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

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Abstract

How are *incumbent* workers impacted by minimum wage increases? Is labour market concentration an important moderating factor? We address these questions by presenting a new model, drawing on comprehensive employer-employee matched panel data from Portugal, and considering individual mandated wage rises from both national and collectively bargained minimum wages. We find that employment protection plays a key role in these effects: workers under permanent contracts are almost completely insulated from the exit hazards associated with 'wage bites' while workers in temporary (fixed-term) contracts are exposed to these hazards. Labour concentration can significantly reduce negative effects on temporary workers, by up to 60%. A model of monopsonistic competition with firing costs rationalises these patterns.

JEL classification

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minimum wages, collective bargaining, temporary contracts, labour market power

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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 try to 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 will 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 recruitments in the relevant occupation and local labour market.

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 monopsonistic competition, adding idiosyncratic shocks to match productivity. Where the minimum wage binds, a firm is pushed onto its labour demand schedule and contracts, so exit hazards rise with further wage rises despite its monopsony power. Where it does not, the firm remains constrained by labour supply and 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 employment rises. 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.

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 theoretical model; and Section 5 concludes.

2 Institutional Background 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 (see Table B17 for a time series). 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]

2.2 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 sur-

[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 bargaing in Portugal.

vey 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):

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

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 ("*wagebite*")

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.

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.

[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.

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 these 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 do not suffer 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:

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

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 $t \rightarrow t+1$ 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	.048	$5,279,224$
0-10% excluded	(38.28)	(-5.21)	(2.09)	(2.80)	
2) Temporary workers	.636	$-.075$	.008	.046	$4,254,935$
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)	(2.59)	
4) Permanent workers	.912	$-.010$	.002	$-.01$	$13,506,095$
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-bargain-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 we may conclude from this 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.[9] 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 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

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

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 0.4% arising from the minimum wage hike although unlike us they do not consider explicitly the effects of monopsony power. 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 product 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 along with some other supplementary sensitivity analyses not discussed here.

First, dropping value added as a control (Table B5) 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 B13) 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 B16). Aside from a slight increase in β the results were largely unchanged.[10] 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 B11). 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.[11] 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 B12) 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. We found no quantitatively or statistically significant changes in the coefficients aside 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

[10]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.

[11]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.

workers in 1986 caused exit hazards to fall but who do not consider the role of labour market power.[12]

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 focused solely on the largest low-wage sector, namely "Wholesale and Retail Trade; Repair of Motor Vehicles". Results for the first exercise (Table B15) show no significant difference in the β and γ estimates for firms with high/low shares of bitten workers. The second exercise (Table B14) whilst reaffirming our core result has estimates that are 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.

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

[12] 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.

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.[13]

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 $1/(1-\alpha)$.[14] 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 floor to R.

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

The second is the markdown itself, acting as a *retention buffer*. Incumbent matches carry match-specific capital $\kappa > 0$, which a new hire lacks in her first period, and receive negative idiosyncratic productivity shocks $-\varepsilon$ ($\varepsilon \ge 0$), $\varepsilon \sim G$, at rate λ. In the supply-determined regime all the labour forthcoming at the wage floor is employed, so a slot

[13] 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.

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

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

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

where $x(b)$ is the log wedge between what the marginal worker produces and what she is paid—the unexhausted markdown. 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.[15] A floor increase shrinks the buffer, and matches whose existing ε lies between the old and new thresholds are immediately severed. 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.

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: an expanding firm retains all its incumbents, but hiring does not undo an incumbent's exit, so that employment and retention can move in opposite directions. A contracting firm lays off *additional* incumbents only where the contraction exceeds the turnover it experiences in any case—quits plus the separations the shock process already generates. 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 almost entirely buffer erosion.

4.3 Predictions

For a single firm at a fixed bite, the profile implied by (3)–(4) is piecewise flat—concentration matters only through which regime the firm occupies—matching the smooth profile in the data comes from aggregation. 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 a mixture over

[15] 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.

the bite distribution. As concentration raises b^*, firms leave the downsizing regime—more firms are supply constrained—and the aggregate response declines smoothly, while the buffer erosion 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.[16]

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 $\beta + \gamma \leq 0$. The first and third hold generally; 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.4). 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. That our estimates respect it says the quit margin may be dominated for bitten temporary workers in our sample.[17]

Figure 1 evaluates the mixture at each decile's empirical bite distribution against the estimated profile.[18] The downsizing channel alone delivers too big a decline at high concentration; with the buffer channel included, the composite tracks the estimates across most of the range, but the response is too large at the highest concentrations. As in Azar et al.'s simulations, the figure illustrates the mechanism rather than calibrates it; parameters and robustness are in Appendix A.

4.4 Permanent contracts

A firing cost dampens both channels: dismissal requires the per-worker loss to exceed the cost of dismissal, and the threshold $\bar{\varepsilon}$ in (4) shifts up. Write F for the amortised, surplus-equivalent dismissal cost, defined in the same log units as κ and x, so that separation

[16] It also widens with concentration, to 0.37 at the most concentrated decile, but the profile does not depend on this: holding dispersion at its least-concentrated value throughout leaves Figure 1 essentially unchanged.

[17] The linear specification is an OLS projection with fixed effects rather than an average over an exogenous bite distribution, 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 $\hat{\gamma}$.

[18] That is, how does the model match the estimated $\hat{\gamma}/\hat{\beta}$ ratio and is the model profile close to linear?

