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MINIMUM WAGES AND WORKER RETENTION: THE ROLES OF LABOUR CONCENTRATION AND EMPLOYMENT PROTECTION

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Minimum Wages and Worker Retention: The Roles of Labour Concentration and Employment Protection*

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Abstract

We examine how minimum wages affect incumbent workers' retention and how this relationship varies with labour market concentration and employment protection. Using 2002-18 matched employer-employee data from Portugal, we construct worker-year-specific wage bites incorporating both statutory and collectively bargained wage floors. Larger bites are associated with reductions in retention for fixed-term workers but virtually no effects for permanent workers. However, labour concentration attenuates the response for fixed-term workers by approximately 60%. A model of oligopsony with match-specific productivity shocks and firing costs rationalises these patterns: higher concentration leads to larger markdowns which cushion incumbent matches against higher wage floors, while employment protection limits adjustment through separations.

JEL Classifications: J23, J38, J42, J63

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1 Introduction

What effect do minimum wages have on employee retention? The answer to this question may depend on the labour market power of each firm. In competitive markets, retention will likely fall with minimum wage increases. Where firms have labour market power, however, the outcome is more nuanced: firms that, despite a minimum wage hike, still pay wages below marginal product will retain their incumbents. On the other hand, if the minimum wage rise puts workers onto the firm's labour demand schedule then we have the competitive outcome and layoffs are expected (see, for example, Azar et al. 2024). On the worker side, theoretical models of on-the-job search and job ladders strongly indicate that low-paid workers are less likely to quit following a rise in the minimum wage - a consequence of the compression of outside offers at the lower end of the pay scale (Burdett & Mortensen 1998, Mortensen 2000). Finally, labour market regulations can have a role to play as they affect the cost of layoffs. Overall, the relationship between incumbent worker retention, minimum wages and labour market power is essentially an empirical issue - one that we explore in this paper.

Specifically, our empirical analysis draws on a comprehensive matched employer-employee panel for Portugal, allowing us to monitor individual workers' retention in each firm and year following minimum wage increases. We compare each worker's wage to the applicable minimum wage of the following year, from 2002 to 2018, creating a *wagebite* measure for each worker. Moreover, we explore whether the effects on retention vary depending on a proxy for the firm's labour market power. This proxy is the worker's Herfindahl-Hirschman Index, computed from all establishment recruitments in the relevant occupation and local labour market in each year.

We also pay particular attention to the institutional context. Similarly to many other countries, Portugal exhibits significant labour market duality or segmentation, in this case between open-ended and fixed-term contracts.[1] Open-ended contracts are subject to much greater legal protection against dismissals, even in the face of steep minimum wage increases. Another important institutional dimension concerns collective bargaining: again as in some other countries, minimum wages in Portugal have both national and collective bargaining dimensions. In a novel contribution, we consider both types of minimum wages and compare their effects. Moreover, rather than focus on a single low wage sector as is often the case in minimum wage studies, we exploit our full cover-

[1] In less developed countries, labour market duality is typically expressed between formal and informal jobs.

age of (private-sector) employment to compare workers that are differentially exposed to minimum wage increases.

We find that employment protection plays a key role in the effect of minimum wages on exit hazards of workers whose wages are directly affected: workers under permanent contracts are almost completely insulated whilst workers in temporary (fixed-term) contracts are significantly exposed to higher exit hazards. Labour market concentration significantly reduces the negative effects of minimum-wage increases on temporary workers but not to the point of increasing their retention chances. Collective bargaining minimum wages tend to have similar effects on exit hazards as their national minimum-wage (NMW) counterparts. Although our data do not directly identify quits from layoffs, if - as most theories suggest - quits of low paid workers do not rise following minimum wage hikes then our estimates of the effects of minimum wages on exit hazards may be interpreted as a lower bound on firms' layoff response.

We view our empirical findings through the lens of an extension of Azar et al.'s 2024 model of oligopsonistic competition, adding idiosyncratic shocks to match productivity. The model accounts for the principal features of our estimates: a retention response that is negative throughout, attenuated by concentration but not eliminated even at high levels of concentration. Where the minimum wage is high enough so that the firm is on its labour demand schedule, employment contracts with further wage rises (and retentions fall) despite monopsony power. Where the floor is lower, the firm remains constrained by labour supply but pays below marginal revenue product; that markdown acts as a buffer, allowing it to retain workers who have received adverse shocks. A higher wage floor erodes the buffer, so incumbents are lost even when total firm employment rises. At higher levels of concentration, however, there are fewer contracting firms (i.e., those constrained by demand), and those constrained by labour supply have larger buffers leading to fewer incumbents being lost due to idiosyncratic shocks.

The remainder of the paper is as follows: Section 2 introduces the institutional setting and the data; Section 3 presents our empirical specification and results, including robustness checks; Section 4 presents our theory; and Section 5 concludes.

2 Institutions and Data

2.1 Institutional Setting

In Portugal, minimum wages are of two types: national statutory minimum wages and collective bargaining wage floors. The former defines a single economy-wide monthly minimum wage, updated in almost all years. The latter are set by collective bargaining agreements (CBAs) and are mainly negotiated through sectoral bargaining. These agreements set minimum wages at the job title level and specify additional working conditions. Initially, the agreements' conditions apply only to the signatories and their members (i.e., union members and firms affiliated with the employer's association); however, extensions to non-signatory parties are commonly granted, thereby increasing the agreement's coverage to all workers in the relevant industry (Martins 2021). The wage floors set by CBAs are usually updated annually through wage-table revisions, although the broader agreement may remain active longer. In cases where the CBA wage floor is below the statutory NMW, the statutory minimum wage is the binding one.[2]

Another important aspect of the Portuguese labour market is the duality of its employment protection legislation. In contrast to permanent (open-ended) contracts, temporary (fixed-term) contracts may be used only to meet temporary company needs, for a period typically not exceeding three years. (When the maximum duration is reached, the firm must convert the worker's contract to a permanent one to retain the worker.) Temporary contracts are also subject to greater flexibility with respect to dismissals, therefore allowing employers to adjust workforce levels at lower cost. Legal uncertainty and costs, as well as the difficulty of proving the validity of dismissal grounds, are other challenges associated with layoffs of workers on permanent contracts. These factors may be particularly important for firms in light of the stringent employment protection of permanent contracts, leading to one of the largest shares of temporary contracts in Europe, at around 25%.[3]

[2]Figure C1 presents the annual share of employees whose relevant (i.e., higher) minimum wage is a collective bargaining wage floor. The share is generally high across most of the sample, indicating that a large fraction of workers are covered by CBA wage floors that exceed the NMW. The decreasing trend may be driven by large NMW increases in recent years and the growing share of workers not covered by CBAs.

[3]See Martins (2009), Cahuc et al. (2026) and Pereira et al. (2025) for more on the employment law, temporary contracts, and collective bargaining in Portugal.

2.2 Data

Our study is based on QP (*Quadros de Pessoal*), a comprehensive matched employer-employee longitudinal dataset. It spans the universe of private-sector firms in Portugal with at least one paid employee. The dataset derives from an annual mandatory survey conducted by the Ministry of Employment to enforce compliance with employment law and collective bargaining. The information refers to October of each year. It does not include public servants, the self-employed or the unemployed. The data contain (a) individual information on each of the firm's employees; (b) information on each firm's establishments; (c) firm-level characteristics; and (d) time-invariant firm, establishment and worker identifiers. For our purposes, we use the base wage, contract type and occupational category, among other worker-level variables, from 2002 to 2018. Additionally, we link QP with SCIE (*Sistema de Contas Integradas das Empresas*), a firm-level accounting database, to obtain firm value added at factor costs. The QP data have been widely used in the literature, including in the monopsony (Bassanini et al. 2026, Martins & Melo 2024) and collective bargaining (Card & Cardoso 2022) contexts.

2.2.1 Labour Market Concentration

We proxy labour concentration following the literature, using a standard Herfindahl-Hirschman Index (HHI), as in Azar et al. (2020):

$$HHI_{mt} = \sum_{j=1}^{N_{mt}} s_{j(mt)}^2,$$

where $s_{j(mt)}$ is the ratio of new hires by firm j in local labour market m and at year t to the total number of new hires in the same market-year pair. A local labour market m is defined by the pair of a 5-digit occupation code and municipality. Therefore, $s_{j(mt)}$ can be interpreted as the labour market share of firm j in market m at time t. N_{mt} represents the number of firms hiring in that local labour market m at year t. The HHI ranges from 0 to 1, with 1 a monopsonistic local labour market and approaching zero (competitive local labour market) as the number of employers grows. Moreover, the HHI time series (Figure C2) is stable across our sample period, with concentration values between 0.2 and 0.3. Labour concentration levels by the type of contract are similar, with workers on temporary contracts experiencing slightly higher labour concentration levels.

