Wage Rigidity and Employment Outcomes: Evidence from Administrative Data *

Gabriel Ehrlich
University of Michigan

Joshua Montes
Federal Reserve Board

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Abstract

This paper examines the relationship between downward nominal wage rigidity and employment outcomes using linked employer-employee data. Wage rigidity prevents 27.1 percent of counterfactual wage cuts, with a standard deviation of 19.2 percent across establishments. An establishment with the sample-average level of wage rigidity is predicted to have a 4.2 percentage point higher layoff rate, a 6.4 percentage point lower quit rate, and a 2.0 percentage point lower hire rate. Estimating a structural model by indirect inference implies that the cost of a nominal wage cut is 30 percent of an average worker's annual compensation.

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> 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?
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> *Letter of David Ricardo to Thomas Malthus, 1821*[1]

1 Introduction

A perennial debate in economics concerns the extent to which difficulty reducing nominal wages affects employment outcomes. This paper uses administrative, linked employer-employee data 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. Establishments with more rigid wages exhibit higher layoff rates and lower quit and hire rates, consistent with the predictions of a theoretical model of establishment decision-making in the face of downward nominal wage rigidity (simply "wage rigidity" hereafter).

The estimates suggest that wage rigidity prevents 27.1 percent of wage cuts at the average establishment, with a standard deviation of 19.2 percent across establishments. Establishments in the construction supersector display the least wage rigidity, with an average of 8.8 percent of wage cuts prevented. Establishments in the public administration and finance supersectors display the most wage rigidity, with average levels of 39.3 and 41.7 percent of wage cuts prevented, respectively. The data are particularly well-suited for the task of estimating establishment-level wage rigidity because they contain total compensation histories for every worker at each of the sampled establishments. Those compensation histories are taken from administrative data and should be free of measurement error.

The paper introduces a measure of wage rigidity that builds on the kernel density approach popular in the literature (e.g., Card and Hyslop, 1997; Knoppik et al., 2007) and is suitable for establishment-level analysis. There are three major advantages to the proposed estimator. First, it uses both 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, the kernel density estimation does not impose a parametric form on the wage change distribution. Third, the estimator performs well regardless of whether the median wage change is

[1] Reprinted in Ricardo (1887).

above or below zero, a situation that arises with non-trivial frequency at the establishment-year level.

The paper also 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 4.2 percentage point higher layoff rate, a 6.4 percentage point lower quit rate, and a 2.0 percentage point lower hire rate than an establishment with no wage rigidity.[2]

Wage rigidity is potentially measured with error and is also potentially endogenous with respect to layoff, quit, and hire rates; thus, the paper uses an instrumental variables strategy to isolate plausibly exogenous variation in wage rigidity at the establishment level. The instrument exploits the extent and stringency of regional sector-level wage floors, and it is strongly predictive of measured wage rigidity.[3] During the paper's sample period, collective bargaining in Germany by and large occurred at the region-sector level, not at individual establishments; only 4.1 percent of establishments covering 8.4 percent of employees had establishment-level collective bargaining agreements in this time. Furthermore, strong institutional norms provided incentives for all establishments within a region-sector to respect those wage floors, even establishments that did not have a formal collective bargaining agreement. Collective bargaining typically focused on wages rather than on employment levels. Therefore, the wage floors should not have affected employment outcomes at the establishment level directly except through their effects on wages.

The instrument is uncorrelated with past and future measures of state-sector level revenue growth and with establishment-level expectations of future employment growth. Conversely, the instrument shows strong path dependence in the sense that its lagged values strongly predict its current values. Therefore, historical and institutional norms appear to be the key determinant of whether establishments faced binding collective bargaining agreements. Thus, the data support that the instrument satisfies the exclusion restriction and suggest that the estimated relationship between wage rigidity and employment outcomes is causal.

The individual-level wage data used in this paper is a measure of total compensation that includes base salary, bonuses, and other forms of compensation, which is a significant advantage relative to much of the previous literature. Due to data limitations, many previous studies focus on wage rigidity in base pay only. However, establishments may circumvent wage rigidity in base pay by altering bonuses and other forms of compensation. Thus, a complete examination of the

[2]For comparison, the sample average layoff rate is 6.8 percent, the sample average quit rate is 11.1 percent, and the sample average hire rate is 22.3 percent.

[3]A supplementary instrumental variables strategy uses lagged participation in collective bargaining agreements to instrument for current participation, leveraging the high persistence and path depedendence in German collective bargaining. The results are similar across all instrumental variables specifications.

relationship between wage rigidity and employment outcomes should include a measure of total compensation, as this paper does.

Estimating the structural model via an indirect inference procedure allows for consistent estimates of the underlying parameters even if the reduced form wage rigidity estimator is misspecified. The estimates suggest that nominal wage cuts cost 9,000 euros on average, or thirty percent of an average worker's annual compensation. The average level of wage rigidity is estimated to increase the layoff rate by 1.7 percentage points, reduce the quit rate by 3.8 percentage points, and reduce the hire rate by 3.1 percentage points within the context of the model. Wage rigidity endogenously generates roughly one-quarter of all layoffs. The effects in the estimated model are similar to the effects estimated in the empirical analysis.

Several previous studies have documented the existence of wage rigidity in microeconomic datasets. Prominent examples using U.S. survey data include Card and Hyslop (1997), Kahn (1997), Lebow et al. (2003), and Daly and Hobijn (2014). Examples using German administrative data include Bauer, Bonin, Goette, and Sunde (2007) and Knoppik and Beissinger (2009). Dickens, Goette, Groshen, Holden, Messina, Schweitzer, Turunen, and Ward (2007) use international data from the United States and Europe.

Using U.S. administrative data from the payroll processor ADP, Grigsby, Hurst, and Yildirmaz (2021) estimate that cuts to base pay among job stayers are much less common than is implied by survey data. They also show that considering additional forms of compensation such as bonuses and fringe benefits significantly increases the flexibility of total pay. There are two main differences between this paper and Grigsby, Hurst, and Yildirmaz (2021). First, they distinguish between base and non-base pay and show that base pay is much more downwardly rigid than non-base pay. The data set used in this paper contains only measures of total compensation. The two papers suggest similar amounts of downward rigidity in terms of compensation (base pay plus bonuses and overtime). Second, this paper additionally links establishment-level variation in wage rigidity to employment flows.

Despite the bulk of this literature pointing to the existence of substantial wage rigidity, it has been difficult 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." 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, Saks, and Wilson (2003) 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".

There are some exceptions to this pattern. Kaur (2019) finds strong causal effects of wage rigidity on employment levels in informal agricultural labor markets in India. De Ridder and Pfajfar (2017) find that contractionary monetary policy and tax shocks increase unemployment and decrease economic activity by more in states with more rigid wages. Schoefer (2015) shows that wage rigidity can interact with financial frictions at the firm level to generate cyclical fluctuations in hiring rates. Makridis and Gittleman (2019) document substantial heterogeneity between performance pay and fixed wage jobs in U.S. data and show that employment is more cyclically sensitive in fixed wage jobs. Finally, Kurmann and McEntarfer (2019) use administrative data from the U.S. Longitudinal Employer-Household Dynamics (LEHD) program to link firm variation in wage rigidity to employment flows. Similarly to this paper, they find that firms that exhibit more wage rigidity display significantly lower hiring rates. The LEHD data does not allow Kurmann and McEntarfer (2019) to distinguish between layoffs and quits. The theoretical model in this paper suggests that wage rigidity should have opposite effects on the two forms of job destruction, complicating any comparison between wage rigidity's effects on total separations and wage rigidity's effects on layoffs and quits separately.

A number of papers focusing on related issues in the context of the German economy are significant to note. Dustmann and Schönberg (2009) show that union firms in Germany display fewer wage cuts than non-union firms. Further, they find that workers completing their apprenticeship training have higher layoff rates and lower quit rates in union firms compared to non-union firms, which they attribute to higher wage rigidity in unionized firms. Dustmann, Schönberg, and Stuhler (2016) show that the heterogeneous responses between old and young workers to a migration shock in Germany are consistent with differential levels of wage rigidity across those workers, while Gathmann, Helm, and Schönberg (2018) show that the regional wage and employment responses to mass layoff events in Germany are also consistent with downward wage rigidity. Merkl and Stüber (2017) estimate substantial heterogeneity in wage cyclicality across German establishments, consistent with the heterogeneous levels of wage rigidity estimated here. They build a model showing that the large fraction of firms with countercyclical wages substantially amplifies shocks to the German labor market.

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. Stüber and

Beissinger (2012) show that this wage compression effect is present in West Germany over the period 1975–2007 in times of low inflation such as this paper's study period.

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.

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 measured wage rigidity across establishments in the sample. Section 5 describes the paper's empirical approach to estimating the empirical relationships between wage rigidity and layoffs, quits, and hires with ordinary least squares and instrumental variables regressions. Section 6 presents the estimated empirical relationships. Section 7 uses those results to estimate the theoretical model by indirect inference. Section 8 concludes.

2 Model of Establishment Decision Making with Wage Rigidity

This section examines the dynamic wage and employment policies of a single establishment in partial equilibrium with heterogeneous worker types facing an imperfectly competitive labor market.[4] The establishment is a monopsonist in the labor market that can unilaterally set wages subject to an asymmetric wage-adjustment cost function and an upward-sloping labor supply curve.[5] Its 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.[6]

[4]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.

[5]For a generalized model of monopsony in the labor market, see Manning (2006).

[6]The establishment's production function features decreasing returns to scale in labor, implicitly an assumption that the capital stock is fixed for the effective duration of its wage and employment decisions. This assumption is necessary for computational tractability. Based on average job separation rates, the expected duration of a job match is 5.6 years.

Appendix A presents an analytical model that delivers the same key predictions regarding wage rigity's relationship with employment flows as the numerical model in this section.

2.A 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, which are given as:

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

where n_j is the stock of type j labor used in production; α governs returns to scale; w_j is the real 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; c_ℓ is the cost per layoff; and a_j is a stochastic process that shifts the marginal revenue product of labor. a_j is the product of an establishment-wide productivity level z and a type j productivity level u_j.[7] The marginal revenue product of each worker type does not depend on the other types, an assumption that simplifies the analysis.[8] However, the marginal revenue products of labor may still be correlated across types by means of the establishment-wide productivity level z. All workers of the same type must be paid the same wage; in particular, new hires must be paid the same wage as incumbent workers of the same type. This "equal treatment" assumption features in several models of wage stickiness, including prominently Gertler and Trigari (2009) and Snell and Thomas (2010). Stüber (2017) and Snell et al. (2018) present empirical evidence that the real wages of new hires and incumbent workers in Germany have approximately equal cyclicalities. Snell et al. (2018) show additionally that real wages behave differently in upswings and downswings; they interpret the evidence as showing that "equal treatment and asymmetric adjustment are a key feature of German wages."

Downward nominal wage rigidity enters the model through the wage adjustment cost function, $g\left(w_j, w_{j,-1}\right)$. Although w_j represents the real wage, the wage adjustment cost function is specified in terms of the change in nominal wages, calculated as $(1+\pi)w_j - w_{j,-1}$, where π is the rate of price inflation.[9] $g\left(w_j, w_{j,-1}\right)$ is specified in per-employee terms as a polynomial in nominal wage

[7] 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.

[8] Adding further correlation among the worker-type productivity levels u_j has little effect on the results.

[9] Denoting the price level in period t as p_t, the nominal wage in period t is then $p_t w_t$ and the nominal wage in period $t-1$ is $p_{t-1}w_{t-1}$. Thus, a worker's nominal wage change is $p_t w_t - p_{t-1}w_{t-1}$, and the worker experiences a nominal wage cut if and only if $p_t w_t < p_{t-1}w_{t-1} \iff \frac{p_t}{p_{t-1}}w_t < w_{t-1} \iff (1+\pi)w_t < w_{t-1}$.

reductions:

$$g\left(w_j, w_{j,-1}\right) \;=\; \lambda_0 \mathbb{1}_{(1+\pi)w_j < w_{j,-1}} + \lambda_1 \left(\frac{w_{j,-1} - (1+\pi)w_j}{w_j}\right) \mathbb{1}_{(1+\pi)w_j < w_{j,-1}} \tag{2}$$

λ_0 represents a fixed menu cost of cutting wages, while λ_1 represents a cost that scales linearly with the size of the nominal wage cut. 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. The function $g(\cdot)$ 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. Multiple sources of wage rigidity have been proposed in the literature, and potential sources remain a topic of discussion. Bewley (1999) emphasizes that wage cuts may reduce morale, thereby lowering worker productivity. Similarly, Elsby (2009) and Kaur (2019) both model wage rigidity as arising from reductions in morale associated with wage cuts. However, the model here focuses on the consequences of wage rigidity rather than its sources.[10]

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

$$\delta\left(w_j\right) \;=\; \bar{\delta}\left(\frac{w_j}{\overline{w}}\right)^{-\gamma}, \quad \gamma > 0 \tag{3}$$

where $\bar{\delta}$ is a parameter that scales the 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 establishment's market power over its incumbent workers vanishes.

[10]The empirical portion of the paper in Sections 5 and 6 takes a less agnostic approach and utilizes another possible source of wage rigidity, the structure of collective bargaining agreements, as a plausibly exogenous source of variation in establishment-level wage rigidity in this context.

The establishment faces a cost per hire c_h as in Manning (2006):

$$c_h(h_j, n_{j,-1}, w_j) \;=\; \bar{c} w_j^{-\phi} h_j^{\upsilon} n_{j,-1}^{-\eta} \tag{4}$$

This functional form allows for either economies or diseconomies of scale in the hiring cost function. Most studies of hiring costs indicate that they are subject to decreasing returns to scale, for instance Shapiro (1986), Manning (2006), Blatter, Muehlemann, and Schenker (2012) and Muehlemann and Pfeifer (2016). Further, the hiring cost function depends on the wage rate w_j: intuitively, an establishment that pays higher wages should find it easier to hire workers.

The layoff cost c_l is a per-worker cost. In Germany, the bulk of this cost is likely to represent direct severance payments to laid-off workers (Hunt, 2000). The German system of employment protection provides limited explicit legal guidance on such severance payments, but they are common in practice.[11] German policies regarding severance pay and job security generally rank in the middle of OECD countries in terms of stringency (Addison and Teixeira, 2005).

2.B 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:

$$\begin{aligned} V_j\left(z, u_j, w_{j,-1}, n_{j,-1}\right) \;=\; & \max_{w_j, h_j, l_j} a_j n_j^{\alpha} - w_j n_j - c_h(h_j, n_{j,-1}, w_j) h_j - c_\ell \ell_j \\ & - g(w_j, w_{j,-1}) n_j + \beta E\left[V_j\left(z', u_j', w_j, n_j\right)\right] \end{aligned} \tag{5}$$

subject to

$$\ln a_j \;=\; \ln z + \ln u_j \tag{6}$$

$$\ln z \;=\; (1-\psi_z)\ln \bar{z} + \psi_z \ln z_{-1} + \varepsilon_z,\; \varepsilon_z \sim N\left(0, \sigma_z^2\right) \tag{7}$$

$$\ln u_j \;=\; (1-\psi_u)\ln \bar{u} + \psi_u \ln u_{j,-1} + \varepsilon_{u_j},\; \varepsilon_{u_j} \sim N\big(0, \sigma_{u_j}^2\big) \tag{8}$$

$$n_j \;=\; \left(1-\delta\left(w_j\right)\right) n_{j,-1} + h_j - \ell_j \tag{9}$$

$$h_j, \ell_j \;\geq\; 0 \tag{10}$$

[11] Grund (2006) argues that, "There is enormous uncertainty for both about the bonuses or respectively the costs of dismissals at least if severance payments have not been fixed in an ex ante bargaining. Due to the unspecified legal situation it seems to be beneficial for employees to insist on severance payments and threaten with a suit in order to increase the chance of a substantial severance payment. Taking into account this behavior, employers are frequently better off with fixing a (generous) amount of severance payment as an explicit part of the employment contract."

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 7 and 8, both log productivity levels evolve according to a mean reverting, $AR(1)$ process. The errors ε_z and ε_{u_j} are assumed to be independent.

Computational details of the model solution and estimation are provided in Appendix E.

2.C Establishment Policy Functions and Simulations

Figures 1 and 2 display the establishment's policy functions for a single worker type with high and low productivity levels, respectively.[12] The blue policy functions illustrate the case with rigid wages, while the yellow policy functions illustrate the case with flexible wages, that is, with the wage cut cost parameters λ_0 and λ_1 set to zero. Panel A in both figures illustrates the establishment's wage policy functions, while panels B, C, and D illustrate the quit, layoff, and employment level policies, respectively.

Note that the establishment will never find it optimal both to endogenously layoff and hire workers of the same type within a period. However, the establishment as a whole may exhibit positive hire and layoff rates within a given period because it employs multiple worker types, each with its own productivity process. Thus, to give a sense of the effect of wage rigidity on the establishment's hire and layoff policies for a single worker type, it is helpful to illustrate the policies under different productivity levels.

The wage policy functions estimated with costly wage cuts in panel A of both figures display the distinctive patterns associated with wage rigidity. When the previous period's wage is relatively low, wage rigidity does not bind, and the establishment increases wages to the optimal level.[13] As the previous period's wage increases, wage rigidity begins to bind, which is shown in the upward-sloping portions of the wage policy functions. For a high enough level of the previous period's wage, however, it becomes worthwhile for the establishment to pay the menu cost λ_0 of cutting nominal wages, at which point the wage policy function jumps downward in Figure 2. Finally, both figures illustrate that the establishment generally pays a higher wage when the previous period's employment level is lower, reflecting higher wages' role in helping to recruit workers in equation (4).

Panels B of Figures 1 and 2 illustrate the establishment's quit rate policy, which the establishment controls deterministically by setting the wage rate. The quit rate varies inversely with the wage rate as reflected in the quit rate function in equation (3). The optimal quit rate is generally

[12] The policy functions are calculated using the estimated parameters described in Section 7.

[13] That optimal level can still be affected by wage rigidity, as seen in the spread between the rigid and flexible policy functions in Figure 1, where for low levels of the previous wage, the new target wage under wage rigidity is less than the target wage without wage rigidity. This pattern is an illustration of the wage compression effect described in Elsby (2009).

lower when the previous period's employment level was higher. The quit rate declines over the regions where wage rigidity binds, as higher wages induce fewer quits. The quit rate jumps up in Figure 2 at the point that the establishment chooses to cut nominal wages.

Panels C of Figures 1 and 2 display the establishment's layoff policy functions under the two different productivity levels. The layoff rate is bounded below by the exogenous layoff rate s_x. In Figure 1, the productivity level is high enough that the establishment does not choose to lay off any workers endogenously. When productivity is low, as in Figure 2, the establishment generally desires to shrink its workforce. With rigid wages, it then faces a trade-off between either paying the layoff cost c_ℓ or paying the cost g of cutting nominal wages, thereby inducing a higher quit rate. In the figure, this trade-off leads to increasing layoffs as the previous wage rises, up to the point that the establishment institutes a large nominal wage cut.[14] It is worth highlighting that in the absence of wage rigidity, the establishment almost never desires to endogenously lay off any workers.[15] Without wage rigidity, it is always more profitable to lower wages to induce quits than to lay off workers when employment is greater than desired.[16]

The employment level policy functions in panel D of Figures 1 and 2 generally increase with the previous period's employment levels. When wage are flexible, the previous period's wage has no effect on the new employment level. When wages are rigid, the previous period's wage has two effects on the new employment level. First, binding wage rigidity lowers optimal employment levels by making each worker more costly to employ. Second, binding wage rigidity lowers quit rates and reduces the cost of hiring, which increases the new optimal employment level. Quantitatively, however, the first effect substantially outweighs the second effect in the simulated model.

Figure 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 expected given a symmetrical wage change distribution. Figure 4 displays a simulated wage change histogram in the case of rigid wages.[17] The distribution of wage

[14] At that point, the establishment would like to pay an even lower wage than shown in Figure 2, but it is constrained by the minimum level of the wage grid described in Appendix E. In simulations using the baseline parameters described in Section 7, the minimum level of the wage grid is binding less than one percent of the time.

[15] The exception arises in cases in which the minimum wage on the wage grid is higher than the optimal level, which is very rare.

[16] This result is consistent with some seminal papers in the efficient turnover literature. McLaughlin (1991) notes that, "Many models of layoffs do not limit the firm's ability to lower wages; if the firm exercised this power, all separations would be quits induced by lowering the wage." That feature, which is common in many search and matching models of the labor market, such as Burdett (1978), does not apply to the model in this paper when the cost of cutting nominal wages is positive. In the presence of wage rigidity, the cost of cutting nominal wages will sometimes induce the establishment to separate from some workers who would have remained on the job at a lower nominal wage rate. Topel (1982) shows that temporary layoffs can be an efficient response to fluctuations in market demand in industries in which it is difficult to use inventories. Approximately one-sixth of layoffs in the data are "temporary" in the sense that the laid-off worker later returns to the same employer without an intervening spell of employment elsewhere. However, fewer than one in three of these temporary layoffs lasts less than six months.

[17] The λ parameters are again set to their estimated values from Section 7.

changes is noticeably compressed relative to the case of flexible wages and clearly asymmetrical. The portion corresponding to wage cuts appears to be "hollowed out" relative to the portion corresponding to wage increases.

Figure 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 by proportionally upscaling λ_0 and λ_1. Panel A shows the estimated fraction of wage cuts prevented due to wage rigidity associated with a given level of λ_0 and λ_1; estimated wage rigidity increases with the cost of imposing nominal wage cuts in the model.[18] Panel B shows the average layoff rate, which increases with wage rigidity. When there is no wage rigidity, the establishment can reduce the size of its workforce entirely by lowering wages and inducing more quits. The more expensive it is to cut nominal wages, the less ability the establishment will have to induce quits through lowering the nominal wage, and the more affordable paying the layoff cost will appear relative to paying the costs of cutting wages. Panels C and D illustrate the average quit 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 also reduces the establishment's need to hire new workers. Additionally, forward-looking establishments 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.[19]

Therefore, the simulations presented in Figure 5 naturally provide three testable predictions of the relationship between wage rigidity and employment outcomes: establishments with more measured wage rigidity should exhibit higher layoff rates, lower quit rates, and lower hire rates. Section 6 tests these three empirical predictions. It is important to stress that the model indicates that wage rigidity should have opposite effects on layoff and quit rates. Estimating the effects of wage rigidity on the total job separation rate or the job destruction rate will conflate these opposite effects, making it difficult to discern a relationship between wage rigidity and employment outflows.

[18]Specifically, the panel shows the estimated level of wage rigidity using the estimator described in Section 4.A, which is the same estimator applied to the actual data in Section 6.

[19]Because the model is stationary, it is theoretically possible that the hire rate will increase with wage rigidity if the increase in the layoff rate is larger than the decrease in the quit rate. However, this situation does not arise when realistic parameter values are used in the model.

3 Data Description

3.A Overview of Dataset

This study uses the longitudinal model of the Linked-Employer-Employee Data (LIAB) (Version LM 2, Years 1993–2007) from the IAB. Data access was provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB) and subsequently remote data access.[20] Jacobebbinghaus (2008) and Fischer et al. (2009) document the data. The LIAB includes 5,293 West German establishments that participated in the annual IAB Establishment Panel employer 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.[21]

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 agencies of the German social security system 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 the findings of Lebow, Saks, and Wilson (2003) and Grigsby, Hurst, and Yildirmaz (2021) 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 such as layoffs, quits, and hires may be imputed with minimal assumptions.

[20]The IAB provides an outline of the data set here.

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

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 employer 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 works council or wage bargaining agreement.

The dataset does not contain employee-level data on hours worked, so in principle, a reduction in hours could appear as a wage cut using the daily average wage rate reported in the data. However, the data suggest that hours worked are not a quantitatively meaningful margin of adjustment during the study period. The IAB Establishment Panel employer survey asks establishments for average weekly hours worked by full-time employees in five of the seven years in the study period. Average weekly hours worked rarely change within establishments across years. Indeed, 80 percent of all establishments report a constant level of hours in every year of the sample. Additionally, 88 percent of establishment-year observations show no year-over-year change in average weekly hours, and for the 12 percent of establishment-years that do show a year-over-year change in hours, the average absolute change is only 1.6 hours. Unconditionally, the standard deviation of weekly hours within an establishment across years averages only 0.2 hours, suggesting minimal changes in hours at most.[22] The consistency of average weekly hours likely reflects that it is costly for establishments to adjust hours given the required negotiations with works councils. Thus, it is unlikely that changes in hours constitute a major source of measurement error in estimated wage rigidity.[23]

To further minimize any potential for changes in hours worked to affect measured wage rigidity, the measured wage change distributions include only workers whose hours status does not change between periods. While the employee-level data do not provide employee-specific hours worked, the dataset does distinguish between part-time workers working less than half of full-time, those working more than half of full-time, and full-time workers. Wage changes for workers whose hours status changes relative to the previous year are discarded when measuring wage rigidity.

Reported compensation in the dataset 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

[22]This statistic is calculated in two steps. The first step calculates the standard deviation of hours across years within each establishment. The second step takes the average of the standard deviations calculated in the first step.

[23]The IAB Establishment Panel employer survey also asks questions related to whether workers are working overtime hours. Roughly three-quarters of establishments report either having no overtime in any year or having some overtime in every year. The average weekly hour calculations reported above include overtime hours. Therefore, fluctuations in overtime hours within establishments are quantitatively unimportant for the purpose of measuring wage rigidity.

not for the purpose of calculating employment flows.[24]

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

Finally, this study also uses the weakly anonymous Sample of Integrated Labour Market Biographies (SIAB, weakly anonymous version 7508, Years 1975–2008).[25] Dorner et al. (2010) and Dorner et al. (2010) document the data. 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 IAB Establishment 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 3.D.