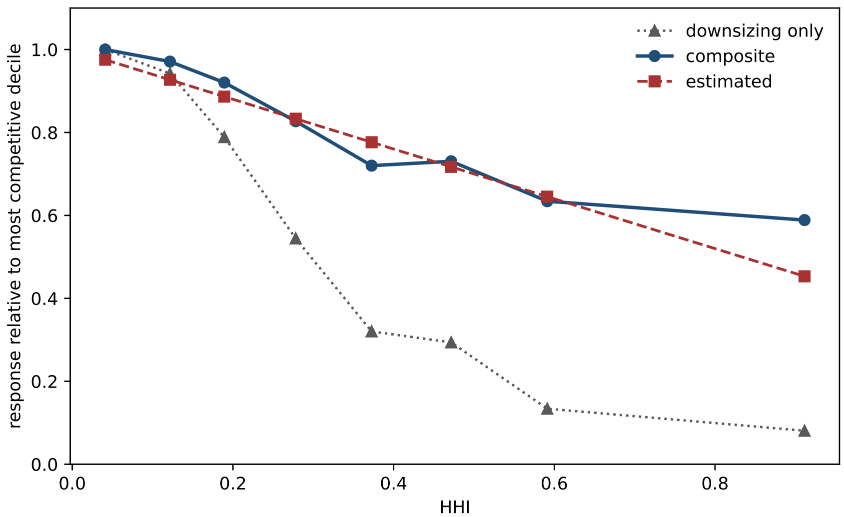

Figure 1: The marginal retention response to the bite as a function of HHI, normalized to its value at $HHI = 0$: downsizing term only (dotted), composite with the buffer term (solid), and the estimated profile $1 + (\hat{\gamma}/\hat{\beta})HHI$ from specification (1) (dashed). Markers evaluate each HHI decile of the estimation sample at its average HHI using its own fitted bite distribution.

requires $V(\varepsilon) < -F$ and the threshold shifts up by exactly F.[19] For 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) shows. Exposure is not the explanation: 29% of permanent employment lies at or below the incoming floor against 36% 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

[19] A dismissal cost changes the continuation comparison rather than simply translating the threshold; F is defined as the surplus-equivalent shift that this induces, which is the object the argument uses.

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 in Azar et al. (2024) to exit hazards is readily able to account for key aspects of the empirical findings.

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Appendix A Theory appendix

This appendix gives the derivations behind Section 4: the regime structure and the location of the switch point (Appendix A.1), the retention problem and its comparative statics (Appendix A.2), the availability of replacements by regime (Appendix A.3), aggregation over heterogeneous bites and the mapping to the estimating equation (Appendix A.4), and the parameters and robustness of the illustration (Appendix A.5). We stress that as in Azar et al. (2024), this is primarily an indicative exercise (and we use their parameterisations); we also stress that the analysis considers only the firm demand side, and does not model worker quit decisions, which are assumed to be independent of the minimum wage.

Appendix A.1 The regimes and the switch point

The switch point—when the bite is of sufficient size to push the firm onto its demand curve—in (3) is $b^* = m/\rho$ rather than m: the wage floor exhausts the markdown buffer strictly before it reaches the full markdown. The reason is that expansion along the supply curve itself erodes the buffer. A firm whose floor lies in the supply-determined range hires the additional labour the higher floor supplies, and in doing so raises the marginal cost of its now higher workforce, so each unit of bite consumes $\rho = 1 + \eta(1 - \alpha) > 1$ units of buffer rather than one. The η term is the wage bill effect of moving along the supply curve and the $1 - \alpha$ term the fall in the marginal product as employment rises.

Two features of the supply-determined regime are worth stating explicitly, since they are what allow us to derive the closed form in (3). The symmetric outcome is well defined.[20] And because b^* is increasing in m and hence in concentration, a floor of given bite is demand-determined in competitive markets and supply-determined in concentrated ones, which is the comparative static the empirical specification exploits.

[20]Firms are Cournot competitors in labour demands so in a symmetric equilibrium each firm will choose $1/J$ of market supply, any higher demand leading to a jump upward in marginal costs, However note that there are a continuum of asymmetric equilibria.

Appendix A.2 The retention problem

We work in log-productivity units, so κ, ε and x are all in logs, as are b and m. An incumbent match yields log output $\text{mrpl} + \kappa$, where $\kappa > 0$ reflects match-specific human capital, less an idiosyncratic component $\varepsilon \geq 0$ drawn from G. The draw is redrawn at Poisson rate λ and is otherwise persistent.[21] A new hire lacks κ in her period of hire and holds it thereafter, and we assume that replacing an incumbent costs a one-off hiring cost $H > 0$. Neither bears on the supply-determined regime, where there is no replacement to be had; Appendix A.3 works out their role in the demand-determined case, and the restriction we make that ties H to κ. Matches also end for reasons unrelated to the buffer—quits, retirements and the like—at rate ς, so a match survives to the next decision with probability $\sigma \equiv 1 - \varsigma$.[22]

Steady state in the supply-determined regime. Consider a fixed wage floor. Since no replacement is available at the floor in this regime (Appendix A.3), the value of retaining a match with current draw ε solves

$$V(\varepsilon) = \big(\kappa + x - \varepsilon\big) + \frac{\sigma}{1+r}\Big[(1-\lambda)V(\varepsilon) + \lambda \int \max\{V(\varepsilon'), 0\}\, dG(\varepsilon')\Big], \tag{5}$$

with discount rate r, the alternative being a vacant slot worth zero. The continuation carries the survival probability σ as well as the discount factor, because the match may end for other reasons before the next period.[23] The firm separates when $V(\varepsilon) < 0$, that

