	(0.009)	(0.063)
Affected State x Private	0.152	0.152
	(0.040)	(0.040)
Crisis x Private	0.215	0.204
	(0.064)	(0.066)
Crisis x Affected State x Private	-0.243	-0.239
	(0.078)	(0.079)
2001-2004 Recession	-0.030	-0.029
	(0.024)	(0.024)
Employment Measures	No	Yes
R-squared	0.021	0.021
N	56,612	56,612

Standard errors in parentheses. Unit of observation is a federal state-year. Output and household income are real variables defined in terms of 2005 euros. Crisis takes a value of 1 during the years of the financial crisis, 2007-2010. The variable Affected State takes a value of 1 for the federal states exposed to the U.S. subprime crisis: Northrine Westfalia, Baden-Wurttemberg, Bavaria, and Saxony. The variable Private takes a value of 1 when the establishment is prviately held, either as a partnership or a limited liability company. The variable (Crisis x Affected State) is the interaction between the Crisis and Affected State variables and thus takes a value of 1 for the exposed states during the years of the financial crisis. The variable (Affected State x Private) is the interaction between the Affected State and Private variables and thus takes a value of 1 for the privately held establishments in exposed states. The variable (Crisis x Private) is the interaction between the Crisis and Private variables and thus takes a value of 1 for the privately held establishments during the years of the financial crisis. The variable (Crisis x Affected State x Private) is the interaction between the Crisis, Affected State, and Private variables and thus takes a value of 1 for the privately held establishments in exposed states during the years of the financial crisis. The variable East takes a value of 1 if the federal state is located in East Germany. The labor market reform dummies consist of three variables: one variable takes a value of 1 in the year 1999 when the first wave of the Hartz reform lead to a large influx of part-time workers; a second variable takes a value of 1 after 1999 to account for a trend break due to the first Hartz reform; a third variable takes a value of 1 after 2004 to account for the final wave of the Hartz reform that changed the unemployment insurance system. The time trend variables include a linear and quadratic time trend. All regressions are weighted by the square root of the population of each federal state.

Table 5.13: Establishment-Level Investment Regressions

Dependent Variable	Investment per Employee (1)	Investment per Employee (2)
Crisis	-693	1,034
	(1,934)	(2,125)
Affected State	4,095	4,144
	(1,063)	(1,063)
Private	1,264	1,310
	(1,193)	(1,194)
Crisis x Affected State	6,719	5,953
	(2,228)	(2,313)
Affected State x Private	-3,509	-3,553
	(1,392)	(1,393)
Crisis x Private	798	-350
	(2,323)	(2,401)
Crisis x Affected State x Private	-6,931	-6,021
	(2,793)	(2,860)
2001-2004 Recession	742	805
	(834)	(835)
Employment Measures	No	Yes
R-squared	0.021	0.021
N	40,678	40,678

Standard errors in parentheses. Unit of observation is a federal state-year. Output and household income are real variables defined in terms of 2005 euros. Crisis takes a value of 1 during the years of the financial crisis, 2007-2010. The variable Affected State takes a value of 1 for the federal states exposed to the U.S. subprime crisis: Northrine Westfalia, Baden-Wurttemberg, Bavaria, and Saxony. The variable Private takes a value of 1 when the establishment is prviately held, either as a partnership or a limited liability company. The variable (Crisis x Affected State) is the interaction between the Crisis and Affected State variables and thus takes a value of 1 for the exposed states during the years of the financial crisis. The variable (Affected State x Private) is the interaction between the Affected State and Private variables and thus takes a value of 1 for the privately held establishments in exposed states. The variable (Crisis x Private) is the interaction between the Crisis and Private variables and thus takes a value of 1 for the privately held establishments during the years of the financial crisis. The variable (Crisis x Affected State x Private) is the interaction between the Crisis, Affected State, and Private variables and thus takes a value of 1 for the privately held establishments in exposed states during the years of the financial crisis. The variable East takes a value of 1 if the federal state is located in East Germany. The labor market reform dummies consist of three variables: one variable takes a value of 1 in the year 1999 when the first wave of the Hartz reform lead to a large influx of part-time workers; a second variable takes a value of 1 after 1999 to account for a trend break due to the first Hartz reform; a third variable takes a value of 1 after 2004 to account for the final wave of the Hartz reform that changed the unemployment insurance system. The time trend variables include a linear and quadratic time trend. All regressions are weighted by the square root of the population of each federal state.