Labor Market Concentration, Wages, and Job Security in Europe

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ABSTRACT

We leverage administrative linked employer–employee data from six European countries to provide the first comparable cross-country evidence on the impact of labor market concentration on wages and job security. We find strikingly similar and relatively low wage elasticities across countries, but greater elasticities for job security, as measured by contract type. We provide suggestive evidence that the similarity of our wage elasticities and the greater sensitivity of job security to labor market concentration may be explained by the fact that sector-level collective bargaining is dominant in the countries we study and that it sets wages but usually not contract type.

[Submitted February 2023; accepted January 2024]; doi:10.3368/jhr.0223-12757R1
JEL Classification: J31, J42, J52, and L41
ISSN 0022-166X E-ISSN 1548-8004
Supplementary materials are available online at: https://jhr.uwpress.org.

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I. Introduction

Labor market concentration is an important source of monopsony power since, when there are fewer employers in a market, it is more difficult for workers to find suitable outside options (for example, Jarosch, Nimczik, and Sorkin 2024). Investigating the effects of concentration on wages and other dimensions of job quality is therefore essential to assess the extent of monopsonistic competition in labor markets and the resulting need for policy intervention.

The issue of monopsony power and labor market concentration has recently attracted a rising interest, both among academics (see, for example, the special issues of the

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Journal of Labor Economics 2010 and the *Journal of Human Resources* 2022)[1] and among policymakers.[2] The impact of local labor market concentration on wages has been estimated both in the US (for example, Arnold 2021; Schubert, Stansbury, and Taska 2021; Azar, Marinescu, and Steinbaum 2022; Rinz 2022; Benmelech, Bergman, and Kim 2022) and in other countries (Martins and Melo 2024; Dodini et al. 2024; Dodini, Salvanes, and Willén 2022; Marinescu, Ouss, and Pape 2021; Bassanini, Batut, and Caroli 2023; OECD 2021; Popp 2024). However, due to heterogeneity in the definition of local labor markets and in the resulting measures of concentration, and to differences in specifications, the estimated elasticities are hardly comparable across studies.

Moreover, this strand of research only considers the impact of labor market concentration on wages.[3] However, there is broad evidence in the literature that workers also value nonwage job attributes and that they may be willing to trade off wages for other dimensions of job quality (Mas and Pallais 2017; Taber and Vejlin 2020; Kesternich et al. 2021). If offering high-quality jobs is costly, employers enjoying monopsony power are likely to offer poorer nonwage attributes (for example, Manning 2003). Hence, considering only the wage effects of labor market concentration is likely to underestimate its true cost for workers.

This work addresses these two limitations by providing comparable evidence of the effects of labor market concentration in six European countries and by considering how such concentration affects not only wages, but also one key dimension of job quality, namely job security. To do so, we leverage rich, administrative linked employer–employee data from Denmark, France, Germany, Italy, Portugal, and Spain in the 2010s. We build comparable measures of concentration, by computing Herfindahl–Hirschman indexes (HHI) for new hires in local labor markets, defined as a combination of four-digit occupations and functional areas. The latter are characterized as the set of all Eurostat functional urban areas (FUAs)—corresponding to a city and its catchment area—and all NUTS-3 regions[4] in which at least 70 percent of the municipalities are not part of a FUA. As a robustness check, we also use two-digit occupations and, alternatively, the method proposed by Arnold (2021), which accounts for transitions between occupations, as well as FUAs only and, alternatively, NUTS-3 regions to build other measures of local labor market concentration.

We have information on wages in Denmark, France, Germany, and Portugal. We first investigate the impact of labor market concentration on daily wages of full-time workers. Our specification includes individual and local labor market fixed effects along with

1. *Journal of Labor Economics*, 28(2), April 2010 and *Journal of Human Resources*, 57(S), April 2022.

2. See the US Horizontal Merger Guidelines, 2010 (https://www.justice.gov/atr/horizontal-merger-guidelines-08192010, accessed October 7, 2025) and its revised version (for example, https://www.justice.gov/opa/press-release/file/1463566/download, accessed October 7, 2025), the US Antitrust Guidance for Human Resource Professionals, 2016 (https://www.justice.gov/atr/file/903511/download, accessed October 7, 2025), and M. Vestager's speech on *A new era of cartel enforcement* in October 2021 (https://ec.europa.eu/commission/commissioners/2019-2024/vestager/announcements/speech-evp-m-vestager-italian-antitrust-association-annual-conference-new-era-cartel-enforcement_en, accessed October 7, 2025).

3. A notable exception is Qiu and Sojourner (2022).

4. The NUTS classification (Nomenclature of Territorial Units for Statistics) is a hierarchical system for dividing up the economic territory of the EU and the UK for the purpose of collection, development, and harmonization of European regional statistics. NUTS-3 is the most disaggregate level of this classification.

time-varying individual characteristics. In this setup, the potential correlation of labor market concentration with unobserved time-varying local or aggregate characteristics is a threat to identification. We address this issue by including, first, firm-by-municipality-by-year fixed effects in our specification. This allows us to control for productivity shocks at the establishment level, but also for product market concentration and productivity shocks at the local and aggregate levels since firm-by-municipality-by-year fixed effects incorporate local areas, as well as industry fixed effects. We can include this very rich set of fixed effects and still identify the coefficient of interest thanks to the fact that, in our data, there are many occupations within firm-by-municipality in any given year, which in turn leads to variation in the concentration indexes even within firm-by-municipality-by-year cells. Second, as concentration could remain endogenous despite these controls, we also use the standard leave-one-out instrument employed, among others, by Azar, Marinescu, and Steinbaum (2022); Rinz (2022); Qiu and Sojourner (2022); and Marinescu, Ouss, and Pape (2021). One worry with this instrument is that it could capture national trends in aggregate labor supply or demand. We provide evidence suggesting that, given our large set of fixed effects, such exogeneity violations are likely to be small. We then use the method proposed by Conley, Hansen, and Rossi (2012) to show that our results are anyway robust to large potential violations of exogeneity of the instrument.

Despite the heterogeneity of most labor market institutions across the countries we study, the wage elasticities we estimate are strikingly similar, ranging from −0.019 in Germany to −0.022 in France, −0.025 in Portugal, and −0.033 in Denmark. These estimates imply that increasing labor market concentration by one standard deviation from the mean reduces daily wages by 3.5 percent in Denmark, 2.4 percent in France, 2.1 percent in Germany, and 2.5 percent in Portugal. Using hourly wages of either full-timers or all employees, in countries where they are available, yields very similar results. Interestingly, when considering separately new hires and incumbent workers, we find a negative effect of labor market concentration on daily wages of full-timers for both groups. This indicates that reduced outside options not only affect the bargaining power of workers at the time of hiring, but also that of incumbents (or their representatives, for instance, trade unions) when negotiating pay raises and/or promotions.