[21] Both polar alternatives fail. With a once-and-for-all draw at match formation the incumbent stock contains no unimpaired matches: every survivor carries a live draw, and survivors are precisely those whose draws lie below the reservation threshold. The workforce is therefore bunched immediately beneath that threshold — at our calibration some fifty times denser there than the underlying distribution — so a small fall in the threshold severs an implausibly large share of it. The relevant object in the aggregation of Appendix A.4 then becomes the reverse hazard (i.e., g/G) rather than the density, and the regularity condition there is met only for an incumbency premium far larger than is credible for these workers. At the other extreme, with an i.i.d. redraw every period, a bad draw carries no information about the future, and since a released slot cannot be refilled in the supply-determined regime the firm rides out transitory losses and almost never separates except for extreme draws. Persistent shocks with stochastic arrival, as in Mortensen & Pissarides (1994), avoid either problem.

[22] Recall we treat the quit rate as being independent of the wage floor.

[23] The flow payoff is stated in log units: exact profit is proportional to $e^{\kappa+x-\varepsilon} - 1$, and (5) is its first-order approximation. Nothing structural turns on this, since the payoff depends on $\kappa + x - \varepsilon$ through a single monotone index, so the exact problem still yields a reservation draw of the form $\bar{\varepsilon} = \kappa + x + \hat{\Omega}$ with $\hat{\Omega} \geq 0$. The zero outside option is specific to this regime; in the demand-determined regime a replacement is available and the alternative is her value rather than zero; see Appendix A.3.

is above a reservation draw $\bar{\varepsilon}$ with $V(\bar{\varepsilon}) = 0$:

$$\bar{\varepsilon} = \kappa + x + \frac{\sigma\lambda}{1+r}\,\Omega, \qquad \Omega = \int \max\{V(\varepsilon'), 0\}\, dG(\varepsilon') \;\geq\; 0. \tag{6}$$

The firm thus retains workers who lose money today ($\kappa + x - \varepsilon < 0$) as long as the option value Ω of a future favourable redraw offsets the current loss. To streamline the exposition we make a simplification by setting $\Omega = 0$ throughout, so that $\bar{\varepsilon} = \kappa + x$ and $d\bar{\varepsilon}/dx = 1$.[24]

A new hire enters with $\varepsilon = 0$ and maintains this until the first shock arrival. Since attrition and shock arrival are both independent of the existing draw, the stationary incumbent stock is a mixture of never-shocked matches and shocked survivors, the latter distributed as G truncated at $\bar{\varepsilon}$:

$$\underbrace{1 - \varphi G(\bar{\varepsilon})}_{\text{mass at } \varepsilon=0} \qquad \text{and} \qquad \underbrace{\varphi\, g(\varepsilon) \text{ on } (0, \bar{\varepsilon}]}_{\text{shocked survivors}}, \qquad \varphi \equiv \frac{\sigma\lambda}{1 - \sigma + \sigma\lambda}. \tag{7}$$

With $\varsigma = 0$ we have $\varphi = 1$ and the mass at zero is simply $1 - G(\bar{\varepsilon})$; faster attrition lowers φ, so the stock is younger and the never-shocked mass larger. A shock arrives with probability λ and is survived with probability $G(\bar{\varepsilon})$, so among matches that survive attrition the buffer-channel separation flow, and the total separation rate, are

$$s = \lambda\big(1 - G(\bar{\varepsilon})\big), \qquad s^{\text{tot}} = \varsigma + \sigma s. \tag{8}$$

It is s^{tot} that is observed, and matching it to the measured hazard is what pins down λ in Appendix A.5.

Separated workers re-enter supply, but in the supply-determined regime the pool is rationed: an extra separation does not buy an extra replacement, and the vacancy it creates goes unfilled. We also assume a separated worker cannot immediately return to her own firm. Both claims are worked out in Appendix A.3. Since $\bar{\varepsilon}$ is increasing in x, a deeper markdown buffers a larger shock before separation, so s falls with x. As $x = \max\{m - \rho b, 0\}$ is decreasing in the bite and increasing in concentration, the separation flow rises in the bite and, for a given bite, falls in concentration. Only the signs are used in the text; the magnitudes in Appendix A.5 require values for G, κ and λ, of which only the shock distribution is fitted.

[24] Solving the fixed point instead gives $\bar{\varepsilon} = \kappa + x + \frac{\sigma\lambda}{(1+r)\Delta}\int_0^{\bar{\varepsilon}} G$, where $\Delta \equiv 1 - \sigma(1-\lambda)/(1+r)$ is an amortisation factor that recurs in Appendix A.3, and $d\bar{\varepsilon}/dx = \big(1 - \frac{\sigma\lambda}{(1+r)\Delta} G(\bar{\varepsilon})\big)^{-1} > 1$: a deeper buffer raises the value of every future match, which raises the option value, which then widens the threshold further. The simplification therefore understates both the level of the threshold and its response to the wedge. Refitting the shock distribution absorbs almost all of this and leaves Figure 1 essentially unchanged; the calculation is available on request.