2.2.2 Collective Bargaining Minimum Wages

We derive collective bargaining minimum wages using the modal base wage approach (Cardoso & Portugal 2005), an accurate proxy for the wage floors set by collective bargaining agreements in Portugal. The approach works as follows: within each Collective Bargaining Agreement × Professional Category × Year cell, we calculate the modal base wage earned and attribute it to all workers in that cell. For each cell, we use the modal $t+1$ base wage among workers who remain with their firm and assign that mode to all workers —stayers and leavers— in the corresponding cell in year t.[4]

2.2.3 Minimum Wage Bite

A key variable in our analysis is "*wagebite*", the wage increase required by law, following any changes to the national or collective bargaining minimum wages. We first define the binding wage floor as:

$$\bar{w}_{i,t+1} = \max\left\{NMW_{t+1},\, CBMW_{i,t+1}\right\},$$

where NMW_{t+1} is the real hourly NMW in year $t+1$, and $CBMW_{i,t+1}$ is the real hourly collective bargaining minimum wage for worker i in year $t+1$. We then define the "*wagebite*" as:

$$wagebite_{it} = \max\left\{0,\ \log \bar{w}_{i,t+1} - \log w_{it}\right\}$$

where w_{it} is the real hourly base wage of the same incumbent i in the current year t. In words, *wagebite* represents the minimum increase in the log wage the firm must pay by law to retain incumbent workers. Workers with a positive *wagebite* are referred to as "bitten", while the "unbitten" incumbents have a *wagebite* of zero. Similar *wagebite* definitions have previously been used in the literature (Card & Krueger 1994, Cengiz et al. 2019). However, to the best of our knowledge, this work is the first to consider collective bargaining wage floors when calculating wage bites, thereby enabling us to precisely define each incumbent's relevant minimum wage in settings in which collective bargaining is an important driver of wage floors.

[4]Because workers who leave the firm between t and $t+1$ may switch collective bargaining agreements or professional categories, their observed base wage in $t+1$ is not informative about the wage floor that would have applied had they remained with their current employer. Our approach establishes the collective bargaining wage floor relevant to the firm's decision regarding each incumbent worker, independently of workers' future outcomes.

2.2.4 Summary Statistics

Table 1 presents descriptive statistics for key variables in our analysis. Considering all workers, tenure within the firm is around 7 years, with high levels of CBA coverage. Almost 80% of workers are employed in the same firm in the following period, and the mean employment-weighted HHI is 0.23. Slightly less than one-third of the full sample is bitten, while almost 78% of all workers are associated with a collective bargaining wage floor above the NMW across the sample period.

The labour market duality arising from the Portuguese employment protection legislation results in sorting of workers across the two main contract types. Temporary workers are younger and less likely to be covered by a CBA. Employees on permanent contracts have higher wages and retention rates, which align with their greater seniority and their stronger protection against dismissal. The higher incidence of positive wage bites for temporary contract workers is in accordance with their lower hourly salary. Additionally, the probability that their relevant minimum wage is the national one is higher than it is for workers on a permanent contract. Lastly, the similarity in the HHI suggests that workers on different types of contracts sort into local labour markets with comparable concentration levels. Hence, differential exposure to labour concentration across types of contract is less of a concern when interpreting the results.

Table 1: Summary Statistics, 2002 to 2018

	All workers		Temporary workers		Permanent workers	
	Mean	(SD)	Mean	(SD)	Mean	(SD)
Age	38.9	(11.2)	34.7	(11)	40.6	(10.9)
Tenure (in years)	7.27	(8.48)	1.60	(3.32)	9.64	(8.85)
Hourly real base wage (€)	5.21	(5.94)	4.12	(6.79)	5.67	(5.48)
Uncovered CB worker (0 or 1)	.087	(.281)	.113	(.317)	.075	(.264)
Retention at firm $(t+1)$	.793	(.405)	.639	(.48)	.855	(.352)
New hire HHI	.230	(.284)	.246	(.289)	.224	(.283)
Wagebite	.050	(.123)	.055	(.122)	.048	(.123)
Extensive *wagebite* (0 or 1)	.319	(.466)	.368	(.482)	.299	(.458)
CB MW exceeds the NMW	.776	(.417)	.692	(.462)	.810	(.392)
Observations	41,903,126		12,327,278		29,496,053	

Source: Authors' analysis based on Quadros de Pessoal.

Notes: The table presents summary statistics from 2002 to 2018. 'Tenure in years' refers to firm-specific tenure. 'Hourly real base wage (€)' is the worker's hourly base wage, deflated to 2012 euros. 'Uncovered CB worker (0 or 1)' is a binary indicator of whether the worker is not covered by a collective bargaining agreement. 'Retention at firm $(t+1)$' is a binary variable indicating if the worker is employed by the same firm in the following year. 'New Hire HHI' is the Herfindahl-Hirschman index based on new hires. '*Wagebite*' is the minimum required increase in the log base wage that the firm must pay an incumbent worker taking into account the highest (national or collective bargaining) minimum wage in $t+1$. Workers earning at or above the next period's relevant minimum wage are assigned 0. 'Extensive *wagebite* (0 or 1)' is 1 if wage bite is positive and 0 otherwise. 'CB MW exceeds NMW' is 1 when the worker's collective bargaining minimum is above the national minimum wage.

3 Empirical Specification and Results

A key issue when analysing minimum wage effects on employment is potential correlation of the former with the firm's profitability. This problem is especially salient when the object of interest is the response of total firm employment to minimum wages (as in, for example, Azar et al., 2024 and Popp, 2024) but it is also an issue for our analysis of exit hazards. It is well documented that firms and workers within collective bargaining agreements share common economic conditions, so to counter the endogeneity problem, we include collective bargaining "agreement-year" fixed effects in our main specifications.[5] To control further for firm economic conditions, we also include the log of real firm value added per worker amongst the regressors. In a sensitivity analysis (below), we drill down yet further into the firm's current economic conditions by replacing firm and agreement-year fixed effects with firm-by-year fixed effects. Although we believe we adequately control for potential endogeneity of minimum wages we do not claim to have a natural experiment of the kind exploited by, for example, Portugal & Cardoso (2006) and Hijzen et al. (2026). However, in deriving results from a non-selective and economy-wide analysis of incremental minimum wage changes, we suffer far less from the external validity issues that arise when variation comes via a one-off special event such as a change in minimum wage law.

We document empirical results for temporary and permanent workers separately. For the latter we find little quantitative impact of *wagebite* on exit hazards irrespective of the levels of monopsony power. However, for temporary workers, the results are very different, and much of the analysis below focuses on this type of worker alone.[6]

We estimate the following Linear Probability model (LPM) separately for temporary and permanent contract workers i in firm j and year t:

$$I_{i,j,t+1} = \alpha + \beta wagebite_{ijt} + \delta HHI_{ijt} + \gamma(wagebite_{ijt} \times HHI_{ijt}) + \mathbf{X}'_{ijt}\theta + \phi_i + \psi_j + \lambda_{ct} + \varepsilon_{ijt} \quad (1)$$

where $I_{i,j,t+1}$ is a dummy variable that equals 1 if the worker is present in the same firm j in year $t+1$ and zero otherwise. The variable $wagebite_{ijt}$ is as defined above. HHI_{ijt} is the Herfindahl Index in worker i's local labour market in year t. $\mathbf{X}_{ijt}$ is a vector of

[5] Workers not covered by a CBA are considered as an $(n+1)$th "residual" group.

[6] We restate the caveat that QP does not distinguish quits from layoffs, so we cannot present our results as arising solely from the latter. However, and as noted above, prior reasoning and most theoretical models, including those built around job search and job ladders, predict that minimum wage rises would not increase quits amongst low-paid workers. Hence we interpret the estimated effects of minimum wage hikes on exit hazards as being a lower bound on the response of layoffs.

time-varying controls, including tenure, tenure squared, potential experience, potential experience squared, and log (real) value added per worker.[7] ϕ_i are worker fixed effects, ψ_j are firm fixed effects, and λ_{ct} are collective-bargaining-agreement-by-year fixed effects. As noted above our main focus is on temporary workers because we expect the effects to be largest there. We present results for permanent workers later.[8]

In any analysis such as ours, there is always a threat to identification from spillover effects of minimum wages to the wages of the unbitten (see for example Brochu et al., 2026). Ignoring spillovers leads to attenuation bias in the estimates (see, for example, Lewbel, 2007, who analyses the case where the spillover group is effectively treated and, more pertinently, Vazquez-Bare, 2023 who looks at the case where the spillover group receives a lower level of treatment). Some studies suggest quantitatively important effects on those workers earning up to about 30% above the minimum wage (a 30% spillover "margin"). By contrast, Autor et al. (2016), for the US, suggest that their own findings of a 20% margin may simply be due to measurement error. In any event, we would not expect these spillover margins to be the same across different labour markets/countries. Here, we make the reasonable assumptions that spillover effects decline with the wage and that there is a group of sufficiently highly paid earners that are completely unaffected by them. With all of this in mind, we estimate three specifications. The first excludes those incumbent workers who in year t earn a wage that lies between the worker's relevant minimum wage and 10% above that (hence allowing for a 10% spillover margin). The second specification extends the margin to 20%. The third ignores spillovers altogether. The attenuation bias that results from ignoring spillover effects leads us to expect that the estimates of γ and β obtained from the first and second specifications will be larger (in absolute value) than those obtained from the third. Furthermore, if the first and second exercises deliver similar estimates whilst the third is substantially lower, we may conclude that the spillovers extend to the 10% margin but probably not above that.

The results for the three specifications are given in Table 2. Specification 1) - our main case - is for temporary workers only and allows for a 10% spillover margin. Specification 2) extends the spillover margin to 20% and 3) ignores spillovers altogether. We see that

[7] Real value added per worker is defined as the firm's value added at factor cost divided by its number of workers. We obtain the variable from the SCIE dataset.