As a second step, we consider the impact of labor market concentration on job security, as proxied by employment contract type (permanent versus temporary).[5] Information on the type of contract at the time of hiring is available—or can be reconstructed—in all our countries except Denmark. We therefore estimate the effect of concentration in a local labor market on the probability of being hired on a permanent rather than a temporary contract. In Italy and Spain, we also know whether individuals hired on a temporary contract in a given year had a permanent contract in the following calendar year with the same employer. This allows us to estimate how labor market concentration affects contract conversions from temporary to permanent status. We find

5. This measure of job security is consistent with evidence provided by OECD (2014) suggesting that, in European countries, people employed on temporary contracts perceive a much higher risk of losing their job within the next six months than people employed on permanent contracts.

that higher labor market concentration reduces the probability of being hired on a permanent contract in France, Germany, and Portugal. The corresponding elasticities are as large as −0.046, −0.053, and −0.233,[6] respectively, that is, at least twice as large as those estimated for wages. This implies that increasing the HHI by one standard deviation from the mean reduces the probability of being hired on a permanent contract by 5 percent in France, 6 percent in Germany, and 24 percent in Portugal. We do not find the same effect in Italy and Spain where the impact of labor market concentration on the type of contract of new hires is not significant at conventional levels. However, in both countries, labor market concentration strongly affects the probability of being converted from a temporary to a permanent contract by the end of the first year following that of hiring. The elasticity of conversions with respect to labor market concentration is as large as −0.245 in Italy and −0.060 in Spain. Increasing labor market concentration by one standard deviation from the mean therefore reduces the probability of conversion by 28 percent in Italy and 8 percent in Spain. Overall, this suggests that when firms have some monopsony power, the cost for workers materializes in various dimensions of job quality—not only in the form of lower wages, but also in terms of poorer job security.

We conjecture that the greater effect of concentration on contract type as compared to wages, and the similarity of wage elasticities across countries, could be explained by the fact that the countries we study have high coverage of collective bargaining, which takes place predominantly at the industry level. As a consequence, wages are largely set by industry-level collective agreements, which likely leaves little room for single employers to adjust pay downward even when they gain market power. By contrast, the type of contract on which employees are hired is largely unregulated by collective agreements. This should make it easier for firms to use contract type, rather than wages, as a margin of adjustment when concentration increases. We provide evidence supporting this conjecture using data for Germany and Portugal, the two countries for which we have information on collective bargaining. We show that the effect of concentration on wages is indeed significantly smaller in industries where coverage of sector-level collective bargaining is larger, while sector-level collective bargaining does not affect the relationship between concentration and contract type.

Our paper contributes to the literature on labor market concentration and wages by providing estimates for four European countries. To the best of our knowledge, this paper is the first providing comparable estimates across countries.[7] The existing literature of single-country studies has found different wage elasticities across EU countries, ranging from −0.014 in Portugal to −0.048 in France—see Marinescu, Ouss, and Pape (2021); Popp (2024); Bassanini, Batut, and Caroli (2023); and Martins and Melo (2024). However, it is hard to disentangle whether these differences are due to truly different effects of concentration on wages across countries or to differences in specifications and variable definitions. The fact that the wage elasticities estimated for the United States also vary substantially from one paper to another[8] suggests that

6. Elasticities are computed at the sample average, unless otherwise indicated.

7. OECD (2021) estimates wage elasticities for six countries. However, as acknowledged in the study itself, the confidence intervals are so large that they do not allow country-by-country comparisons. Country-specific estimates are therefore only used to derive an average cross-country elasticity.

differences in methods play an important role. As a matter of fact, we show that, when using the same methodology for Denmark, France, Germany, and Portugal, we find very similar wage elasticities, which are all quite small.

We also contribute to the surprisingly small literature focusing on the impact of labor market concentration on nonwage job attributes. To our knowledge, the only two papers doing so are Qiu and Sojourner (2022) and Meiselbach et al. (2022), who find a negative effect of concentration on employer-provided health insurance in the United States. We consider another dimension of job quality, namely job security, a very important nonwage job attribute in dual labor markets where permanent contracts providing a high degree of employment protection coexist with temporary ones. To our knowledge, we are the first to show that higher labor market concentration reduces the probability of being hired on a permanent contract or converted to such a contract after being hired on a temporary one. This finding suggests that considering only the effect of labor market concentration on wages underestimates its overall impact on job quality and hence the resulting welfare loss for workers.

Finally, our paper speaks to the literature investigating the role of the countervailing power of organized labor in shaping the effect of employers' market power on labor market outcomes. Early work shows that bilateral monopolies in the labor market can yield efficient bargaining outcomes since monopolistic unions offset the market power of monopsonists—see MaCurdy and Pencavel (1986) and Espinosa and Rhee (1989). More recently, Marinescu, Ouss, and Pape (2021); Benmelech, Bergman, and Kim (2022); and Dodini, Salvanes, and Willén (2022) provide evidence that in industries where the unionization rate is higher, the wage elasticity to labor market concentration is smaller. We complement this literature by showing that, in countries with prevalence of sectoral collective bargaining, such as many European countries, wage elasticities to labor market concentration are significantly smaller where coverage by sectoral agreements is larger. This might explain why previous research has tended to find smaller elasticities in many European countries than in the US, where collective bargaining coverage is low, and sectoral bargaining does not exist. By contrast, we find that the effect of labor market concentration on the probability of being hired on a permanent contract does not vary with sector-level collective bargaining coverage, which is consistent with the fact that the choice of contract type is rarely regulated by sectoral agreements.

In the following, Section II presents the data and some descriptive statistics. Section III lays out our empirical specification. Section IV presents the results, and Section V concludes.

II. The Data

We use near-universe national administrative data spanning the 2010s for Denmark, France, Germany, Italy, Portugal, and Spain; see Online Appendix 1 for a detailed description of the data. We drop agriculture and industries where the public administration is dominant.[9] We also exclude self-employed and household employees and only keep workers with at least one month of tenure with their current employer.

8. When using microdata, as we do, the wage elasticities estimated in the USA vary between −0.05 and −0.22 (Benmelech, Bergman, and Kim 2022; Qiu and Sojourner 2022; Arnold 2021).

A. *Labor Market Concentration*

A local labor market $l = (o, z)$ is defined as the intersection between a four-digit occupation o and a geographical area z. We measure concentration with a Herfindahl–Hirschman index (HHI) based either on hirings or on employment:

$$HHI_{l,t} = \sum_{e=1}^{E} s_{e,l,t}^{2}$$

where E denotes the total number of employers with positive hirings (respectively employment) in each local labor market, and $s_{e,l,t}$ is the share of employer e in the total number of hirings—respectively employment—(H) in local labor market l at time t:

$$s_{e,l,t} = \frac{H_{e,l,t}}{\sum_{m=1}^{E} H_{m,l,t}}$$

In our baseline specification, we use an HHI based on hirings, but we also present robustness checks using an HHI based on employment, in all countries for which this information is available.[10] In a static Cournot model of oligopsony, wages are indeed inversely related to the HHI measured in terms of employment (Boal and Ransom 1997). In contrast, in dynamic monopsony models, such as Manning (2003), wages positively depend on the availability of outside options and their value. Jarosch, Nimczik, and Sorkin (2024) show that, in a stationary search and matching model with granular search where concentration affects wages by changing workers' outside options, HHIs based on either employment or hirings can be used interchangeably to obtain a measure of labor market concentration that is relevant for wage determination. However, if the environment is not stationary, downsizing firms may have a positive share in employment while not contributing to create outside options for workers (Marinescu, Ouss, and Pape 2021). This is why our preferred specification uses an HHI based on hirings.

In order to harmonize our units of observation across countries, we define employers using a firm-by-municipality concept (Online Appendix 1.1.1); that is, we consider that an employer is composed of all the establishments belonging to a given firm that are located in a given municipality. New hires are defined as individuals who are in a firm-by-municipality couple at time t and were not there at $t - 1$. In the remainder of the paper, f will index a firm-by-municipality couple.