Impact of a wage floor increase. In the supply-determined-regime, consider a one-off increase in the wage floor from b_0 to b_1. The buffer drops from x_0 to $x_1 < x_0$ and the reservation draw from $\bar{\varepsilon}_0$ to $\bar{\varepsilon}_1$. Matches whose existing ε lies in $(\bar{\varepsilon}_1, \bar{\varepsilon}_0]$ are severed at once, without waiting for a fresh shock, since the firm re-evaluates its existing matches against the new floor. Never-shocked matches are not "culled", since their draw is zero, so by (7) the fraction of the stock culled is

$$c = \varphi\big[G(\bar{\varepsilon}_0) - G(\bar{\varepsilon}_1)\big] \ \in [0,1] \tag{9}$$

— an impact effect rather than a flow effect.[25]

In the demand-determined regime the reservation draw is $\bar{\varepsilon} = \kappa$, independent of the floor (Appendix A.3): replacing an incumbent costs H and starts the new match without κ, so a small adverse draw is still worth absorbing. For a firm already past the boundary $c = 0$ and s is unchanged, leaving downsizing as the only response. A firm the increase carries across the boundary culls in the ordinary way, against a threshold falling from $\kappa + x_0$ to κ, which vanishes continuously as the boundary is approached.

Combining the two channels. Retention of incumbents in the first-period after the floor increase is

$$R_1 = \min\big\{\sigma(1-c)(1-s_1),\ e^{\Delta \ln L}\big\}, \qquad s_1 = \lambda\big(1 - G(\bar{\varepsilon}_1)\big), \tag{10}$$

the surviving fraction of the incumbent stock, unless that exceeds the firm's new employment target, in which case the firm sheds the difference. Two separate groups leave through the buffer channel. The cull c re-evaluates the draws incumbents are already carrying; s_1 is the flow of fresh draws that fail the new threshold, and it responds to the floor because $\bar{\varepsilon}$ has fallen. We treat the cull, attrition and fresh arrivals as independent within the period. The minimum operator rather than a sum is what prevents double counting: shock-driven separations count toward any required contraction instead of adding to it. Equivalently, a downsizing firm sheds its lowest-surplus incumbents first. Setting $c = 0$ and $s_1 = s_0$ recovers the pure downsizing case, measured against the counterfactual retention $\sigma(1 - s_0)$.

[25] λ enters c only through φ. Conditional on the thresholds and with $\varsigma = 0$ it would drop out entirely, since the composition of shocked survivors is a G draw conditioned on $\varepsilon \le \bar{\varepsilon}$ whatever the arrival rate; with $\varsigma > 0$ the arrival rate also governs how much of the stock has been shocked at all. In the general model the thresholds depend on λ as well, through the option value; under the simplification $\Omega = 0$ adopted here they do not, so φ and s are the only routes by which λ enters.

Retention cannot exceed its no-bite value at any level of concentration: $c \geq 0$, (10) is bounded above by $\sigma(1 - s_0)$, and $x(b) \leq m$ ensures the buffer is never wider than in the absence of a floor. This is the source of the restriction $\beta + \gamma \leq 0$ in the text.

Appendix A.3 Replacement availability by regime

The retention decision compares a shocked incumbent against the firm's alternative to retaining her, and the alternative differs by regime.

In the supply-determined regime labour is scarce at the wage floor: the firm already wants more workers than it can obtain, so a released slot stays empty and the loss is the wedge between what the worker produces and what she is paid. The firm absorbs a large shock before separating, and the bigger the wedge the larger the shock it absorbs. In the demand-determined regime there is a queue at the floor, a replacement is hired at once, and all that is forgone is the cost of hiring her and the fact that she is new.

Dobbin et al. (2026), not in a minimum wage context, but where firms post wages, obtain a related result: 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. In Arabzadeh et al. (2024) the wedge is the cost of financing the wage bill and 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.

Formally, recall that $H > 0$ is the cost of hiring a replacement. In this regime the firm sits on its demand curve, so the buffer is exhausted and $x = 0$ in (4); the flow payoffs below therefore carry no wedge. A replacement produces her wage in her period of hire and is an incumbent thereafter, so in surplus units

$$V(\varepsilon) = (\kappa - \varepsilon) + \frac{\sigma}{1+r}\big[(1-\lambda)V(\varepsilon) + \lambda\tilde{\Omega}\big], \tag{11}$$

$$V_H = -H + \frac{\sigma}{1+r}\big[(1-\lambda)V(0) + \lambda\tilde{\Omega}\big], \qquad \tilde{\Omega} = \int \max\{V(\varepsilon'), V_H\}\, dG(\varepsilon'). \tag{12}$$

Here V is the counterpart of (5) for this regime rather than the same function: the alternative to retention is the replacement's value V_H and not zero, and the flow payoff carries no wedge. Nor is the option value set to zero, as it was after (6); here it need not be, since it cancels.

Rearranging (11) gives $\Delta \times V(\varepsilon) = (\kappa - \varepsilon) + \sigma\lambda\tilde{\Omega}/(1+r)$, with Δ the amortisation factor of footnote 24, so V is linear in the draw, $V(\varepsilon) = V(0) - \varepsilon/\Delta$. Evaluating (11) at $\varepsilon = 0$ and subtracting (12), the continuation terms are identical—both run through $V(0)$ and $\tilde{\Omega}$—so they cancel, leaving $V(0) - V_H = \kappa + H$: the incumbent's advantage over a replacement is her premium plus the hiring cost she saves. The firm separates where the draw has eroded exactly that advantage, $V(\bar{\varepsilon}_D) = V_H$, which by linearity is $\bar{\varepsilon}_D/\Delta = \kappa + H$, so

$$\bar{\varepsilon}_D = (\kappa + H)\,\Delta. \tag{13}$$

Here $H > 0$ is essential: at $H = 0$ the threshold is $\kappa\Delta$ and a firm facing a queue replaces almost every worker who is shocked. Adding a common wedge to both flow payoffs would leave (13) unchanged, since the replacement would enjoy the same wedge: it cancels from $V(0) - V_H = \kappa + H$ and hence from the threshold. It is the unavailability of a replacement, not the markdown, that protects the incumbent. The invariance is more than interpretive. Setting $x = 0$ is exact only if the boundary is located exactly, and the treatment of the hiring margin below locates it only approximately; but since (13) carries no wedge, the demand-regime threshold—and with it s and the absence of a cull—is unaffected by that imprecision.