3.B Defining Key Concepts

Layoff, quit, and hire rates are measured as fractions of the establishment's total workforce as of December 31st of the preceding year. Because the model predicts that wage rigidity should have opposite effects on layoffs and quits, it is essential to identify the two accurately in the data.[26] Following a convention for distinguishing involuntary layoffs and voluntary quits in the worker biographies similar to that of Blien and Rudolph (1989), Haas (2000), and Dustmann and Schönberg (2009), a layoff is defined as an interruption between employment spells that results in the employee flowing into unemployment before the beginning of another employment spell, as indicated by receipt of unemployment assistance during the intervening period. 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.[27] In the data,

[24] 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.

[25] Data access was again provided via on-site use at the Research Data Centre (FDZ) of the German Federal Employment Agency (BA) at the Institute for Employment Research (IAB) and subsequently remote data Access. The IAB provides an outline of the data set here.

[26] To the extent that layoffs are mistaken for quits or vice versa, the empirical results in Section 6 will attenuate the relationships between wage rigidity and layoffs, and quits.

[27] A fourth possibility for employment adjustment is that of a "spin", which can take the form of either an inflow

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 occurrences 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).[28] This situation arises in less than one percent of separations.

Establishment revenue and value added are measured from the IAB Establishment Panel employer survey questions. The theoretical model abstracts from intermediate inputs, which empirically can account for a large proportion of revenues. An establishment's value added is calculated as total revenues minus intermediate inputs and external costs.[29]

3.C Sample Selection and 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 fewer than 30 employees that year or fewer than 5 continuing employment relationships for which a valid wage change can be calculated; the establishment is excluded altogether if it averages fewer than 20 such continuing employment relationships over the years it is in the sample. These requirements effectively exclude very small establishments from the analysis. Thus, the full sample empirical results discussed in Section 6.A–6.C apply only to establishments with at least 30 employees. Additionally, the analysis requires data on establishment revenues in both the current and previous years in order to calculate the

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.

[28]The establishment-level analysis 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.

[29]Each year, the IAB Establishment Panel employer survey includes a question regarding total turnover and a question regarding the share of revenue attributable to external costs. For instance, in the 2002 survey the question regarding turnover read: "What was your turnover in the last fiscal year (normally the year 2001)?" The question regarding intermediate inputs and external costs 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 trade and commerce and professional associations)?

establishment's change in revenue. These restrictions leave 2,628 establishments for the analysis.

Table 1 shows the descriptive statistics for the layoff, quit, and hire rates for the sample of establishments from 1997 through 2003. The average annual layoff rate over the period is 6.8 percent with a standard deviation of 11.5 percent across establishment-years. The average annual quit rate over the period is 11.1 percent with a standard deviation of 15.2 percent. The average annual hire rate is 22.3 percent with a standard deviation of 33.6 percent. The average establishment employs 466 workers, versus 168 workers for the median establishment. These relatively large establishment sizes are the result of the sample restrictions regarding establishment size that exclude very small establishments. The average nominal wage is 87.1 euros per day, with a standard deviation of 28.5 euros per day. The average wage expressed in year 2000 euros was 86.7 euros per day, with a standard deviation of 28.2 euros per day.[30]

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).[31] The average establishment-year revenue growth in the sample is 4.3 percent with a standard deviation of 25.4 percent.

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

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

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

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. Previous studies using German worker-level administrative data from the IAB, such as Beissinger and Knoppik (2001), Knoppik and Beissinger (2003), and Stüber and Beissinger (2012), provide similar evidence for the existence of aggregate wage rigidity in Germany. Studies that provide evidence of wage rigidity in Germany and other countries include Dickens et al. (2006), Dickens et al. (2007), Holden and Wulfsberg (2008), and Knoppik and Beissinger (2009). The paper now turns to measuring the degree and extent of wage rigidity across German establishments.

4 Estimating Wage Rigidity

4.A 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 7 and 8 illustrate the approach.[32]

For each establishment i and year t, estimate the distribution of observed wage changes using kernel density estimation.[33] The estimate of the density at a point x is

$$\hat{f}_{it}(x) = \frac{1}{n} \sum_{j=1}^{n} \frac{1}{h_j} K\left(\frac{x - x_j}{h_j}\right) \tag{11}$$

where n is the number of observations, x_j for $j \in \{1, ..., n\}$ denotes a point in the observed distribution, h_j is an adaptive bandwidth following the procedure of Van Kerm et al. (2003), and

[32] Appendix D.1 presents results from an alternative method of measuring wage rigidity based on an approach that uses more parametric assumptions than the approach in the main body of the text.

[33] The estimation procedure focuses on wage changes within 15 percentage points of the median wage change each year to avoid the influence of outliers.

K is a kernel function.[34] Using adaptive bandwidths is helpful in the context of establishment-level density estimates, where the data can be sparse in some regions of the distribution. The specific kernel function used in the estimation is an Epanechnikov kernel of the form

$$K(z) = \begin{cases} \frac{3}{4}\left(1 - z^2\right) & \text{if } |z| < 1 \\ 0 & \text{otherwise.} \end{cases} \tag{12}$$

Denote the estimated distribution of observed wage changes as $\hat{f}_{it}^{obs}$, and let m_{it} represent the median wage change at establishment i from time $t-1$ to time t expressed in percentage points.

Next, construct a counterfactual wage change distribution $\hat{f}_i^{cf}$ for establishment i by averaging the upper tails of the estimated observed distributions $\hat{f}_{it}^{obs}$ across each year. In constructing the average, first normalize the observed distribution for each year around its median.[35] Then, reflect the averaged distribution of the upper tails around the median each year as illustrated in Figure 7.

The estimated proportion of wage cuts prevented by wage rigidity is then calculated by comparing the implied proportion of counterfactual wage cuts to the number observed. For establishment i and year t, denote the proportion of wage cuts in the estimated observed wage change distribution as $\hat{F}_{it}^{obs}(0^-)$, illustrated as the lightly shaded areas in Figure 8.[36] Denote the proportion of wage cuts in the estimated counterfactual distribution as $\hat{F}_{it}^{cf}(0^-)$, illustrated as the darkly shaded areas in Figure 8. Let the sum across years of these proportions be denoted $\hat{F}_i^{obs}(0^-)$ and $\hat{F}_i^{cf}(0^-)$. The measure of establishment-level wage rigidity is then the proportion of counterfactual wage cuts that are "missing" from the data and is calculated as

$$\widehat{wr}_i = 1 - \frac{\hat{F}_i^{obs}(0^-)}{\hat{F}_i^{cf}(0^-)}. \tag{13}$$

Therefore, the wage rigidity estimate in equation (13) is time-invariant for each establishment.[37] $\widehat{wr}_i$ 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.[38]

[34]The global bandwidth is set to be 0.005. The adaptive bandwidths are calculated as the product of the global bandwidth and a local bandwidth factor that is proportional to the square root of the underlying density function at the sample points. The adaptive bandwidths have the property that their geometric average equals the global bandwidth.

[35]In practice, in situations in which the observed median is negative and there are more observed wage cuts than wage increases, recalculating the median by excluding observed wage changes between -0.25% and 0.25% helps to correct for the "sweep-up" of counterfactual wage cuts to zero. This adjustment improves the accuracy of the procedure in the Monte Carlo simulations discussed in Appendix B. Those years are then excluded when averaging the upper tails, but are included when calculating the counterfactual wage cuts prevented by wage rigidity.

[36]The notation 0^- indicates that the measured proportion does not include wage changes of exactly zero.

[37]Appendix D.1 presents an alternative parametric estimator that varies over time within establishments.

[38]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

This approach to estimating wage rigidity has three main advantages in an establishment-level context. 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, the kernel density estimation does not impose a parametric form on the wage change 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, p. 86) 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 C 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.

This approach 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, as well. Monte Carlo simulations of the estimator presented here suggest that it performs well in practice given the estimated level of wage compression in the data.

The Monte Carlo simulations suggest that there is some sampling error associated with the estimator. This sampling error will lead to attenuation bias in the ordinary least squares estimates of the relationships between wage rigidity and employment outcomes presented in Section 6, which is one motivation for the instrumental variable approach also presented in that section. Please see Appendix B for a discussion of the Monte Carlo simulations.

25 percent more wage cuts in the data than would be predicted by the upper tail of the wage change distribution.

4.B The Distribution of Wage Rigidity in West Germany

Table 2 shows the mean, median, and standard deviation of the distribution of wage rigidity estimates for individual establishments within the sample. The average establishment-level measure of wage rigidity is 27.1 percent, implying that wage rigidity prevents 27.1 percent of counterfactual wage cuts at the average establishment. The standard deviation of the estimates across establishments is 19.2 percent and the median estimate is 24.6 percent. Thus, there is both a notable degree of estimated wage rigidity among establishments and significant variation across establishments.

Table 2 also shows the mean, median, and standard deviation 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 9 percent to 22 percent. Among supersectors, finance and public administration exhibit the highest degree of wage rigidity, with an average of 41.7 and 39.3 percent of wage cuts prevented by wage rigidity across establishments in those supersectors, respectively. Construction exhibits the smallest degree of average wage rigidity, with 8.8 percent of nominal wage cuts prevented.

5 Empirical Approach to Measuring Wage Rigidity's Effects on Employment Adjustment

This section presents the paper's empirical approach to measuring wage rigidity's effects on employment outcomes. Section 5.A outlines a simple ordinary least squares (OLS) approach to testing the hypotheses regarding the predictions generated by the model in Section 2 for wage rigidity's effects on layoff, quit, and hire rates. It also describes why the OLS approach might lead to inconsistent estimates and proposes a complementary instrumental variables strategy that leverages institutional features of German collective bargaining. Section 5.B describes the institutional background of collective bargaining in West Germany during the analysis period to provide intuition for why the proposed instrumental variables are likely to be valid. Section 5.C provides evidence suggesting that the instrumental variables are likely to be excludable, and Section 5.D provides evidence that the instrumental variables are relevant for predicting wage rigidity.

5.A Empirical Approach

The predictions from the theoretical model in Section 2.C imply empirical regressions of the form:

$$y_{it} = \beta_0 + \beta_1 wr_i + X_{it}'\Upsilon + \epsilon_{it} \tag{14}$$

where the unit of observation is an establishment-year. y_{it} represents an employment flow 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 4.A. X_{it} represents a vector of control variables, including a dummy for the presence of a works council, the median year-over-year percentage wage change, a set of year and state fixed effects, dummies for establishment size groups, the fraction of the workforce that is female, controls for workforce educational attainment and occupation, and indicators for large-scale relocations of workers across establishments within the same firm. Alternative specifications include sector fixed effects and controls for establishment-level revenue growth specified as a linear spline function with a kink at zero, permitting disparate associations between revenue growth and employment adjustment depending on whether revenue growth is positive or negative.

One identifying assumption necessary for OLS estimation of equation 14 to provide consistent estimates of wage rigidity's effects on employment outcomes is that the random error ϵ_{it} is uncorrelated with wr_i conditional on the other covariates. That assumption may be violated if an omitted variable causes both wage rigidity and employment outcomes at the establishment level. For instance, it is possible that implicit norms influence the prevalence of wage rigidity across establishments (Bewley, 1999) and also influence employment outcomes such as layoff, quit, and hire rates.

Another identifying assumption necessary for OLS estimation to be consistent is that wage rigidity is measured without error. As Appendix B shows, however, the measure of wage rigidity described in Section 4 is likely to have some noise. To the extent that wage rigidity is measured with error, OLS estimates of equation 14 will provide estimates of wage rigidity's effects on employment outcomes that are attenuated toward zero.

In both cases, a valid instrumental variables approach can identify the causal effect of wage rigidity on employment outcomes. To be valid, an instrumental variable must predict wage rigidity (instrument relevance) but may not affect employment outcomes except through its effect on wage rigidity (the exclusion restriction).[39] To address the potential confounding effect of implicit norms described above, such an instrument must predict wage rigidity through a mechanism unrelated to such norms.

German wage-setting institutions provide intuitively appealing instrumental variables for wage

[39]Formally, a valid instrument z_{it} must be uncorrelated with the second-stage regression error ϵ_{it}.

rigidity that are likely to satisfy the exclusion restriction for individual establishments and that are measurable in the IAB Establishment Panel employer survey data set. Section 5.B describes the institutional background in detail. The primary instrumental variable used in the analysis is the proportion of all workers at establishments in each sector-state that pay at the collectively-bargained wage floor. Each year, the IAB Establishment Panel employer survey asks whether the establishment is bound by a sector-wide wage agreement, and if so, whether the establishment pays wages at or above the collectively agreed upon sector-wide wage floor. Thus, it is straightforward to construct a measure of the proportion of employees that are at establishments paying at the wage floor within each sector-state as a whole.[40] Intuitively, the instrument should be positively correlated with measured wage rigidity.

Table 3 shows means and standard deviations of this "proportion at the wage floor" instrument, both for the entire sample and at the supersector level. For the entire sample, 37.6 percent of workers are in establishments that pay at the collectively bargained wage floor. As with measured wage rigidity, the instrument displays considerable variation across supersectors. For example, only 34.0 percent of workers in the construction supersector work at establishments that pay at the wage floor; conversely, in the public administration supersector, 80.3 percent of workers work at establishments that pay at the wage floor. Thus, collectively bargained wage floors appear to be more binding in the public administration supersector than in the construction supersector.[41]

As described in Section 5.B, the institutional features of collective bargaining in Germany indicate that an important source of exogenous variation in collectively bargained wages occurs at the state-sector level. Thus, the primary instrument for wage rigidity, the "proportion at the wage floor" instrument, varies at the state-sector level as well.

It is also possible to construct a secondary instrument for wage rigidity that varies at the establishment-year level. This secondary instrument leverages the strong persistence of establishment participation in collective bargaining by using one-year lagged participation in a collective bargaining agreement to predict current-period wage rigidity. This secondary instrument provides a finer level of variation than the primary instrument. Section 5.B provides some information on the strong persistence of collective bargaining in the data.

For these instruments to satisfy the exclusion restriction, they must affect employment out-

[40]The analysis uses two-digit sector codes taken from the IAB Establishment Panel employer survey and harmonized across years by the authors. Both measures include all establishments that responded, not only those that are in the LIAB sample in the main analysis. Using all of the establishment responses increases the sample size substantially and thus provides a more complete picture of the impact that collective bargaining has on wage setting within each state-region.

[41]It is worth noting that not all workers at establishments with collective bargaining agreements are covered by those agreements. Further, some workers at establishments that report paying at the collectively bargained wage floor may receive wages above the floor. The survey responses reflect the predominant policy at each establishment but do not apply to every worker within those establishments.

comes only through their influence on wage rigidity, conditional on the other covariates. In particular, the instruments may not affect employment outcomes directly. The institutional background of collective bargaining in Germany provides a good deal of theoretical support for the proposed instruments' validity. Section 5.C provides data supporting the argument that the instruments are excludable, and Section 5.D provides evidence showing that the instruments are relevant.

5.B Institutional Background of Collective Bargaining in Germany

Historically, collective bargaining in Germany has been characterized by a "dual system" under which wages were bargained at the sector-region level between employer associations and worker unions, while employment levels were determined at the establishment level, often via bargaining with the establishment's works council.[42] As Hübler and Jirjahn (2003, p. 471) summarize:

> The German system of industrial relations is characterised by a dual structure of employee representation through works councils and unions. Works councils provide a highly developed mechanism for establishment-level participation while collective bargaining agreements are negotiated between unions and employers' associations on an industrial level.

Ellguth, Gerner, and Stegmaier (2014, pp. 96–97) provide additional detail:[43]

> German employment relations are characterized by a distinct dual system. First, working conditions (especially working hours) and wages are typically determined by industrywide regional CBAs that are negotiated between unions and employer associations... Second, working conditions are also negotiated at the establishment level. In addition to company agreements or individual contracts, works councils are the crucial mechanism for employer-employee negotiations at the establishment level in Germany... works councils are dedicated mainly to production issues (e.g. working hours or overtime) and personnel affairs. They usually have minimal influence on distribution issues (e.g. wages or payment schemes) because the latter are typically regulated by industry-wide agreements in Germany.

Addison et al. (2017, p. 195) note further that,

[42] Sisson et al. (2012) summarizes the motivations and elements of European collective bargaining arrangements during this period.

[43] For more background on the German system of collective bargaining, see for instance Bauer, Bonin, Goette, and Sunde (2007). As documented by Dustmann, Fitzenberger, Schönberg, and Spitz-Oener (2014), German labor market institutions and collective bargaining have evolved considerably from the mid-1990s to the present day. The changes were an ongoing process during the period of this study, 1997 to 2003.

> German legislation formally prohibits establishment-level agreements between works councils and management that bypass industry-wide contracts. That is, works councils cannot conclude plant agreements (*Betriebsvereinbarungen*) on issues covered by collective bargaining unless expressly authorized to do so by the relevant sectoral agreement. For this reason it is conventional to describe their function as integrative (focusing on issues related to the size of the pie rather than its distribution).

Even establishments without a formal collective wage agreement typically followed the region-sector wage floors during thr analysis period. Fitzenberger, Kohn, and Lembcke (2008, p. 3) note that in Germany,

> Wages set at the firm level as well as individually bargained wages are adapted towards collective bargaining agreements, be it in order to reduce transaction costs or not to create incentives for employees to join a union. Collective agreements can also be declared generally binding by the Minister for Labor and Social Affaires. In light of the traditionally high coverage, researchers in the past ... often assumed that collective bargaining agreements apply to all employees.[44]

Germany's historical dual system of industry-sector wage bargaining and works council-level negotiations over employment and other items therefore suggests that the exclusion restriction is reasonable in this context. To control for the possibility of establishment-specific negotiations, all of the employment outcome regressions in the sections below include a dummy for whether an establishment has a works council.

The institutional features of wage setting in Germany suggest that exogenous variation in collectively bargained wages occurs at the state-sector level, so it is necessary to include an instrument that varies at that level. The secondary instrument, which uses establishments' past participation in a collective bargaining agreement to predict current-period wage rigidity, also finds support in the instutional features of German collective bargaining. An unreported regression of an establishment's current-year participation in a wage agreement on lagged participation suggests that an establishment that participated in an agreement the previous year has a 96% probability of participating in the current year, implying strong path dependence of participation in collective bargaining. Thus, this "one-year lagged participation" instrument is highly predictive of contemporaneous participation but is less likely to be correlated with contemporaneous economic conditions. This secondary instrument provides a finer level of variation than the primary instrument, and it is also likely to be excludable given that it is pre-determined relative to contemporaneous economic conditions and employment decisions.

[44] For further discussion, see Fitzenberger, Kohn, and Lembcke (2013).

Appendix D.3 displays additional results using an establishment's participation in a collective bargaining agreement in years prior to the beginning of the main analysis sample in 1997 as an instrument. The results are qualitatively similar to those using one-year lagged participation, although the sample size is smaller because there are fewer survey responses in the earlier years.

5.C Instrument Excludability

Although the exclusion restriction is not formally testable, it is possible to examine whether the primary instrument is correlated with measures of state-sector-level business conditions and establishments' expectations of future business conditions. The IAB Establishment Panel employer survey allows for the measurement of total revenue at the state-sector-year level.[45] The survey also asks establishments about their expectations for future employment levels.

If binding collective bargaining agreements are systematically related to business conditions, then observable measures such as revenue or employment growth should be correlated with the instrumental variables. However, the data do not support that case. Table 4 shows regressions of the primary instrument's values on its own lagged values as well as current, lagged, and future state-sector revenue growth and establishment expectations regarding future employment. Column 1 shows that, on its own, the instrument's lagged value is a very strong predictor of its current value, with an R^2 of above 0.8 in a regression that also includes state, sector, and year fixed effects. Column 2 shows that contemporaneous revenue growth does not add any statistically or economically predictive power to this regression, while columns 3 and 4 show that neither do lagged or future revenue growth. Past, present, and future business conditions are all uncorrelated with the instrument.

Alternatively, it may be expectations of future business conditions, rather than realized conditions, that affect whether collective bargaining agreements bind. Although the IAB Establishment Panel employer survey does not regularly include questions about establishments' expectations of future revenue, it does regularly ask about expectations of future employment levels. Column 5 shows the results of a regression that adds to the lagged instrument value the proportion of establishments in a state-sector-year that expect their employment levels to increase, the proportion that expect their employment levels to decrease, and the proportion that are unsure about the direction of their employment levels.[46] Establishments' employment expectations are also not predictive of the instrument. Columns 6 through 8 show that including both revenue movements and employment expectations does not change these results.

[45]This measure is constructed as the sum of all establishments' reported revenue in the IAB Establishment Panel employer survey within a given state-sector-year.

[46]The ommitted category in this regression is the proportion of establishments that expect their employment levels to remain "approximately constant."

Therefore, neither realized revenues nor expected future employment levels are correlated with the primary instrument's value, as one would expect if the instrument were systematically related to business conditions. In fact, the only meaningful predictor of the instrument's current value is its lagged value, suggesting strong path dependence in the value of the instrumental variable that is unrelated to economic conditions. These results strongly support the notion that the instrument provides exogenous variation in wage rigidity across establishments.

A second concern regarding the primary instrument's validity is that establishments may choose to discontinue their participation in collective bargaining in response to a changing institutional setting or establishment-level business conditions. However, the data suggest that participation in collective bargaining is strongly persistent at the establishment level regardless of business conditions. As noted in Section 5.A, a regression of an establishment's current-year participation in a wage agreement on its lagged participation suggests that an establishment that participated in an agreement the previous year has a 96% probability of participating in the current year, consistent with the results in Table 4. Furthermore, Dustmann, Fitzenberger, Schönberg, and Spitz-Oener (2014, p. 178) note that, "After opting out of a collective agreement, firms still have to pay wages for the incumbent employees according to the collective agreement until a new agreement at the firm level has been reached..." Therefore, the effects of an industry-wide collective bargaining agreement may persist for the small number of establishments that choose to opt out.

Another potential complication regarding the instruments is that establishments may choose to forego the industry-wide collective bargaining agreement and enter a firm- or establishment-specific agreement. However, the IAB Establishment Panel employer survey asks separately about the two types of agreements. To avoid including those establishment-level agreements, only establishments with industry-level agreements are used to construct the instruments. Therefore, the instruments should capture whether establishments are in industries that are heavily affected by collective bargaining agreements.

These arguments also suggest that the secondary instrument, lagged participation in a collective bargaining agreement, is likely to be excludable. First, because the instrument is lagged, it is predetermined relative to contemporaneous shocks and decisions regarding employment levels. Second, the strong establishment-level path dependence of participation in collective bargaining suggests that such participation is not responsive to short-term fluctuations in business conditions. Appendix D.3 extends the idea of using a lagged instrument further by using establishment participation in collective bargaining agreements prior to 1997, the beginning of the main analysis period. The results are similar to those using one-year lagged participation, although they are limited by a smaller sample.

5.D Instrument Relevance and First-Stage Results

Figure 9 plots measured wage rigidity on the vertical axis against the "proportion at the wage floor" instrument on the horizontal axis. Each point on the scatterplot represents the average across establishments in a sector.[47] There is a visually apparent positive correlation between measured wage rigidity and the proportion at the wage floor instrument, suggesting the instrument's relevance for wage rigidity.

Table 5 shows the first-stage results from the two-stage least squares regressions used to analyze wage rigidity's effects on employment flows. Column (1) shows the first-stage regression using the primary instrument and including all baseline covariates. Consistent with the relationship in Figure 9, the proportion at the wage floor instrument predicts more rigid wages. The relationship is highly statistically significant. Column (2) shows the first-stage regression adding the spline in revenue growth as a covariate. As in column (1), the instrument has the expected, statistically significant, relationship with wage rigidity. The instrument appears to be strong in both columns, as evidenced by large first-stage F-statistics. Column (3) shows the first-stage regression including both the primary and secondary instruments. Both are statistically signifcant predictors of measured wage rigidity in this specification. Note, however, that the sample size in this specification is almost 20 percent smaller than in columns (1) and (2), because some establishments do not have recorded responses to the relevant questions in the IAB Establishment Panel employer survey. Finally, column (4) displays first-stage regression results using the primary instrument and restrcting the sample to large establishments. The results are broadly consistent across specifications.

6 Empirical Evidence of Wage Rigidity's Effects on Employment Adjustment

This section describes the results of the empirical approach described in Section 5. Sections 6.A, 6.B, and 6.C document the empirical relationships between wage rigidity and layoff, quit, and hire rates, respectively. Those relationships are consistent with the model predictions in Section 2: more rigid wages lead to higher layoff rates and lower quit and hire rates.

Tables 6, 7, and 8 display results from regressions corresponding to equation 14, which are described in Sections 6.A, 6.B, and 6.C below. The tables follow a parallel format. Columns (1) through (4) show the results of the OLS regressions, while columns (5) through (8) show the results of the IV regressions. Column (1) presents the OLS estimates using the baseline covariates, while column (2) adds a spline in establishment revenue. Column (3) adds sector fixed effects

[47] Due to disclosure requirements, only sectors with 20 or more establishments in the analysis are included in the figure. However, sectors with fewer than 20 establishments are still included in the empirical analysis.

to the covariates in column (1), while column (4) restricts the sample in column (1) by including only establishment-years with employment levels above the sample median of 168 (that is, larger establishments). Columns (5) and (6) display results from instrumental variables regressions corresponding to columns (1) and (2) but instrumenting for wage rigidity using the proportion paying at the wage floor instrument. Column (7) adds the lagged participation in a collective bargaining agreement instrument to the specification in column (5). Column (8) displays IV results correpond-ing to the specification in column (4) using the proportion paying at the wage floor instrument. All standard errors are clustered at the establishment level.