Our main definition of local labor markets is based on four-digit occupations. There is evidence that workers who change occupations often undergo a wage penalty (for example, Gathmann and Schönberg 2010; Kambourov and Manovskii 2009). However, if workers move across a range of four-digit occupations (and not only within each of them) with a limited wage penalty, the local labor markets that are relevant to compute the HHI are larger than those based on four-digit occupations. To make sure that our results are stable with respect to the occupational size of local labor markets, we

9. In practice, we keep NACE Rev.2 industries ranging from 05 to 82, as well as 90 and 92 to 96.

10. Information on the level of employment is available in Denmark, France, Germany, and Portugal, but neither in Italy nor in Spain.

alternatively use two-digit occupations and the method proposed by Arnold (2021). The latter considers a Herfindahl–Hirschman index where the share of each employer in a given local labor market $l = (o, z)$ is computed as a weighted average of the employment of that employer in all those occupations o' of the local area z, which could represent valuable outside options for a worker in occupation o.[11]

Our baseline HHI is built using an original concept of geographical area z that we call Functional Areas. It is based on a mixture of Functional Urban Areas (FUAs) and NUTS-3 regions. FUAs consist of a city and its catchment area, which is a commuting zone whose labor market is highly integrated with the city (OECD 2012; Dijkstra, Poelman, and Veneri 2019). FUAs are constructed using the same algorithm in all countries and hence provide a harmonized definition of cities and their areas of influence in international perspective.[12] By definition, FUAs do not cover rural areas. To overcome this limitation, we define functional areas (FAs) as the set of all FUAs and all NUTS-3 regions[13] (excluding the municipalities that are part of a FUA) in which at least 70 percent of the municipalities are not part of a FUA.[14]

Table 1 reports the distribution of country-specific HHIs based on four-digit occupations and FAs and weighted by the number of new hires in each local labor market. The level and distribution of labor market concentration appear to be rather similar across the six countries we consider, despite their different industrial structures and labor market institutions. More than 75 percent of the new hires are employed in local labor markets with an HHI below 0.15, the 2010 threshold for moderate concentration defined by the US antitrust authorities. Moreover, the 90th percentile of the HHI distribution is higher than 0.25—the 2010 threshold for high concentration—in only one country, namely, Portugal.

B. Dependent and Control Variables

We have information on wages for Denmark, France, Germany, and Portugal. We construct monthly wages for Portugal, daily wages for Denmark, France, and Germany, and hourly wages for Denmark, France, and Portugal (see Online Appendix 1.1.2 for details). To ensure that our results on monthly and daily wages are not affected by the incidence of short part-time employment, we restrict our sample to full-time workers. For the sake of comparability, we also do so when using hourly wages, although in this case we also run additional estimates on the whole population of full- and part-timers.

11. The value of each occupation o' is obtained from the actual average transition matrix from occupation o to occupations o', taking into account the actual size of each occupation. The weight of each occupation o' is therefore computed as the ratio of the estimated probability of transition from o into o' to the estimated probability of a within-occupation transition (that is, from o to o) divided by the ratio of employment in occupation o' to that of occupation o. Arnold (2021) sets forth the theoretical foundations of this measure of concentration in a Cournot model of oligopsony, therefore based on employment shares and not hiring shares. As this method requires employment stock data, we cannot perform this robustness check for Italy and Spain.

12. FUAs have been used by Ascheri et al. (2021) to characterize labor market concentration in European urban areas.

13. In Portugal, we use districts (distritos) instead of NUTS-3 regions since the latter are smaller than in other countries while the former are of comparable size.

14. As a robustness check, we also run our analysis using HHIs based alternatively on FUAs and NUTS-3 regions.

Table 1
Labor Market Concentration

HHI	Mean	SD	P25	P50	P75	P90
Denmark	0.0589	0.1093	0.0097	0.0217	0.0546	0.1437
France	0.0657	0.1297	0.0055	0.0194	0.0604	0.1710
Germany	0.0591	0.1213	0.0057	0.0171	0.0511	0.1508
Italy	0.0636	0.1372	0.0047	0.0145	0.0530	0.1648
Portugal	0.0957	0.1686	0.0108	0.0311	0.0921	0.2654
Spain	0.0702	0.1490	0.0051	0.0169	0.0600	0.1806

Notes: Herfindahl–Hirschman index based on hiring, computed using firm-by-municipality identifiers. Local labor markets are defined based on four-digit occupations and functional areas (FAs). FAs are defined as FUAs and NUTS-3 regions (excluding the municipalities that are part of a FUA) for which at least 70 percent of the municipalities are not part of a FUA. Districts instead of NUTS-3 regions are used for Portugal. Statistics are weighted by the number of new hires. Labor market concentration is computed over the following time periods: 2011–2018 for Denmark, 2010–2017 for France, 2012–2018 for Germany, 2013–2018 for Italy, 2011–2019 for Portugal, and 2011–2017 for Spain.

Information on the type of contract upon hiring (permanent versus temporary) is available in all countries except Denmark. For France, Germany, Italy, Spain, and Portugal, we then define a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract.[15] In Italy and Spain, the nature of the data also allows to identify conversions from temporary to permanent contracts. We define a dummy variable for conversion that is equal to one if the individual was hired on a temporary contract at year *t* and had a permanent contract with the same employer in the following calendar year and zero otherwise.

Our data also contain information on individuals' age, gender, and education, as well as whether individuals work part-time or full-time and in which industry they are employed. For new hires, we also know whether they were in employment the year before. In Portugal and Germany, we have information on collective bargaining coverage, at least at the industry level. We define four alternative indicators. First, in each industry, we compute the proportion of employees covered by any type of collective agreement, that is, either at the sector or at the firm level. Second, we define a dummy variable equal to one if an industry has a proportion of employees covered by any type of collective agreement larger than the median and zero otherwise. Finally, we define two similar indicators based on the proportion of employees covered by sectoral collective agreements only.

Although our measures of labor market concentration are constructed using the entire population, the regressions are conducted on random subsamples of the population in

15. We do so rather than the opposite, that is, define a dummy variable equal to one if the individual is hired on a temporary contract and zero otherwise, since permanent contracts are very similar in all countries—they are essentially open-ended contracts—while there exist a large variety of temporary contracts, both within and across countries.

three out of six countries.[16] Descriptive statistics of our data are provided in Online Appendix Table 1.1.

III. Empirical Specification

A. Baseline Model

We first estimate the impact of labor market concentration on wages using the following specification:

(1) $\log(w_{i,j,f,l,s,t}) = \beta\log(HHI_{l,t}) + \gamma\mathbf{X}_{i,j,f,l,s,t} + \mu_i + \mu_{f,t} + \mu_l + \varepsilon_{i,j,f,l,s,t}$

where i indexes the worker, j the establishment, f the firm-by-municipality couple, l the local labor market, s the industry, and t is the year. w alternatively denotes the daily or hourly wage. *HHI* is the concentration index computed using firm-by-municipality identifiers. **X** is a vector of individual controls including yearly dummies for the worker's age, whether or not the individual is a new hire in the firm and if so, whether or not they were in employment the year before. **X** also includes establishment and/or industry fixed effects when not collinear with firm-by-municipality fixed effects, as well as a dummy variable for working part-time versus full-time, whenever our regression sample is not restricted to full-timers. Our specification also includes individual and local labor market fixed effects (μ_i and μ_l, respectively). Finally, firm-by-municipality-by-year fixed effects ($\mu_{f,t}$) control for firm productivity at a very disaggregate level.[17] This is particularly important to identify the effect of labor market concentration on wages going through a reduction of workers' outside options. If high-productivity firms drive low-productivity competitors out of the market, an increase in concentration could be accompanied by an increase in the proportion of high-quality jobs, as suggested, for example, by Beaudry, Green, and Sand (2012). Controlling for firm-by-municipality-by-year fixed effects allows us to net this effect out. Firm-by-municipality-by-time fixed effects also control for both aggregate and local product market competition, as well as aggregate productivity shocks and business cycle fluctuations.