In the supply-determined regime there is no one to hire at the floor. Every firm is quantity-constrained at the kink: the marginal product exceeds the wage, so each would hire more at $w_{\min}$ and cannot. Separations re-enter the pool, but the pool is in excess demand, so an additional vacancy goes unfilled. No matching friction is required. The alternative to retention is a vacant slot, H is never incurred at the margin, and the threshold is $\bar{\varepsilon}_S = \kappa + x$ as in (6).

The two thresholds are constructed against different alternatives and need not agree at $x = 0$, in which case a firm crossing the boundary would re-evaluate its workforce against a different standard and shed or retain a block of matches at once. We restrict H to remove this,

$$H = \kappa\,\frac{1 - \Delta}{\Delta}, \tag{14}$$

so that $\bar{\varepsilon}_D = \kappa$ and the reservation draw runs continuously from κ at the boundary to $\kappa + x$ in the interior. This restricts a parameter not otherwise pinned down, but the profile in Appendix A.5 is (visually) insensitive to departures from it by a factor of two or three. The sharp switch is in any case an artefact of deterministic rationing: a refill probability rising continuously to one as excess demand vanishes would deliver continuity for any H, at the cost of a vacancy value we cannot discipline.

The assumption that a separated worker does not immediately return to her own firm is not an additional restriction: shocks are match-specific and persistent, so a worker dismissed on a bad draw carries it with her, and rehiring her is pointless by construction.

Both κ and H affect the hiring margin, in opposite directions. Either impacts the no-floor equilibrium—and with it w^*, against which the bite is defined—but leaves the regime structure and both elasticities of (3) intact up to terms of order $\kappa + H$, which we neglect; in the supply-determined regime the hiring margin is moot, since employment is pinned down by supply.

Appendix A.4 Aggregation

At a given HHI the estimated coefficient is a mixture over the bite distribution $F(\cdot \mid HHI)$:

$$\beta(HHI) \;\approx\; \underbrace{-\frac{1}{1-\alpha}\big[1 - F\big(b^*(HHI)\big)\big]\,\Pr\{z > \bar{z}\}}_{\text{downsizing}} \;-\; \underbrace{(\varphi+\lambda)\rho \int_0^{b^*} g\big(\bar{\varepsilon}(b)\big)\, dF(b)}_{\text{buffer erosion}}, \qquad (15)$$

with $\bar{\varepsilon}(b) = \kappa + x(b)$ from (6). The coefficient has the two pieces of (10): φ for the culling of pre-existing draws, λ for the fresh draws that fail the lowered threshold.

The factor $\Pr\{z > \bar{z}\}$ carries the turnover margin of (10) into the mixture. A firm past the boundary sheds no incumbent unless its contraction exceeds the turnover it experiences anyway, which for a demand-regime firm ($c = 0$, s fixed) requires an increment above

$$\bar{z} = -(1-\alpha)\ln\big[(1-\varsigma)(1-s)\big], \qquad (16)$$

because a match ended by a bad draw vacates a position just as a quit does. Since $(1-\varsigma)(1-s)$ is one minus the observed separation rate, (16) is fixed by measurement; at our estimates it sits just above the mean bite. We treat the increment as independent of the level bite; were they perfectly correlated the effective boundary would be $\max\{b^*, \bar{z}\}$ and the profile non-monotone at low concentration.

The first term is a survival function of the boundary: as concentration raises b^*, firms drop out of the downsizing regime and the layoff response declines smoothly. The second keeps the response strictly negative where none remain. Differentiating (15), $\gamma(HHI) > 0$ requires g declining over the buffer range and

$$\frac{\Pr\{z > \bar{z}\}}{1-\alpha} \;\geq\; (\varphi+\lambda)\rho\, g(\kappa), \qquad (17)$$

so that firms crossing the boundary lose more from the lapsing layoff channel than they gain in renewed buffer erosion. The layoff channel enters net of the turnover margin, since it lapses only for those firms whose increment exceeds (16). No shape restriction on F is needed.

Differentiating (15) also picks up terms in $\partial F/\partial HHI$, and a monotone profile requires that these composition effects not dominate. The mean share of temporary employment at or below the incoming floor is flat across deciles (0.37 to 0.35) while a Kaitz-type index rises modestly; the illustration evaluates (15) at decile-specific bite distributions and so carries whatever composition is present.

Two caveats attach to the correspondence with specification (1). The model's b is a level whereas z is the current increment, and $z \leq b$, so the illustration understates the share of firms past b^*; the bound $\beta + \gamma \leq 0$ is unaffected. And (15) averages over an exogenous bite distribution whereas the estimate is an OLS projection with fixed effects. We ask the model to rationalize the sign, the bound and the direction of the heterogeneity, not to deliver a structural reading of $\hat{\beta}$ and $\hat{\gamma}$.

Appendix A.5 Illustration

Figure 1 sets $\alpha = 0.6$ and $\eta = 1.6$ (so $\rho = 1.64$), following Azar et al. (2024). Bite distributions are lognormals fitted within HHI decile to the empirical moments of firm-level average z among bitten workers, employment weighted, each decile evaluated at its average HHI; the top three deciles are collapsed in the source data. The renewal shock is exponential with rate ζ.