Note again that the sample excludes very small establishments, as described in Section 3.C, so the results should be interpreted as applying to establishments with at least 30 employees. Appendix D.2 splits the sample into establishments with employment levels smaller and larger than the median of 168 employees. Tables D.4 through D.7 display results analogous to Tables 6 through 8 split by sample size.

6.A Layoffs

Table 6 shows the results of regressions using the layoff rate as the dependent variable. The estimated coefficients on wage rigidity in the OLS regressions are generally economically small and are not statistically distinguishable from zero. The estimated coefficients on wage rigidity in columns (1) and (2) are nearly zero. The coefficient in column (3) is similar to those in columns (1) and (2), suggesting that isolating within-sector variation does not change the results from the OLS regressions.[48] One possible reason for this result may be that the establishment-level coavriates in columns (1) and (2) already account for much of the cross-sectoral variation in employment flows without the inclusion of sector fixed effects. The coefficient in column (4) is positive, consistent with the predictions of the quantitative model. Indeed, the coefficient estimate of 3.2 percent implies that an establishment with the sample average level of measured wage rigidity for large establishments would have a layoff rate nearly one percentage point higher than an establishment with no measured wage rigidity. The standard error of the estimate is substantially larger than in columns (1) and (2), consistent with a sample size that is half as large. To the extent that wage rigidity is measured more accurately in establishments with larger employment levels, the estimate in column (4) suggests that measurement error in wage rigidity at smaller establishments may be biasing the estimates in columns (1) and (2) downward.

The estimated coefficients on wage rigidity in the IV regressions in columns (5) through (8) show that wage rigidity has economically large, and statistically significant, effects on layoffs. The

[48]The R^2 of the regression in column (3) is higher than in columns (1) and (2), suggesting that the sector fixed effects capture cross-sector variation in layoff rates. A similar pattern holds in the results for quits and hires in Tables 7 and 8.

coefficient of 15.4 percent in column (5) implies that an establishment with sample average measured wage rigidity of 27.1 percent would be expected to have a layoff rate that is 4.2 percentage points higher than an establishment with no measured wage rigidity. That increase corresponds to three-fifths of the 6.8 percent sample average layoff rate. The point estimate in column (6), which includes a spline in establishment revenue growth, is slightly smaller than the point estimate in column (5), as is the point estimate in column (7), which uses the lagged collective bargaining instrument and features smaller sample sizes. Finally, the point estimate in column (8), which uses only large establishment-years, is a bit larger than in column (5), suggesting that the use of an instrumental variable reduces the potential problem of measurement error seen in the OLS results in columns (1) through (4). The Wooldridge (1995) score test of regressor exogeneity in columns (5) through (8) rejects the null hypothesis that wage rigidity is exogenous with respect to layoffs at the 5-percent confidence level in all specifications, suggesting that the IV estimates are preferable to the OLS estimates.

Figure 10 illustrates the relationships between average measured wage rigidity and layoff rates at the state-sector-year level, as disclosure requirements prevent plotting establishment-level outcomes. Both measures have been residualized using all covariates, including state, sector, and year fixed effects. Figure 10 shows a positive correlation between wage rigidity and layoff rates, although given the different levels of aggregation necessitated by the disclosure requirements, it does not capture the exact establishment-level relationships that the regressions in Table 6 do. Figure 11 illustrates the relationship between measured wage rigidity, instrumented with the proportion paying at the wage floor instrument, and layoff rates, again at the state-sector-year level. All measures have been residualized using the additional covariates specified in column (6) of Table 6 to remove the influence of other observables. Figure 11 shows a clear positive correlation between instrumented wage rigidity and layoff rates, consistent with the theoretical model's predictions in Section 2.C.

Overall, the results in this section suggest that higher wage rigidity causes an economically large and statistically significant increase in establishment-level layoff rates.

6.B Quits

The regression results in Table 7 show a significant negative relationship between wage rigidity and quits across both the OLS and IV specifications, consistent with the model's predictions. The specifications in each column parallel the pattern in Table 6 for layoffs. In column (1), the estimated coefficient on the wage rigidity term is -7.0 percent and is statistically significant. That estimate implies that an establishment with sample-average measured wage rigidity would be predicted to have a quit rate 1.9 percentage points lower than an establishment with no measured wage rigidity, or roughly 17 percent of the sample average quit rate of 11.1 percent. The results in column (2) are similar. In column (3), which includes sector fixed effects, the coefficient is again similar, suggesting that within-sector variation in wage rigidity rather than cross-sector variation drives the result. In column (4), which uses large establishments, the estimated coefficient is nearly twice as large in magnitude as in columns (1) and (2), again consistent with the idea that there is likely less measurement error in wage rigidity for larger establishments.

In column (5), the estimated coefficient on the wage rigidity term is -23.5 percent, substantially larger in magnitude than the OLS estimate in column (1). The IV estimate implies that an establishment with sample-average measured wage rigidity would be predicted to have a quit rate 6.4 percentage points lower than an establishment with no measured wage rigidity, or roughly 58 percent of the sample average quit rate. The estimated coefficients on wage rigidity in columns (6) through (8) are similar. The exogeneity tests all strongly reject the null hypothesis that wage rigidity is exogenous in the quit rate regressions at the 1-percent confidence level. Again, the IV estimates are very consistent between the full sample and the large establishment sample, suggesting that the use of the IV strategy corrects for measurement error in addition to potential endogeneity.

Figure 12 illustrates the relationship between average measured wage rigidity and quit rates at the state-sector-year level, while Figure 13 illustrates the relationship between instrumented wage rigidity and quit rates.[49] Both Figure 12 and Figure 13 show a clear negative correlation between wage rigidity and quit rates.

Thus, the results in this section provide strong evidence that wage rigidity reduces quit rates, consistent with the predictions of the theoretical model. Those results are robust across specifications.

6.C Hires

Table 8 shows a consistently negative relationship between wage rigidity and establishment hire rates; the pattern of specifications across columns again parallels that in Table 6. In column (1), the estimated coefficient on the wage rigidity term is -7.3 percent and is statistically significant. That estimate implies that an establishment with sample-average measured wage rigidity would be predicted to have a hire rate 2.0 percentage points lower than an establishment with no measured wage rigidity, or roughly 8.9 percent of the sample average hire rate of 22.3 percent. The results in column (2) are again similar. In column (3), which includes sector fixed effects, the coefficient is slightly larger in absolute value, again suggesting that within-sector variation in wage rigidity rather than cross-sector variation drives the result. The sample of large establishments in column

[49]The measures have again been aggregated to the state-sector-year level. In Figure 12, both measures have been residualized using all covariates, including state, sector, and year fixed effects. In Figure 13, the measures have been residualized using the additional covariates specificed in column (6) of Table 7.

(4) yields a similar point estimate to the one in column (3), but the smaller sample leads to less precise estimates.

The IV results in columns (5) through (7) show larger point estimates for the coefficients on the wage rigidity term than the OLS estimates in columns (1) through (4), but with substantially larger standard errors as well. The point estimate in column (8) is smaller than in columns (1) through (4) but also has a large standard error. Therefore, the estimated coefficients in the IV hires regressions are not statistically different from zero. Consistent with the larger standard errors in the IV estimates, the regressor exogeneity tests in columns (4) through (8) do not reject the null hypothesis that wage rigidity is exogenous, suggesting that the more efficient OLS estimates are to be preferred for this set of regressions.

Figure 14 illustrates the relationship between average measured wage rigidity and hire rates at the state-sector-year level, while Figure 15 illustrates the relationship between instrumented wage rigidity and hire rates.[50] Both Figure 14 and Figure 15 show a clear negative correlation between wage rigidity and hire rates.

These results are consistent with the model's prediction that establishments with more wage rigidity should exhibit lower hire rates: forward-looking establishments realize that if they hire workers in good times, they may have to pay costs associated with rigid wages in response to future negative shocks.

7 Model Results

7.A Model Estimation

This paper employs an indirect inference approach (e.g. Gourieroux et al. 1993, Smith 1993, Gallant and Tauchen 2010) to estimate 17 out of the 20 parameters in the theoretical model described in Section 2, which are listed in Table 10. Of the other three parameters, the deterministic price inflation rate, π, is the average consumer price inflation rate in Germany in the period 1997-2003 as reported in the World Bank's World Development Indicators. Average worker-type productivity $\overline{u}$ is normalized to 1, and the average daily wage $\overline{w}$ is normalized to approximately 105 year-2000 euros.[51] Table 9 displays the 19 empirical moments used as targets in the indirect inference procedure along with the simulated moments from the model.

Model data is generated by computing the establishment's optimal policy functions for a given guess of parameters and simulating a series of wage change distributions and employment out-

[50]The measures have again been aggregated to the state-sector-year level. In Figure 14, both measures have been residualized using all covariates, including state, sector, and year fixed effects. In Figure 15, the measures have been residualized using the additional covariates specificed in column (6) of Table 8.

[51]$\overline{w}$ is a normalization because it cannot be estimated separately from $\overline{\delta}$ in equation (3).

comes under a set of random shocks.[52] Simulated moments are taken from the model data and compared to the empirical moments. The model period is taken to be one year. With 19 target moments from the data and 17 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.

Ten target moments are estimates from auxiliary models: the wage rigidity estimator described in Section 4; the coefficients on measured wage rigidity in the layoff, quit, and hire regressions; the wage elasticity of the quit rate; the wage elasticity of the hire rate, empirical economies of scale of the hire rate, and the root mean square error of the hire rate elasticity regression; and the persistence coefficient and the root mean square error from an $AR(1)$ regression of establishment value added per worker. Additionally, several descriptive statistics from the empirical data are used as target moments. The target moments are described in detail in Appendix F.

Table 9 shows the empirical and simulated moments using the estimated parameters. The model generally matches the target moments well. In particular, the simulated wage change percentiles match their empirical targets almost exactly, suggesting that the estimated model captures the asymmetry in the wage change distribution. The estimated model also closely matches the average layoff rate, the empirical elasticities of the quit and hire rates with respect to wages, the economies of scale of the hire rate, and root mean square error of the hire rate elasticity regression. Likewise, the model closely matches the moments related primarily to productivity: the real average daily wage rate, real average value added per worker, and the persistence and root mean square error of the value added per worker *AR(1)* regression.

Next, the model does reasonably well in matching the average measured level of wage rigidity and the coefficients on estimated wage rigidity in the simulated layoff and quit regressions. The magnitudes of the layoff and quit regression coefficients are smaller in the simulated data than in the actual data, whereas the average level of measured wage rigidity is higher. However, the model does match the coefficient on estimated wage rigidity in the hire rate regression. Finally, the estimated labor's share of value added is a little lower than in the data.

Table 10 shows the estimated parameter values. The persistences and standard deviations of the establishment-wide and worker-type productivity shocks imply that shocks to the individual worker types are smaller but more persistent than shocks to establishment-wide productivity. The unconditional variance of worker-type productivity is about twice as large as the unconditional variance of establishment-level productivity.

The estimates of the hiring cost parameters, $\overline{c}$, ϕ, υ, and η, are reasonable. Applying the estimated values to the hiring cost function in equation (4) yields an average per-employee hiring

[52] See Appendix E for details of this procedure.

cost of 11 weeks of total compensation.[53] Manning (2006) shows that there are diseconomies of scale in hiring if $\upsilon - \eta > 0$. The estimates suggest that value is 0.3, implying diseconomies of scale in hiring. That result is consistent with the studies cited in Section 2.A. The wage elasticity of the hiring cost, ϕ, is estimated to be 1.2, suggesting that per-worker recruitment costs are slightly more than unit elastic with respect to the wage rate. The estimated cost of laying off a worker, c_ℓ, is 1,055 euros, or nearly two weeks' pay.[54]

The estimated wage elasticity of the quit rate, γ, is 1.6. Manning (2006, p. 98) reports that, "... studies of the sensitivity of quits to wages rarely find an elasticity above 1." Using German data, however, Hirsch, Schank, and Schnabel (2010) estimate elasticities of between 1.5 and 1.9 for men and between 1.0 and 1.3 for women. Regardless, the estimated $\hat{\gamma}$ implies a high degree of establishment monopsony power in the labor market.

The estimate of the establishment discount rate, β, is 0.552. That rate implies that the establishment discounts the future by substantially more than would be implied by observed interest rates.[55] A lower rate of discounting in the model (i.e., a higher β), however, would lead to an unrealistically compressed wage change distribution, as the establishment would moderate wage increases in good times to avoid the costs of cutting wages in bad times. Note that even with the high rate of discounting in the estimated model, wage rigidity produces a clearly compressed wage change distribution, consistent with Elsby (2009) and as seen in Figures 3 and 4.

The estimates of primary interest are for the wage cut cost parameters, λ_0 and λ_1. The estimated values imply that a nominal wage cut costs 8,999 year 2000 euros on average, or over 15 weeks of average worker compensation. Roughly 86 percent of the average cost stems from the fixed cost λ_0. The large costs of cutting wages lead to endogenous layoffs in the model of 1.7% per year on average. That rate is roughly one-quarter of total layoffs in the model, with the exogenous layoff rate, s_x, of 5.2% per year accounting for the remainder.[56]

An advantage of using the indirect inference procedure to estimate the model is that it will identify the structural parameters even if the reduced form estimates are misspecified. In particular, the estimates of the structural parameters will be consistent even if the wage rigidity estimation procedure presented in Section 4 is misspecified. In that case, model simulations can be used to estimate the effects of wage rigidity on layoffs, quits, and hires.

[53] For comparison, Muehlemann and Pfeifer's (2013) estimate more than 8 weeks of wages for skilled German workers.

[54] The estimated layoff cost is lower than redundancy costs for Germany reported in the World Bank's Doing Business project, which averaged 4,695 euros for workers with 1 year and 5 years of tenure.

[55] For instance, the World Bank's World Development Indicators suggest a nominal bank rate of 9.5 percent per year for private sector lending in the period 1997-2002, or roughly 8.2 percent per year in real terms.

[56] Michaillat (2012) estimates in U.S. data that 2.1 percentage points out of a total unemployment rate of 5.8 percent are due to rationing unemployment, or slightly more than one-third.

7.B Model Explorations

A simple counterfactual exercise of simulating the model with the wage rigidity parameters λ_0 and λ_1 set to zero implies that moving from zero measured wage rigidity to the estimated measured level of 35 percent increases the layoff rate by 1.3 percentage points, reduces the quit rate by 3.8 percentage points, and reduces the hire rate by 3.1 percentage points.

The estimated model sheds further light on the relative importance of the menu cost of cutting wages, λ_0, versus the linear cost, λ_1 for employment outcomes. Table 11 examines each parameter's contribution to the simulated layoff, quit, and hire rates at different points in the productivity distribution. The columns show the simulated employment flows with λ_0 and λ_1 set either to their baseline values or to zero in each of four possible combinations. Each panel shows results at various levels of the establishment's productivity distribution. The menu cost λ_0 always accounts for a greater share of wage rigidity's effects on employment outcomes than the linear cost λ_1. For layoffs, the menu cost generates roughly half of wage rigidity's effect on its own, while the linear cost generates only a small portion. There is complementarity in the two costs' effects on layoffs, however, so that the two costs together generate a larger increase in layoffs than the sum of their individual effects. At high productivity levels, the establishment does not desire to cut wages, so neither cost contributes to an increase in layoffs. For quits and hires, the menu cost is again more important than the linear cost, although the linear cost accounts for roughly one-fifth of wage rigidity's effects on its own. One interesting pattern is that wage rigidity affects quit and hire rates even at high productivity levels, because the establishment anticipates that wage rigidity may bind in the future even if the establishment does not desire to cut wages today.

Table 12 explores whether wage rigidity is a necessary feature for the model to match the moments in the data. The table compares the data to the model moments simulated with and without wage rigidity for various levels of the establishment's monopsony power. Each entry in the table shows the ratio of the sum of squared errors in a model simulation without wage rigidity (i.e., with λ_0 and λ_1 set to zero) to the sum of squared errors in a model simulation with wage rigidity (i.e., with λ_0 and λ_1 set to their baseline values). A table entry greater than one implies that the simulated data without wage rigidity produced a worse fit to the target moments than the simulated data with wage rigidity.

Table 12 considers alternative values for the parameters γ (the wage elasticity of the quit rate) and ϕ (the wage elasticity of the hiring rate). The elasticities γ and ϕ are the key parameters that govern the shape of the effective labor supply curve facing the establishment in monopsony models and thus wage determination. Table 12 considers a range of multiples for the monopsony power parameters γ and ϕ from one-half to twice their baseline values. For the wage elasticity of the quit rate γ, that range corresponds to 0.8 to 3.2, which is substantially wider than the range in Hirsch et al. (2010) of 1.0 to 1.9.

Wage rigidity improves the model's fit to the data for a wide range of parameters governing the establishment's monopsony power. Each row in the table corresponds to a different set of target empirical moments used to calculate the sums of squared errors from the simulations. The first row considers the full set of target moments from Table 9. The full set of moments includes average measured wage rigidity and the estimated coefficients from the layoff, quit, and hire regressions on measured wage rigidity, which the model without wage rigidity cannot replicate successfully. Therefore, subsequent rows in the table consider more limited sets of moments that the model without wage rigidity should in principle be able to replicate if wage rigidity is unimportant for explaining employment flows. All table entries are much greater than one, suggesting that wage rigidity is necessary for the model to match the target moments under a wide range of parameter values.

8 Conclusion

This paper explores the relationship between downward nominal wage rigidity and employment outcomes theoretically and empirically using establishment-employee linked administrative data to measure wage rigidity and employment adjustment at the establishment level. The estimates suggest that wage rigidity prevents 27.1 percent of counterfactual nominal wage cuts. The paper introduces a theoretical model of an establishment's wage and employment decisions in the face of real resource costs for cutting nominal wages. The model predicts that establishments with more rigid wages should have higher layoff rates and lower quit and hire rates.

The empirical results are consistent with the predictions of the theoretical model. An establishment with the sample average level of measured wage rigidity is predicted to have a 4.2 percentage point higher layoff rate, a 6.4 percentage point lower quit rate, and a 2.0 percentage point lower hire rate than an establishment with no measured wage rigidity. The instrumental variable strategy suggests that these relationships are likely to be causal. The results therefore help to resolve the "micro-macro" puzzle highlighted by Lebow, Saks, and Wilson (2003). In particular, the use of administrative linked establishment-employee data indicates that downward nominal wage rigidity does affect employment outcomes as predicted by theory.

The theoretical model highlights the opposite effects that wage rigidity should have on layoff and quit rates: wage rigidity increases layoffs and reduces quits. This prediction finds support in the data. Thus, estimating the effects of wage rigidity on total separations or job destruction is likely to understate the true effects of wage rigidity by conflating these opposite effects.

Finally, the indirect inference procedure allows the effects of wage rigidity on employment outcomes to be interpreted through the lens of the structural model. The estimates imply that the average cost of a nominal wage cut is 9,000 euros, or roughly 30% of a worker's average annual compensation. Wage rigidity causes approximately one-quarter of all layoffs in the model.

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Table 1: Establishment-Level Descriptive Statistics

Number of Establishments	2,628
Sample Size, Establishment-Years	10,906
Mean Layoff Rate	0.068
	(0.115)
Mean Quit Rate	0.111
	(0.152)
Mean Hire Rate	0.223
	(0.336)
Mean Employees per Establishment	466
	(1,120)
Median Employees per Establishment	168
Mean Daily Wage, Nominal	87.124
	(28.537)
Median Daily Wage, Nominal	87.237
Mean Daily Wage, Year 2000 Euros	86.652
	(28.164)
Average Revenue Growth	0.043
	(0.254)

Standard deviations are in parentheses where applicable. Data come from the longitudinal model of the Linked-Employer-Employee Data (LIAB) (Version LM 2, Years 1993–2007) from the IAB. Sample restrictions and definition are described in Section 3.C.

Table 2: Establishment-Level Wage Rigidity Estimates

	Estimated Wage Rigidity		
	Mean	Median	Standard Deviation
	(1)	(2)	(3)
All Establishments	0.271	0.246	0.192
Supersector:			
Agriculture	0.190	0.162	0.171
Manufacturing	0.193	0.168	0.147
Mining/Energy/Water	0.339	0.347	0.169
Construction	0.088	0.072	0.089
Trade/Foodservice	0.259	0.238	0.138
Transportation	0.184	0.154	0.138
Finance	0.417	0.431	0.150
Real Estate	0.275	0.228	0.222
Public Administration	0.393	0.402	0.173
Other Services	0.341	0.367	0.216

Wage rigidity is estimated for the years 1997–2003 as described in Section 4. The wage rigidity estimator is a fixed characteristic of the establishment and estimates which fraction of nominal wage cuts were prevented due to downward nominal wage rigidity. Data come from the longitudinal model of the Linked-Employer-Employee Data (LIAB) (Version LM 2, Years 1993–2007) from the IAB.

Table 3: Supersector-Level Wage Collective Bargaining Statistics

	Percent of Workers at Establishments with Wages Set at Collectively Bargained Floor	
	Mean	Standard Deviation
	(1)	(2)
All Establishments	37.6	29.5
Supersector:		
Agriculture	27.0	10.3
Manufacturing	12.1	3.1
Mining/Energy/Water	54.3	22.3
Construction	34.0	9.2
Trade/Foodservice	14.2	6.5
Transportation	47.5	14.5
Finance	41.9	9.8
Real Estate	28.7	9.0
Public Administration	80.3	11.3
Other Services	65.7	10.9

Columns (1) and (2) show the means and standard deviations across states, respectively, of the percent of workers within each supersector who are employed at establishments bound by the sector-wide, collectively bargained wage agreements that pay at the collectively agreed upon wage floor from 1997 through 2003. Data come from the longitudinal model of the Linked-Employer-Employee Data (LIAB) (Version LM 2, Years 1993–2007) from the IAB.

Table 4: Instrument Validity

Dependent Variable	Proportion of Workers at Establishments Paying at Wage Floor							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
One-Year Lagged Depend. Var.	0.387***	0.387***	0.387***	0.365***	0.387***	0.387***	0.388***	0.365***
	(0.055)	(0.055)	(0.055)	(0.056)	(0.056)	(0.056)	(0.056)	(0.057)
Revenue Growth		0.016	0.016	0.013		0.016	0.017	0.013
		(0.041)	(0.041)	(0.042)		(0.041)	(0.041)	(0.042)
One-Year Lagged Revenue Growth			0.000	-0.012			0.002	-0.013
			(0.051)	(0.052)			(0.051)	(0.052)
One-Year Forward Revenue Growth				-0.01				-0.007
				(0.045)				(0.044)
Prop. Estabs Expect Emp Increase					-0.157	-0.157	-0.156	-0.134
					(0.110)	(0.110)	(0.110)	(0.124)
Prop. Estabs Expect Emp Decrease					0.04	0.042	0.042	-0.03
					(0.067)	(0.068)	(0.068)	(0.065)
Prop. Estabs Not Sure Emp Growth					0.135	0.135	0.135	0.039
					(0.130)	(0.130)	(0.131)	(0.138)
R-Squared	0.834	0.834	0.834	0.848	0.836	0.836	0.836	0.848
N	596	596	596	527	596	596	596	527

The unit of observation is a state-sector-year. Standard errors are clustered at the state-sector level and are in parentheses. One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively. All regressions include state, sector, and year fixed effects. The dependent variable is the proportion of workers in a state-sector-year working at establishments that a collectively-bargained wage floor, which is used as an instrumental variable in Tables 5-8. Revenue growth is defined as the percentage change in total revenue across all establishments in a state-sector-year. The proportion of establishments expecting employment increases and decreases are forward-looking variables calculated as the fraction of establishments in a state-sector-year that report expecting their levels of employment to increase or decrease, respectively, one year ahead. The proportion of establishments not sure about employment growth are those that reported “not sure at present” as to whether their employment level would increase, decrease, or be approximately constant one year ahead. The “approximately constant” response is the ommitted employment expectation category in the regressions in columns 5-7.

Table 5: Instrumental Variables First Stage Regressions for Estimated Wage Rigidity

Dependent Variable	Measured Wage Rigidity			
	(1)	(2)	(3)	(4)
Fraction Workers at Establishments Paying Collectively Bargained Floor	0.196***	0.198***	0.180***	0.213***
	(0.024)	(0.024)	(0.027)	(0.030)
Lagged Establishment-Level Collective Bargaining Agreement			0.031***	
			(0.007)	
Revenue Spline	No	Yes	No	No
Large Establishments Only	No	No	No	Yes
F-Statistic	65.5	66.9	39.4	51.9
R-Squared	0.364	0.365	0.369	0.455
N	10,906	10,906	8,989	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels. Results from columns 1 through 4 are first-stage estimates for columns 5 through 8, respectively, of Tables 6, 7, and 8. The lagged establishment-level collective bargaining agreement variable is a binary indicator for the presence or absence of an agreement in the previous year. Column 4 restricts the sample to establishment-years with employment above the median level.