With this large set of fixed effects, the model is identified only if there exist several occupations in firm-by-municipality couples. This is the case in our data since the average number of occupations per firm by municipality (weighted by the number of observations) ranges from 2.9 in Spain to 17.9 in Denmark.[18] Identification then stems from three different sources of variations. The main one lies in the different changes over time in the HHIs of the various occupations within each firm by municipality. In combination with changes in the HHI over time, two additional

16. The full population is used in Denmark, Italy, and Portugal. To save computational time, random subsamples of the population are used in Germany (10 percent of the population) and Spain (15 percent). In France, panel data are available only for a random subsample covering one-twelfth of the population (Online Appendix 1.1.2).

17. Since, in the countries of our sample, establishments are not accounting centers, no linked employer–employee data set could be used to measure productivity at a more disaggregate level.

18. The average number of occupations per firm-by-municipality couple is 6.4 in Italy, 7.5 in France, 9.1 in Germany, and 11.2 in Portugal.

sources of variation are provided by individuals changing occupation—even if remaining in the same firm by municipality—or changing firm by municipality—even if remaining in the same occupation.

We also estimate the impact of labor market concentration on job security. To do so, we restrict our sample to new hires and estimate the following specification:

$$(2)\quad S_{i,j,f,l,s,t} = \beta \log(HHI_{l,t}) + \gamma \mathbf{X}_{i,j,f,l,s,t} + \mu_{f,t} + \mu_l + \varepsilon_{i,j,f,l,s,t}$$

where S alternatively denotes a dummy variable equal to one when the individual is hired on a permanent contract, and zero otherwise, or a dummy variable equal to one when the individual was hired on a temporary contract and had a permanent contract with the same firm-by-municipality in the following year (an outcome that we refer to as contract conversion) and zero otherwise. In the latter case, the regression sample is further restricted to those initially hired on a temporary contract. We do not include individual fixed effects in Equation 2 to avoid confining the analysis to individuals who changed job several times over the period under study. Individual heterogeneity is accounted for by augmenting the vector of individual controls $\mathbf{X}$ with invariant or quasi-invariant characteristics (gender and education), in addition to the individual's age, a dummy variable for being part-time, and another one for whether they were in employment the year before. Standard errors are clustered at the local-labor-market-by-year level for all estimates.

B. Instrumental Variable Strategy and Violations of Exogeneity of the Instrument

In this setup, identification may be under threat if an omitted variable varies across local labor markets and over time and is correlated with both the HHI and our dependent variables. This may occur if positive or negative shocks to local labor supply or demand affect both the wage (and/or the contract type) that workers are willing to accept and the number of firms that find it attractive to operate in the local labor market. For instance, if a school specialized in training students in skills that are particularly useful in certain occupations opens in a given geographical area, this will increase the supply of labor with those specific skills and likely reduce the corresponding wage and nonwage job attributes that firms are willing to offer. At the same time, it may increase the number of firms that find it attractive to operate in this area. If this reduces labor market concentration, it will generate a positive spurious correlation between concentration on the one hand and wages and contract type on the other hand.

As is standard in the literature (for example, Marinescu, Ouss, and Pape 2021; Azar, Marinescu, and Steinbaum 2022; Rinz 2022), we tackle this issue by using a quasi-leave-one-out instrumental variable (IV) strategy. We instrument log(HHI) in local labor market $l = (o, z)$ at time t with the average of $\log(1/N_{o,z',t})$, where $N_{o,z',t}$ is the number of firms with a positive number of hires—alternatively, employees—in all other geographical areas z' for the same occupation o and time period t.[19] This

19. Instrumenting a variable in one zone using the average of this variable in other zones (that is, using a Hausman or leave-one-out instrument) is standard in international economics and industrial organization (for example, Hausman, Leonard, and Zona 1994; Autor, Dorn, and Hanson 2013; Bai et al. 2017; Azar, Berry, and Marinescu 2022).

instrument aims at capturing changes in labor market concentration taking place at the national level as a result of, for example, mergers or divestitures of large national companies.[20] When such firms merge or split, this modifies the number of companies operating in all local labor markets where they are present without being correlated with idiosyncratic shocks in these markets. As a consequence, it generates an exogenous shock that allows identification of the causal impact of labor market concentration on wages and/or contract type.

One worry with this instrument is that its variations could also capture national trends in occupational labor supply and demand, which could also affect our dependent variables. To alleviate this concern, we proceed in the following way. First, we augment Equations 1 and 2 by including the time-varying share of each four-digit occupation in new hires at the national level to capture occupation-specific national trends in labor supply and demand.[21] One caveat is that the potential endogeneity of this variable could bias the estimates of the effect of $\log(HHI_{l,t})$ even in the IV specification. However, if our instrument substantially captured national occupational trends in labor supply and demand, we would expect that the introduction of the share of each four-digit occupation in new hires would significantly modify our estimates. We will show that this is not the case in our data, so that, if any, the violations of exogeneity of the instrument are likely to be small.

As a second step, we quantify the exogeneity violation that our models may tolerate using the plausibly exogenous instrument regression method proposed by Conley, Hansen, and Rossi (2012). We consider the following model, in which the instrument (Z) is not fully exogenous and therefore may have a direct effect on the dependent variable:

$$(3)\quad Y_{i,j,f,l,s,t} = \beta\log(HHI_{l,t}) + \gamma Z_{l,t} + \delta\mathbf{C}_{i,j,f,l,s,t} + \varepsilon_{i,j,f,l,s,t}$$

where Y is the outcome variable, and $\mathbf{C}$ is the vector of controls including fixed effects.

If the true value of γ (denoted by γ^*) were known, Z would be a valid instrument in the following equation, obtained by subtracting the known effect of Z from the dependent variable in Equation 3, since it would no longer be correlated with the error term:

$$(4)\quad Y_{i,j,f,l,s,t} - \gamma^* Z_{l,t} = \beta\log(HHI_{l,t}) + \delta\mathbf{C}_{i,j,f,l,s,t} + \varepsilon_{i,j,f,l,s,t}$$

In practice, the direct effect of Z is unknown, but we can still determine how large it should be to make the coefficient of interest, β, insignificant. This is done in three steps. First, we show that Z is positively correlated to $\log(HHI)$. As a consequence, if γ^* were positive, incorrectly using Z as an instrument for $\log(HHI)$ would generate a bias towards positive values when estimating β by two-stage least squares (2SLS) in Equation 1. The risk of overestimating the magnitude of the (negative) elasticity hence only exists if γ^* is negative. As a consequence, we can take zero as the upper bound of the support of γ, which corresponds to the situation in which Z is exogenous. We then consider decreasing potential values of γ^* one by one, take them as given and

20. We also run robustness checks using a leave-one-out HHI as the instrument.

21. We do so since we cannot control for occupation-by-time fixed effects, which would be almost collinear with our instrument. This is because, in practice, the relevant variation in the instrument is at the occupation-by-time national level.

estimate Equation 4 by 2SLS, instrumenting log(HHI) with Z and storing confidence intervals for the estimate of β ($\hat{\beta}$) at every step. This allows us to determine the lowest value of γ^* that would still make $\hat{\beta}$ significant at the 10 percent level. Finally, we estimate the following reduced form equation by ordinary least squares (OLS):

$$(5) \quad Y_{i,j,f,l,s,t} = \alpha Z_{l,t} + \delta \mathbf{C}_{i,j,f,l,s,t} + u_{i,j,f,l,s,t}$$

where u is a standard disturbance term. α captures the overall effect of Z on the dependent variable Y, that is, both the direct effect, independent of log(HHI), and the indirect effect going through log(HHI). We can then express the lowest value of γ^* that would still make $\hat{\beta}$ significant at the 10 percent level in percentage of the reduced-form effect $\hat{\alpha}$, estimated in Equation 5. This ratio provides an order of magnitude of the violations of instrument exogeneity that our model may tolerate, that is to say, that still yield a point estimate $\hat{\beta}$ significant at the 10 percent level.