[8] In all specifications, we exclude worker-year observations corresponding to the last year in which the employee appears in QP (using an extended observation window up to 2023), ensuring that the $t \to t+1$ outcomes are observed and are not mechanically driven by right-censoring at the worker level. Additionally, we condition the analyses on firms remaining in business in $t+1$, abstracting from potential effects of minimum wages on firms' survival.

Table 2: Summary of Retention Regression Estimates

Specification	α	β	δ	γ	N
1) Temporary workers	.611	−.080	.008	.047	5,279,224
0-10% excluded	(38.28)	(−5.21)	(2.09)	(2.80)	
2) Temporary workers	.636	−.075	.008	.039	4,254,935
0-20% excluded	(36.49)	(−4.16)	(1.77)	(2.06)	
3) Temporary workers	.597	−.062	.013	.028	7,200,127
All included	(37.36)	(−5.14)	(3.40)	(1.86)	
4) Permanent workers	.912	−.010	.002	−.001	13,506,095
0-10% excluded	(102.68)	(−5.18)	(1.61)	(−.21)	

Notes: The table presents results for different specifications based on Equation 1 and its parameters. The first three specifications focus on temporary contract workers whilst the fourth addresses workers on permanent contracts. Specification 1) allows for a spillover margin of 10% by excluding those workers who earn a base wage in t that lies between their relevant $t+1$ minimum wage and 10% above that. Specification 2) extends the spillover margin to 20%. Specification 3) includes all workers and hence ignores spillovers altogether. Specification 4) is based on permanent contract workers only, allowing for a spillover margin of 10% above. All specifications include firm, worker and collective-bargaining-agreement-year fixed effects, log firm (real) value added per worker, as well as worker tenure, potential experience and their squared terms. The analyses are restricted to continuing firms, and we exclude worker observations if the current observation is the last year the employee appears in QP where the terminal date is 2023. The years analysed are from 2002 to 2018. t-ratios are in parentheses below the estimates. Standard errors are clustered at the firm level.

the estimates in 1) and 2) are extremely close whilst estimates of the key parameters β and γ in 3) are respectively around 25% and 40% below their counterparts in 1) and 2). Given the discussion above on attenuation bias this is strong evidence [9] that minimum wage spillovers are unlikely to extend much beyond the 10% margin. In subsequent specifications, we therefore omit those workers whose wages in t put them in the 10% "spillover margin".

Setting $HHI = 0$ in 1) gives us the impact of *wagebite* on exit hazards in competitive markets. Roughly speaking a 10% rise in *wagebite* here increases the exit hazard by just under 1 percentage point - quantitatively modest but highly statistically significant.[10] There is a substantial offset of this effect due to monopsony power; under perfect monopsony this offset is nearly two-thirds of the original effect. Specification 4) gives the

[9] Whilst the evidence is strong it is not conclusive because we are comparing estimates across different subsamples of workers.

[10] Probabilities here and below are obviously conditional in nature, i.e. here they are conditional on the worker being on a temporary contract.

results for workers on permanent contracts. Here we see that the retention rate amongst these workers exceeds 90% and the increase in exit hazard due to a positive *wagebite* whilst statistically significant is (quantitatively) unimportant. This reinforces the argument that layoffs impinge on temporary workers and probably occur via the lowering of the conversion rate of temporary workers to permanent status. Henceforth we discuss results only for temporary workers.

A Comparison With Results in the Literature.
Several papers have looked at the effect of minimum wage rises on exit hazards. The closest to ours in terms of results and labour market similarity is Hijzen et al. (2026). They study the 22% increase in the Spanish statutory minimum wage in 2019, although not considering the role of labour market power. In Spain, as in Portugal, there is a dual labour market and like us they too focus on employees on fixed-term contracts (about 27% of workers). With regard to exit hazards, they find a modest increase of about 0.5 percentage points arising from the minimum wage hike, although differently than us they do not consider the effects of monopsony power explicitly. Some other studies find decreases in exits in response to minimum wage hikes. In particular Dube et al. (2016) using Quarterly Workforce Indicators from the US and a cross-border regression-discontinuity design find - inter alia - that exit hazards (separations) fall as minimum wages rise. They argue this arises because higher minimum wages compress the outside wage offer distribution and reduce quits. They illustrate this using a theoretical job-ladder model where reductions in the firm's labour force are deemed to occur through a fall in hiring. In a study of fast-food restaurants, Wiltshire et al. (2026) find that increasing minimum wages had no effect on employment but increased employee retention. These last two papers do not find the increase in exit hazards that we do, possibly because the institutional setting and empirical designs are very different from ours. Even so, they both claim that quits fall following a rise in the minimum wage. This reinforces our earlier argument that despite not being able to identify quits in our data our results on exit hazards may be interpreted as a lower bound on the layoff response of firms to minimum wage increases.

Sensitivity Analyses
Several sensitivity analyses beyond those reported above were carried out, and we report here the most salient of these. In what follows, the discussion focuses purely on heterogeneity in the key parameters β and γ. In general when we look at subsamples of the full dataset we should expect to obtain estimates with different magnitudes. For example in sectors where capital and labour are close substitutes and/or the price elasticity of prod-

uct demand is high, we may expect to see larger responses to *wagebite* than we do in the whole sample. Our main interest however is to see if *wagebite* has a significantly negative impact on bitten worker retention ($\beta < 0$) and that monopsony power ameliorates this effect ($\gamma > 0$) - the core results of our paper. Full details of all of the estimates are in Online Appendix B.

First, dropping value added as a control (Table B1) resulted in a lowering of γ but virtually no difference in β. However, dropping agreement-year fixed effects (moving from (3) to (2) in Table B2) reduced the size of both estimates by over 20%. To home in more on the firm's current economic conditions we replaced agreement-year, value added per worker and firm fixed effects with firm-by-year fixed effects (see Table B3). Aside from a slight increase in β the results were largely unchanged.[11] These results suggest that CBA-Year effects are the main workhorse in controlling for potentially endogenous components of *wagebite*.

Second, we estimated incremental effects for those whose binding wage was the NMW (Table B4). The key takeaway is that once all of our fixed effects are included there is no significant difference between the impacts of the National and CBA Minimum Wages on exit hazards.[12] Another interesting angle here is the increase (in absolute value) in β and γ of around 20% when we add CBA-Year Fixed Effects (i.e. when we move from column (2) to column (3)). This suggests that CBA minimum wages are to some extent endogenous and, further, that they are negotiated in such a way as to stabilise the economic environment of firms in the agreement.

In a third exercise we split the sample into high and low unemployment periods. This may tell us whether or not firm profitability and the state of outside options for workers affect our estimates. The results (see Table B5) were broadly unchanged, suggesting that - once all of our controls are added - these factors do not separately affect the responses of exit hazards to minimum wages and monopsony power.

Fourth, we address potential heterogeneity of treatment effects with respect to age. Here we allowed separate regression coefficients for those aged 17 to 19 inclusive (Ta-

[11] There is a cost to including such detailed fixed effects as firm-year. Most of our firms are small, so that including firm-year fixed effects effectively places greater weight on larger firms. For this reason, we use agreement-year fixed effects as our baseline control - the robustness with respect to switching to firm-year fixed effects reassures us that this is not a misstep.

[12] The significantly negative dummy term suggests that NMW workers have higher steady state churn rates than CBA workers. This may be due to differing job characteristics associated with the two groups of workers but we do not wish to pursue this further in the current paper.

ble B6). We found no quantitatively or statistically significant changes in the coefficients except possibly for the effect of HHI itself. These findings contrast with those of Portugal & Cardoso (2006) who found that a large across-the-board increase in the minimum wage for all young workers in 1986 caused exit hazards to fall but who do not consider the role of labour market power.[13]

Finally, and following Azar et al. (2024), who focus solely on high bite sectors, we present two further results relating to lower paid workers. Firstly we allow for different coefficients for those workers whose firm has a lower than median proportion of bitten workers in year t and secondly we focus solely on a single low-wage sector, namely "Wholesale and Retail Trade; Repair of Motor Vehicles". Results for the first exercise (Table B7) show a significant, negative difference in β and a positive, insignificant increase in the size of γ for firms with low shares of bitten workers. The second exercise (Table B8), while reaffirming our core result, yields estimates nearly four times larger than their baseline counterparts. These two results tell us that it is hazardous to draw general conclusions about low wage firms from a single low wage sector; the extent of capital-labour substitution, the responsiveness of the sector's market demand to price rises, etc., may all affect the size of the response of the exit hazard to minimum wage rises, and it is hard to generalise about these.

4 Theory

We adapt the oligopsony model of Azar et al. (2024) from the employment margin to the retention of incumbents, adding idiosyncratic match shocks. Details are in Appendix A. This is indicative of a potential mechanism consistent with our results and is not intended to be a full calibration.

[13] These contrasting results may be due to the macro effects of a one-off large natural experiment such as theirs. After the minimum wage hike, young workers as a bloc became more expensive than their older counterparts. Outside options for these workers may well have dried up and hence search intensity and quit rates may have declined. The substitution of older for younger workers is supported by their finding that new hires of the young also fell. Whilst in our data a positive *wagebite* may also reduce search activity and quits, it is largely an individual event with arguably few or no macro effects.