Table 6: Wage Rigidity and Layoffs – Regression Results

Dependent Variable	Layoff Rate as a Fraction of Establishment Workforce							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Wage Rigidity	-0.004	-0.002	-0.020	0.032	0.154**	0.144**	0.122*	0.196*
	(0.020)	(0.020)	(0.018)	(0.040)	(0.074)	(0.073)	(0.073)	(0.102)
Positive Revenue Growth		0.014**				0.014**		
		(0.006)				(0.006)		
Negative Revenue Growth		-0.088***				-0.094***		
		(0.015)				(0.016)		
Works Council	-0.045***	-0.046***	-0.042***	-0.092***	-0.050***	-0.050***	-0.053***	-0.096***
	(0.010)	(0.010)	(0.009)	(0.030)	(0.010)	(0.010)	(0.012)	(0.030)
Specification	OLS	OLS	OLS, Sector FE	OLS	IV	IV	IV, 2 Insts.	IV
Sample Restriction	Full Sample	Full Sample	Full Sample	Large Estabs	Full Sample	Full Sample	2 Insts. Sample	Large Estabs
P-value of Exogeneity Test					0.008	0.013	0.037	0.046
R-Squared	0.148	0.154	0.199	0.247	0.100	0.113	0.121	0.211
N	10,906	10,906	10,906	5,439	10,906	10,906	8,989	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Layoffs are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, and year dummies as controls. Establishment characteristics include a set of controls for occupational mix, dummies for federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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 cover the period 1997 to 2003 and include establishments with at least 30 employees in a given year. The instrumental variable regressions in columns 5 through 8 are estimated by two-stage least squares, with first stages shown in Table 5. Columns 5 through 8 instrument for wage rigidity using state-sector level wage agreements as described in section 5.A. Column 7 supplements the first instrument with a second instrument based on lagged establishment-level collective bargaining agreements, also as described in section 5.A. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table 7: Wage Rigidity and Quits – Regression Results

Dependent Variable	Quit Rate as a Fraction of Establishment Workforce							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Wage Rigidity	-0.070***	-0.069***	-0.070***	-0.113***	-0.235***	-0.243***	-0.221***	-0.260***
	(0.017)	(0.017)	(0.020)	(0.029)	(0.056)	(0.056)	(0.052)	(0.081)
Positive Revenue Growth		0.044***				0.044***		
		(0.009)				(0.009)		
Negative Revenue Growth		-0.069***				-0.061***		
		(0.018)				(0.017)		
Works Council	-0.056***	-0.056***	-0.039***	-0.055***	-0.051***	-0.051***	-0.053***	-0.052***
	(0.008)	(0.008)	(0.006)	(0.015)	(0.008)	(0.008)	(0.008)	(0.014)
Specification	OLS	OLS	OLS, Sector FE	OLS	IV	IV	IV, 2 Insts.	IV
Sample Restriction	Full Sample	Full Sample	Full Sample	Large Estabs	Full Sample	Full Sample	2 Insts. Sample	Large Estabs
P-value of Exogeneity Test					0.001	0.001	0.001	0.024
R-Squared	0.144	0.148	0.214	0.236	0.114	0.115	0.206	0.218
N	10,906	10,906	10,906	5,439	10,906	10,906	8,989	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Quits are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, and year dummies as controls. Establishment characteristics include a set of controls for occupational mix, dummies for federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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 cover the period 1997 to 2003 and include establishments with at least 30 employees in a given year. The instrumental variable regressions in columns 5 through 8 are estimated by two-stage least squares, with first stages shown in Table 5. Columns 5 through 8 instrument for wage rigidity using state-sector level wage agreements as described in section 5.A. Column 7 supplements the first instrument with a second instrument based on lagged establishment-level collective bargaining agreements, also as described in section 5.A. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table 8: Wage Rigidity and Hires – Regression Results

Dependent Variable	Hire Rate as a Fraction of Establishment Workforce							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Wage Rigidity	-0.073**	-0.074**	-0.092**	-0.090	-0.128	-0.135	-0.139	-0.043
	(0.036)	(0.036)	(0.036)	(0.062)	(0.116)	(0.116)	(0.106)	(0.172)
Positive Revenue Growth		0.144***				0.144***		
		(0.023)				(0.023)		
Negative Revenue Growth		0.005				0.008		
		(0.031)				(0.031)		
Works Council	-0.130***	-0.129***	-0.109***	-0.170***	-0.129***	-0.127***	-0.137***	-0.171***
	(0.018)	(0.018)	(0.036)	(0.044)	(0.018)	(0.018)	(0.021)	(0.044)
Specification	OLS	OLS	OLS, Sector FE	OLS	IV	IV	IV, 2 Insts.	IV
Sample Restriction	Full Sample	Full Sample	Full Sample	Large Estabs	Full Sample	Full Sample	2 Insts. Sample	Large Estabs
P-value of Exogeneity Test					0.601	0.549	0.409	0.753
R-Squared	0.113	0.121	0.146	0.318	0.112	0.120	0.105	0.318
N	10,906	10,906	10,906	5,439	10,906	10,906	8,989	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Hires are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, and year dummies as controls. Establishment characteristics include a set of controls for occupational mix, dummies for federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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 cover the period 1997 to 2003 and include establishments with at least 30 employees in a given year. The instrumental variable regressions in columns 5 through 8 are estimated by two-stage least squares, with first stages shown in Table 5. Columns 5 through 8 instrument for wage rigidity using state-sector level wage agreements as described in section 5.A. Column 7 supplements the first instrument with a second instrument based on lagged establishment-level collective bargaining agreements, also as described in section 5.A. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table 9: Empirical and Simulated Moments

Moment	Sample Value	Simulated Value
Layoff Regression Coefficient on Wage Rigidity	0.15	0.08
Quit Regression Coefficient on Wage Rigidity	-0.24	-0.16
Hire Regression Coefficient on Wage Rigidity	-0.07	-0.09
Real Average Daily Wage Rate	86.65	83.95
Wage Change Distribution – 10th-50th Percentiles	-0.05	-0.05
Wage Change Distribution – 25th-50th Percentiles	-0.02	-0.02
Wage Change Distribution – 75th-50th Percentiles	0.03	0.03
Wage Change Distribution – 90th-50th Percentiles	0.07	0.07
Average Level of Measured Wage Rigidity	0.27	0.41
Average Layoff Rate	0.07	0.06
Average Hire Rate	0.22	0.35
Labor's Share of Value Added	0.60	0.49
Real Average Value Added per Worker	73,497	68,565
Empirical Wage Elasticity of the Quit Rate	1.01	1.22
Empirical Wage Elasticity of the Hire Rate	2.54	1.83
Empirical Economies of Scale of the Hire Rate	-0.18	-0.15
RMSE of the Empirical Hire Rate Elasticity Regression	0.57	0.60
AR(1) Persistence of Establishment Value Added per Worker	0.24	0.28
AR(1) RMSE of Establishment Value Added per Worker	0.36	0.34

The coefficients on wage rigidity in the layoff, quit, and hire regressions are from column 3 of Table 6, column 3 of Table 7, and column 1 of Table 8, respectively. Measured level of wage rigidity is mean wage rigidity for all establishments from Table 2, calculated as described in section 4. The wage change distribution percentiles are censored to be within 15 percentage points of the establishment-year median wage change, consistent with the wage rigidity estimation procedure.

Table 10: Estimated Parameter Values

Parameter	Description	Estimated Value	Standard Error
λ_0	Menu Cost of Downward Wage Adjustment	8,616	981
λ_1	Linear Cost of Downward Wage Adjustment	7,109	2,138
ψ_u	Persistence of Worker Type Productivity	0.794	0.014
σ_u	Standard Deviation of Worker Type Productivity	0.458	0.012
$\overline{c}$	Hiring Cost Function Scale Term	16,126	414
ϕ	Hiring Cost Function Wage Elasticity	1.247	0.060
υ	Hiring Cost Function Hire Rate Elasticity	0.642	0.010
η	Hiring Cost Function Employment Level Elasticity	0.309	0.002
c_l	Firing Cost per Worker	977	31
α	Returns to Scale in Production	0.703	0.019
γ	Wage Elasticity of the Quit Rate	1.692	0.016
$\overline{\delta}$	Quit Rate Scale Parameter	0.154	0.004
β	Establishment Discount Rate	0.530	0.002
$\overline{z}$	Average Establishment-Wide Productivity	140,425	7,714
ψ_z	Persistence of Establishment-Wide Productivity	0.388	0.010
σ_z	Standard Deviation of Establishment-Wide Productivity Shock	0.419	0.003
s_x	Exogenous Layoff Rate	0.053	0.003

Standard errors are in parentheses. Parameters are estimated by the indirect inference procedure described in section 7. The deterministic inflation rate, π, is set to be 1.3 percent per year, the average consumer price inflation rate from the World Bank's World Development Indicators for Germany.

Table 11: Decomposing the Costs of Cutting Nominal Wages at Different Productivity Levels

λ_0 λ_1	Baseline Baseline	Baseline Zero	Zero Baseline	Zero Zero
	A. Unconditional Productivity Distribution			
Layoff Rate	0.063	0.059	0.055	0.054
Quit Rate	0.272	0.281	0.301	0.308
Hire Rate	0.344	0.351	0.369	0.376
	B. Median Productivity Level			
Layoff Rate	0.062	0.057	0.053	0.053
Quit Rate	0.315	0.328	0.352	0.362
Hire Rate	0.186	0.192	0.206	0.213
	C. 25th Percentile Productivity Level			
Layoff Rate	0.073	0.062	0.054	0.053
Quit Rate	0.312	0.329	0.359	0.37
Hire Rate	0.166	0.17	0.191	0.197
	D. 75th Percentile Productivity Level			
Layoff Rate	0.053	0.053	0.053	0.053
Quit Rate	0.261	0.267	0.284	0.288
Hire Rate	0.419	0.428	0.447	0.457

This table decomposes the effects of each of the wage cutting cost parameters (λ_0 and λ_1) on the simulated establishment's layoff, quit, and hire rates. The first column displays results from model simulations in which both parameters are set to the values estimated in Table 10. The second column shows results in which the menu cost of cutting nominal wages (λ_0) is set at its estimated value, but the linear cost (λ_1) is set to zero. The third column shows results in which λ_0 is set to zero and λ_1 is set at its estimated value. The fourth column shows results in which both λ_0 and λ_1 are set to zero. Panel A shows simulations over the entire estimated productivity distribution. Panel B shows results at the median productivity level, panel C shows results at the 25th percentile productivity level, and panel D shows results at the 75th percentile productivity level.

Table 12: Wage Rigidity's Effect on Model Fit Under Various Parameter Values

	Ratio of Simulated Sum of Squared Errors with and without Wage Rigidity				
	Market Power Parameters Scale Factor				
Set of Target Moments	2	1.5	1	2/3	1/2
Full Set	123.0	7,294.1	8.7	2,289,613.8	162,658.6
Employment Flows Only	3.3	2.9	2.5	2.3	2.4
Employment Flows and Wage Change Percentiles	3.8	3.9	4.1	3.8	4.0
Employment Flows, Wage Change Percentiles, and Wage Elasticities	2.6	3.4	4.0	3.9	4.1
Full Set Except Measured Wage Rigidity and Wage Rigidity Coefficients	3.2	6.9	16.4	13.4	10.0

This table compares the data to the simulated model fit with and without wage rigidity for various levels of the establishment's market power over its employees. Each entry in the table shows the ratio of the sum of squared errors in a model simulation without wage rigidity (i.e., with the cost of wage cut parameters λ_0 and λ_1 set to zero) to the sum of squared errors in model simulation with wage rigidity (i.e., with the cost of wage cut parameters λ_0 and λ_1 set to their baseline values from Table 10). Each column specifies a ratio for the market power parameters γ (the wage elasticity of the quit rate) and ϕ (the wage elasticity of the hiring rate) relative to their baseline values, where higher values correspond to less market power. Each row corresponds to a different set of target empirical moments used to calculate the sums of squared errors from the simulations. In the first row, the full set of target moments from Table 9 is used. The second row uses the average layoff, quit, and hire rates as target moments. The third row adds the wage change percentiles shown in Table 9. The fourth row adds additionally the empirical quit and hire rate elasticities with respect to the wage rate. The final row uses the full set of moments from Table 9 except for measured wage rigidity and the layoff, quit, and hire regression coefficients on wage rigidity. A table entry greater than one implies that the simulated data without wage rigidity produced a worse fit to the target moments than the simulated data with wage rigidity. Because all entries are greater than one, the simulations suggest that wage rigidity improves the model's fit to the data for a wide range of parameter values and target moments.

Figure 1: Establishment Policy Functions - High Productivity

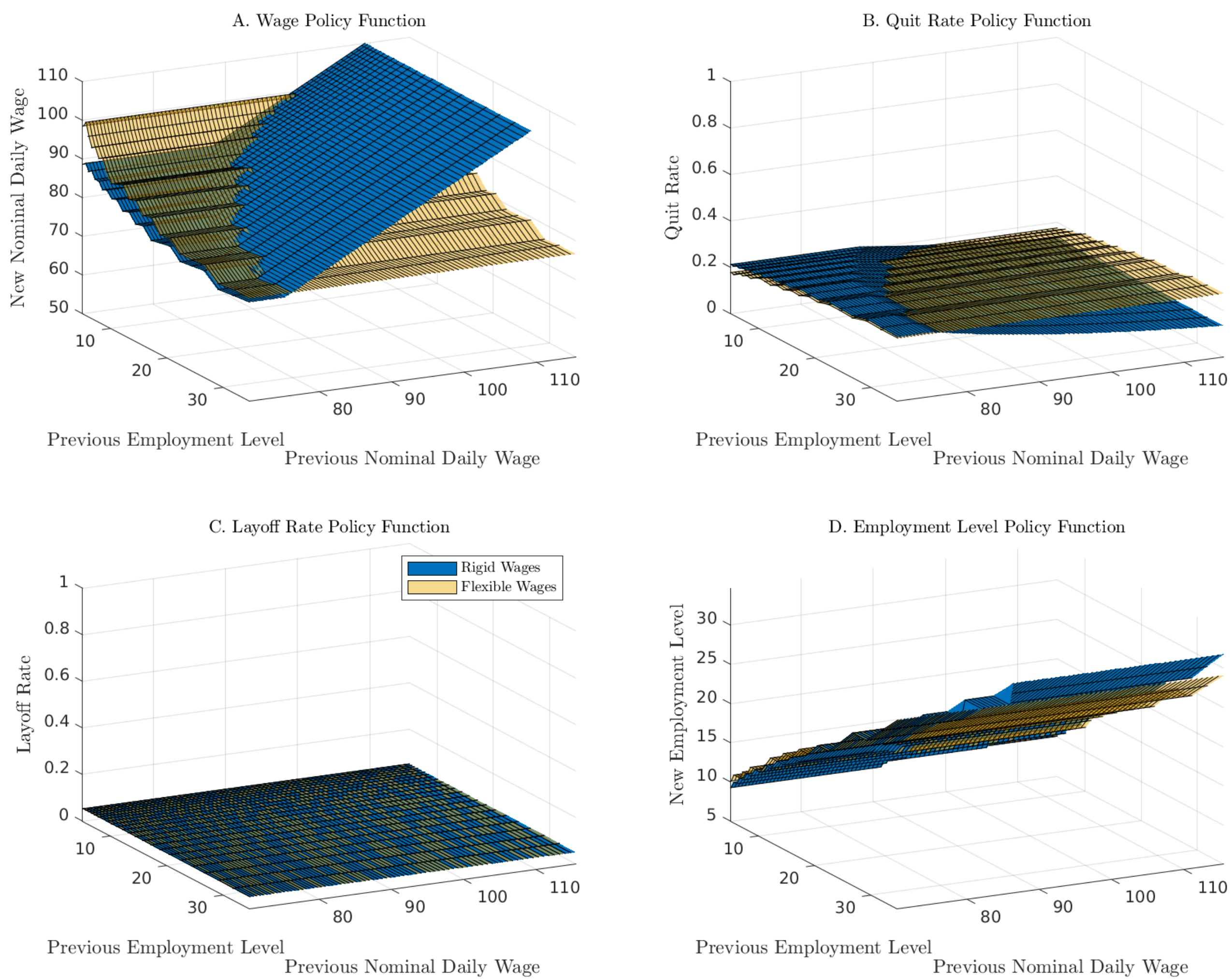


Figure 2: Establishment Policy Functions - Low Productivity

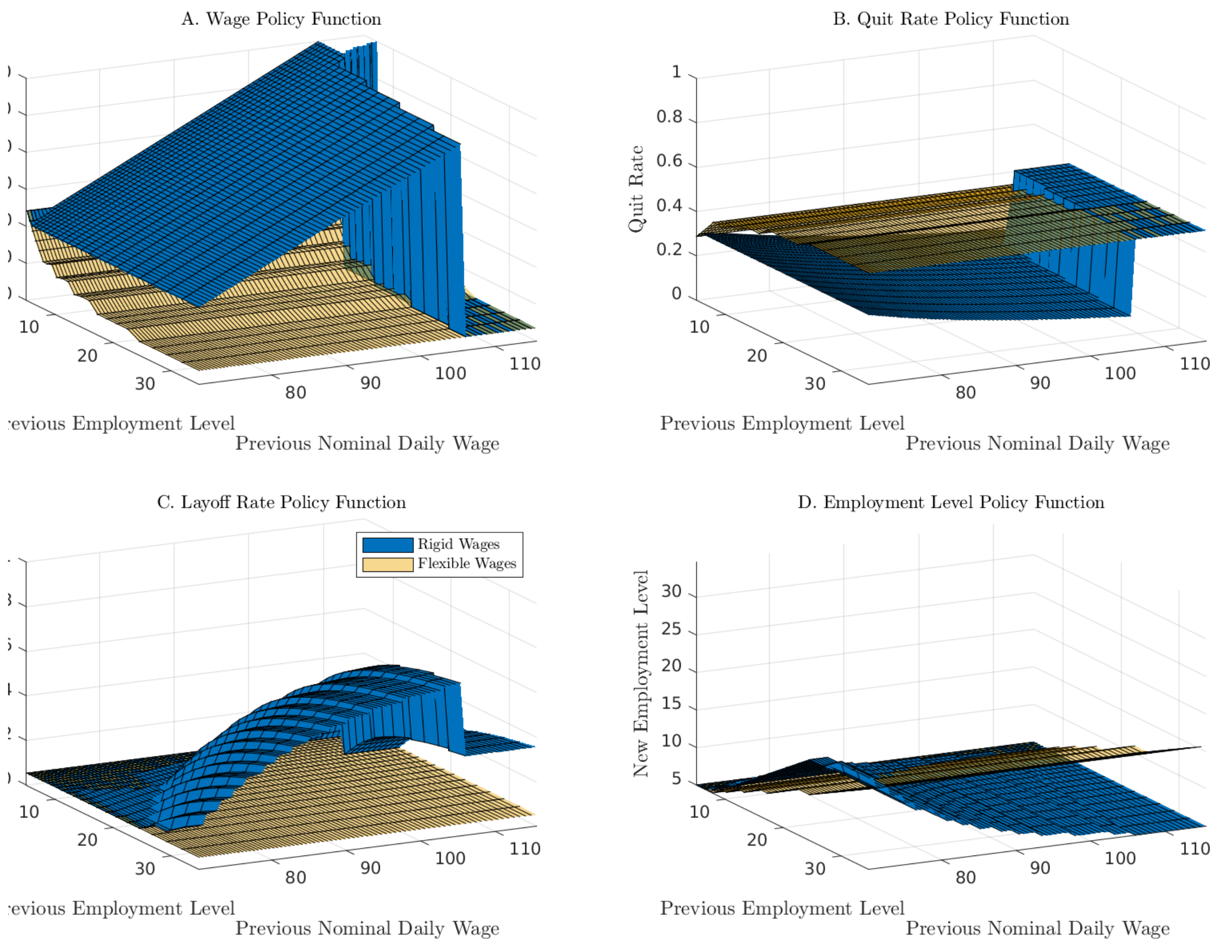
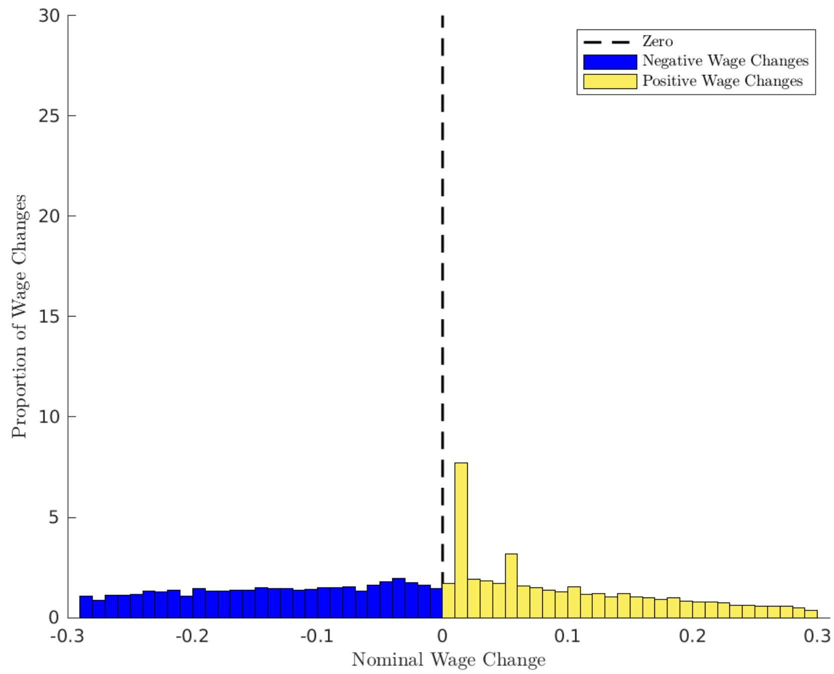
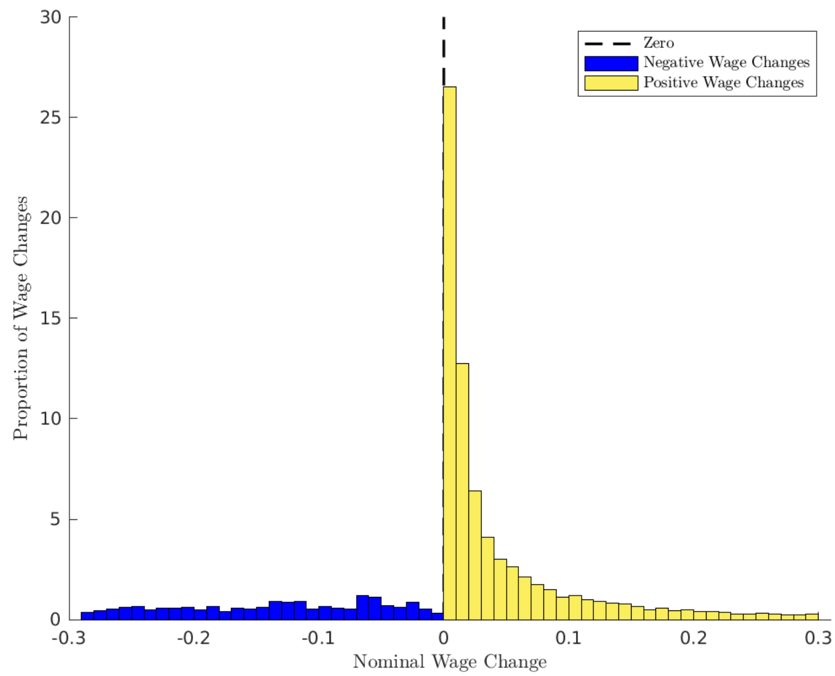


Figure 3: Wage Change Distributions with No Wage Rigidity



Note: Wage change distribution is censored at -30% and 30%.

Figure 4: Wage Change Distributions with Wage Rigidity



Note: Wage change distribution is censored at -30% and 30%.