We set $\kappa = 0.05$ and $r = 0.05$, and take $\varsigma = 0.165$, the measured firm-wide annual separation rate across contract types. Then λ follows from matching s^{tot} in (8) to the measured annual exit hazard of bitten temporary workers, 0.331, giving $\lambda = 0.22$; and H follows from the continuity restriction (14), $H = \kappa(1 - \Delta)/\Delta = 0.08$, within the range of replacement-cost estimates for low-wage workers (Dube et al. 2010).[26] That leaves $\zeta = 0.42$ as the single fitted object, and (17) holds at every decile, $(\varphi + \lambda)\rho g(\kappa) = 0.50$ against a left-hand side running between 0.70 and 0.84.

Deciles are weighted by market counts, employment by decile being unavailable; bounding between those and weights proportional to $1/HHI$ moves λ between 0.21 and 0.22, ζ between 0.42 and 0.47, and the profile by less than 0.01 anywhere. The fit is

[26] Those estimates include the cost of learning on the job, which here is the new hire's missing κ rather than part of H, so the comparable figure is below the headline.

more exposed to ς, which enters the discounting, the composition of the stock and the split of the observed hazard, and deteriorates in both directions away from the measured value. At the most competitive decile the response is entirely downsizing and at the most concentrated 86% buffer erosion, so the profile's decline is a residual of two large opposing movements; the model tracks the first six deciles and under-attenuates at the top, predicting 0.59 against an estimated 0.45. At elasticities below the value assumed by Azar et al. (2024) the required contraction is smaller relative to ordinary turnover, so fewer firms shed additional incumbents and the profile ceases to decline monotonically. Working the other way, treating workers as perfect substitutes allows turnover to be used one-for-one to shrink; where they are not, separations fall on positions the firm still needs filled, only part of turnover is available, and (16) is an upper bound. We report the profile as indicative of the mechanism, not as evidence on its magnitude.

Figure 2 reproduces the simulation design of Azar et al. (2024)—floors set 3% above the market-specific oligopsony wage, then raised by 12%, then by 40%—for the retention object. 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 dotted line shows the change in retention for the middle increase; in a competitive market labour demand contracts about 24% and in oligopsonistic—supply-constrained—markets firms expand so all incumbents are retained; the upward sloping section reflects a move from supply-constrained to demand constrained such that employment falls.) Solid lines aggregate the same steps over a distribution of market-level bites, a normal location shifter with $\sigma = 0.12$: the kinks smooth out and the profile declines everywhere. Panel (b) shows that baseline attrition trims the levels without changing the shape.

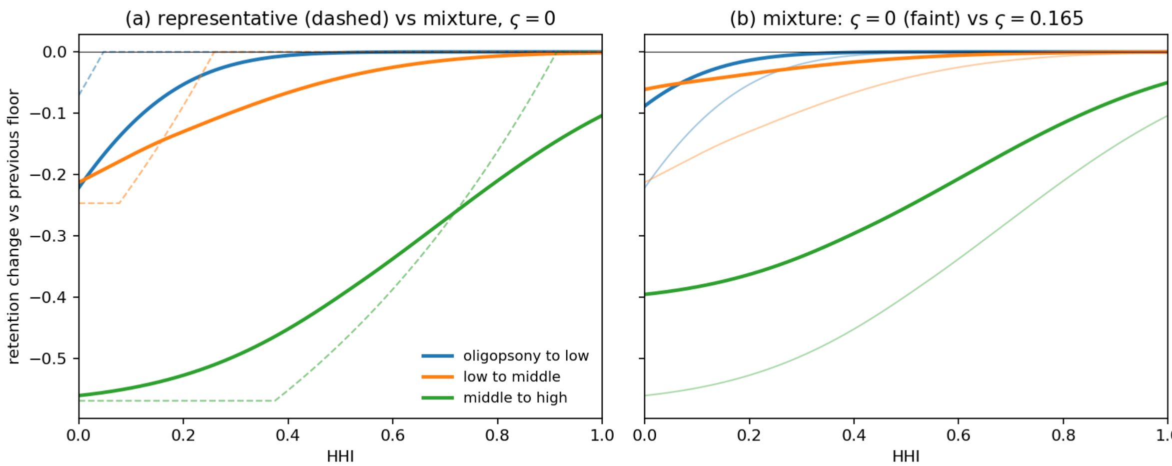


Figure 2: Retention change per minimum-wage step (relative to the previous floor), downsizing channel, by HHI, following the simulation design of Azar et al. (2024). Panel (a): fixed bite across markets (dashed) versus aggregation over a distribution of market-level bites (solid), $\varsigma = 0$. Panel (b): the aggregated curves with $\varsigma = 0.165$ (solid, using the natural wastage in the simulations for illustrative purposes) 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, extending the spillover margin to 120% of the relevant minimum wage, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0152	-0.0553***	-0.0752***
	(0.0130)	(0.0172)	(0.0181)
HHI (δ)	0.0071	0.0123***	0.0081*
	(0.0079)	(0.0044)	(0.0046)
wagebite $\times$ HHI (γ)	0.0354	0.0261	0.0389**
	(0.0294)	(0.0192)	(0.0189)
Firm VA per worker		0.0104***	0.0114***
		(0.0017)	(0.0017)
Constant (α)	0.4989***	0.5771***	0.6357***
	(0.0084)	(0.0183)	(0.0174)
F-test: $\beta + \gamma = 0$	0.41	4.47	6.87
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2042	0.2796	0.2842
Observations	5,509,298	4,255,521	4,254,935