4.1 A wage floor in a concentrated labour market

A local labour market contains J identical firms, so $HHI = 1/J$. Following Azar et al., each produces revenue AL^{α}, $\alpha \in (0,1)$, and market supply is $N(w) = \Phi w^{\eta}$, where L is labour employed and w the wage; firms compete à la Cournot in employment. In the absence of a wage floor, the symmetric equilibrium wage w^* lies below the marginal revenue product by the (log) markdown

$$m(HHI) \;=\; \ln\!\Big(1+\frac{HHI}{\eta}\Big) \;\approx\; \frac{HHI}{\eta}, \tag{2}$$

which is increasing in concentration. Define the *bite* of a wage floor $w_{\min}$ as $b \equiv \ln(w_{\min}/w^*)$, a level: the floor relative to the wage the firm would pay unconstrained.[14]

A binding floor produces three regimes. Writing $\ell(b)$ for log employment per firm relative to its no-floor level,

$$\ell(b;HHI) = \begin{cases} 0, & b \le 0, \\ \eta b, & 0 < b \le b^*, \\ \dfrac{m-b}{1-\alpha}, & b > b^*, \end{cases} \qquad b^* = \frac{m}{\rho}, \quad \rho \equiv 1+\eta(1-\alpha) > 1. \tag{3}$$

For moderate bites the floor raises the wage above its marked-down level and employment expands along the supply curve; once the bite exhausts the markdown buffer—at $b^* < m$, since expansion along the supply curve itself erodes it—the firm moves onto its labour demand curve and contracts at elasticity $1/(1-\alpha)$.[15] Since b^* rises with concentration, a floor of given bite is demand-determined in competitive markets and supply-determined in concentrated ones.

4.2 Retention of incumbents

Let R denote the probability that an incumbent is retained over the year; the empirical exit hazard is $1-R$. Two channels connect the wage floor to R.

The first is *downsizing*. A floor rising from b_0 to b_1 changes target employment by $\Delta \ln L = \ell(b_1) - \ell(b_0)$; if this falls short of current employment net of attrition, incumbents are shed.

[14] Not the empirical bite z of the results section, which is the increment mandated this year; under the model $z \le b$, so moments of z bound the bite distribution from below.

[15] Equation (3) renders in closed form the model Azar et al. (2024) solve numerically, and reproduces their Figure 1.

The second is the markdown itself, acting as a *retention buffer*. We assume ongoing matches receive negative idiosyncratic productivity shocks $-\varepsilon$ ($\varepsilon \geq 0$), $\varepsilon \sim G$, with annual probability λ; a new match starts without one. In the supply-determined regime all the labour forthcoming at the wage floor is employed, so a slot released by a separation cannot be refilled without bidding above the floor. The firm therefore retains a shocked worker whenever her output still covers her cost, up to a reservation draw

$$\bar{\varepsilon} \;=\; x(b) + (\text{option value}), \qquad x(b) \;=\; \max\{m - \rho b,\, 0\}, \quad b \geq 0, \tag{4}$$

where $x(b)$ is the log gap between what the marginal worker produces and what she is paid. Retention is increasing in x. The buffer is largest in concentrated markets, erodes as the floor rises, and is exhausted once the firm reaches its demand curve, where the marginal worker is paid her product.[16] A floor increase shrinks the buffer, and matches whose existing ε lies between the old and new thresholds are laid off.[17]

This is where the retention margin departs from Azar et al.'s. While total employment can rise with the floor in concentrated markets, retention falls because of increased incumbent exits due to compression of the buffer. Both channels are bounded above by zero: expansion removes downsizing-induced separations, but buffer erosion can still cost incumbents, so hiring and retention can move in opposite directions. A contracting firm may be laying off workers if there is insufficient natural wastage to cover the contraction. Monopsony can reduce the exit response toward zero but not reverse its sign.

In competitive markets the retention response is essentially all downsizing—since ordinary turnover absorbs much of any required contraction, downsizing is of smaller magnitude on this margin relative to its role in total employment. By contrast, in concentrated ones the response is mostly buffer erosion.

4.3 Predictions

At a fixed bite, the profile in HHI implied by (3)–(4) jumps at the regime boundary; to match the smooth profile in the data we aggregate. At a given *HHI* the sample pools firms whose bites vary widely—a common floor meets heterogeneous counterfactual wages, and sectoral floors cross our occupation–municipality markets—so the estimated coefficient is

[16] In the supply-determined regime $x(b)$ is (minus the log of) the shadow markdown of Berger et al. (2025), whose static model has no separation margin.

[17] Dobbin et al. (2026) obtain similar buffer logic outside the minimum wage setting—markdowns cushion matches from adverse shocks—and find supporting evidence in Brazilian matched data.

a mixture over the bite distribution. As concentration raises b^*, firms "leave" the downsizing regime—more firms are supply-constrained—and the aggregate response declines smoothly, while buffer erosion, itself weakening as buffers widen, keeps it strictly negative. The dispersion this requires is present in the data: across cells, the standard deviation of the bitten share of temporary employment is at least 0.18 at every concentration decile, on a flat mean of 0.36.[18]

Three predictions follow: minimum wage increases raise incumbent exit hazards; the effect is attenuated in more concentrated markets; and attenuation cannot reverse the sign, so in the notation of specification (1), $\beta < 0$, $\gamma > 0$ and $\beta + \gamma \leq 0$. In the model the first and third hold at every level of concentration; the second requires mild regularity—the shock density must decline over the relevant range and the downsizing response at the regime boundary must exceed the buffer response there, so that firms crossing the boundary as concentration rises lose more than they gain (Appendix A). For the fitted linear specification the third is not automatic, since a straight line fitted to a convex response could cross zero before $HHI = 1$; with both channels it does not in any of our simulations. The bound is a restriction rather than an identity: a model with voluntary quits may not satisfy it, and, e.g., Portugal & Cardoso (2006), as discussed above, find contrary evidence for substantial youth minimum wage increases. The fact that the estimates satisfy this restriction suggests the quit margin may be dominated for bitten temporary workers in our sample.[19]

Figure 1 shows specification (1) estimated on data simulated from the model, for a population of temporary incumbents resembling our sample. The simulated coefficients satisfy all three predictions, and the simulated $\hat{\gamma}/\hat{\beta}$ is -0.57, close to the estimated -0.59; with the downsizing channel alone it would be -0.79, attenuating too fast. Plotted by concentration, the model's response is convex, which a linear interaction cannot pick up. At very low concentration almost all firms with bitten workers, who make up 37% of incumbents, are pushed into the demand regime by the increase, but only around a quarter of them face a bite large enough that target employment falls by more than the workers they would lose over the year anyway, so that they must lay off incumbents. As

[18] It also widens with concentration, to 0.37 at the most concentrated decile. The bite itself is dispersed too: the moments in Table 1 imply that among bitten temporary workers it has a mean of 0.15 and a standard deviation of 0.16. Our simulation holds this distribution fixed across levels of concentration; Appendix A reports what happens if its dispersion widens with concentration.

[19] The linear specification includes fixed effects that our simulation cannot reproduce, so we ask the model to rationalize the sign, the bound and the direction of heterogeneity, not to deliver a structural interpretation of $\hat{\beta}$ and $\hat{\gamma}$.

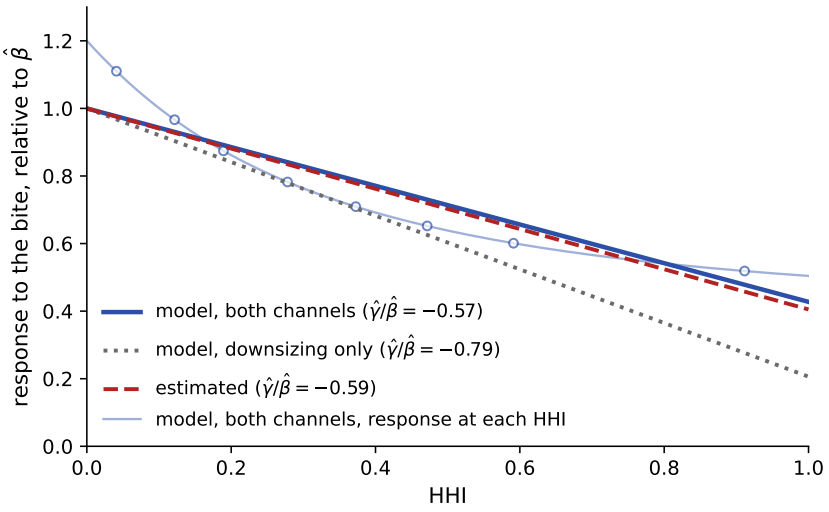


Figure 1: The retention response to the bite by concentration, from specification (1) estimated on data simulated from the model and on our data. Lines show $1 + (\hat{\gamma}/\hat{\beta})HHI$ for the model with both channels (solid) and with the downsizing channel only (dotted), and as estimated (dashed). The light line is the model's response at each level of *HHI* with both channels: the least-squares slope of retention on the bite, divided by the simulated $\hat{\beta}$. Its markers are at the average HHI of each decile of our sample, and the solid line is the straight line that specification (1) fits through them; since the response is convex, the line lies below it at both ends.

concentration rises, fewer firms end up in the demand regime, and those that do contract by less, because the markdown absorbs the first part of the increase. The bite needed to trigger layoffs therefore rises, and since it starts near the middle of the bite distribution, the share of firms laying off falls fast, from about 29% at $HHI = 0$ to 12% at $HHI = 0.19$: hence the initial steep decline. Further up, that bite lies in the thin right tail of the distribution, so the share falls only slowly, and buffer erosion, which weakens only gently with concentration, takes over: hence the flattening. As in Azar et al.'s simulations, the figure illustrates the mechanism rather than calibrates it: the ratio depends on the shock distribution, which we set rather than measure, on how far firms were already bound before the increase, which we cannot observe, and on how the dispersion of bites varies with concentration. Nor do we target levels: the simulated coefficients are several times the estimates. Parameters and robustness are in Appendix A.