Figure 5: Simulated Moments with Different Levels of Wage Rigidity

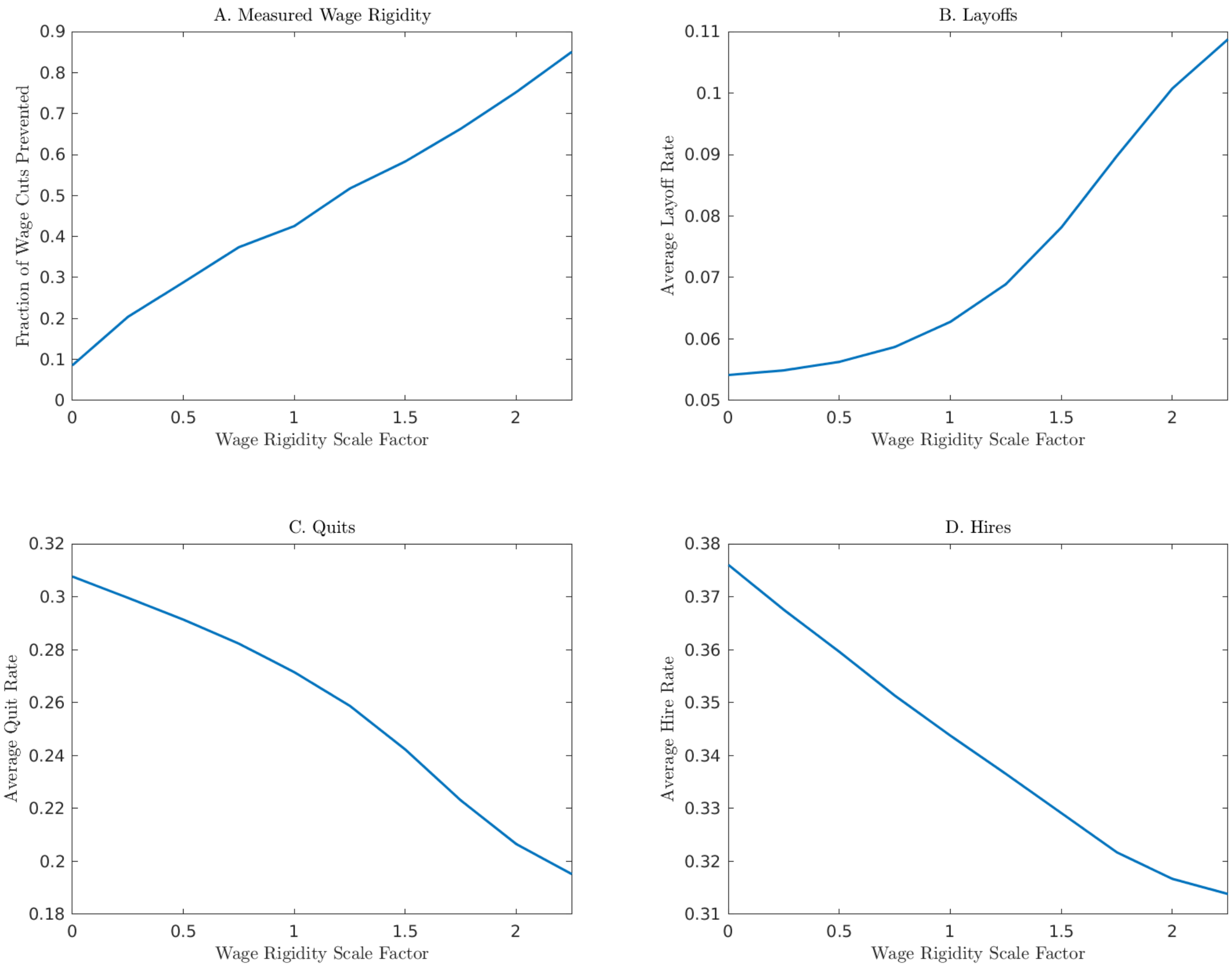


Figure 6: Aggregate Wage Change Distributions 1997 to 2003

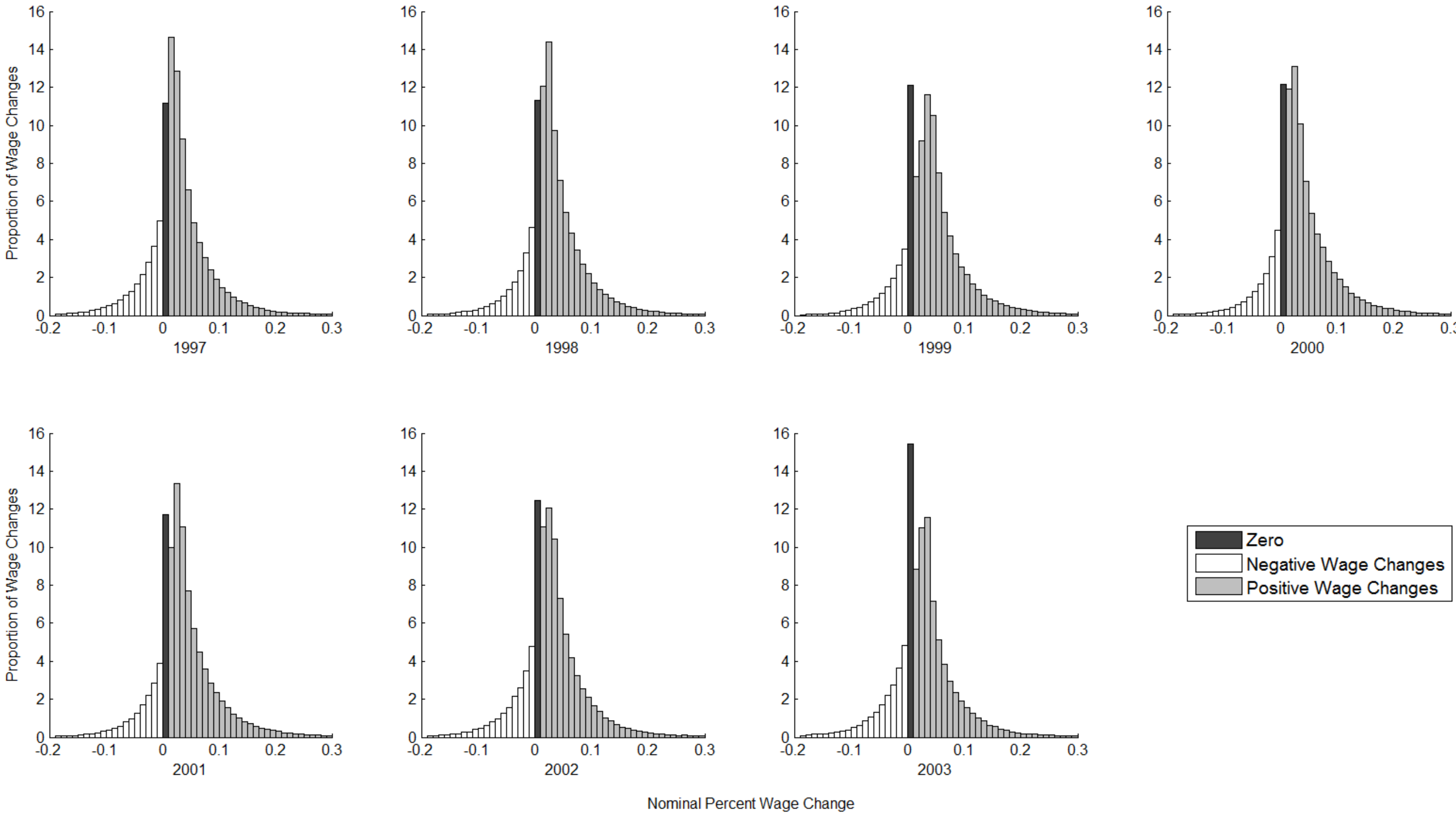
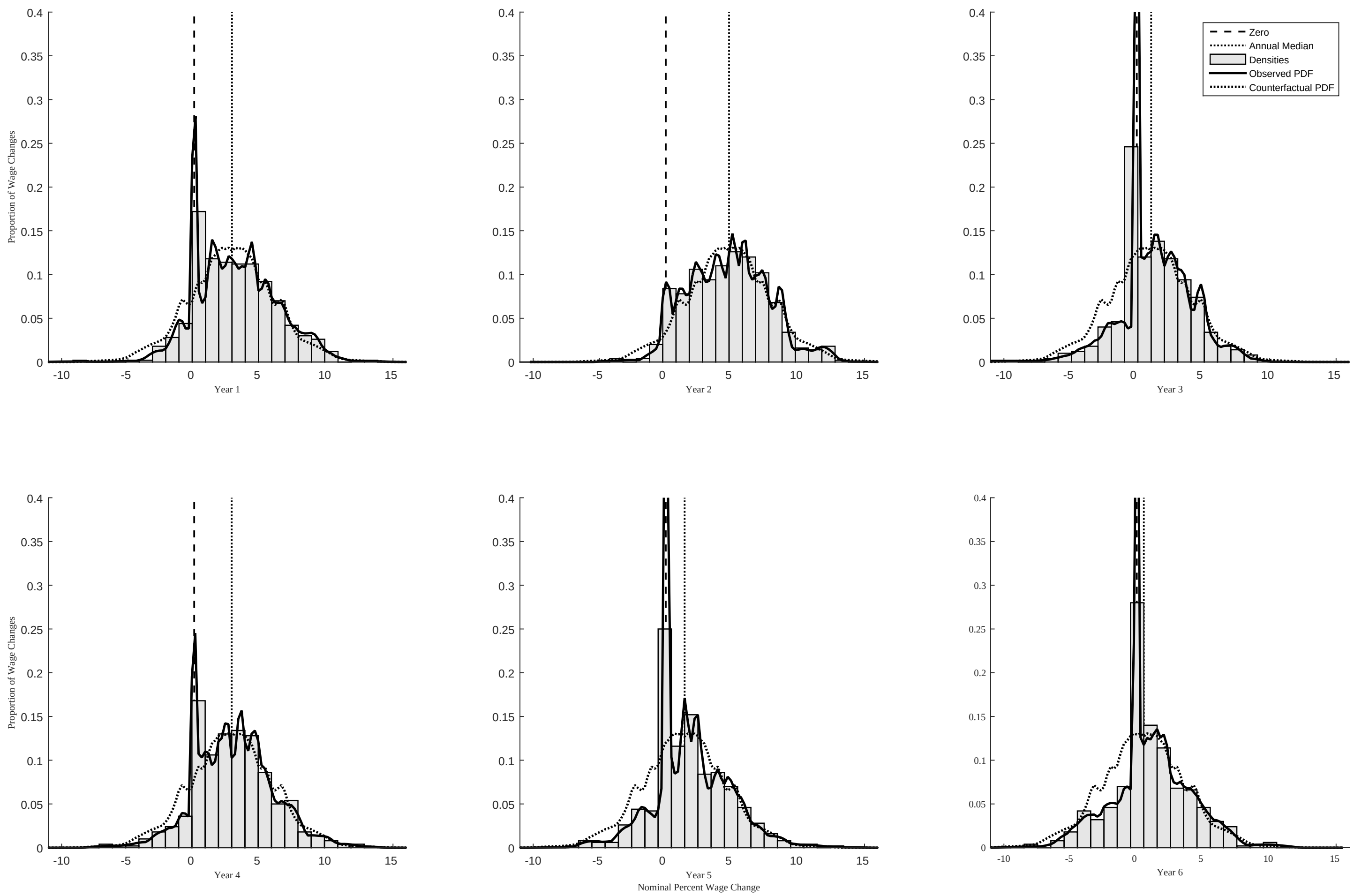
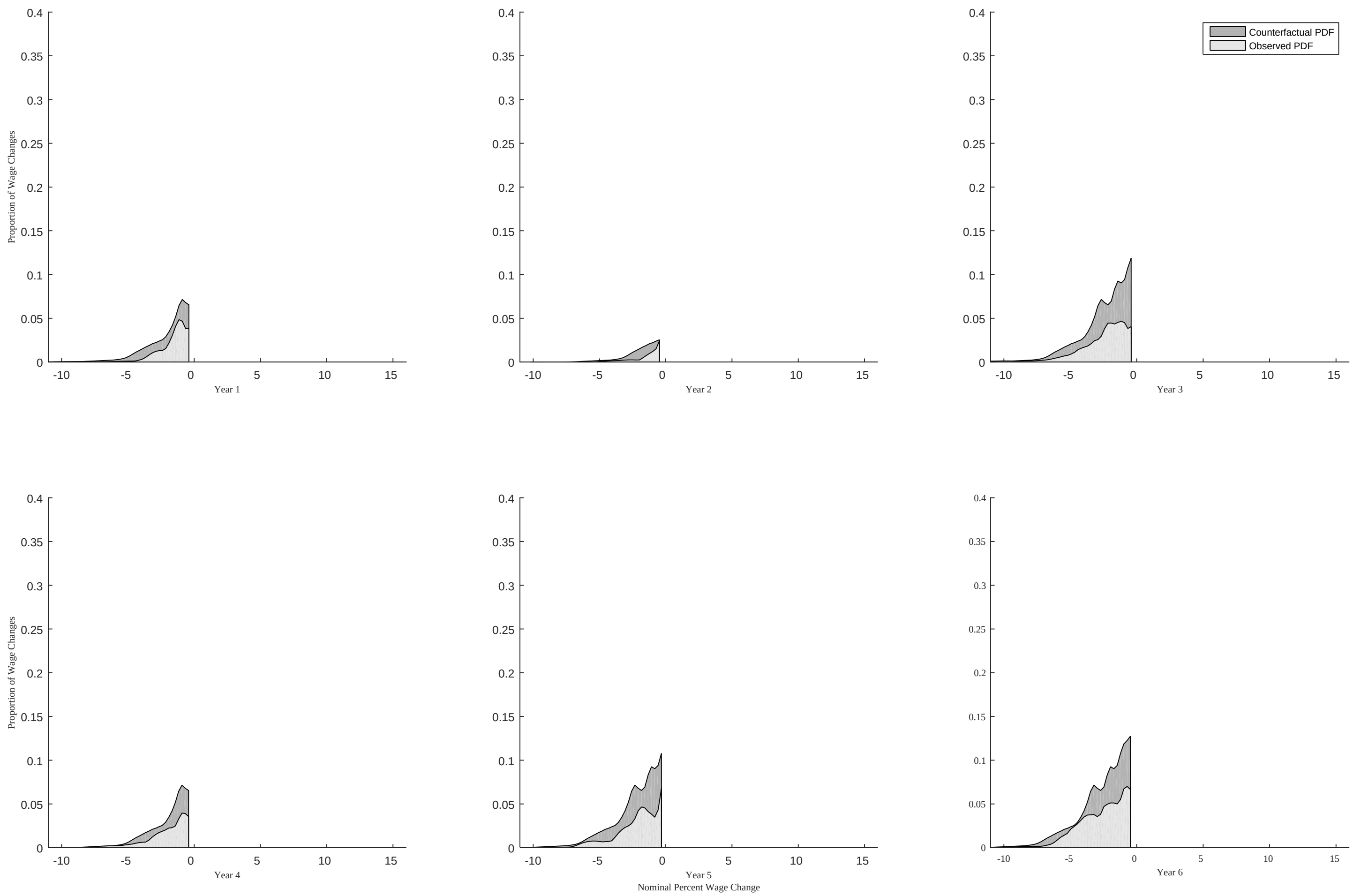


Figure 7: Illustration of Wage Rigidity Estimator: Estimated Kernel Densities



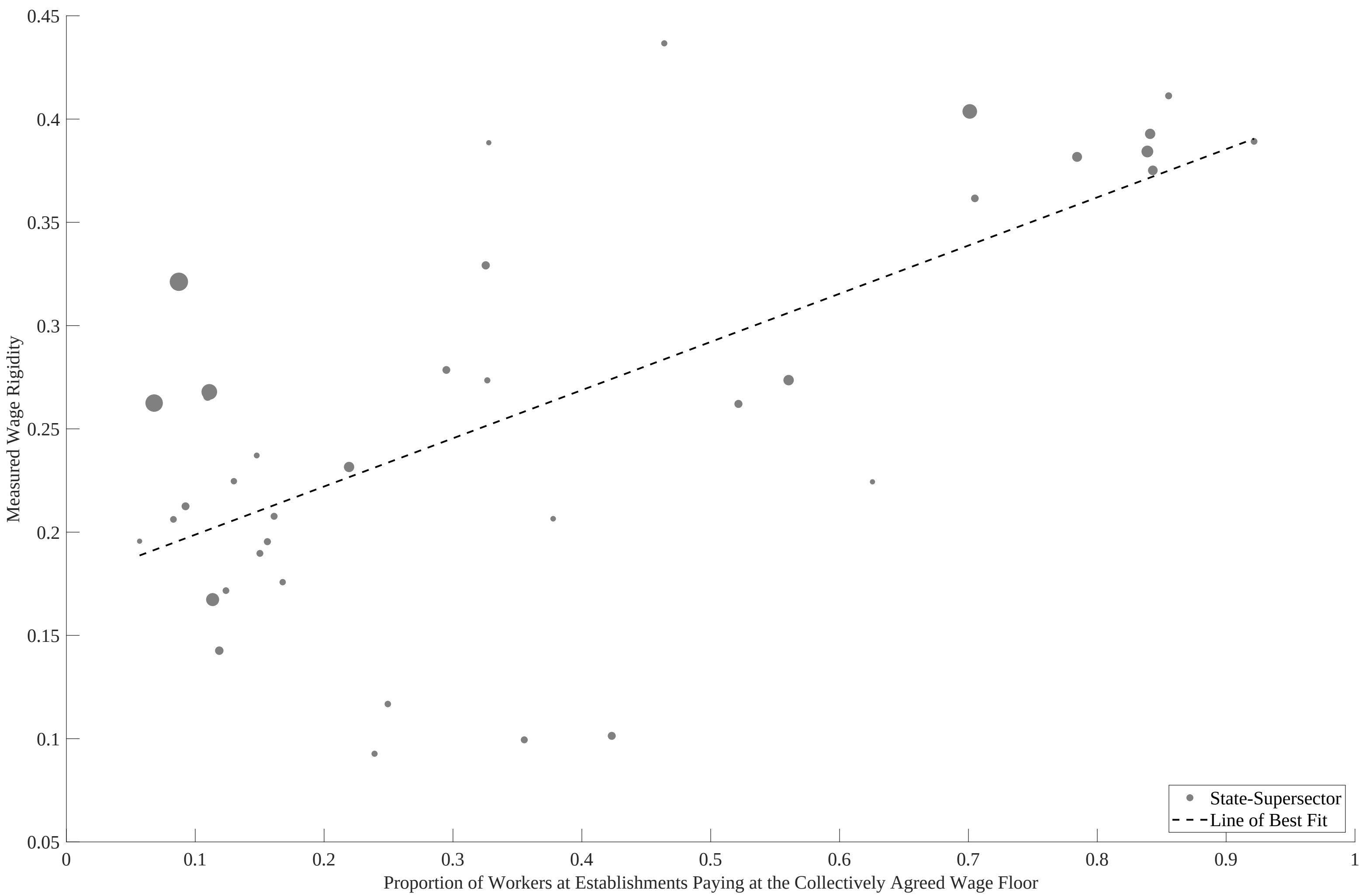
Wage change distributions from a simulated establishment over six years. Distributions estimated using kernel density estimation. The counterfactual distribution is constructed by reflecting the upper half of the estimated distribution, averaged over all years, around the median wage change in each year.

Figure 8: Illustration of Wage Rigidity Estimator: Estimating Missing Wage Cuts



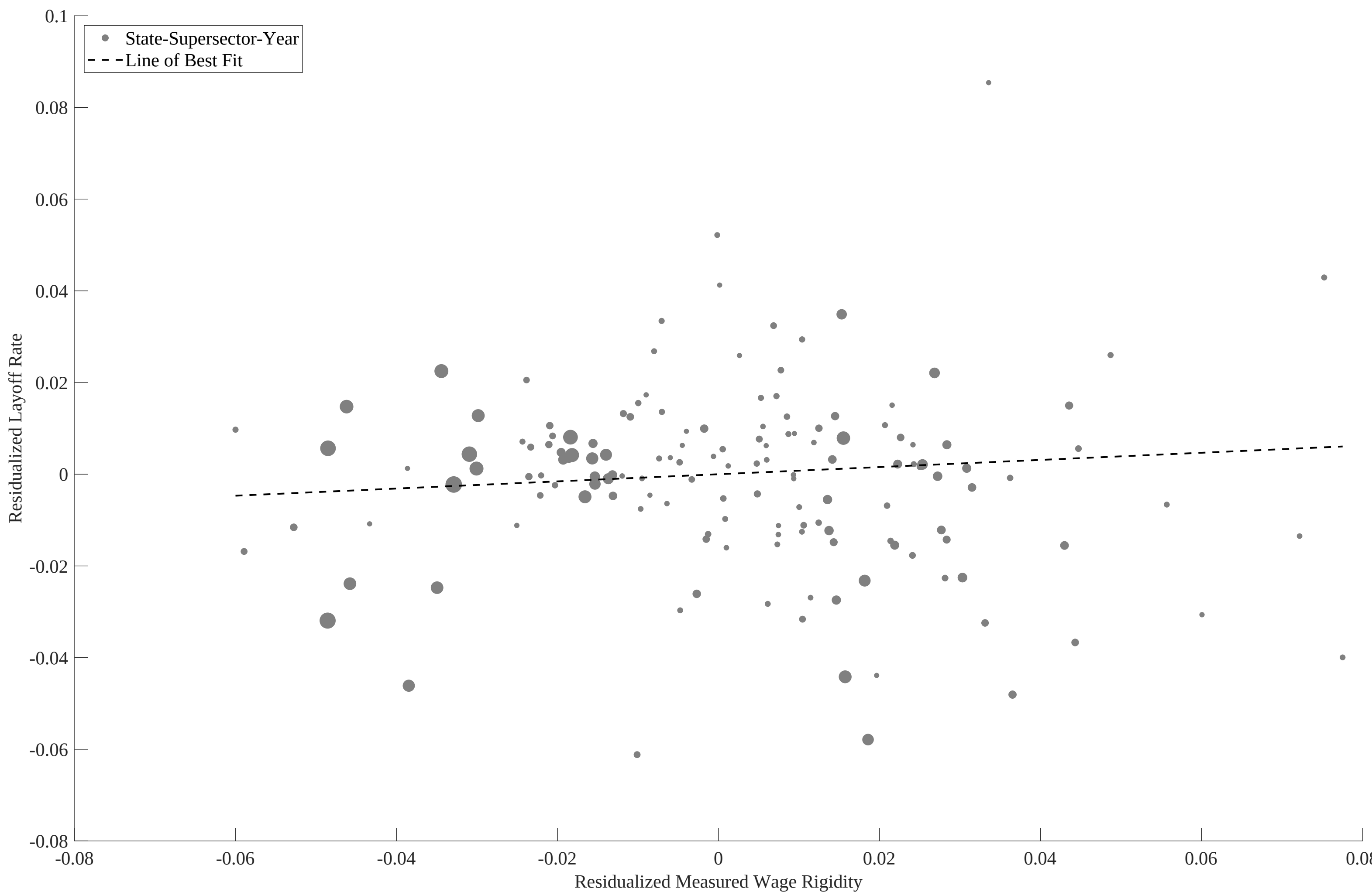
Wage change distributions from a simulated establishment over six years. The fraction of missing wage cuts is calculated as one minus the total area across years of the observed PDFs divided by the total area across years of the counterfactual PDFs.

Figure 9: Measured Wage Rigidity vs. Share of Workers at Establishments Paying at Collectively Bargained Floor



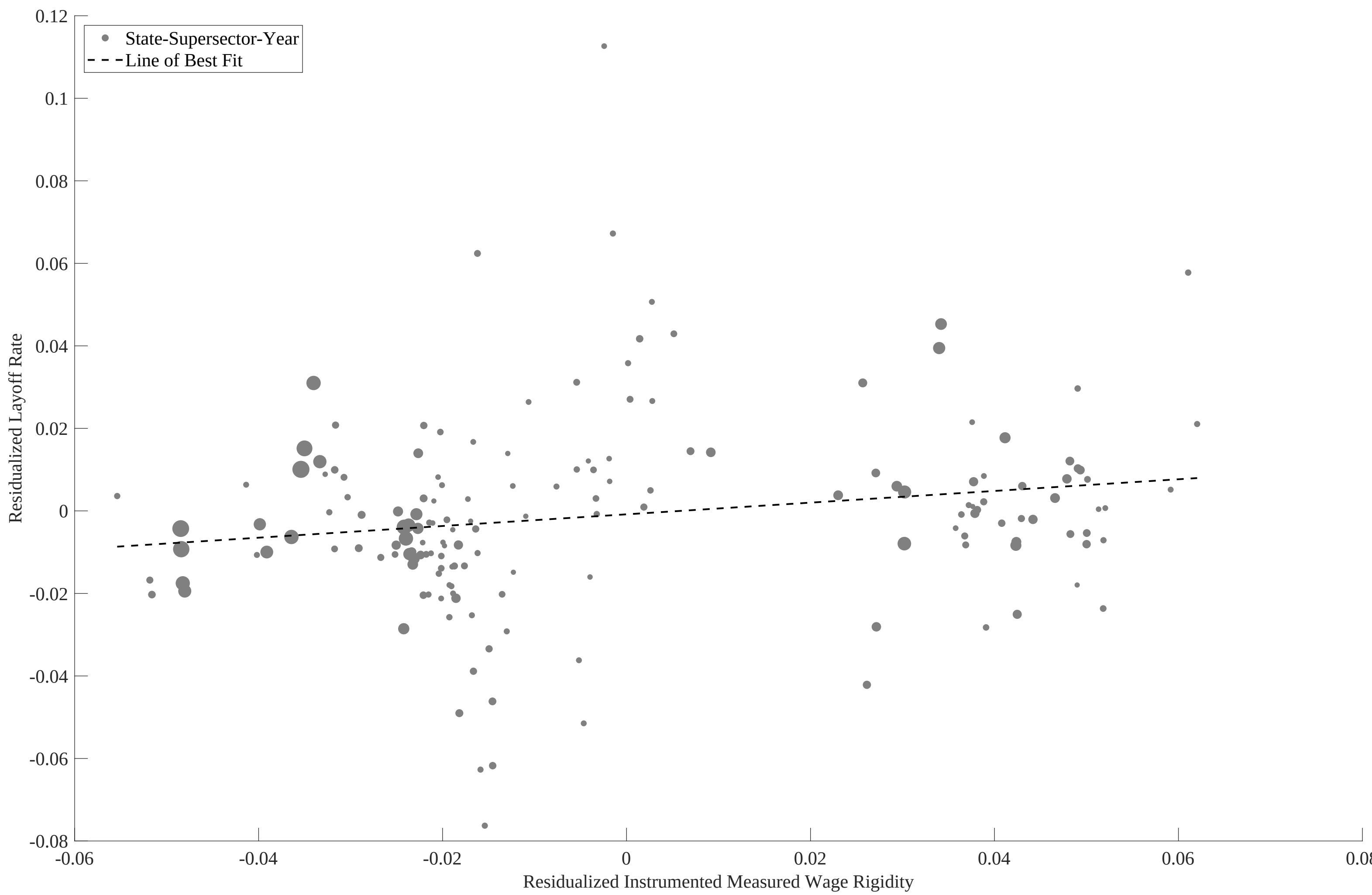
Points on the scatterplot are state-supersector level averages. Due to disclosure requirements, only state-sectors with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector.