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 temporary workers and continuing firms, excluding workers earning between the minimum wage in t+1 and 20% 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, ignoring potential spillovers of the minimum wage increase, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0159	-0.0458***	-0.0615***
	(0.0116)	(0.0114)	(0.0120)
HHI (δ)	0.0041	0.0141***	0.0130***
	(0.0111)	(0.0039)	(0.0035)
wagebite $\times$ HHI (γ)	0.0409	0.0189	0.0275*
	(0.0353)	(0.0155)	(0.0148)
Firm VA per worker		0.0123***	0.0127***
		(0.0017)	(0.0016)
Constant (α)	0.4727***	0.5093***	0.5966***
	(0.0087)	(0.0174)	(0.0160)
F-test: $\beta + \gamma = 0$	0.43	4.83	9.03
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.1984	0.2767	0.2809
Observations	9,355,969	7,200,505	7,200,127

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 temporary workers and continuing firms, including workers earning from the minimum wage in t+1 up to 10% above it. These are considered unbitten workers. 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 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 B4: Effects of minimum wage and labour market concentration on retention, considering bitten workers those earning up to 110% of their relevant minimum wage, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0139	-0.0495***	-0.0651***
	(0.0127)	(0.0120)	(0.0126)
HHI (δ)	0.0025	0.0129***	0.0119***
	(0.0114)	(0.0038)	(0.0035)
wagebite $\times$ *HHI* (γ)	0.0541	0.0328**	0.0391**
	(0.0376)	(0.0162)	(0.0151)
Firm VA per worker		0.0123***	0.0126***
		(0.0017)	(0.0016)
Constant (α)	0.4727***	0.5101***	0.5978***
	(0.0088)	(0.0175)	(0.0160)
F-test: $\beta + \gamma = 0$	0.95	1.75	5.49
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.1984	0.2767	0.2809
Observations	9,355,969	7,200,505	7,200,127

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 temporary workers and continuing firms, including workers earning from the minimum wage in t+1 up to 10% above it. These are considered bitten workers. 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 B5: 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 B6: Effects of minimum wage and labour market concentration on retention, including the change in firm VA per worker instead of the level of firm VA per worker, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0255**	-0.0629***	-0.0794***
	(0.0125)	(0.0148)	(0.0156)
HHI (δ)	0.0018	0.0115***	0.0091**
	(0.0096)	(0.0039)	(0.0039)
wagebite $\times$ HHI (γ)	0.0533	0.0351**	0.0464***
	(0.0326)	(0.0176)	(0.0171)
Δ Firm VA per worker		-0.0418***	-0.0415***
		(0.0027)	(0.0026)
Constant (α)	0.4908***	0.6559***	0.7385***
	(0.0086)	(0.0086)	(0.0057)
F-test: $\beta + \gamma = 0$	0.60	4.40	6.58
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2020	0.2823	0.2867
Observations	6,831,528	5,191,131	5,190,615

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. Δ Firm VA per worker refers to log change in 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 F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B7: Effects of minimum wage and labour market concentration on retention, including an extensive wagebite variable interacted with the firm VA per worker, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0255**	-0.0098	-0.0240***
	(0.0125)	(0.0064)	(0.0058)
HHI (δ)	0.0018	0.0112***	0.0091**
	(0.0096)	(0.0039)	(0.0038)
wagebite $\times$ HHI (γ)	0.0533	0.0254	0.0364**
	(0.0326)	(0.0155)	(0.0150)
Firm VA per worker		0.0133***	0.0141***
		(0.0016)	(0.0016)
Dummy bite $\times$ Firm VA per worker		-0.0029***	-0.0033***
		(0.0006)	(0.0006)
Constant (α)	0.4908***	0.5373***	0.6094***
	(0.0086)	(0.0172)	(0.0161)
F-test: $\beta + \gamma = 0$	0.60	1.36	0.91
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2020	0.2807	0.2853
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. Dummy bite $\times$ Firm VA per worker interacts the dummy-bite measure with 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 F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B8: Effects of minimum wage and labour market concentration on retention, including an extensive wagebite variable interacted with the change in firm VA per worker, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0255**	-0.0632***	-0.0797***
	(0.0125)	(0.0148)	(0.0156)
HHI (δ)	0.0018	0.0115***	0.0091**
	(0.0096)	(0.0039)	(0.0039)
wagebite $\times$ HHI (γ)	0.0533	0.0351**	0.0464***
	(0.0326)	(0.0176)	(0.0171)
Δ Firm VA per worker		-0.0433***	-0.0428***
		(0.0039)	(0.0037)
Dummy bite $\times$ Δ Firm VA per worker		0.0031	0.0026
		(0.0033)	(0.0031)
Constant (α)	0.4908***	0.6559***	0.7385***
	(0.0086)	(0.0086)	(0.0057)
F-test: $\beta + \gamma = 0$	0.60	4.47	6.66
Worker FEs	Yes	Yes	Yes
Firm FEs	No	Yes	Yes
Collective Bargaining Agreement $\times$ Year FEs	No	No	Yes
Adj. R^2	0.2020	0.2823	0.2867
Observations	6,831,528	5,191,131	5,190,615