IV. Results

A. Labor Market Concentration and Wages

We first estimate the impact of labor market concentration on daily wages of full-timers in Denmark, France, Germany, and Portugal. Instrumental variable estimates are presented in Table 2, while OLS estimates are provided in Online Appendix Table 2.1.[22] As shown by the values of the Kleibergen–Paap (KP) F-statistics and first-stage estimates (Online Appendix Table 2.2), the instrument is strongly positively correlated with labor market concentration in all countries. The F-statistic is lower in Portugal and Denmark than in Germany and France,[23] but remains higher than standard critical thresholds in all countries. The point estimates of the coefficient on labor market concentration are strikingly similar across the four countries, ranging from −0.019 in Germany to −0.022 in France, −0.025 in Portugal, and −0.033 in Denmark, all significant at the 1 percent level.[24] Since we found similar HHI distributions across countries (Table 1), this implies that the effective range of wage variation induced by labor market concentration is also very similar. In fact, increasing labor market concentration by one standard deviation from the mean reduces daily wages by 3.5 percent in Denmark, 2.4 percent in France, 2.1 percent in Germany, and 2.5 percent in Portugal.

This similarity is all the more striking given that these countries have different labor market institutions, for example, as regards employment protection legislation, the minimum

22. The OLS point estimates are negative, but much smaller than those obtained when instrumenting labor market concentration. This suggests that they are possibly upward biased.

23. This is likely due to the smaller country size of Denmark and Portugal, which implies a smaller number of functional areas in these countries.

24. The results obtained when using the leave-one-out HHI (rather than the leave-one-out $1/N$) instrument are presented in Online Appendix Table 2.3. Point estimates tend to be smaller, but they are still significant at conventional levels in all countries. Results are also stable if using an HHI based on employment rather than on hirings (Online Appendix Table 2.4).

Table 2
Labor Market Concentration and Daily Wages of Full-Timers—IV Estimates

Dep. Var. Daily Wages	Denmark	France	Germany	Portugal
Log HHI	−0.033***	−0.022***	−0.019***	−0.025***
	(0.008)	(0.002)	(0.002)	(0.009)
KP F-test	28.6	708.9	271.4	14.6
Observations	6,299,179	8,269,375	11,050,435	15,086,998

Notes: 2SLS estimates. The dependent variable is log(wage). Local labor markets are defined based on four-digit occupations and functional areas (FAs). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(HHI) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Monthly wages instead of daily wages are used for Portugal. Standard errors are clustered at the labor-market-by-year level. KP F-test: Kleibergen–Paap Wald F-test. $*p < 0.1$, $**p < 0.05$, $***p < 0.01$. Observations include singletons.

wage, and the importance of active and passive labor market policies.[25] However, Denmark, Germany, France, and Portugal have in common strong collective bargaining at the industry level (OECD 2019) with large coverage among employees.[26] As a consequence, wages are largely set by collective agreements signed at the industry level, which leaves little room for single firms to adjust pay downward even when they gain market power. This may explain not only why the wage elasticity with respect to labor market concentration is similar across the four countries we consider, but also why it is small when compared to what has been found in the US, where sectoral collective bargaining does not exist, and firm-level bargaining covers only 12 percent of employees.[27]

In the case of Germany and Portugal, we have access to industry-level data on collective bargaining that we can leverage to further explore the role of sector-level collective bargaining. We reestimate Equation 1 adding an interaction term between log(HHI) and

25. In 2015, the OECD index of employment protection legislation on regular contracts was 1.53 in Denmark, 2.50 in France, 2.60 in Germany, and 3.14 in Portugal on a scale ranging from zero to six. The Kaitz index (ratio of the minimum to the median wage) was 0.62 in France, 0.48 in Germany, and 0.55 in Portugal while there was no national minimum wage in Denmark. The share of spending on active labor market programs in GDP amounted to 1.65 percent in Denmark, 0.68 percent in France, 0.48 percent in Portugal, and 0.28 percent in Germany. The share of passive labor market policy spending (including unemployment benefits) was 1.27 percent of GDP in Denmark, 2.05 percent in France, 0.88 percent in Germany, and 1.36 percent in Portugal. (Source: https://www.oecd.org/employment/emp/employmentdatabase-labormarketpoliciesandinstitutions.htm, accessed October 7, 2025.)

26. In 2015, collective agreements covered 83 percent of the wage and salary employees in Denmark, 98 percent in France, 57 percent in Germany, and 74 percent in Portugal, while the OECD average was 33 percent (OECD/AIAS Collective Bargaining Database, https://www.oecd.org/employment/ictwss-database.htm, accessed October 7, 2025).

27. See Footnote 8.

each of our indicators of collective bargaining coverage, alternatively.[28] The results regarding coverage by any type of collective agreement, either firm or sector level, are presented in Columns 1 and 2 of Table 3. As evidenced in Panel A, in Germany, the point estimate on the interaction between log(*HHI*) and coverage by any type of collective agreement is positive and significant at conventional levels, whatever the indicator we use. It is also positive in Portugal, although not significant when using the dummy variable for high coverage. Consistent with the literature,[29] this suggests that the negative effect of labor market concentration on wages decreases as the proportion of employees covered by any type of collective agreement increases. Interestingly, results in Columns 3 and 4 of Table 3 show that in both countries this dampening effect is driven by sector-level, rather than firm-level collective bargaining. The point estimates on the interaction terms between log(*HHI*) and sector-level coverage indeed tend to be larger than those on the interactions between concentration and coverage by any type of agreement.[30] As a matter of fact, when including both firm- and sector-level collective bargaining separately, which stretches our data to their limits, the only point estimates that turn out positive are those on the interaction terms between log(*HHI*) and sectoral collective bargaining coverage.[31] The key role of sector-level collective bargaining in countries where it is dominant—like Germany and Portugal, but also Denmark and France—may help explain why we find similar, but also rather low, wage elasticities in the four countries. To the extent that wages are one of the main job attributes negotiated at the industry level and subsequently set by sector-level collective agreements, it is unlikely that individual firms would be able to exert a strong downward pressure on wages even when concentration increases in their local labor market.