Figure 10: Layoffs vs. Measured Wage Rigidity



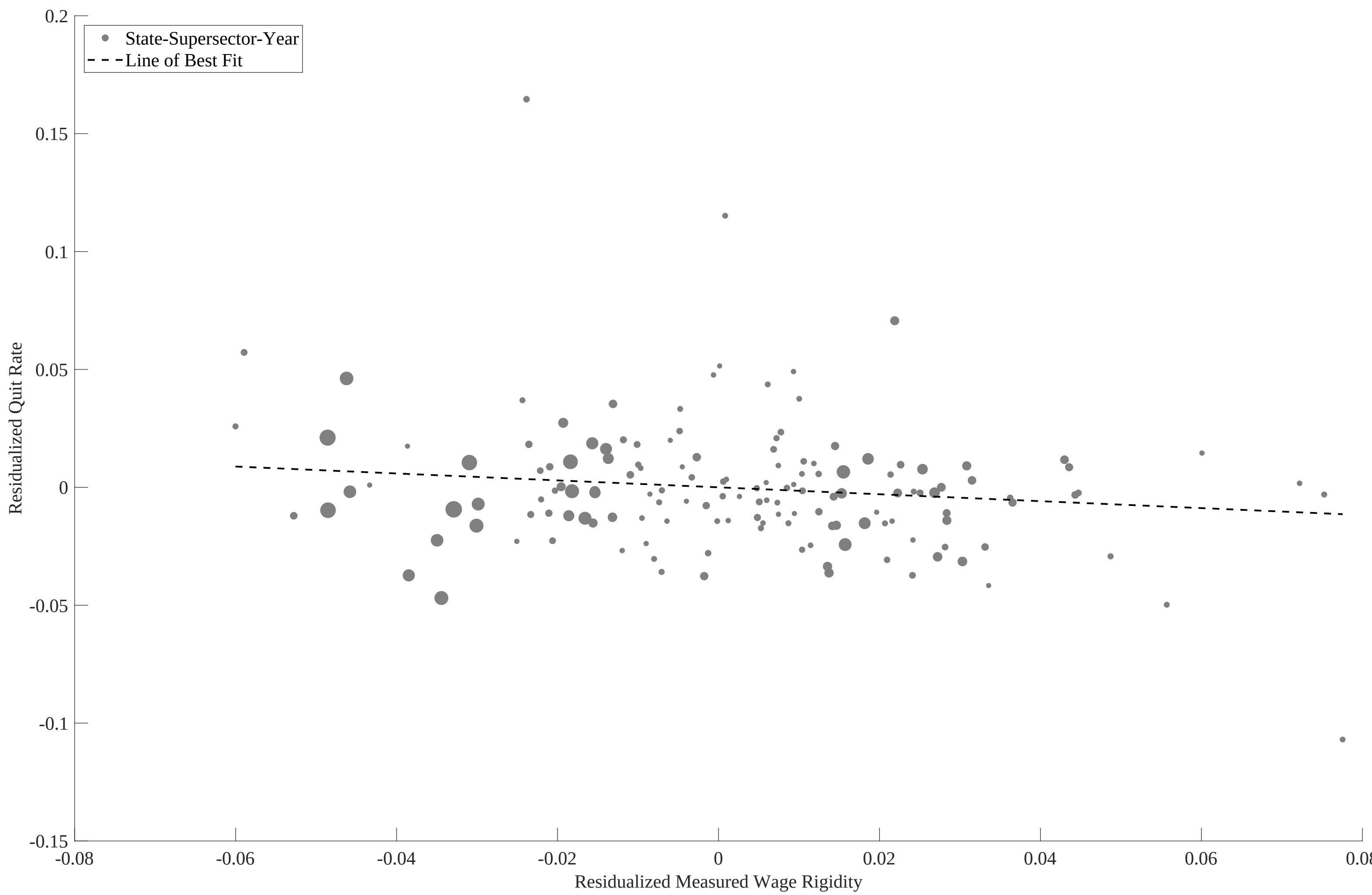
Points on the scatterplot are state-supersector-year level averages. Due to disclosure requirements, only state-sector-years with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector-year. Layoff rate and measured wage rigidity are residualized using the covariates described in Table 6.

Figure 11: Layoffs vs. Instrumented Measured Wage Rigidity



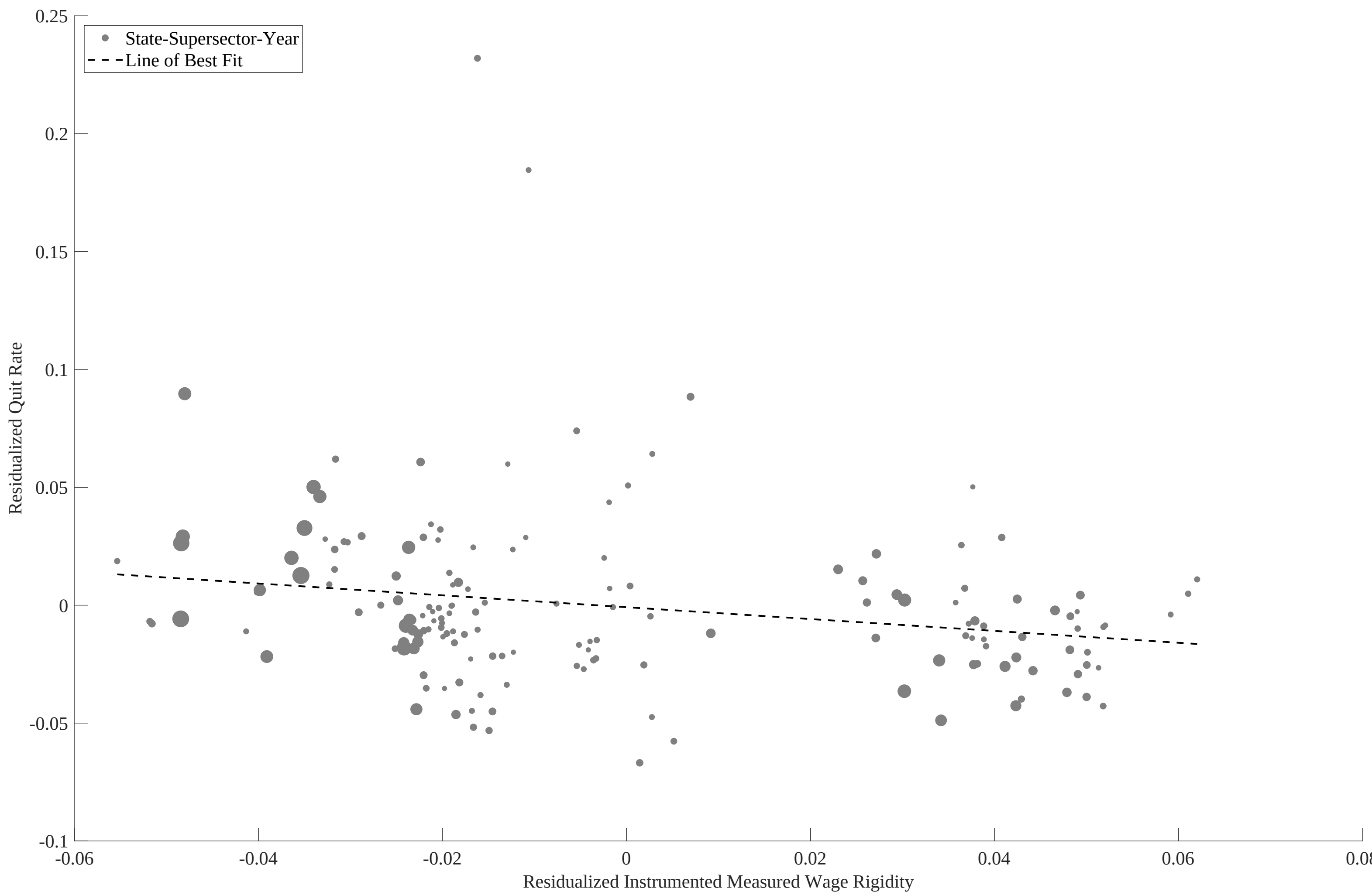
Points on the scatterplot are state-supersector-year level averages. Due to disclosure requirements, only state-sector-years with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector-year. Layoff rate is residualized using the covariates described in Table 6. Instrumented wage rigidity is the level of measured wage rigidity predicted by those covariates and the proportion paying at the wage floor instrument in Table 5, which is residualized using the covariates.

Figure 12: Quits vs. Measured Wage Rigidity



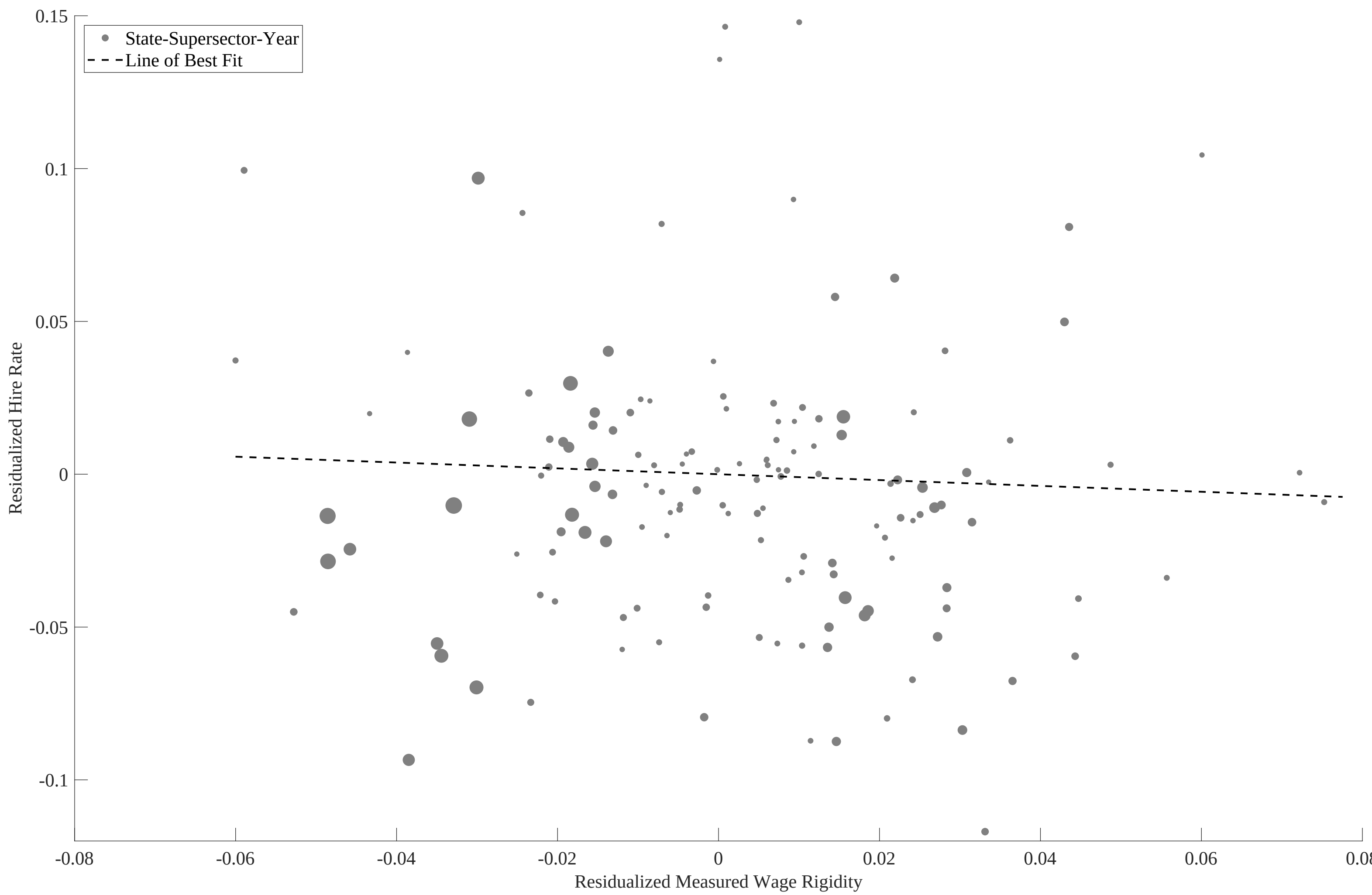
Points on the scatterplot are state-supersector-year level averages. Due to disclosure requirements, only state-sector-years with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector-year. Quit rate and measured wage rigidity are residualized using the covariates described in Table 6.

Figure 13: Quits vs. Instrumented Measured Wage Rigidity



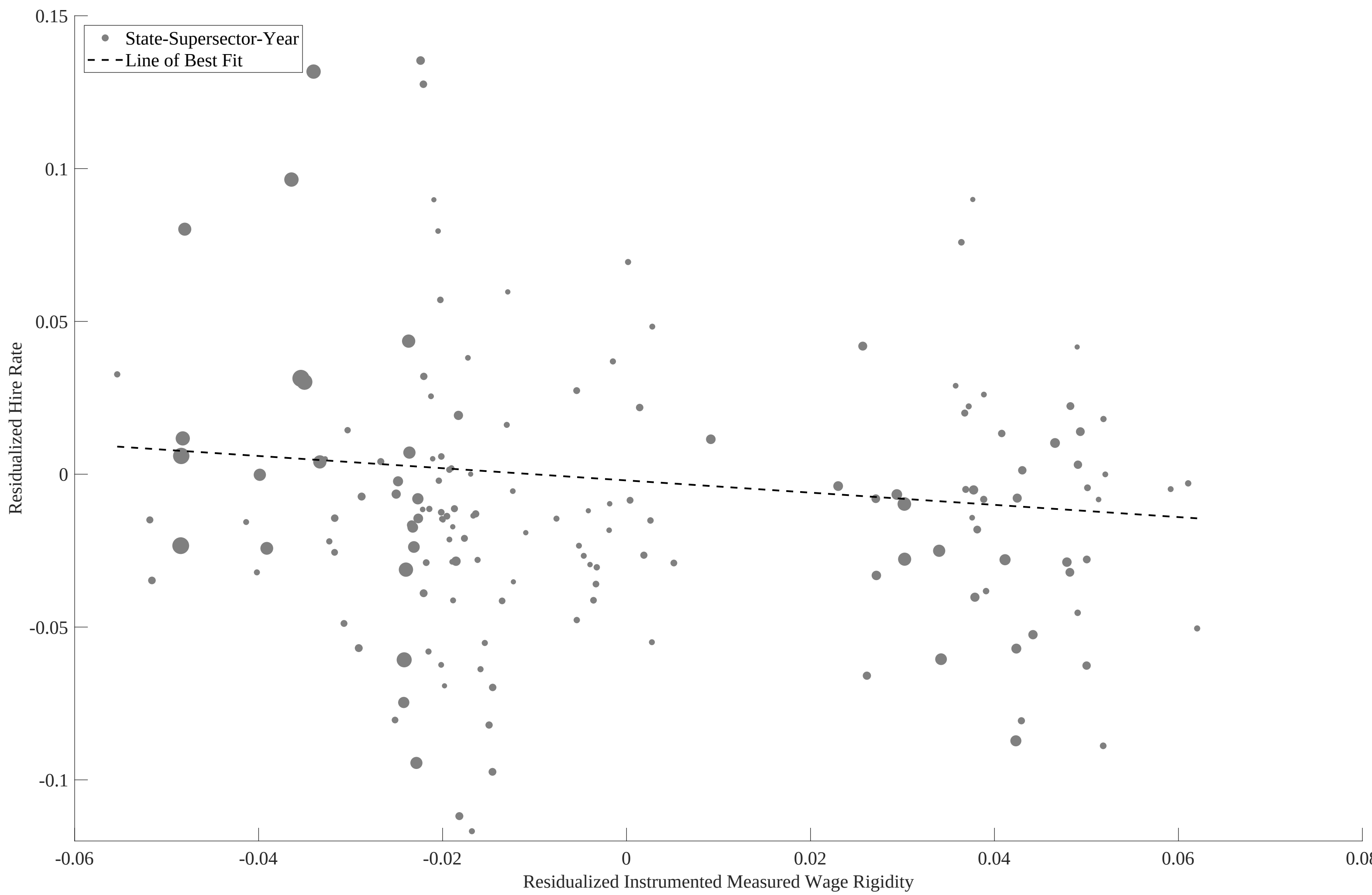
Points on the scatterplot are state-supersector-year level averages. Due to disclosure requirements, only state-sector-years with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector-year. Quit rate is residualized using the covariates described in Table 7. Instrumented wage rigidity is the level of measured wage rigidity predicted by those covariates and the proportion paying at the wage floor instrument in Table 5, which is residualized using the covariates.

Figure 14: Hires vs. Measured Wage Rigidity



Points on the scatterplot are state-supersector-year level averages. Due to disclosure requirements, only state-sector-years with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector-year. Hire rate and measured wage rigidity are residualized using the covariates described in Table 6.

Figure 15: Hires vs. Instrumented Measured Wage Rigidity



Points on the scatterplot are state-supersector-year level averages. Due to disclosure requirements, only state-sector-years with 20 or more establishments in the analysis are included in the figure. The size of each point is proportional to the numbers of establishments in the state-supersector-year. Hire rate is residualized using the covariates described in Table 8. Instrumented wage rigidity is the level of measured wage rigidity predicted by those covariates and the the proportion paying at the wage floor instrument in Table 5, which is residualized using the covariates.

Appendix for Online Publication

A An Analytical Model of Downward Nominal Wage Rigidity's Effects on Employment Flows

This appendix presents an analytical version of the quantitative model developed in Section 2 to derive the theoretical effects of wage rigidity on an establishment's layoff, quit, and hiring rates. Similarly to the numerical model in Section 2, the analytical model augments the model in Manning (2006) to include a resource cost of adjusting wages. It thereby incorporates a notion of wage rigidity and allows for a distinction among separations between layoffs and quits. The analytical model ignores the role of price inflation in the analysis and therefore treats nominal and real wage rigidity interchangeably.

Specifically, consider an establshment that begins period t with productivity a_t, an employment stock N_t, and which paid wage w_{t-1} last period. It chooses levels of hires H_t, layoffs L_t, and the wage w_t, to maximize the current period's flow profits and a continuation value according to the following problem:

$$
\begin{aligned}
V(a_t, N_t, w_{t-1}) = \max_{H_t, L_t, w_t} & F(a_t, N_t) - w_t N_t - C^H(H_t, N_t, w_t)H_t - c_l L_t - \\
& g(w_{t-1}, w_t)N_t + \beta E_t\left[V(a_{t+1}, N_{t+1}, w_t)\right] \qquad \text{(A.1)}
\end{aligned}
$$

subject to:

$$
\begin{aligned}
N_{t+1} &= N_t + H_t - L_t - \delta(w_t)N_t && \text{(A.2)}\\
a_{t+1} &= f(a_t, \epsilon_{t+1}) && \text{(A.3)}\\
H_t, L_t, N_t, w_t &\geq 0. && \text{(A.4)}
\end{aligned}
$$

The function $F(a_t, N_t)$ in equation A.1 is the production function, $C^H(H_t, N_t, w_t)$ is the per-hire recruitment cost function, $g(w_{t-1}, w_t)$ is the per-employee cost of adjusting wages, and $\delta(w_t)$ is the quit rate. Next period's productivity a_{t+1} is a function of current productivity a_t plus a random error ϵ_{t+1}. In some places, the analysis assumes certain functional forms for C^H, $\delta(w_t)$, and $g(w_{t-1}, w)$. In those cases, the following functional forms are assumed:

$$
\begin{aligned}
C^H(H_t, N_t, w_t) &= \bar{c} w_t^{-\phi} H_t^{\nu} N_t^{-\eta} && \text{(A.5)}\\
\delta(w_t) &= \bar{\delta}\left(\frac{w_t}{\overline{w}}\right)^{-\gamma} && \text{(A.6)}\\
g(w_{t-1}, w_t) &= \lambda\left(\frac{\pi}{2} - \arctan\left(\theta\left(w_t - w_{t-1}\right)\right)\right) && \text{(A.7)}
\end{aligned}
$$

All of the parameters in equations (A.1)–(A.7) are assumed to be positive unless noted otherwise. The functional forms of C^H and $\delta(w_t)$ are the same as in the numerical model in Section 2. The equation of motion for a_t is not specified but is also consistent with the model in Section 2.

The functional form of $g(w_{t-1}, w_t)$ merits further discussion because it differs from the cost function for nominal wage adjustments in in Section 2. $g(w_{t-1}, w_t)$ in equation (A.7) is continuous in the wage change and has a uniformly positive derivative with respect to current-period wages. One way to conceptualize this alternative cost function is to note that establishments face an ongoing cost of negotiating wages with their employees; the functional form of $g(w_{t-1}, w_t)$ asserts that this cost is lower when the newly offered wage is higher. The functional form preserves the property from the numerical model in the main text that the gradient is especially steep in the vicinity of the previous period's wage. Fitzenberger, Kohn, and Lembcke (2008) note that such transaction costs factor into establishments' wage setting behavior: "Wages set at the firm level as well as individually bargained wages are adapted towards collective bargaining agreements, be it in order to reduce transaction costs or not to create incentives for employees to join a union."

Nonetheless, the functional form of $g(w_{t-1}, w_t)$ is very similar to the wage adjustment cost function in the numerical model in the body of the text. Note that, in the limit as $\theta \to \infty$, $g(w_{t-1}, w_t)$ converges to a step function that takes value λ when $w_t < w_{t-1}$ and value 0 otherwise. In that sense, the wage adjustment cost function in equation (A.7) is a continuous approximation to the function in equation (2), with $\lambda_0 = \lambda$ and $\lambda_1 = 0$. The use of a continuous and differentiable function simplifies the analysis considerably by eliminating the need to consider several separate cases depending on whether or not the establishment is raising, freezing, or cutting wages.[57]

The definitions of $\delta(w)$ and C_H correspond to their definitions in equations (3) and (4) from Section 2; they also correspond to the functional forms in Manning (2006). Nonetheless, those functional forms are not essential to the key analytical results.

It will be analytically helpful to work with what Manning (2006) terms the "labor cost function" $\mathcal{C}(N, w)$, which represents the firm's per-worker hiring cost of hiring exactly enough workers to offset quits when it pays wage w:

$$\mathcal{C}(N, w) = C^H(H, N, w)\delta(w). \tag{A.8}$$

As the analysis will show, the key necessary assumption is that the firm's labor cost function, $\mathcal{C}(N, w)$, is convex in the wage rate.[58] The functional forms in equations (A.5) and (A.6) are consistent with this assumption.[59]

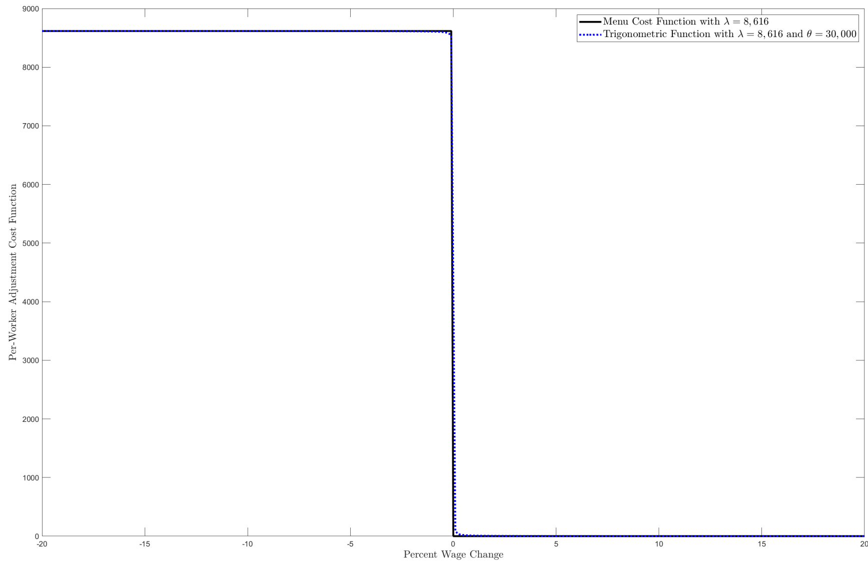
The plan of analysis in this section is to show that, in this stylized model, an increase in the wage rigidity parameter λ increases the establishment's layoff rate and decreases its quit and hire rates. Those are the three predictions generated by the quantitative model in Section 2 and tested in Section 6. The analysis begins by showing that the latter two results hold in steady state in the neighborhood of the flexible wage case ($\lambda \approx 0$) under the simplifying approximation that establishments are perfectly patient ($\beta \approx 1$). The latter assumption is innocuous but simplifies the algebra. The layoff rate will be zero in the model's steady state regardless of the level of wage rigidiy, so a non-steady state analysis is necessary. The analysis shows that, outside of steady state, the layoff rate must be an increasing function of the wage adjustment cost parameter λ in a

[57]The exact functional form of g is not essential; its essential features for the analysis are that it is continuous in w_t; its first derivative with respect to w_t is always negative ($g_{w_t} < 0$); in the neighborhood of $\lambda \approx 0$, it and its first and second derivatives with respect to w_t are approximately 0 ($\lambda \approx 0 \to g \approx g_{w_t} \approx g_{w_t w_t} \approx 0$); and its cross-derivative with respect to w_t and λ is negative ($g_{w\lambda} < 0$, i.e., it is more expensive to cut wages when λ is larger). Finally, note that π in equation (A.7) refers to the mathematical constant, not to the rate of price inflation.

[58]An additional necessary condition is that the firm's quit rate is decreasing in the wage it pays, which is a natural assumption.

[59]Manning (2006) also assumes this condition in its analysis.

Figure A.1: Illustration of $g(w_{t-1}, w_t)$ for large θ



This figure illustrates the nominal wage adjustment cost function in equation (2) of the main text for $\lambda_0 = 8,618$ and $\lambda_1 = 0$, as well as the continuous approximation $g(w_{t-1}, w_t)$ used in the analytical model for a large value of θ.

neighborhood of flexible wages.