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. Δ in firm VA per worker refers to the change in log firm value added per worker. Dummy bite $\times$ Δ in firm VA per worker interacts the dummy-bite measure with the change in 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 F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B9: Effects of minimum wage and labour market concentration on retention, including a firm-level variable capturing the % of bitten workers, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0090	-0.0647***	-0.0793***
	(0.0121)	(0.0151)	(0.0156)
HHI (δ)	0.0016	0.0102***	0.0080**
	(0.0096)	(0.0039)	(0.0038)
wagebite $\times$ HHI (γ)	0.0495	0.0368**	0.0473***
	(0.0340)	(0.0174)	(0.0169)
Share bitten workers in firm-year	-0.0193**	0.0017	-0.0006
	(0.0090)	(0.0050)	(0.0037)
Firm VA per worker		0.0123***	0.0126***
		(0.0017)	(0.0016)
Constant (α)	0.4977***	0.5339***	0.6110***
	(0.0089)	(0.0179)	(0.0162)
F-test: $\beta + \gamma = 0$	1.43	4.40	6.28
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. Share bitten workers in firm-year is the firm-year share of bitten workers. 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 F-test for the null that $\beta + \gamma = 0$. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B10: Effects of minimum wage and labour market concentration on retention, including firm size dummy, temporary workers only

	(1)	(2)	(3)
wagebite (β)	-0.0258***	-0.0066	-0.0219***
	(0.0049)	(0.0065)	(0.0068)
HHI (δ)	0.0067**	-0.0066*	-0.0091**
	(0.0028)	(0.0037)	(0.0037)
wagebite $\times$ HHI (γ)	-0.0171	0.0011	0.0077
	(0.0112)	(0.0146)	(0.0148)
Large firm	-0.0122*		
	(0.0066)		
Large firm $\times$ HHI	-0.0048	0.0193***	0.0196***
	(0.0107)	(0.0057)	(0.0057)
Large firm $\times$ *wagebite*	-0.0026	-0.0751***	-0.0756***
	(0.0170)	(0.0187)	(0.0186)
Large firm $\times$ *wagebite* $\times$ HHI	0.0948**	0.0547**	0.0596**
	(0.0453)	(0.0267)	(0.0258)
Firm VA per worker		0.0123***	0.0127***
		(0.0016)	(0.0016)
Constant (α)	0.5003***	0.5338***	0.6098***
	(0.0069)	(0.0171)	(0.0160)
Joint significance F-test	1.52	11.24	11.55
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. Large firm is an indicator for firms above the analysis-specific employment threshold (10). Large firm $\times$ HHI interacts the large-firm indicator with the new hires HHI. Large firm $\times$ *wagebite* interacts the large-firm indicator with the bite measure. Large firm $\times$ *wagebite* $\times$ HHI is the triple interaction of the large-firm 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 large-firm interactions: Large firm $\times$ *wagebite*, Large firm $\times$ HHI, and Large firm $\times$ *wagebite* $\times$ HHI. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B11: 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 B12: 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 B13: 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 B14: 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: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table B15: 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 B16: 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 B17: National Minimum Wage in Mainland Portugal, 2002–2023

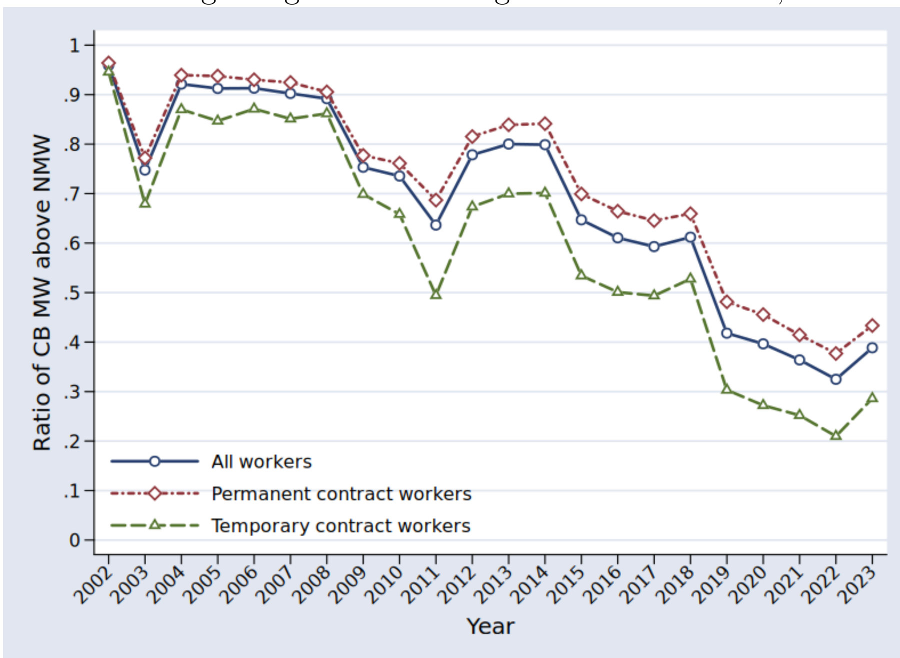
Year	NMW (€)	Year	NMW (€)
2002	348.0	2013	485.0
2003	356.6	2014	485.0
2004	365.6	2015	505.0
2005	374.7	2016	530.0
2006	385.9	2017	557.0
2007	403.0	2018	580.0
2008	426.0	2019	600.0
2009	450.0	2020	635.0
2010	475.0	2021	665.0
2011	485.0	2022	705.0
2012	485.0	2023	760.0

Source: PORDATA.

Notes: The table reports the nominal monthly national minimum wage (NMW) in euros. Azores and Madeira autonomous regions have supplements added to the national wage floor.

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.