Our instrument for $\log(HHI_{l,t})$, being the average of $\log(1/N_{o,z',t})$ in all other geographical areas z' for the same occupation o and time period t, provides a source of variation in labor market concentration based on national rather than local changes in the occupation we consider. If changes in an occupation result, for example, from mergers and divestitures of large national companies, they will likely affect concentration in the local labor markets where these companies are present, without being

28. In Germany, we also control for the share of employees covered by a works council in each industry interacted with log(*HHI*). We do so because works councils cover a large proportion of employees in this country—39 percent in 2017 (Ellguth and Kohaut 2018)—and can influence, among other things, how bonuses are attributed to workers (Fulton 2021). In contrast, this is not the case in Portugal. According to the Ministry of Labor (Ministério do Trabalho 2016), there were only 191 works councils in the entire country in 2015, and they were essentially responsible for health and safety matters.

29. See Marinescu, Ouss, and Pape (2021); Benmelech, Bergman, and Kim (2022); and Dodini, Salvanes, and Willén (2022).

30. The differences in point estimates on the interaction terms in Columns 1 and 3 (respectively 2 and 4) may not seem very large. However, in interpreting these differences, one has to take into account the fact that few workers are covered by firm-level collective agreements in both countries (Online Appendix Table 1.2), which implies that the indicators for sector-level coverage and coverage by any type of agreement are not much different one from the other.

31. The point estimate on log(*HHI*) interacted with the share of employees covered by a sector-level collective agreement is 0.0077 in Germany (respectively 0.0074 in Portugal), with standard error 0.0037 (respectively 0.0024 in Portugal). The point estimate on log(*HHI*) interacted with the dummy variable indicating whether the share of employees covered by sector-level collective agreements is higher than the median is 0.0024 in Germany (respectively 0.0011 in Portugal), with standard error 0.0010 (respectively 0.0009 in Portugal).

Table 3

Labor Market Concentration and Daily Wages of Full-Timers as a Function of Collective Bargaining Coverage—IV Estimates

Dep. Var. Daily Wages	(1)	(2)	(3)	(4)
Panel A: Germany				
Log HHI	−0.0224***	−0.0164***	−0.0229***	−0.0161***
	(0.0029)	(0.0026)	(0.0029)	(0.0026)
Log HHI*Coverage share, Any CBA	0.0112***			
	(0.0036)			
Log HHI*High coverage dummy, Any CBA		0.0037***		
		(0.0011)		
Log HHI*Coverage share, Sector CBA			0.0137***	
			(0.0033)	
Log HHI*High cover. dummy, Sector CBA				0.0044***
				(0.0011)
KP *F*-test	90.5	90.5	90.5	90.5
Observations	11,050,435	11,050,435	11,050,435	11,050,435
Panel B: Portugal				
Log HHI	−0.0310***	−0.0257***	−0.0308***	−0.0263***
	(0.0094)	(0.0090)	(0.0092)	(0.0090)
Log HHI*Coverage share, Any CBA	0.0070***			
	(0.0024)			
Log HHI*High coverage dummy, Any CBA		0.0009		
		(0.0008)		
Log HHI*Coverage share, Sector CBA			0.0083***	
			(0.0016)	
Log HHI*High coverage dummy, Sector CBA				0.0020**
				(0.0008)
KP *F*-test	7.3	7.3	7.4	7.3
Observations	15,086,998	15,086,998	15,086,998	15,086,998

Notes: 2SLS estimates. The dependent variable is log(wage). Local labor markets are defined based on four-digit occupations and functional areas (FAs). "Coverage share, Any (respectively Sector) CBA" is the proportion of employees covered by any (respectively a sector-level) collective agreement in the industry where the individual is employed. "High coverage dummy, Any (respectively Sector) CBA" is a dummy variable equal to one if the individual is employed in an industry with a proportion of employees covered by any (respectively a sector-level) collective agreement above the median and zero otherwise. Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. In Germany, control variables also include the share of employees covered by a works council in the industry interacted with log(*HHI*). Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Log(*HHI*) interacted with a collective bargaining indicator (respectively the works council variable) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation interacted with the collective bargaining indicator (respectively the works council variable). Monthly wages instead of daily wages are used for Portugal. Standard errors are clustered at the labor-market-by-year level. KP *F*-test: Kleibergen–Paap Wald *F*-test. *$p < 0.1$, **$p < 0.05$, ***$p < 0.01$. Observations include singletons.

correlated to idiosyncratic shocks in these markets. However, variations in the instrument could, in principle, also capture national trends in supply and demand that could also affect the dependent variables. To alleviate this concern, we first augment Equation 1 by including the share of each four-digit occupation in new hires at the national level to capture occupation-specific national trends in labor supply and demand. If our instrument were in practice capturing these trends, the introduction of this control should substantially modify our estimates. As shown in Online Appendix Table 2.5, this is not the case in France, Germany, and Portugal, where the wage elasticities are very similar to those presented in Table 2, ranging from −0.021 in France to −0.022 in Portugal and −0.025 in Germany. In Denmark, controlling for four-digit occupational shares in new hires at the national level lowers the magnitude of the point estimate from −0.033 to −0.015, although the coefficient remains statistically significant. This suggests that, in this country, the instrument may partly capture national trends in labor supply and demand. In contrast, in France, Germany, and Portugal, the violation of exogeneity of the instrument, if any, is likely to be small.

As a second step, we quantify the exogeneity violation that our models may tolerate using the method presented in Section III.B. As evidenced in Online Appendix Table 2.6, our IV estimates are robust to large violations of exogeneity of the instrument: $\hat{\beta}$ would still be significant at the 10 percent level if the direct effect of the instrument on wages were as large as 87 percent of the reduced-form estimate in France, 80 percent in Germany, and 57 percent in Portugal. We consider such large violations as unlikely given the stability of our estimates when controlling for the four-digit national occupational shares in new hires. In Denmark, our estimate is robust to violations of exogeneity as large as 74 percent of the effect obtained with the reduced form, which still allows us to draw conclusions—although cautiously—from our findings.

For the sake of cross-country comparability, we measure labor market concentration using a firm-by-municipality concept, that is, considering that an employer is composed of all its establishments located in a given municipality (Online Appendix 1.1).[32] Since this definition is not quite standard in the literature, for countries in which we can identify establishments and aggregate them nationwide into firms (that is, in all countries except Germany), we rerun our estimates using this more standard definition of employers. Online Appendix Table 2.8 presents the results. The elasticity of wages with respect to labor market concentration is similar to that estimated in Table 2 for France (−0.024) and Portugal (−0.021). It is even larger for Denmark (−0.055), although less precisely estimated. This suggests that our main results are not driven by the specificity of the employer concept that we use.

Our estimates could also be sensitive to the definition of local labor markets. In particular, if workers move across a range of four-digit occupations with a limited wage penalty, local labor markets based on four-digit occupations may be too small. To make sure that our results are not sensitive to the size of local labor markets, we reestimate Equation 1 with HHIs computed for local labor markets defined as the intersection between a two-digit occupation o and a

32. We show in Online Appendix Table 2.7 that, for France, the only country where we can do so, our results are unchanged when measuring labor market concentration using an HHI based on firm-by-canton or firm-by-intercommunalité (that is, larger administrative areas) rather than on firm-by-municipality identifiers.

geographical area z. Online Appendix Table 2.9 presents the results. They are stable in all countries, with point estimates even larger than in Table 2. For Portugal, our estimate is not significant at conventional levels, but the instrument is weak, with a KP F-test no higher than 0.17, which makes it impossible to interpret the second stage. Note however that the vast majority of job changes within two-digit occupations occur within four-digit occupations,[33] so that our preferred definition of local labor markets is based on the latter since it provides greater variability in labor market concentration within countries. In another robustness check, we use the measure of the HHI proposed by Arnold (2021), which takes into account the actual transitions across occupations; see Section II.A. Our results are also stable to this alternative measure of labor market concentration (Online Appendix Table 2.10).