To show that the steady-state quit and hire rates are decreasing in λ, start by substituting the law of motion for employment and taking the first derivatives of the value function with respect to H_t and w_t:[60]

$$(H_t): \quad \beta E_t \left[V_{N_{t+1}}\right] = C_t^H + H_t C_{H_t}^H \tag{A.9}$$
$$(w_t): \quad -N_t - H_t C_{w_t}^H - g_{w_t} N_t - \delta_{w_t} N_t \beta E_t \left[V_{N_{t+1}}\right] = 0. \tag{A.10}$$

The envelope condition for N_t is:

$$V_{N_t} = F_{N_t} - w_t - C_{N_t}^H H_t - g_t + \beta E_t \left[V_{N_{t+1}}\right](1 - \delta_t). \tag{A.11}$$

Using the fact that in steady state, $V_{N_t} = E_t\left[V_{N_{t+1}}\right] = V_N$ and the approximation $\beta \approx 1$ gives the steady-state relationship:

$$V_N = \frac{F_N - w - C_N^H H - g}{\delta}. \tag{A.12}$$

[60]This portion of the analysis will ignore the first-order condition with respect to L_t, which must equal zero in steady state.

Combining the envelope condition for N with the first-order condition for H gives:

$$\frac{F_N - w - C_N^H H - g}{\delta} = C^H + HC_H^H. \tag{A.13}$$

The envelope condition for w_{t-1} is:

$$V_{w_{t-1}} = -g_{w_{t-1}} N_t \tag{A.14}$$

which can be rolled one period forward to give the relationship $V_{w_t} = -g_{w_t} N_{t+1}$. Combining the first-order conditions for H and w and substituting into the envelope condition for w_t gives:

$$-N_t - H_t C_{w_t}^H - g_{w_t} N_t - \delta_{w_t} N_t (C_t^H + H_t C_{H_t}^H) - \beta g_{w_t} N_{t+1} = 0. \tag{A.15}$$

As mentioned above, it is useful to work with the labor cost function $\mathcal{C}(N, w)$ defined in equation (A.8), which represents the firm's per-worker hiring cost of hiring exactly enough workers to offset quits when it pays wage w. Note that, in steady state, $H = \delta(w)N$, so hires are equal to quits. Taking the derivative of $\mathcal{C}$ with respect to N and manipulating gives:

$$\begin{aligned}
\mathcal{C}_N &= \delta C_N^H + \delta C_H^H \delta \\
N\mathcal{C}_N &= \delta N C_N^H + \delta N C_H^H \delta = HC_N^H + HC_H^H \delta \\
\mathcal{C} + N\mathcal{C}_N &= C^H \delta + HC_N^H + HC_H^H \delta \\
\frac{\mathcal{C} + N\mathcal{C}_N - HC_N^H}{\delta} &= C^H + HC_H^H \\
\frac{F_N - w - HC_N^H - g}{\delta} &= \frac{\mathcal{C} + N\mathcal{C}_N - HC_N^H}{\delta} \text{ (substituting from A.13)} \\
F_N - w - g &= \mathcal{C} + N\mathcal{C}_N.
\end{aligned} \tag{A.16}$$

Taking the derivative of $\mathcal{C}$ with respect to w and manipulating gives:

$$\begin{aligned}
\mathcal{C}_w &= C_w^H \delta + C_H^H N \delta_w \delta + C^H \delta_w \\
\mathcal{C}_w &= C_w^H \delta + C_H^H H \delta + C^H \delta_w.
\end{aligned} \tag{A.17}$$

Returning to equation (A.15), using the approximation $\beta \approx 1$ and focusing on steady state gives:

$$\begin{aligned}
&- N - HC_w^H - g_w N - \delta_w N (C^H + HC_H^H) - g_w N = 0 \\
&- 1 - 2g_w = \delta C_w^H + \delta_w C^H + \delta_w HC_H^H \\
&- 1 = \mathcal{C}_w - 2g_w
\end{aligned} \tag{A.18}$$

Taking the total derivative of equation (A.18) gives:

$$\begin{aligned}
&- \mathcal{C}_{ww} dw - 2g_{ww} dw - 2g_{w\lambda} d\lambda = 0 \\
&\frac{dw}{d\lambda} = \frac{-2g_{w\lambda}}{\mathcal{C}_{ww} + 2g_{ww}}.
\end{aligned} \tag{A.19}$$

$g_{w\lambda}$ can be derived as:

$$g_{w\lambda} = -\frac{\theta}{\pi\,\left(\theta^2\,(w_{t-1} - w_t)^2 + 1\right)} < 0, \tag{A.20}$$

while g_{ww} can be derived as:

$$g_{ww} = -\frac{\lambda\,\theta^3\,\left(2\,w_{t-1} - 2\,w_t\right)}{\pi\,\left(\theta^2\,(w_{t-1} - w_t)^2 + 1\right)^2}. \tag{A.21}$$

In the neighborhood of flexible wages ($\lambda \approx 0$), $g_{ww} \approx 0$.

Using the functional forms assumed in equations (A.5) and (A.6), $\mathcal{C}_{ww}$ can be derived as:

$$\mathcal{C}_{ww} = \frac{c_h\,\bar{\delta}\,N^{\nu-\eta}\left(\frac{\bar{\delta}}{\left(\frac{w}{\bar{w}}\right)^{\gamma}}\right)^{\nu}\left(\gamma^2\,\nu^2 + 2\,\gamma^2\,\nu + \gamma^2 + 2\,\gamma\,\nu\,\phi + \gamma\,\nu + 2\,\gamma\,\phi + \gamma + \phi^2 + \phi\right)}{w^{\gamma+\phi+2}\left(\frac{1}{\bar{w}}\right)^{\gamma}} > 0. \tag{A.22}$$

It thus follows that, in the neighborhood of flexible wages, wages rise with the wage adjustment cost factor:

$$\boxed{\frac{dw}{d\lambda} > 0.} \tag{A.23}$$

Given the definition of the quit rate function in equation (A.6), it is straightforward to show that:

$$\delta_w = -\bar{\delta}\gamma\left(\frac{w}{\bar{w}}\right)^{-\gamma} < 0, \tag{A.24}$$

or in other words, that the quit rate is a decreasing function of the wage rate, as it ought to be. Then by the chain rule,

$$\boxed{\frac{d\delta}{d\lambda} = \frac{d\delta}{dw}\frac{dw}{d\lambda} < 0.} \tag{A.25}$$

To complete the steady-state analysis, note that in steady state, the hire rate $h = \frac{H}{N} = \delta(w)$, or in other words, hires offset quits to maintain a stable employment level. Therefore,

$$\boxed{\frac{dh}{d\lambda} = \frac{d\delta}{d\lambda} < 0.} \tag{A.26}$$

Therefore, in steady state, the quit rate and hire rate are decreasing with the level of wage rigidity, at least in the neighborhood of flexible wages, as seen in the quantitative model in Section 2.

In steady state, the establishment's layoff rate will be zero, so a non-steady state analysis is necessary to show that the layoff rate is increasing with the cost of adjusting wages. The plan of analysis to establish this final empirical prediction from the quantitative model in Section 2 is to show that, when wages are perfectly flexible, the layoff rate is zero in the analytical model, both in and out of steady state. If the layoff rate is positive for some positive level of the wage adjustment cost parameter, there must therefore exist a region in the neighborhood of flexible wages in which the layoff rate is increasing with the cost of adjusting wages. Furthermore, along the intensive margin, when the establishment does lay off workers, it can be shown that the layoff rate is an increasing function of the wage adjustment cost parameter λ.

It is straightforward to see that the layoff rate in the analytical model will be zero when wages are perfectly flexible. Intuitively, any time the establishment wishes to reduce its employment level, it can lower the wage rate that it pays, thereby increasing quits and reducing its wage bill at the same time that it avoids paying the cost of laying off employees. To see this result analytically, substitute the equation of motion for the employment stock into the establishment's value function and take the first-order condition with respect to layoffs L_t:[61]

$$\beta V_{N_{t+1}} = -c_l < 0. \tag{A.27}$$

When wages are flexible, however, $V_{N_{t+1}} > 0$, generating a contradiction. Recall the envelope condition for w_t from equation (A.14), $V_{w_t} = -g_{w_t} N_t$. Setting $\lambda = g = g_{w_t} = 0$ implies that $V_{w_t} = 0$; simply put, when wages are flexible, the value of the state does not depend on the wage. Then simplifying and re-arranging the first-order condition for w_t from equation (A.10) gives:

$$V_{N_{t+1}} = \frac{-1 - C^H_{w_t} \frac{H_t}{N_t}}{\delta_{w_t} \beta} > 0, \tag{A.28}$$

where the latter inequality follows because $C^H_{w_t}, \delta_{w_t} < 0$, and $H_t, N_t \geq 0$.

The result that the value function must be increasing in the stock of employment when wages are flexible is intuitive, because a higher employment stock allows the establishment to pay lower wages to maintain the same desired future employment level. This result also shows that, with flexible wages, the establishment's first-order condition to lay off workers can never be satisfied: it is always more profitable to reduce wages further instead.

The presence of wage rigidity is a necessary but not sufficient condition to produce a positive layoff rate in the model. Whether the firm will ever lay off workers will depend on the model's parameters. As $c_l \to \infty$, the layoff rate will go to zero, even in the presence of wage adjustment costs. Nonetheless, suppose that the model's parameter values are such that the establishment does lay off workers in some states. The foregoing analysis shows that if all other model parameters were held constant, but wages were perfectly flexible (i.e., λ were set to 0), the layoff rate would be zero. In such cases, the layoff rate must be an increasing function of the wage adjustment cost parameter λ in the neighborhood of flexible wages.

Finally, it can be shown that, in state where the establishment lays off workers, more rigid wages will lead to higher layoffs. Re-writing equation (A.27) outside of steady-state gives:

$$\beta E_t \left[V_{N_{t+1}} \right] = -c_l. \tag{A.29}$$

Taking a total derivative of this first-order condition yields:

$$-\beta E_t \left[\frac{\partial V_{N_{t+1}}}{\partial L_t} dL_t + \frac{\partial V_{N_{t+1}}}{\partial \delta(w_t)} \frac{\partial \delta(w_t)}{\partial w_t} \frac{\partial w_t}{\partial \lambda} d\lambda \right] = 0. \tag{A.30}$$

Re-arranging and taking terms known at time t out of the expectations operator yields:

$$E_t \left[\frac{\partial V_{N_{t+1}}}{\partial L_t} \right] dL_t = -E_t \left[\frac{\partial V_{N_{t+1}}}{\partial \delta(w_t)} \right] \frac{\partial \delta(w_t)}{\partial w_t} \frac{\partial w_t}{\partial \lambda} d\lambda. \tag{A.31}$$

[61] Note that this analysis remains in steady state, so an expectations operator is unnecessary,

Because a layoff this period reduces next period's employment level by one employee, the expected derivative of the value of a marginal employee next period with respect to layoffs this period equals the negative of the expected value of a marginal employee next period with respect to the next period's employment stock:

$$E_t\left[\frac{\partial V_{N_{t+1}}}{\partial L_t}\right] = -E_t\left[\frac{\partial V_{N_{t+1}}}{\partial N_{t+1}}\right]. \tag{A.32}$$

An increase in the quit rate reduces the next period's employment level proportionally to the current stock of employment. Therefore, by similar logic, the expected derivative of the value of a marginal employee next period with respect to the quit rate equals the current employment level times the negative of the expected value of a marginal employee next period with respect to the next period's employment stock:

$$E_t\left[\frac{\partial V_{N_{t+1}}}{\partial \delta(w_t)}\right] = -N_t E_t\left[\frac{\partial V_{N_{t+1}}}{\partial N_{t+1}}\right]. \tag{A.33}$$

Substituting these relationships into equation (A.31) and canceling the remaining expectation terms gives:

$$\boxed{\frac{d\left(\frac{L_t}{N_t}\right)}{d\lambda} = -\frac{\partial \delta(w_t)}{\partial w_t}\frac{\partial w_t}{\partial \lambda} > 0.} \tag{A.34}$$

In other words, on the intensive margin, the derivative of the establishment's layoff rate with respect to the wage adjustment cost parameter equals the negative of the derivative of the quit rate with respect to wages times the derivative of wages with respect to the cost of adjustment.[62]

B Monte Carlo Simulations of Wage Rigidity Estimator

This section tests the performance of the estimator of wage rigidity proposed in Section 4 using Monte Carlo simulations. Wage change distributions for 500 establishments facing different levels of wage rigidity are simulated, with the number of years' worth of wage changes observed in the sample for each establishment generated as a random integer uniformly distributed between three and seven. Next, the number of employees per establishment is generated as a random integer uniformly distributed between 15 and 500; the number of employees is fixed over the simulation period. For each establishment the proportion of nominal wage cuts that will be prevented by downward nominal wage rigidity is also simulated 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: $\mu_{it} \sim N(.025, .04^2)$. Then, draw the standard deviation of the counterfactual wage change distribution from a uniform distribution over the interval [0.02 .07] ($\sigma_{it} \sim U[0.02, .07]$) and draw the counterfactual flexible wage changes for each year from the normal distribution $\Delta \ln w_{ijt}^{cf} \sim N(\mu_{it}, \sigma_{it}^2)$,

[62] It is straightforward to see that equation (A.19), which shows that wages increase with the wage adjustment cost parameter, holds outside of steady state.

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.

Fraction wr_i of counterfactual wage cuts are affected by wage rigidity each period. Wage cuts are chosen to replace randomly: there is no tendency for smaller wage cuts to be more likely to be prevented. Of the wage cuts affected by wage rigidity, 98 percent are set to zero, and 2 percent are replaced with missing values to reflect the proportion of endogenously generated layoffs in the structural model.[63] Finally, compression in the wage change distribution in the face of wage rigidity is introduced by multiplying counterfactual wage changes by a compression factor 0.57.[64]

Figure B.1 displays the estimated and actual proportions of counterfactual wage cuts prevented by wage rigidity in these simulations. A regression of the form

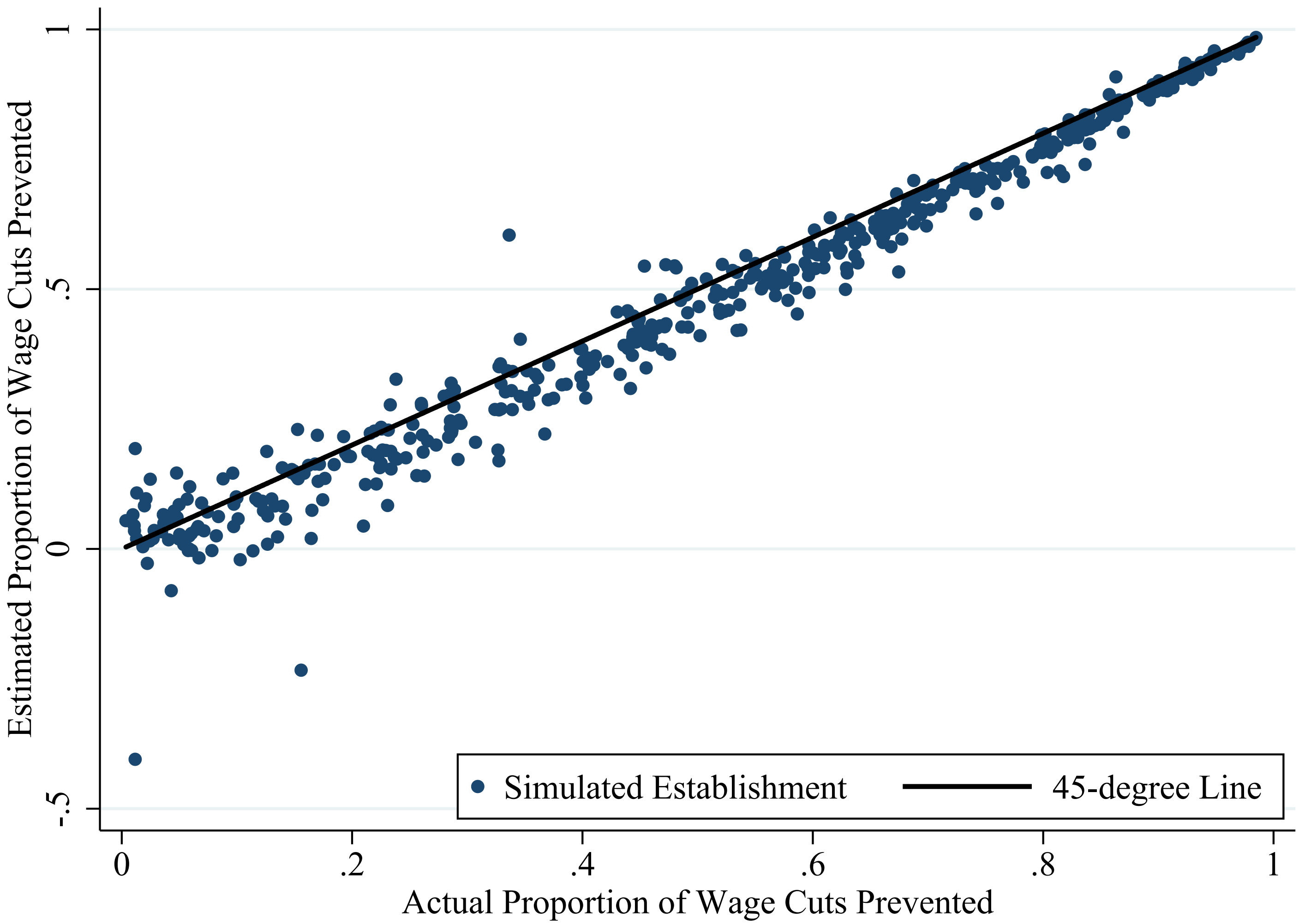
$$\widehat{wr}_i = \alpha + \beta wr_i + u_i \tag{B.35}$$

gives an estimate for $\hat{\alpha}$ of -0.03 with a standard error of 0.005 and an estimate for $\hat{\beta}$ of 1.00 with a standard error of 0.008. Therefore, estimated wage rigidity moves essentially one-for-one with true wage rigidity.

[63] A further five percent of wage changes are selected randomly from the entire counterfactual wage change distribution to reflect exogenous layoffs.

[64] The compression factor was chosen as the ratio of the simulated standard deviation of wage changes with wage rigidity and without wage rigidity in the structural model.

Figure B.1: Monte Carlo Simulations of Wage Rigidity Estimates



xi

C 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 3.A. 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 C.1: Aggregate Wage Change Distributions 1976 to 1990

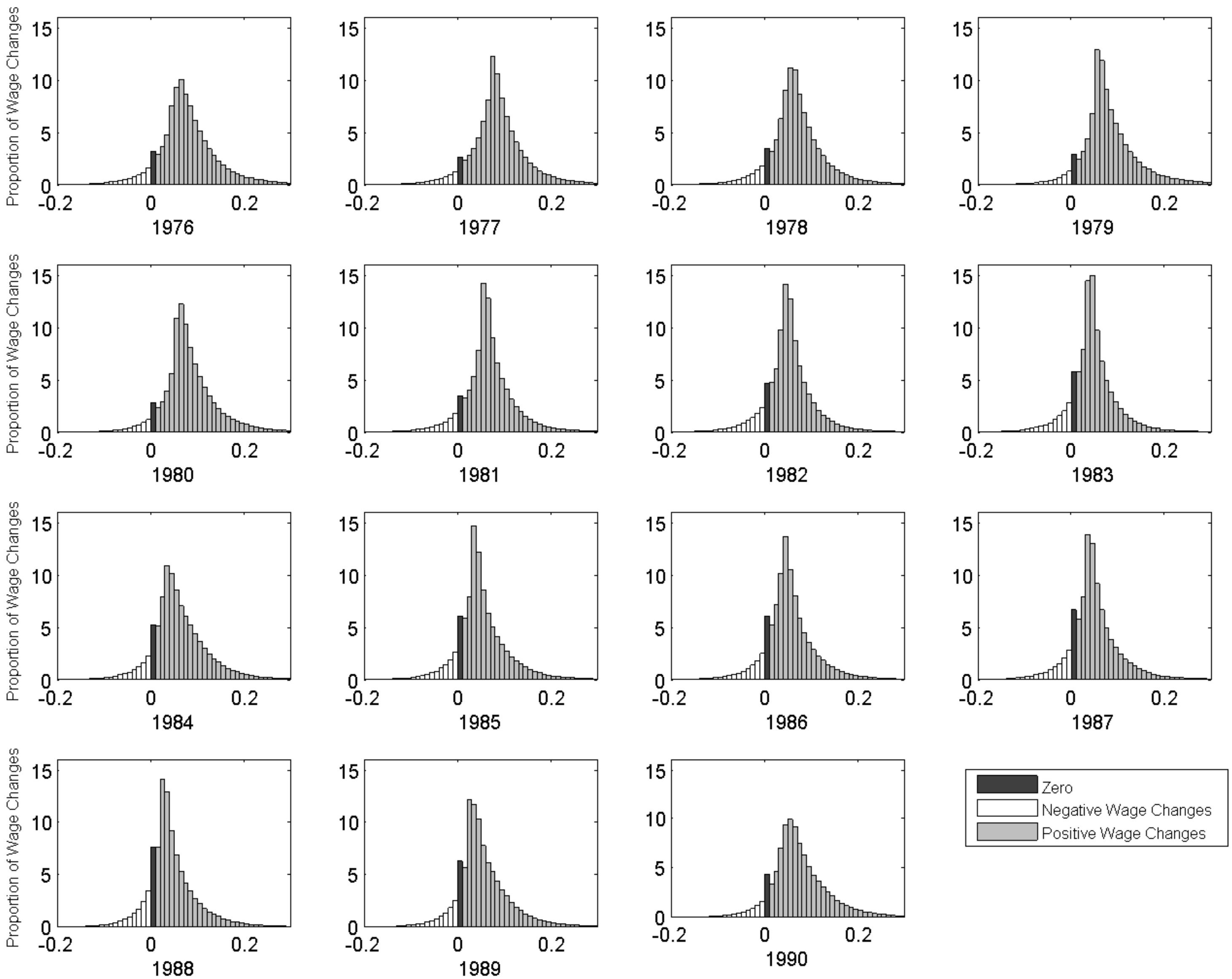
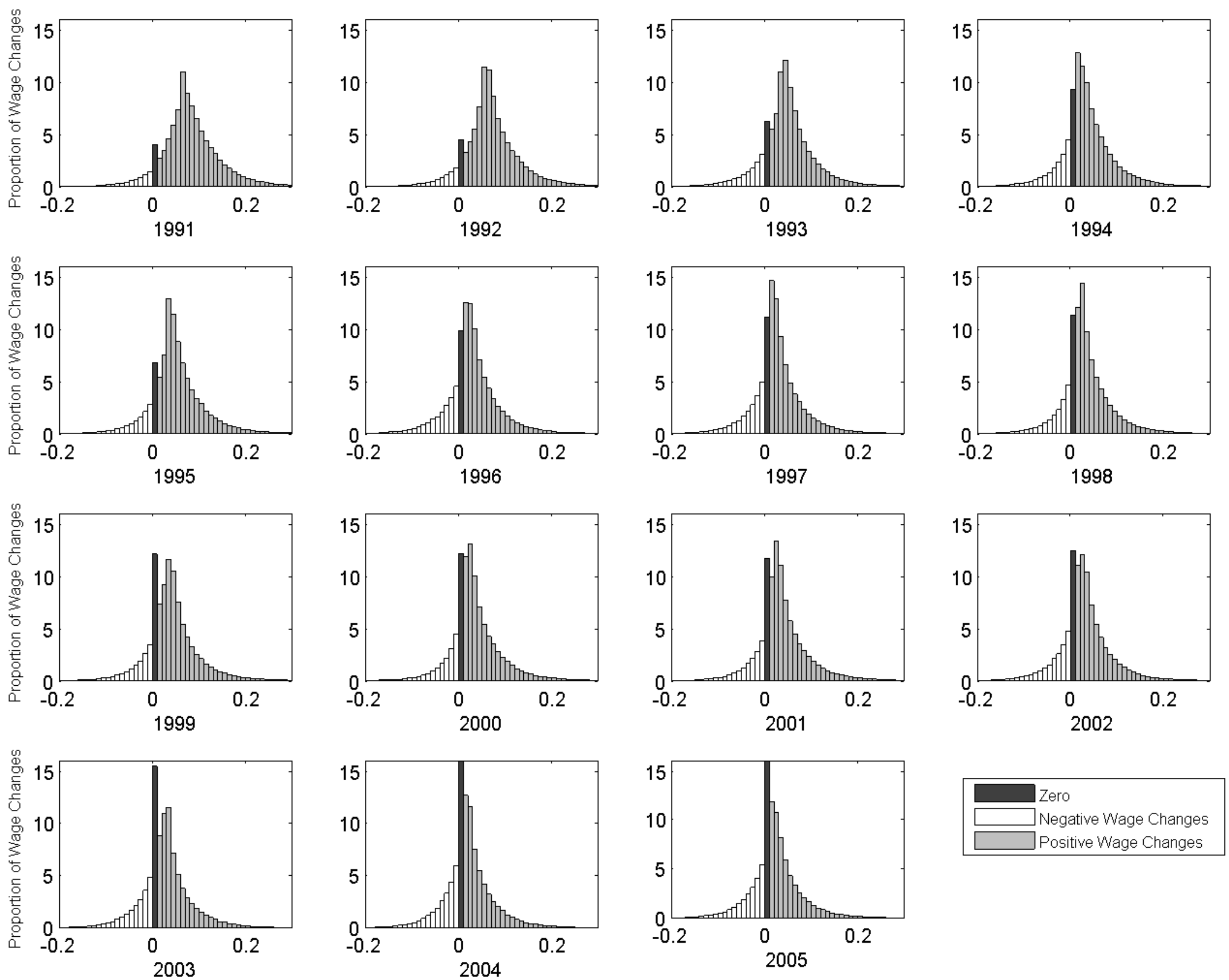


Figure C.2: Aggregate Wage Change Distributions 1991 to 2005



D Robustness

This appendix assesses the robustness of the main empirical results in two ways. First, an alternative estimator of wage rigidity is proposed, and employment flows are related to wage rigidity using the alternative measure. Second, the sensitivity of the main results to the sample selection is assessed by splitting the samples into groups of establishments with employment levels below and above the median.