4.4 Permanent contracts

A firing cost dampens both channels. Dismissal now requires the loss per worker to exceed this cost, which shifts the threshold $\bar{\varepsilon}$ in (4) upward. Letting $\mathcal{F}$ denote the amortised cost of dismissal, measured in the same log units as x, the shift is exactly $\mathcal{F}$. It also reduces the downsizing response: The firm need not dismiss anyone ordinary turnover will remove in any case, and it dismisses the remainder only where the loss from carrying a worker over her expected remaining tenure exceeds the cost of shedding her. For $\mathcal{F}$ large relative to the match surplus—the relevant case for permanent contracts in Portugal—the model predicts $\beta \approx \gamma \approx 0$, as specification (4) in Table 2 shows. Exposure is not the explanation: 30% of permanent employment lies at or below the incoming floor against 37% of temporary, so permanent incumbents are bitten nearly as often.

5 Conclusions

In this paper we have revisited the question of minimum wage effects on employment from the point of view of incumbent exit hazards. Using the universe of Portuguese private-sector workers in the QP data, we find quantitatively modest but statistically significant increases in exit hazards for temporary contract workers but virtually no effects for permanent contract workers. The effects are substantially ameliorated by the existence of monopsony power. The results are robust when we allow for heterogeneous responses in several dimensions and when we control for current economic conditions within the firm in different ways. Under the theoretically reasonable assumption that quits do not rise in response to minimum wage hikes, we may interpret our results as a lower bound on the layoff response of firms. Our results suggest that the small but significant negative impact of mandated wage increases on incumbents impinges almost entirely on temporary contract workers. An extension of the model of Azar et al. (2024) to exit hazards, in which markdowns act as a buffer that protects shocked incumbents until a rising floor erodes it, illustrates a mechanism consistent with these qualitative findings.

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Appendix A Theory appendix

This appendix sets out the model behind Section 4 and the simulation behind Figure 1. As in Azar et al. (2024), whose parameter values we use, the exercise is meant to illustrate a mechanism, not to calibrate it. We model only the firm side: quits are treated as independent of the minimum wage.

The regimes. The switch point $b^* = m/\rho$ in (3) lies below the markdown m. A firm in the supply-determined regime hires the extra labour that a higher floor brings forth, and in doing so lowers its marginal product, so each unit of bite consumes $\rho = 1 + \eta(1 - \alpha)$ units of markdown rather than one: the η term reflects the move along the supply curve and the $1 - \alpha$ term the fall in the marginal product as employment rises. In logs, (3) is exact for the symmetric Cournot equilibrium.[20] Because b^* rises with concentration, a floor of given bite is demand-determined in competitive markets and supply-determined in concentrated ones, which is the comparison our empirical specification exploits.

Shocks and the stock of incumbents. Each year a match is hit by a productivity shock with probability λ. A shock lowers log output by $\varepsilon \geq 0$, drawn from a continuous distribution G, so that $G(0) = 0$, and the draw persists until the next shock replaces it.[21] A new hire starts without a shock. Matches also end with annual probability ς for reasons unrelated to the buffer (quits, retirements and the like), so each year a match survives attrition with probability $\sigma \equiv 1 - \varsigma$.

Since attrition and shocks are independent of the current draw, the stationary stock of incumbents is a mixture of matches never shocked, all at $\varepsilon = 0$, and shocked survivors

[20] In a symmetric equilibrium each firm takes $1/J$ of market supply, since any higher demand would raise its marginal cost discontinuously. There is also a continuum of asymmetric equilibria, which we do not consider.

[21] Both polar alternatives fail. With a single draw at match formation every survivor carries a live draw and survivors are bunched just inside the threshold, so a small fall in the threshold severs an implausibly large share of the workforce. A symmetric single draw avoids this but makes every replacement a fresh draw, which a firm facing a queue can screen, so that its standard is no longer an average match. With an independent redraw every period, a bad draw carries no information about the future, and a firm that cannot replace anyone simply rides it out. Persistent shocks with stochastic arrival, as in Mortensen & Pissarides (1994), avoid these problems.

distributed as G truncated at the reservation draw $\bar{\varepsilon}$:

$$\underbrace{1-\varphi G(\bar{\varepsilon})}_{\text{mass at } \varepsilon=0} \qquad \text{and} \qquad \underbrace{\varphi\, g(\varepsilon) \text{ on } (0,\bar{\varepsilon}]}_{\text{shocked survivors}}, \qquad \varphi \equiv \frac{\sigma\lambda}{1-\sigma+\sigma\lambda}. \tag{5}$$

Here φ is the probability that a match is shocked before attrition ends it. Faster attrition makes the stock younger, with more of it never shocked and a smaller pool for the buffer channel to act on. Among matches that survive attrition, a shock arrives with probability λ and is survived with probability $G(\bar{\varepsilon})$, so the shock-driven and total separation rates are

$$s = \lambda\big(1-G(\bar{\varepsilon})\big), \qquad s^{\text{tot}} = \varsigma + \sigma s. \tag{6}$$

It is s^{tot} that we observe.

Who is kept. Whether a firm keeps a shocked worker depends on its alternative, and the alternative differs by regime.

In the supply-determined regime the firm already wants more workers than it can obtain at the floor: its marginal product exceeds the wage, so it would hire more if it could. A separated worker re-enters the pool, but the pool is in excess demand, so the vacancy she leaves goes unfilled; no matching friction is needed for this. The alternative to keeping a worker is therefore an empty slot, worth zero, and the value of a match with current draw ε is

$$V(\varepsilon) = (x-\varepsilon) + \frac{\sigma}{1+r}\Big[(1-\lambda)V(\varepsilon) + \lambda\int \max\{V(\varepsilon'),0\}\, dG(\varepsilon')\Big], \tag{7}$$

where r is the discount rate.[22] The integral is the option value in (4): the value of a later, better draw. We set it to zero for simplicity, so that the firm keeps a worker while her output covers her wage and $\bar{\varepsilon} = x$.[23] A deeper markdown thus protects a larger shock. Since x falls with the bite and rises with concentration, the separation rate s rises with the bite and, for a given bite, falls with concentration.

[22] The flow payoff is a first-order approximation, since exact profit is proportional to $e^{x-\varepsilon}-1$. Because the payoff depends on $x-\varepsilon$ through a single monotone index, the exact problem gives a reservation draw of the same form.

[23] Since V is linear in ε, the exact reservation draw solves $\bar{\varepsilon} - k\int_0^{\bar{\varepsilon}} G(t)\,dt = x$, with $k \equiv \sigma\lambda/\big(1+r-\sigma(1-\lambda)\big) < 1$, so $\bar{\varepsilon}$ rises with x. It equals x at the switch point, where $x=0$ and a worker who has not yet been shocked can only get worse, and exceeds it elsewhere, but only slightly: with $r=0.05$, by at most 0.014 at the parameters we use below. Using the exact threshold changes the simulated $\hat{\gamma}/\hat{\beta}$ reported below by less than 0.01.

In the demand-determined regime there is a queue at the floor, and a separated worker is replaced straight away. The replacement starts without a shock and, by assumption, costs nothing to hire, so she is exactly as good as an incumbent who has not been shocked, and the firm replaces any incumbent who has. The reservation draw is therefore zero, the value (4) takes at the switch point, so it varies continuously with the bite across the two regimes.[24] What protects a shocked incumbent in the supply-determined regime is that no replacement is available. The markdown alone would not protect her: the firm would earn the same markdown on a replacement hired at the floor, so the markdown would cancel from the comparison and she would be replaced, just as in the demand-determined regime.

Two features of this set-up deserve comment. A separated worker does not return to her own firm. This is not an extra assumption if a match-specific shock stays with the worker–firm pair, since she would then bring her bad draw back with her. And replacement is impossible in one regime and immediate in the other. A replacement probability rising smoothly to one as excess demand disappears would be more realistic, but it would give an empty slot a value we cannot discipline. It would narrow the buffer but also change how fast the buffer erodes as the bite rises, so we cannot sign its effect on the response to the floor.

Dobbin et al. (2026) obtain a related result outside of the minimum wage setting, where firms post wages: firms with larger markdowns lay off fewer workers after adverse match-specific shocks, since a posted wage cannot be cut when the draw is bad. Their markdown thus plays the role of our buffer. What a rising wage floor adds is the erosion of that buffer, which is the source of the retention response here. In Arabzadeh et al. (2024) the wedge is the cost of financing the wage bill. It enters the reservation output multiplicatively, raising the bar a match must clear rather than lowering it, and their wage floor reduces separations through hiring selectivity rather than retention.