So far, the geographical definition of local labor markets we have used was based on functional areas. Online Appendix Tables 2.11 and 2.12 provide robustness checks using alternatively FUAs and NUTS-3 regions to define local labor markets. The results are stable, thus suggesting that the choice of functional areas does not affect our findings in a major way.[34]

For a subset of countries, we can compute hourly wages, in addition to daily wages. Estimating the impact of labor market concentration on the former yields similar results to those obtained with daily wages. When the sample is restricted to full-timers only (Table 4, Panel A) point estimates range from −0.016 in France to −0.023 in Portugal and −0.033 in Denmark. Using hourly wages also allows us to consider the entire population of salaried workers and not only full-timers, since the results are immune from any bias due to short part-time work. Panel B of Table 4 presents the corresponding results. The elasticities we find are extremely close to those estimated for full-timers (−0.014 for France, −0.024 for Portugal, and −0.026 for Denmark), thereby suggesting that restricting our sample to full-timers when considering daily wages does not substantially affect our results.

The impact of labor market concentration on wages that we have estimated so far is the aggregation of the impact on two different groups of employees: those who have been hired over the past year (the new hires) and those who were already employed in the firm the year before (the incumbents). The literature conjectures that the depressing effect of labor market concentration on wages could be larger for new hires since they are more sensitive to market conditions (Haefke, Sonntag, and Van Rens 2013; Kudlyak 2014; Marinescu, Ouss, and Pape 2021). To investigate this potential source of heterogeneity in the effect of labor market concentration, we rerun our regressions interacting log(HHI) with two dummy variables, for new hires and incumbents, separately (Table 5). In all countries, we do find that labor market concentration negatively affects daily wages of newly hired full-timers. But we also find a negative effect on incumbents' wages, with elasticities ranging from −0.020 in Germany to −0.022 in

33. On average, 89.0 percent in Denmark, 70.4 percent in France, 86.6 percent in Germany, and 84.9 percent in Portugal, of all transitions within the same two-digit occupational category have the same four-digit origin and destination.

34. When using *districts* instead of functional areas in Portugal (Online Appendix Table 2.12), the first stage becomes relatively weak, and the point estimate is lower and noisier. It still remains within one standard deviation off the estimate presented in Table 2, but it is no longer significant at conventional levels.

Table 4
Labor Market Concentration and Hourly Wages—IV Estimates

Dep. Var. Hourly Wages	Denmark	France	Portugal
Panel A: Full-Timers			
Log HHI	−0.033***	−0.016***	−0.023***
	(0.008)	(0.002)	(0.008)
KP *F*-test	29.1	721.3	13.9
Observations	5,989,967	8,233,169	14,678,273
Panel B: All Employees			
Log HHI	−0.026***	−0.014***	−0.024***
	(0.005)	(0.002)	(0.008)
KP *F*-test	46.7	891.6	15.3
Observations	8,552,216	11,435,329	16,227,766

Notes: 2SLS estimates. The dependent variable is log(wage). Local labor markets are defined based on four-digit occupations and functional areas (FAs). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. In Panel B, controls also include a dummy variable for whether the individual works full-time or not. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Standard errors are clustered at the labor-market-by-year level. KP *F*-test: Kleibergen–Paap Wald *F*-test. *$p < 0.1$, **$p < 0.05$, ***$p < 0.01$. Observations include singletons.

France, −0.025 in Portugal, and −0.031 in Denmark—all significant at the 1 percent level.[35] This finding is consistent with results from Arnold (2021), Thoresson (2021), and Bassanini, Batut, and Caroli (2023), who suggest that labor market concentration generates downward pressure on incumbents' earnings, too. The estimates reported in Table 5 suggest that the negative effect of labor market concentration on wages is stronger for new hires than for incumbents in Denmark and, marginally, in France, but not in Germany and Portugal, pointing to the absence of a universally applicable pattern. This finding is also consistent with a recent literature showing that the wages of new hires are actually no more flexible than those of job stayers (Grigsby, Hurst, and Yildirmaz 2021). Overall, these findings suggest that labor market concentration not only affects the bargaining power of workers at the time of hiring, but also that of incumbents (or their representatives, for instance, trade unions) when negotiating pay raises and/or promotions.

35. One could worry that the effect we find on incumbents could actually be due to the downward pressure exerted by labor market concentration on wages at the time of hiring, which would persist over time. Yet, using a specification very similar to ours, Bassanini, Batut, and Caroli (2023) show that the effect they estimate on French incumbents remains remarkably stable when controlling for a job spell fixed effect, thereby netting out the effect of labor market concentration at the time of hiring.

Table 5
Labor Market Concentration and Daily Wages of Full-Timers—IV Estimates, New Hires Versus Incumbents

Dep. Var. Daily Wages	Denmark	France	Germany	Portugal
Log HHI*New hire	−0.041***	−0.024***	−0.016***	−0.024***
	(0.008)	(0.002)	(0.002)	(0.009)
Log HHI*Incumbent	−0.031***	−0.022***	−0.020***	−0.025***
	(0.008)	(0.002)	(0.002)	(0.008)
KP F-test	14.3	354.5	135.5	7.3
Observations	6,299,179	8,269,375	11,050,435	15,086,998

Notes: 2SLS estimates. The dependent variable is log(wage). Local labor markets are defined based on four-digit occupations and functional areas (FAs). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Interactions between log(*HHI*) and another variable are instrumented by this variable interacted with the instrument for log(*HHI*). Monthly wages instead of daily wages are used for Portugal. Standard errors are clustered at the labor-market-by-year level. KP F-test: Kleibergen–Paap Wald F-test. $*p < 0.1$, $**p < 0.05$, $***p < 0.01$. Observations include singletons.

B. Labor Market Concentration and Job Security

As a second step, we investigate the impact of labor market concentration on a second job attribute, namely, job security. First, we capture job security by the probability that an individual be hired on a permanent rather than temporary contract. To do so, we restrict our samples to new hires.[36] To avoid identifying the effect only on individuals who changed job several times over our relatively short time period, we do not include individual fixed effects. We control for individual heterogeneity in the best way we can, by adding gender and education to our vector of covariates. As shown in Table 6, our IV estimates[37] suggest that a higher level of concentration in the local labor market significantly reduces the probability of new workers to be hired on a permanent rather than a temporary contract, both in France and Germany.[38] When computed at sample average, the corresponding elasticities[39] are large: −0.046 (respectively

36. We do not investigate the effect of labor market concentration on the probability of being *employed* on a permanent contract separately for new hires and incumbents. The reason for not doing so is that, for incumbents, being on a permanent contract is generally an absorbing state, so that our model (Equation 2) would be misspecified.

37. The corresponding OLS estimates are provided in Online Appendix Table 2.13. The first-stage estimates are presented in Online Appendix Table 2.14.

38. We show in Online Appendix Table 2.15 that, for France, the only country where we can do so, our results are unchanged when measuring labor market concentration using an HHI based on firm-by-*canton* or firm-by-*intercommunalité* (that is, larger administrative areas) rather than on firm-by-municipality identifiers.