D.1 An Alternative Method of Estimating Wage Rigidity

This section briefly describes an alternative method for measuring wage rigidity and the results it produces. The major difference between the method in this section and the method used in the body of the paper is that the method in this section imposes an underlying functional form for the counterfactual wage change distribution. Using a parametric estimation technique adds additional structure to the procedure and has the potential to produce more precise results in the relatively small establishment-level samples. In practice, the results of the two methods are very similar.

The estimation of the observed wage change distribution proceeds as in the fully nonparametric method used in the main results. The counterfactual wage change distribution is estimated by assuming that the distribution is symmetric about its median, with the both the upper and lower tails distributed exponentially.[65] Specifically, the counterfactual wage change distribution is assumed to be:

$$f(x) = \begin{cases} \frac{1}{2}\lambda e^{-\lambda(m-x)} & x < m \\ \frac{1}{2}\lambda e^{-\lambda(x-m)} & x \geq m \end{cases} \tag{D.36}$$

where x is a nominal wage change and m is the observed median nominal wage change.

Then for each establishment i in year t, λ may be estimated as:

$$\hat{\lambda} = \frac{n_u - 1}{n_u}(\overline{x}_u - m) \tag{D.37}$$

where n_u is the number of observations in the upper tail of the wage distribution and $\overline{x}_u$ is the sample average wage change in the upper tail of the distribution. Then the counterfactual wage change distribution $\hat{f}_{it}^{cf}$ for establishment i in year t is given by equation D.36. As in the kernel estimation approach used in the main analysis, the estimated proportion of wage cuts prevented by wage rigidity is calculated by comparing the implied proportion of counterfactual wage cuts to the number observed. One advantage of the parametric approach assumed here is that it allows estimated wage rigidity to vary year-by-year, which given the small sample sizes is impractical using the kernel density approach.

The parametric approach produces an average level of estimated wage rigidity of 0.31 across establishments, only slightly higher than the average level of 0.27 produced by the kernel density estimation procedure. The standard deviation of estimated wage rigidity is 0.22, also a bit higher than in the kernel density procedure.

[65] Dickens, Goette, Groshen, Holden, Messina, Schweitzer, Turunen, and Ward (2007) report that, "Exponential distributions provide a much better fit to wage changes above the median in our wage change distributions than do normal distributions." They find that the Weibull distribution, which generalizes the exponential distribution, provides a further improvement. For simplicity, however, this section uses the exponential distribution.

Appendix Tables D.1-D.3 display results analogous to the results in Tables 6-8, but using the parametric estimator for wage rigidity.[66] The results are qualitatively and quantitatively similar to the results from the non-parametric regressions, and indicate a substantial role for downward nominal wage rigidity in affecting employment outcomes.

D.2 Split Sample Analysis

As described in Section 3.C, establishment-years are excluded if the establishment has fewer than 30 employees or 5 valid wage changes in the year. Establishments are excluded altogether if they have fewer than 20 valid wages changes over all of the years they are in the data. These restrictions exclude very small establishments from the sample; the mean establishment size in the main analysis is 466 employees, with a median of 168. To examine whether the main results are driven entirely by large establishments, this section splits the sample in the main analysis into establishment-years with employment levels above the sample median and establishment-years with employment levels lower than or equal to the sample median.

Tables D.4 and D.7 analyze the layoff results for large and small establishments, respectively. The point estimates for the wage rigidity coefficients in Table D.4, which contains large establishments, are larger than in the full sample, although the standard errors are larger also, reflecting the smaller sample sizes. Conversely, the point estimates for the wage rigidity coefficients in Table D.7, which contains small establishments, are smaller than in the full sample. These results suggest that wage rigidity has a larger effect on increasing layoffs in larger establishments.

Tables D.5 and D.8 analyze the quits results for large and small establishments, respectively. The estimates largely mirror the results for the full sample.

Tables D.6 and D.9 analyze the hires results for large and small establishments, respectively. The standard errors are large for all regressions, once again reflecting the smaller sample sizes from splitting the sample. However, both the OLS and IV results have statistically significant estimated coefficients on wage rigidity for the smaller establishments, suggesting that wage rigidity more clearly affects hires for smaller establishments than for larger establishments. This result stands in contrast to the results for layoffs above, which indicate that wage rigidity more clearly affect layoffs at the larger establishments.

Columns (2) and (4) of Tables D.4 through D.6 replicate columns (3) and (6), respectively, of Tables 6 through 8 in the main text. They are included here for completeness and ease of comparison.

D.3 Pre-Period Instrument

This section describes an alternative establishment-level instrument for downard nominal wage rigidity. The establishment-level instrument described in Section 5 and used in Section 6 relies on one-year lagged establishment participation in colective bargaining agreements. In principle,

[66]The first stage results of the instrumental variables regressions are omitted for brevity but show a strong relationship between the instrument and measured wage rigidity and are quantitatively similar to the first stage results for the nonparametric estimator, with first-stage coefficients on the instrument of approximately 0.2. Additionally, the parametric assumptions in this estimator allow for the inclusion of smaller establishments. The regressions include establishment-years with a minimum of 20 wage change observations.

it is possible that one-year lagged participation in a collective bargaining agreement could correlate with contemporaneous business conditions, which could in turn affect employment flows by channels other than wage rigidity, thus violating the exclusion restriction.

The alternative instrument considered in this section uses establishment participation in collective bargaining agreements in the years 1993, 1995, and 1996, prior to the main analysis period of 1997–2003.[67] The sparsity of the survey responses prior to the main analysis period motivates the use of a "waterfall technique" to construct the instrument: if an establishment has a valid answer for participation in a collective bargaining agreement for 1996, that data is used for the instrument. If an establishment has no valid answer for 1996, but does have an answer for 1995, that answer is used. Finally, if an establishment has no valid answer for 1995 or 1996, but does have an answer for 1993, that answer is used. A limitation of this "pre-period" instrument is that the sample of establishments with valid survey responses is roughly half as large as the main sample.

Table D.10 presents instrumental variables regression results in which this alternative pre-period instrument is used in place of the one-year lagged participation in a collective bargaining agreement instrument. Therefore, column (1) of Table D.10 corresponds to column (7) of Table 6, showing results for a regression of layoffs on wage rigidity using the state-sector level "proportion paying at the wage floor" instrument along with the establishment-level "pre-period collective bargaining" instrument. Columns (2) and (3) likewise correspond to column (7) in Tables 7 and 8, showing results for quits and hires, respectively.

The estimated coefficient on wage rigidity in column (1) of Table D.10, showing results for layoffs, is in line with the instrumental variable results in columns (5) through (8) of Table 6. The standard error is larger than in the parallel specification in column (7) of Table 6, consistent with the smaller sample. Nonetheless, this result suggests that using an instrument that is less likely to be correlated with contemporaneous business conditions does not change the qualitative results.

Columns (2) and (3) of Table D.10 display similar patterns. The results are largely in line with the instrumental variables regression results in Tables 7 and 8, although the standard errors are larger.

Therefore, this analysis suggests that using one-year lagged participation in a collective bargaining agreement, as in column (7) of Tables 6–8, is unlikely to introduce confounding through correlation with contemporaneous business conditions.

[67]The IAB Establishment Panel employer survey did not ask about participation in collective bargaining agreements for the year 1994.

Table D.1: Wage Rigidity and Layoffs – Regression Results (Parametric Estimator)

Dependent Variable	Layoff Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.009	-0.008	0.119**	0.113*
	(0.019)	(0.018)	(0.060)	(0.060)
Positive Revenue Growth		0.015***		0.016***
		(0.005)		(0.006)
Negative Revenue Growth		-0.071***		-0.074***
		(0.014)		(0.014)
Works Council	-0.040***	-0.041***	-0.046***	-0.046***
	(0.009)	(0.009)	(0.009)	(0.009)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.007	0.010
R-Squared	0.139	0.143	0.105	0.113
N	12,064	12,064	12,064	12,064

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Layoffs are defined as the fraction of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated using a parametric estimator assuming an exponential distribution as described in Section D.1 and varies year-by-year. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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 have at least 20 employees in a given year and cover the period 1997 to 2003. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.2: Wage Rigidity and Quits – Regression Results (Parametric Estimator)

Dependent Variable	Quit Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.089***	-0.088***	-0.206***	-0.213***
	(0.015)	(0.015)	(0.046)	(0.047)
Positive Revenue Growth		0.039***		0.039***
		(0.008)		(0.008)
Negative Revenue Growth		-0.064***		-0.061***
		(0.016)		(0.015)
Works Council	-0.048***	-0.048***	-0.043***	-0.042***
	(0.007)	(0.007)	(0.007)	(0.007)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.002	0.001
R-Squared	0.136	0.140	0.120	0.121
N	12,064	12,064	12,064	12,064

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Quits are defined as the fraction of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated using a parametric estimator assuming an exponential distribution as described in Section D.1 and varies year-by-year. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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 have at least 20 employees in a given year and cover the period 1997 to 2003. The instrumental variable regressions in columns 3 and 4. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.3: Wage Rigidity and Hires – Regression Results (Parametric Estimator)

Dependent Variable	Hire Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.113***	-0.112***	-0.115	-0.126
	(0.033)	(0.033)	(0.095)	(0.094)
Positive Revenue Growth		0.125***		0.125***
		(0.021)		(0.021)
Negative Revenue Growth		0.002		0.002
		(0.029)		(0.029)
Works Council	-0.113***	-0.112***	-0.113***	-0.112***
	(0.015)	(0.015)	(0.016)	(0.016)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.983	0.866
R-Squared	0.106	0.113	0.106	0.113
N	12,064	12,064	12,064	12,064

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Hires are defined as the fraction of the establishment's total workforce on December 31 of the previous year. Wage rigidity is calculated using a parametric estimator assuming an exponential distribution as described in Section D.1 and varies year-by-year. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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 have at least 20 employees in a given year and cover the period 1997 to 2003. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.4: Wage Rigidity and Layoffs – Regression Results (Large Establishments)

Dependent Variable	Layoff Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	0.032	0.033	0.196*	0.189*
	(0.040)	(0.040)	(0.102)	(0.101)
Positive Revenue Growth		0.022**		0.023**
		(0.009)		(0.010)
Negative Revenue Growth		-0.078***		-0.082***
		(0.018)		(0.017)
Works Council	-0.092***	-0.092***	-0.096***	-0.096***
	(0.030)	(0.030)	(0.030)	(0.030)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.046	0.056
R-Squared	0.247	0.251	0.211	0.219
N	5,439	5,439	5,439	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Layoffs are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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. The regression sample is the same as in Table 6, but only establishments with an average number of employees above the sample median are included. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.5: Wage Rigidity and Quits – Regression Results (Large Establishments)

Dependent Variable	Quit Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.113***	-0.112***	-0.260***	-0.266***
	(0.029)	(0.029)	(0.081)	(0.082)
Positive Revenue Growth		0.052***		0.051***
		(0.014)		(0.014)
Negative Revenue Growth		-0.062**		-0.058**
		(0.024)		(0.024)
Works Council	-0.055***	-0.054***	-0.052***	-0.051***
	(0.015)	(0.014)	(0.014)	(0.014)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.024	0.018
R-Squared	0.236	0.241	0.218	0.221
N	5,439	5,439	5,439	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Quits are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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. The regression sample is the same as in Table 7, but only establishments with an average number of employees above the sample median are included. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.6: Wage Rigidity and Hires – Regression Results (Large Establishments)

Dependent Variable	Hire Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.09	-0.088	-0.043	-0.051
	(0.062)	(0.062)	(0.172)	(0.171)
Positive Revenue Growth		0.150***		0.151***
		(0.029)		(0.029)
Negative Revenue Growth		-0.012		-0.013
		(0.031)		(0.031)
Wage Rigidity x Positive Revenue Growth				
Wage Rigidity x Negative Revenue Growth				
Works Council	-0.170***	-0.164***	-0.171***	-0.165***
	(0.044)	(0.043)	(0.044)	(0.044)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.753	0.797
R-Squared	0.318	0.326	0.318	0.325
N	5,439	5,439	5,439	5,439

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Hires are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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. The regression sample is the same as in Table 8, but only establishments with an average number of employees above the sample median are included. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.7: Wage Rigidity and Layoffs – Regression Results (Small Establishments)

Dependent Variable	Layoff Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.026*	-0.025*	0.062	0.051
	(0.014)	(0.014)	(0.073)	(0.071)
Positive Revenue Growth		0.009		0.009
		(0.006)		(0.006)
Negative Revenue Growth		-0.080***		-0.082***
		(0.017)		(0.018)
Works Council	-0.027***	-0.028***	-0.031***	-0.031***
	(0.005)	(0.005)	(0.006)	(0.006)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.169	0.230
R-Squared	0.124	0.13	0.103	0.115
N	5,467	5,467	5,467	5,467

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Layoffs are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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. The regression sample is the same as in Table 6, but only establishments with an average number of employees below the sample median are included. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.8: Wage Rigidity and Quits – Regression Results (Small Establishments)

Dependent Variable	Quit Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.049***	-0.048***	-0.228***	-0.240***
	(0.017)	(0.017)	(0.056)	(0.056)
Positive Revenue Growth		0.036***		0.035***
		(0.010)		(0.011)
Negative Revenue Growth		-0.071***		-0.065***
		(0.024)		(0.024)
Works Council	-0.052***	-0.052***	-0.046***	-0.046***
	(0.008)	(0.007)	(0.008)	(0.008)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.002	0.001
R-Squared	0.142	0.146	0.099	0.097
N	5,467	5,467	5,467	5,467

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Quits are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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. The regression sample is the same as in Table 7, but only establishments with an average number of employees below the sample median are included. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.9: Wage Rigidity and Hires – Regression Results (Small Establishments)

Dependent Variable	Hire Rate as a Fraction of Establishment Workforce			
	(1)	(2)	(3)	(4)
Wage Rigidity	-0.070**	-0.069**	-0.159	-0.174*
	(0.030)	(0.031)	(0.105)	(0.104)
Positive Revenue Growth		0.131***		0.130***
		(0.037)		(0.037)
Negative Revenue Growth		-0.005		-0.002
		(0.046)		(0.045)
Works Council	-0.117***	-0.117***	-0.114***	-0.114***
	(0.015)	(0.014)	(0.015)	(0.015)
Specification	OLS	OLS	IV	IV
P-value of Exogeneity Test			0.399	0.322
R-Squared	0.118	0.126	0.116	0.123
N	5,467	5,467	5,467	5,467

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Hires are defined as the fraction 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, 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 federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. 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. The regression sample is the same as in Table 8, but only establishments with an average number of employees below the sample median are included. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. The instrumental variable regressions in columns 3 and 4 are estimated by two-stage least squares. Test of regressor exogeneity is from Wooldridge (1995). One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

Table D.10: Wage Rigidity and Employment Flows – IV Regression Results with Pre-Period Instrument

Dependent Variable	Layoffs	Quits	Hires
	(1)	(2)	(3)
Wage Rigidity	0.138	-0.239***	-0.203
	(0.087)	(0.075)	(0.155)
Works Council	-0.046***	-0.043***	-0.111***
	(0.015)	(0.011)	(0.029)
Specification	IV	IV	IV
R-Squared	0.093	0.104	0.091
N	5,748	5,748	5,748

Standard errors, clustered at the establishment level, are in parentheses. The unit of observation is the establishment-year. Layoffs, Quits, and Hires are defined as fractions 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 the sample period. Each regression includes a set of establishment characteristics, individual characteristics, and year dummies as controls. Establishment characteristics include a set of controls for occupational mix, dummies for federal state, establishment size, and large-scale relocations of workers across establishments within the same firm. Individual characteristics include controls for gender and workers' education. All regressions have at least 30 employees in a given year and cover the period 1997 to 2003. The instrumental variables in the regressions are estimated by two-stage least squares. These regressions use two instruments for wage rigidity. The first instrument uses the state-sector level wage agreements as described in Section 5.A. The second instrument is a dummy variable for whether or not an establishment was bound by a collective bargaining agreement prior to the beginning of the sample period, as described in section D.3. One star, two stars, and three stars denote statistical significance at the 10-, 5-, and 1-percent confidence levels, respectively.

E Computational Appendix

The computational approach in the paper is to solve the establishment's optimal policy functions on a discrete grid using standard value function iteration techniques for a given set of parameters. Simulated data is generated from those policy functions using a set of random productivity shocks. The parameters of the model are estimated via an indirect inference procedure by matching moments of the simulated data to corresponding empirical target moments to minimize a quadratic objective function. Formally, let θ be a vector of the structural parameters to be estimated and μ be a vector of the 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 a diagonal weighting matrix with the squared target moments as its entries. The minimization is conducted using the Matlab routine "fminsearch."

The processes for the establishment-wide and worker-type productivity levels z and u are discretized using the method of Tauchen (1986) to have 5 nodes each. Those grids are then interpolated to have 25 points each for the purposes of model simulation. The employment space is discretized to have 35 log-linearly spaced nodes, which are also interpolated for the purposes of simulation to have 150 nodes. The wage space is discretized to have 225 log-linearly spaced nodes, which are interpolated to 600 nodes for the purposes of simulation. It is necessary to have a finer grid for the wage space than for the employment space given the model's focus on estimating wage rigidity accurately and matching the wage change distribution. The wage rigidity estimation procedure uses bins with a width of 0.25 percent. The value function iteration proceeds until the value functions have converged to within a tolerance of 0.05 percent.

The simulations uses 600 periods and drops the first 200 periods to allow for burn-in. The initial employment, wage, and productivity levels all start at the values associated with the middle node in the grid (in the case of an odd number of nodes, the node number is rounded down). The burn-in period is sufficiently long and the shocks are sufficiently large and transitory that these initial conditions do not influence the simulated data after the burn-in period is dropped. There are 75 types of workers within the establishment, each with its own type-specific productivity process u_j. The establishment-wide productivity process z applies to all types.

The simulated moments and auxiliary models used in the indirect inference procedure are constructed to mimic the empirical targets as closely as possible, but there are some minor deviations. For instance, revenues in the model are matched to value added in the empirical data to account for intermediate inputs and external costs, which are absent from the model. Additionally, residualized values are used in the empirical data in some regressions as described in the text to correct for heterogeneity that is not present in the model. The establishment's policy functions are calculated and model data is simulated eight times for each guess of the other parameters, by multiplying the wage rigidity parameters by a multiplicative factor increasing linearly from 0 to 2.25, at which level wage rigidity is estimated to be nearly one. The simulated measure of wage rigidity is then calculated along with the simulated layoff, quit, and hire rates for each of the eight simulations. Regressing the simulated employment flows on the simulated wage rigidity measures gives the three regression parameters that are matched to the regressions results from Section 6.

F Target Moments

Ten target moments are estimates from auxiliary models. First, the wage rigidity estimator described in Section 4 is applied to the simulated wage change distributions. To the extent that estimator is misspecified, using an identical estimation procedure on the simulated data will correct for the misspecification. One example of possible misspecification comes from Elsby (2009), who notes that forward-looking wage setting establishments may compress wage increases in the presence of wage rigidity. That effect is embedded in the dynamic nature of the model in Section 2. Although there is not a one-to-one correspondence between the target moments and the estimated parameters, in practice, the cost of wage cut parameters λ_0 and λ_1 directly influence estimated wage rigidity.

Second, and relatedly, the simulated layoff, quit, and hire rates are regressed on the simulated estimates of wage rigidity to estimate the regression coefficients on wage rigidity from the simulated data. The corresponding empirical moments are the coefficients on measured rigidity from the IV estimates in column (4) in Tables 6 and 7 for layoffs and quits and the coefficient from the OLS estimate in column (1) from Table 8 for hires.[68]

A third auxiliary model targets the parameter γ, the elasticity of the quit rate with respect to wages in equation (3). 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, $\overline{w}$:

$$\delta(w) - \overline{\delta} \approx -\overline{\delta}\gamma\left(\frac{w-\overline{w}}{\overline{w}}\right) \tag{F.38}$$

Equation (F.38) 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.[69] Taking equation (F.38) to the data requires accounting for worker and establishment heterogeneity that is not present in the theoretical model.[70] A Mincer regression of individual log wages on worker and establishment observable characteristics allows for the removal of observable heterogeneity.[71] 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 term $\left(\frac{w-\overline{w}}{\overline{w}}\right)$ in equation (F.38).

To estimate equation (F.38), 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 $\hat{\gamma}$ 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 empirically estimated $\hat{\gamma}$ is

[68] The choice of IV versus OLS estimates is guided by the exogeneity tests described in Section 6.

[69] Although the quit rate function in equation 3 is linear in logs, estimating the equation in logs is not feasible because quits are zero in some establishment-years.

[70] 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.

[71] The covariates included in the Mincer regression are a set of dummies for level of education, a quartic polynomial in work experience, and a dummy for female workers.

1.01.

A fourth auxiliary targets the parameters $\overline{c}$, ϕ, ν, and η. Given that hiring costs are not observed, the hiring cost function (4) cannot be estimated directly in the data. Re-arranging the equation to express it in terms of new hires h yields the following function:

$$h = \left(\frac{c_h}{\overline{c}}\right)^{\frac{1}{\upsilon}} w^{\frac{\phi}{\upsilon}} n_{-1}^{\frac{\eta}{\upsilon}} \tag{F.39}$$

Taking logs and re-arranging further gives:

$$\ln h = -\frac{1}{\upsilon}\ln\overline{c} + \frac{\phi}{\upsilon}\ln w + \frac{\eta}{\upsilon}\ln n_{-1} + \frac{1}{\upsilon}\ln c_h \tag{F.40}$$

Equation (F.40) has four terms, which bear on the four structural parameters of the hiring cost function. The following equation can be estimated in the data to recover moments that identify those four parameters:

$$\ln h = \zeta_0 + \zeta_1 \ln w + \zeta_2 \ln n_{-1} + \xi. \tag{F.41}$$

To estimate equation (F.41), establishment-year hire rates were regressed on the same Mincer residuals described above, lagged establishment-year employment levels, and a set of establishment and year fixed effects. The inclusion of establishment fixed effects again identifies the parameters off of time series variation in wages and employment levels within establishments, rather than cross-sectional variation across establishments. Three empirical moments from equation (F.41), $\hat{\zeta}_1$, $\hat{\zeta}_2$, and $\hat{\sigma}_\xi$, were used in the indirect inference procedure.[72] $\hat{\zeta}_0$ was not included because of the establishment fixed effects; instead, the average hire rate was included as a target moment. $\hat{\zeta}_1$ bears on ϕ and υ, $\hat{\zeta}_2$ bears on η and υ, and $\hat{\sigma}_\xi$ bears on υ. In principle, ζ_0 in equation (F.41) could be used to help identify $\overline{c}$, but it cannot be estimated due to the inclusion of establishment fixed effects. Instead, the average hire rate helps to identify the hiring cost scale parameter $\overline{c}$.

Hiring costs are absorbed into the error term ξ in equation (F.41), and they are not observed in the empirical data. Targeting $\hat{\sigma}_\xi$ helps to discipline the variance of hiring costs. One concern that could arise from the use of equation (F.41) is that hiring costs, and therefore the residual ξ, are a function of wages and employment levels, leading to endogeneity in equation (F.41). However, that same endogeneity is present in the model when running the analogous regression on the simulated model data. The indirect inference procedure should therefore recover the true parameter estimates even in the presence of this endogeneity (Gourieroux et al., 1993).

The final auxiliary model targets the persistence, ψ_z, and variance, σ_z^2, of establishment-wide productivity shocks. In the empirical data, log value added per worker at the establishment-year level, $\ln(\text{VA}_{it})$, is regressed on its own lag and a set of establishment fixed effects.[73] The $AR(1)$ coefficient and the root mean-squared error from the regression are taken as target moments in the indirect inference procedure. The theoretical model abstracts from intermediate inputs, so there is no distinction between establishment revenues and establishment value added. Empirically, however, the establishment's value added is calculated as total revenues minus intermediate inputs

[72] They are labeled 'Empirical Wage Elasticity of the Hire Rate,' 'Empirical Economies of Scale of the Hire Rate,' and 'RMSE of the Empirical Hire Rate Elasticity Regression' in Table 9.

[73] More specifically, residualized log value added per worker is used, where the residualized values are from a regression of log value added per worker on a set of year, sector, and establishment size dummies, the proportion of female workers, and a set of establishment fixed effects.

and external costs.

In addition to the estimates from the auxiliary models, several descriptive statistics from the empirical data are used as target moments. First is real average value added per worker, which bears directly on the average productivity level $\overline{z}$. Second is labor's empirical share of value added, calculated as the average total wage bill divided by establishment value added. The empirical labor share bears closely on the returns-to-scale parameter α. Third is the real average daily wage, which bears on several model parameters.

Fourth is a set of statistics that describes the shape of the wage change distribution: the standard deviation of percentage wage changes, and the differences between the 10th and 50th percentiles of the distribution, the 25th and 50th percentiles, the 75th and 50th percentiles, and the 90th and 50th percentiles. Targeting these moments ensures that the distribution of shocks and frictions in the model combine to generate a realistic wage change distribution separately from the summary wage rigidity measure.

Fifth and finally, the average layoff rate is included as a target moment. The average layoff rate bears closely on the cost of firing parameter c_ℓ and the exogenous separation rate s_x.