A rise in the floor. Suppose the floor rises, taking a firm from bite b_0 to $b_1 = b_0 + z$, where the increment z is the empirical bite of specification (1). If it starts in the supply-determined regime, its buffer falls from x_0 to x_1, and the reservation draw with it. The firm re-evaluates its existing matches against the new floor, so those whose current draw lies between the two thresholds are dismissed straight away; matches never shocked are

[24] A hiring cost, or a productivity premium for experienced workers, would not change this, provided the switch point is placed where a new hire is worth exactly what she costs: both shift the hiring margin, in opposite directions, and the two thresholds still meet.

unaffected. By (5) the share of the stock culled is

$$c = \varphi\big[G(x_0) - G(x_1)\big]. \tag{8}$$

Because the reservation draw varies continuously with the bite, the cull grows continuously with the size of the increase. Once the bite passes the switch point every shocked worker is dismissed, $c = \varphi G(x_0)$, but there is no jump at the switch point itself. In addition, shocks arriving during the year are now survived only below the new threshold, so the shock-driven separation rate rises to $s_1 = \lambda[1 - G(x_1)]$. A firm that starts in the demand-determined regime has no buffer to lose: $x_0 = x_1 = 0$, so $c = 0$ and $s_1 = s_0$, and only downsizing, below, can cost it incumbents. Retention over the year is

$$R_1 = \min\big\{\sigma(1-c)(1-s_1),\ e^{\Delta \ln L}\big\}. \tag{9}$$

The first term is the share of incumbents who survive attrition, the cull and fresh shocks. The second is a cap: the firm's new employment target relative to its old size. Retention is whichever is smaller. Taking the minimum rather than a sum avoids double counting: a firm that must shrink by 20% and has already lost 15% through selection sheds 20% in total, not 35%, since its worst matches go first. In the supply-determined regime the target rises with the floor and the cap never binds, so downsizing appears only beyond the switch point. Setting $c = 0$ and $s_1 = s_0$ for every firm gives the pure downsizing case.

A contracting firm lays off no additional incumbents unless its contraction exceeds the turnover it experiences anyway, since it can shrink by not replacing those who leave. A firm already beyond the switch point replaces every shocked worker, so its turnover rate is $1 - \sigma(1-\lambda)$, and it lays off incumbents only if z exceeds

$$\bar{z} = -(1-\alpha)\ln\big[\sigma(1-\lambda)\big]. \tag{10}$$

A match ended by a bad draw vacates a position just as a quit does, so both count as turnover here. A firm that starts in the supply-determined regime faces a higher hurdle. Once its bite passes the switch point, its target employment changes by $\Delta \ln L = (x_0 - z)/(1-\alpha)$, so it falls only after the increase has used up the buffer x_0, and the firm lays off incumbents only if z exceeds roughly $\bar{z} + x_0$ (slightly more, since the cull already removes some workers).

Retention can never rise above its level without the increase, since $c \geq 0$ and $s_1 \geq s_0$, so the response to the bite is non-positive at every level of concentration: high concentration shrinks the response but cannot reverse it. Specification (1) approximates this response by a straight line, whose value at $HHI = 1$ is $\beta + \gamma$. The restriction $\beta + \gamma \leq 0$

on the fitted line does not follow automatically, since a straight line fitted to a convex profile could cross zero before $HHI = 1$; with both channels at work it does not in any of our simulations.

Why the response attenuates. As concentration rises, two things happen. Firms already in the supply-determined regime have wider buffers, and since the shock density declines, each unit of bite dislodges fewer of their matches; this always reduces the response. And firms move from the demand-determined to the supply-determined regime, so that for each of them downsizing gives way to buffer erosion. In terms of average responses, this second effect also reduces the response if, at the switch point, downsizing removes more incumbents per unit of bite than buffer erosion would:

$$\frac{\Pr\{z > \bar{z}\}}{1-\alpha} \;\geq\; (\varphi + \lambda)\,\rho\, g(0). \tag{11}$$

On the left, $1/(1-\alpha)$ is the employment response beyond the switch point and $\Pr\{z > \bar{z}\}$ the share of firms whose increase outruns ordinary turnover. On the right, ρ converts bite into buffer, and φ and λ count the matches exposed through the cull and through fresh shocks. With the parameters set out below, the two sides are 0.71 and 0.30.

The left-hand side also shows why the dispersion of bites matters. With the same parameters, $\bar{z} = 0.17$, just above the average bite among bitten workers, 0.15, so the amount of downsizing in competitive markets depends on how much of the bite distribution lies beyond $\bar{z}$, and a least-squares slope gives those firms particular weight. With little dispersion, downsizing is weak even in competitive markets, buffer erosion dominates once concentration rises, and the response stops falling.

Simulating the regression. Figure 1 is based on estimating specification (1) on data generated by the model. We construct a population of temporary incumbents resembling our estimation sample. Concentration takes the average HHI of each decile of the sample, each decile weighted by its share of observations (the top three deciles are collapsed in our data). A share 0.37 of incumbents is bitten (Table 1). Their bite z is lognormal with mean 0.15 and standard deviation 0.16, the moments among bitten temporary workers implied by Table 1, and we hold this distribution fixed across levels of concentration, much as Azar et al. (2024) hold the size of the floor increase fixed across markets. Firms pay the unconstrained wage before the increase, $b_0 = 0$, so the level bite after it equals the increment, $b_1 = z$. This is the smallest level consistent with $z \leq b_1$, and so puts the fewest

firms beyond the switch point; we report larger starting bites below. Unbitten incumbents work at firms where the floor does not bind, so their buffer is the full markdown m.

Each incumbent's retention probability then follows from (9), and we regress it on the bite, HHI and their product by least squares, weighting by population shares. Using probabilities rather than simulated exits removes sampling noise. The simulation cannot reproduce the fixed effects of specification (1): in effect it treats the bite as good as randomly assigned conditional on them.

Parameters and results. We use Azar et al.'s values $\alpha = 0.6$ and $\eta = 1.6$, so that $\rho = 1.64$ and the labour demand elasticity is 2.5, and take $\varsigma = 0.165$, the measured annual separation rate across all contracts, as the probability of attrition unrelated to the buffer. Shocks are exponential with rate ζ. We set λ so that the steady-state separation rate of bitten incumbents before the increase, s^{tot} in (6), equals the measured exit rate of bitten temporary workers, 0.331, which gives $\lambda = 0.21$. The one parameter we cannot take from Azar et al. or from our data is ζ, which we set at 0.25: then about one shock in forty lowers output by less than a tenth, and most shocks end the match.

The simulated coefficients are $\hat{\beta} = -0.45$, $\hat{\gamma} = 0.26$ and $\hat{\delta} = 0.011$, against estimates of -0.080, 0.047 and 0.008, so the simulated ratio $\hat{\gamma}/\hat{\beta}$ is -0.57, close to the estimated -0.59. The signs match, including that of $\hat{\delta}$, which is positive because even without a bite the wider buffer of a concentrated market lowers separations. With downsizing alone the ratio would be -0.79. It does not reach -1 because, even at $HHI = 1$, under 2% of bitten firms have increases large enough to exceed both ordinary turnover and the markdown they first use up, and least squares gives these firms, whose bites are far from the mean, considerable weight. The light line in Figure 1 shows the response at each level of concentration. It is convex: it falls steeply over the first deciles, as firms leave the downsizing regime, and then flattens as buffer erosion takes over. In the most concentrated decile buffer erosion accounts for about 80% of the incumbents lost but under 40% of the least-squares slope, for the reason just given. Specification (1) summarises the response as the solid line, the straight line it fits through the light one.

The simulated coefficients are considerably larger than the estimates. We do not target levels, and several things the model leaves out would make the estimates smaller: measurement error in the bite, which fixed effects tend to amplify; quits, which may fall when wages rise and so offset some layoffs; and contractions spread over more than a year. The model also treats a bite as raising the cost of every worker in a firm, when only

around a third of temporary workers are bitten, and uses Azar et al.'s labour demand elasticity of 2.5. Measurement error shrinks $\hat{\beta}$ and $\hat{\gamma}$ in the same proportion, leaving their ratio unchanged, as long as its share in the variance of the bite does not vary with concentration; this is one reason we compare ratios rather than levels.[25]

Robustness. Three choices matter most. The first is the shock parameter: for ζ between 0.15 and 0.4 the simulated ratio runs from -0.67 to -0.43. The second is the starting bite b_0. If firms were already bound before the increase, with b_0 an independent draw from the distribution of the increment, more firms in concentrated markets end up beyond the switch point and attenuation weakens: the ratio is -0.43, and -0.55 if b_0 is drawn from that distribution scaled by one half. The third, and the most important, is the dispersion of bites, for the reason given above. A single distribution with a standard deviation between 0.10 and 0.25 gives ratios between -0.43 and -0.61, but with a standard deviation of 0.07 the attenuation all but disappears. In our data the dispersion of firm-level average bites widens with concentration, from 0.07 in the least concentrated decile to 0.16 in the most. If worker-level bites widened in the same proportion, more firms in concentrated markets would face increases large enough to require layoffs, and the ratio would be -0.37 (between -0.44 and -0.25 for ζ between 0.15 and 0.4). The fixed effects in specification (1) absorb much of the cross-sectional variation in bites, so we cannot tell which pattern of dispersion the estimates respond to. This is the main reason we read Figure 1 as an illustration.

Other choices matter less. Excluding the spillover margin raises the bitten share in our estimation sample; with a share of one half the ratio is -0.58, and dropping the unbitten altogether gives -0.63. Attrition rates of 0.10 and 0.25 give -0.48 and -0.70. A lower labour demand elasticity weakens downsizing relative to ordinary turnover, but the response still declines throughout: with $\alpha = 0.5$, an elasticity of 2, the ratio is -0.47. Finally, treating workers as perfect substitutes lets turnover absorb contraction one for one. Where separations fall on positions the firm still needs filled, only part of turnover is available, so (10) overstates the threshold and the downsizing channel is understated.

[25] Classical measurement error multiplies both coefficients by the reliability ratio: the share of the bite's variance, net of the fixed effects, that is signal rather than noise. If a fifth of that variance were noise, for example, both coefficients would shrink by a fifth; adding such noise to the simulated bites leaves $\hat{\gamma}/\hat{\beta}$ at -0.57. Accounting for the whole gap would take just over four-fifths noise.

The downsizing channel in Azar et al.'s design. Figure 2 asks what the downsizing channel alone delivers for retention in the design Azar et al. (2024) use for employment: floors set 3% above the market-specific oligopsony wage, then raised by 12%, then by 40%. Dashed lines hold the bite fixed across markets, as in their design: each step's retention loss is largest in competitive markets and falls to zero once the step is absorbed by the markdown buffer, but does so along an implausibly kinked path. (For example, in panel (a) the orange dashed line shows the change in retention for the middle increase: in a competitive market labour demand contracts by about 25%; in oligopsonistic—supply-constrained—markets firms expand, so all incumbents are retained; and the upward-sloping section in between reflects a move from the supply-constrained to the demand-constrained regime in which employment falls.) Solid lines aggregate the same steps over a distribution of market-level bites, a normal location shifter with standard deviation 0.12: the kinks smooth out and the loss shrinks steadily with concentration. Panel (b) shows that baseline worker attrition reduces the losses without changing their shape. Even with mixing, though, the downsizing channel alone attenuates too fast: the retention loss goes to zero once the markdown absorbs the step, although for the largest step this does not happen even at $HHI = 1$. That is what the downsizing-only profile of Figure 1 shows against the estimated one. The buffer channel remedies this: as modelled here, a supply-constrained firm goes on shedding incumbents as the floor rises, at a rate that falls with concentration but stays positive.

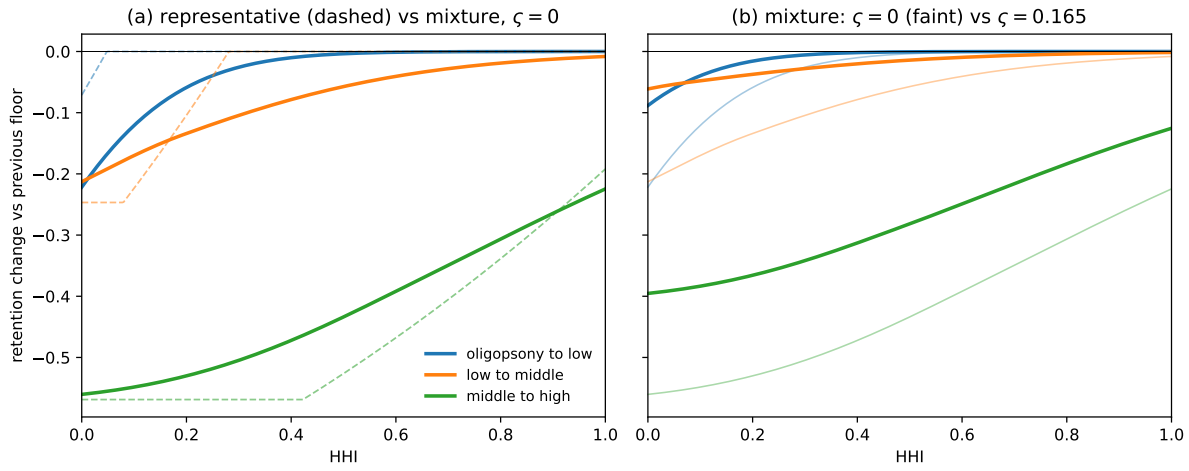


Figure 2: Retention change per minimum-wage step (relative to the previous floor), downsizing channel only, by HHI, following the simulation design of Azar et al. (2024). Panel (a): bite fixed across markets (dashed) against an average over market-level bites, a normal location shifter with standard deviation 0.12 (solid), with $\varsigma = 0$. Panel (b): the averaged curves with $\varsigma = 0.165$ (solid) against $\varsigma = 0$ (faint).

Appendix B Tables

The following tables present summary statistics or regression results. Across specifications, the tables vary by sample, controls and fixed-effects structure, and, in some cases, introduce additional interactions and F-tests.

Table B1: Effects of minimum wage and labour market concentration on retention, excluding firm VA per worker, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0255**	-0.0684***	-0.0752***
	(0.0125)	(0.0106)	(0.0114)
HHI (δ)	0.0018	0.0065**	0.0066**
	(0.0096)	(0.0031)	(0.0030)
wagebite $\times$ HHI (γ)	0.0533	0.0288**	0.0323**
	(0.0326)	(0.0138)	(0.0128)
Constant (α)	0.4908***	0.6429***	0.7358***
	(0.0086)	(0.0063)	(0.0046)
F-test: $\beta + \gamma = 0$	0.60	12.17	18.89
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2020	0.2775	0.2818
Observations	6,831,528	6,760,167	6,759,519

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. The constant is the regression intercept, with coefficient α. We consider years 2002 through 2018. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to temporary workers and continuing firms, excluding workers earning from the minimum wage in t+1 up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test rows report the F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B2: Effects of minimum wage and labour market concentration on retention by contract type

	Temporary workers			Permanent workers		
	(1)	(2)	(3)	(4)	(5)	(6)
wagebite (β)	-0.0255**	-0.0638***	-0.0795***	-0.0250***	-0.0031	-0.0102***
	(0.0125)	(0.0144)	(0.0153)	(0.0019)	(0.0022)	(0.0020)
HHI (δ)	0.0018	0.0102***	0.0080**	0.0097***	0.0041***	0.0018
	(0.0096)	(0.0039)	(0.0038)	(0.0012)	(0.0013)	(0.0011)
wagebite $\times$ HHI (γ)	0.0533	0.0367**	0.0473***	0.0119***	-0.0008	-0.0009
	(0.0326)	(0.0174)	(0.0169)	(0.0039)	(0.0047)	(0.0042)
Firm VA per worker		0.0122***	0.0126***		0.0107***	0.0120***
		(0.0016)	(0.0016)		(0.0010)	(0.0008)
Constant (α)	0.4908***	0.5351***	0.6107***	0.8465***	0.8911***	0.9119***
	(0.0086)	(0.0170)	(0.0160)	(0.0025)	(0.0101)	(0.0089)
F-test: $\beta + \gamma = 0$	0.60	4.16	6.33	15.92	0.78	9.55
Worker FEs	Yes	Yes	Yes	Yes	Yes	Yes
Firm FEs	No	Yes	Yes	No	Yes	Yes
CBA $\times$ Year FEs	No	No	Yes	No	No	Yes
Adj. R^2	0.2020	0.2804	0.2849	0.1417	0.1973	0.2038
Observations	6,831,528	5,279,730	5,279,224	18,640,686	13,506,298	13,506,095

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. We consider years 2002 through 2018. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to continuing firms, excluding workers earning from the minimum wage in $t+1$ up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test rows report the F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B3: Effects of minimum wage and labour market concentration on retention, including firm-by-year fixed effects, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0255**	-0.0638***	-0.0933***
	(0.0125)	(0.0144)	(0.0172)
HHI (δ)	0.0018	0.0102***	0.0081**
	(0.0096)	(0.0039)	(0.0034)
wagebite $\times$ HHI (γ)	0.0533	0.0367**	0.0515***
	(0.0326)	(0.0174)	(0.0173)
Firm VA per worker		0.0122***	
		(0.0016)	
Constant (α)	0.4908***	0.5351***	0.7230***
	(0.0086)	(0.0170)	(0.0052)
F-test: $\beta + \gamma = 0$	0.60	4.16	10.96
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	No
Firm $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2020	0.2804	0.3207
Observations	6,831,528	5,279,730	6,114,945

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. We consider years 2002 through 2018. Firm-by-year FEs are included instead of collective-agreement-by-year FEs. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to temporary workers and continuing firms, excluding workers earning from the minimum wage in t+1 up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test rows report the F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B4: Effects of minimum wage and labour market concentration on retention, including a dummy indicating whether the national minimum wage is binding, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0279**	-0.0666***	-0.0792***
	(0.0122)	(0.0152)	(0.0158)
HHI (δ)	0.0032	0.0063*	0.0050
	(0.0102)	(0.0037)	(0.0037)
wagebite $\times$ HHI (γ)	0.0476	0.0432**	0.0525***
	(0.0334)	(0.0185)	(0.0175)
NMW binding	-0.0499***	-0.0408***	-0.0394***
	(0.0092)	(0.0067)	(0.0057)
NMW binding $\times$ HHI	-0.0111	0.0444***	0.0367***
	(0.0163)	(0.0145)	(0.0129)
NMW binding $\times$ *wagebite*	-0.2650***	-0.0614	-0.0568
	(0.0656)	(0.0552)	(0.0520)
NMW binding $\times$ *wagebite* $\times$ HHI	0.1386	-0.1281	-0.1431
	(0.1313)	(0.1497)	(0.1205)
Firm VA per worker		0.0123***	0.0126***
		(0.0017)	(0.0016)
Constant (α)	0.4918***	0.5357***	0.6158***
	(0.0085)	(0.0172)	(0.0160)
Joint significance F-test	6.10	5.28	4.77
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2030	0.2806	0.2851
Observations	6,831,528	5,279,730	5,279,224

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. NMW binding is an indicator that the national minimum wage is the binding minimum wage. NMW binding $\times$ HHI interacts the NMW-binding indicator with the new hires HHI. NMW binding $\times$ *wagebite* interacts the NMW-binding indicator with the bite measure. NMW binding $\times$ *wagebite* $\times$ HHI is the triple interaction of the NMW-binding indicator, the bite measure, and the new hires HHI. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. We consider years 2002 through 2018. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to temporary workers and continuing firms, excluding workers earning from the minimum wage in t+1 up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test rows report the joint significance F-test for the NMW-binding interactions: NMW binding $\times$ *wagebite*, NMW binding $\times$ HHI, and NMW binding $\times$ *wagebite* $\times$ HHI. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B5: Effects of minimum wage and labour market concentration on retention, including a dummy for high unemployment periods (2009-2018), temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0276**	-0.0545***	-0.0768***
	(0.0111)	(0.0107)	(0.0088)
HHI (δ)	-0.0031	-0.0062	0.0016
	(0.0090)	(0.0052)	(0.0045)
wagebite $\times$ HHI (γ)	-0.0030	0.0363	0.0341*
	(0.0235)	(0.0264)	(0.0206)
High unemployment period	0.0195***	0.0132***	
	(0.0037)	(0.0033)	
High unemployment period $\times$ HHI	0.0048	0.0198***	0.0088*
	(0.0098)	(0.0057)	(0.0052)
High unemployment period $\times$ *wagebite*	0.0106	-0.0143	-0.0050
	(0.0149)	(0.0180)	(0.0179)
High unemployment period $\times$ *wagebite* $\times$ HHI	0.0865**	0.0093	0.0226
	(0.0435)	(0.0349)	(0.0286)
Firm VA per worker		0.0122***	0.0126***
		(0.0016)	(0.0016)
Constant (α)	0.5091***	0.5477***	0.6110***
	(0.0087)	(0.0168)	(0.0160)
Joint significance F-test	7.22	5.19	1.83
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2022	0.2805	0.2849
Observations	6,831,528	5,279,730	5,279,224

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. High unemployment period is an indicator for 2009–2018. High unemployment period $\times$ HHI interacts the high-unemployment indicator with the new hires HHI. High unemployment period $\times$ *wagebite* interacts the high-unemployment indicator with the bite measure. High unemployment period $\times$ *wagebite* $\times$ HHI is the triple interaction of the high-unemployment indicator, the bite measure, and the new hires HHI. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. We consider years 2002 through 2018. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to temporary workers and continuing firms, excluding workers earning from the minimum wage in t+1 up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test row reports the joint significance F-test for the high-unemployment interactions: High unemployment period $\times$ *wagebite*, High unemployment period $\times$ HHI, and High unemployment period $\times$ *wagebite* $\times$ HHI. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B6: Effects of minimum wage and labour market concentration on retention, including an age dummy for young workers, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0252**	-0.0621***	-0.0781***
	(0.0124)	(0.0128)	(0.0135)
HHI (δ)	0.0004	0.0090**	0.0072*
	(0.0097)	(0.0037)	(0.0037)
wagebite $\times$ HHI (γ)	0.0532	0.0359**	0.0458***
	(0.0339)	(0.0167)	(0.0159)
Young	-0.0437***	-0.0138	-0.0100
	(0.0066)	(0.0127)	(0.0126)
Young $\times$ HHI	0.0164	0.0376	0.0343*
	(0.0124)	(0.0235)	(0.0203)
Young $\times$ *wagebite*	0.0333	-0.0340	-0.0468
	(0.0411)	(0.0703)	(0.0703)
Young $\times$ *wagebite* $\times$ HHI	0.0032	0.0094	0.0360
	(0.0726)	(0.1182)	(0.1103)
Firm VA per worker		0.0121***	0.0126***
		(0.0016)	(0.0016)
Constant (α)	0.4930***	0.5327***	0.6042***
	(0.0090)	(0.0173)	(0.0178)
F-test: Young main effects = 0	0.03	0.00	0.03
Joint significance F-test	1.67	1.87	2.57
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2023	0.2804	0.2848
Observations	6,819,231	5,270,838	5,270,332

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. Young is an indicator for workers aged 17 to 19. Young $\times$ HHI interacts the young-worker indicator with the new hires HHI. Young $\times$ *wagebite* interacts the young-worker indicator with the bite measure. Young $\times$ *wagebite* $\times$ HHI is the triple interaction of the young-worker indicator, the bite measure, and the new hires HHI. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. We consider years 2002 through 2018. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to temporary workers and continuing firms, excluding workers earning from the minimum wage in t+1 up to 10% above it and workers aged 16 or less. Standard errors are clustered at the firm level and shown in parentheses. F-test rows report the F-test for the null that the combined Young main effects equal zero; the joint significance F-test for the Young interactions: Young $\times$ *wagebite*, Young $\times$ HHI, and Young $\times$ *wagebite* $\times$ HHI. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B7: Effects of minimum wage and labour market concentration on retention, including a dummy for firm-years with a low share of bitten workers, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0032	-0.0427**	-0.0592***
	(0.0118)	(0.0181)	(0.0189)
HHI (δ)	0.0126*	0.0148**	0.0142**
	(0.0070)	(0.0063)	(0.0057)
wagebite $\times$ HHI (γ)	0.0235	0.0247	0.0292
	(0.0249)	(0.0231)	(0.0230)
Low share of bitten workers	0.0128***	0.0085**	0.0072***
	(0.0029)	(0.0035)	(0.0027)
Low share of bitten workers $\times$ HHI	-0.0180	-0.0067	-0.0098*
	(0.0114)	(0.0062)	(0.0051)
Low share of bitten workers $\times$ *wagebite*	-0.0382***	-0.0410***	-0.0399***
	(0.0120)	(0.0115)	(0.0107)
Low share of bitten workers $\times$ *wagebite* $\times$ HHI	0.0470	0.0147	0.0268
	(0.0341)	(0.0282)	(0.0227)
Firm VA per worker		0.0123***	0.0126***
		(0.0016)	(0.0015)
Constant (α)	0.4833***	0.5283***	0.6064***
	(0.0088)	(0.0164)	(0.0159)
Joint significance F-test	8.12	5.69	6.21
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2020	0.2804	0.2849
Observations	6,831,528	5,279,730	5,279,224

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. 'Low share of bitten workers' refers to a dummy variable which takes the value one if the share of the firm workforce being bitten in t is at or below the median share, across firms, in t. We consider years 2002 through 2018. All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to continuing firms, excluding workers earning from the minimum wage in $t+1$ up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test row reports the joint significance F-test for the low share of bitten workers interactions: Low share of bitten workers $\times$ *wagebite*, Low share of bitten workers $\times$ HHI, and Low share of bitten workers $\times$ *wagebite* $\times$ HHI. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B8: Effects of minimum Wage and labour market concentration on retention in the wholesale and retail, repair of motor vehicles and motorcycles industry, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.2059**	-0.2605**	-0.2955**
	(0.0803)	(0.1127)	(0.1179)
HHI (δ)	0.0040	-0.0071	-0.0097
	(0.0063)	(0.0079)	(0.0084)
wagebite $\times$ HHI (γ)	0.1171**	0.1430**	0.1488***
	(0.0520)	(0.0562)	(0.0543)
Firm VA per worker		0.0128***	0.0136***
		(0.0038)	(0.0038)
Constant (α)	0.6604***	0.6635***	0.7413***
	(0.0182)	(0.0452)	(0.0444)
F-test: $\beta + \gamma = 0$	3.55	2.30	2.93
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2172	0.2435	0.2470
Observations	767,756	704,528	704,182

Notes: The dependent variable is worker staying at the same firm between t and $t+1$. *wagebite* is the bite measure, with coefficient β. HHI is the new hires HHI, with coefficient δ. *wagebite* $\times$ HHI is the interaction between the bite measure and the new hires HHI, with coefficient γ. Firm VA per worker refers to log firm value added per worker. The constant is the regression intercept, with coefficient α. We consider years 2007 through 2018 and focus solely on the Wholesale and Retail; Repair of Motor Vehicles industry (Classification of Economic Activity codes 45, 46 and 47). All specifications additionally include tenure, tenure squared, potential experience, and potential experience squared. The analysis is restricted to continuing firms, excluding workers earning from the minimum wage in $t+1$ up to 10% above it. Standard errors are clustered at the firm level and shown in parentheses. F-test rows report the F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Appendix C Figures

Figure C1: Collective Bargaining Minimum Wages Above the NMW, Fraction of Workers



Notes: The figure reports the fraction of workers whose relevant minimum wage is a collective bargaining agreement minimum wage set above the NMW, from 2002 to 2023. The fraction is weighted by the number of workers attached to each CBA. Separate series are shown for all workers, permanent-contract workers and temporary-contract workers.

Figure C2: HHI Time Series by Contract Type



Notes: The figure plots the HHI time series by contract type from 2002 to 2023. Values are weighted by employment.