Table 6
Labor Market Concentration and Probability of Being Hired on a Permanent Contract—IV Estimates, New Hires Only

Dep. Var. Perm. Contract	France	Germany	Italy	Portugal	Spain
Log HHI	−0.0206***	−0.0353***	−0.0031	−0.0503+	0.0007
	(0.0046)	(0.0119)	(0.0065)	(0.0313)	(0.0018)
KP F-test	592.6	255.4	102.3	13.4	1875.4
Observations	3,530,660	4,167,918	16,781,457	1,039,792	4,895,949

Notes: 2SLS estimates. The dependent variable is a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract. Local labor markets are defined based on four-digit occupations and functional areas (FAs). Control variables include gender, education (four categories), yearly dummies for workers' age, whether the individual was employed the year before, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Standard errors are clustered at the labor-market-by-year level. KP F-test: Kleibergen–Paap Wald F-test. $+p < 0.11$, $*p < 0.1$, $**p < 0.05$, $***p < 0.01$. Observations include singletons.

−0.053) for France (respectively Germany). In Portugal, the effect of labor market concentration is significant only at the 11 percent level. However, the elasticity of the probability of being hired on a permanent contract is very large: −0.233, at sample average. Taking into account the distributions of HHIs, these estimates imply that increasing the HHI by one standard deviation from the mean reduces the probability of being hired on a permanent contract by 5 percent in France, 6 percent in Germany, and 24 percent in Portugal.[40]

39. Elasticities are computed by dividing the point estimate by the average incidence of permanent contracts among new hires (Online Appendix Table 1.1).

40. Results obtained with the leave-one-out HHI instrument are provided in Online Appendix Table 2.16. Point estimates tend to be smaller than with the baseline leave-one-out $1/N$ instrument. However, wherever significant at conventional levels with the latter, they are also significant with the former. In the case of Spain, the effect of labor market concentration on contract type is even negative and significant, while it is insignificant in our baseline estimate. Online Appendix Table 2.17 presents the results obtained when using HHIs based on employment rather than on hirings. They are equally robust in France and Germany, while the point estimate is slightly weaker in Portugal, although in the same range of magnitude as the baseline estimate. Robustness checks on the baseline leave-one-out $1/N$ instrument are provided in Online Appendix Tables 2.18 and 2.19 for the countries where the point estimates on log(*HHI*) are statistically significant in Table 6. When controlling for the share of four-digit occupations in new hires at the national level, the point estimates are of smaller magnitude in France and Portugal, but they are still within the confidence intervals of the baseline estimates. Moreover, as shown in Online Appendix Table 2.19, in France—and Germany—the results are robust to large violations of exogeneity of the instrument. This test would not be meaningful for Portugal since the baseline results are significant only at the 11 percent level. Online Appendix Tables 2.20, 2.21, 2.22, 2.23, and 2.24 provide robustness checks conducted using alternatively firm (instead of firm-by-municipality) identifiers, two-digit occupations and Arnold's (2021) method, as well as FUAs and NUTS-3 regions, respectively (instead of functional areas), when computing HHIs. Since implementing Arnold's method requires data on employment stocks, Italy and Spain cannot be included in Online Appendix Table 2.22. These robustness checks confirm the negative estimates for France and Germany but yield somewhat noisier results for Portugal.

In contrast, labor market concentration does not seem to affect the probability of being hired on a permanent contract in a noticeable way either in Italy or in Spain. This is likely due to the fact that, in these countries, most workers are hired on temporary contracts even in local labor markets where concentration is low. In fact, in markets that belong to the first decile of the concentration distribution, the proportion of new hires on temporary contracts is already very high in Italy (62.2 percent) and Spain (80.4 percent).[41] Therefore, there is not much room for further increasing the probability of being hired on a temporary contract when local labor markets become more concentrated. To the extent that Portugal also has a large proportion of new hires on temporary contracts, it is particularly remarkable that we do find a large and borderline significant effect in this country.

Our findings for Italy and Spain do not imply, however, that concentration in the labor market does not affect job security in these countries. In Table 7, we provide IV estimates of the effect of log(*HHI*) on the probability that an employee hired on a temporary contract at year t be converted to a permanent contract by the end of the following calendar year.[42] Our findings suggest that in local labor markets where concentration is higher, the probability of conversion to a permanent contract is indeed significantly lower, with an elasticity of -0.060 in Spain and -0.245 in Italy, when computed at sample average.[43] Increasing labor market concentration by one standard deviation from the mean therefore reduces the probability of conversion by 28 percent in Italy and 8 percent in Spain. The estimate for Italy looks large. However, it should be considered with some caution since it is sensitive to the specification we use. When we control for four-digit occupational shares in new hires at the national level, the point estimate on log(*HHI*) is three times lower (Online Appendix Table 2.32). This suggests that the instrument could be partly endogenous. Yet, as shown in Online Appendix Table 2.33, our estimate is robust to large violations of exogeneity (67 percent of the effect estimated in the reduced form), which makes it possible to draw clear, albeit cautious, conclusions from our findings. In Spain, the instrument does not seem to capture national trends in labor supply and demand since the results in Online Appendix Table 2.32 are similar to those presented in Table 7. They are also robust to violations of exogeneity of the instrument as large as 46 percent of the effect estimated in the reduced form (Online Appendix Table 2.33). Taken together, these findings suggest that when most workers are hired on temporary contracts (as in Italy and Spain), the negative effect of labor market concentration on job security materializes through a lower chance of subsequent conversion to a permanent contract, rather than through more precarious conditions at the time of hiring.

41. By contrast, this proportion is only 40.8 percent in France and 30.9 percent in Germany.

42. OLS estimates are provided in Online Appendix Table 2.25. The first-stage estimates are presented in Online Appendix Table 2.26.

43. Online Appendix Tables 2.27, 2.28, 2.29, and 2.30 provide robustness checks conducted using firm (instead of firm-by-municipality) identifiers, two-digit occupations, and FUAs and NUTS-3 regions, respectively (instead of functional areas), when computing HHIs. These robustness checks confirm our results. Results obtained with the leave-one-out HHI (rather than leave-one-out $1/N$) instrument are provided in Online Appendix Table 2.31. They are very similar to our baseline estimates.

Table 7
Labor Market Concentration and Conversions from Temporary to Permanent Contracts—IV Estimates, New Hires on Temporary Contracts Only

Dep. Var. Conversion	Italy	Spain
Log HHI	−0.0390***	−0.0034***
	(0.0089)	(0.0011)
KP F-test	61.6	1743.5
Observations	9,012,917	4,105,317

Notes: 2SLS estimates. The dependent variable is a dummy variable equal to one when the individual was hired on a temporary contract at year t and had/started a permanent contract with the same employer in the following calendar year; it is equal to zero otherwise. Local labor markets are defined based on four-digit occupations and functional areas (FAs). Control variables include gender, education (four categories), yearly dummies for workers' age, whether the individual was employed the year before hiring, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Standard errors are clustered at the labor-market-by-year level. KP F-test: Kleibergen–Paap Wald F-test. *$p < 0.1$, **$p < 0.05$, ***$p < 0.01$. Observations include singletons.

The magnitude of the effect on job security appears to be larger than the effect on wages in all countries where we can estimate the two of them—that is, France, Germany, and Portugal.[44] A possible explanation for this difference may be that, in contrast to what occurs for wages, the type of contract on which employees are hired is largely unregulated by sectoral collective agreements. This makes it easier for firms to use contract type, rather than wages, as a margin of adjustment when concentration increases. Consistent with this explanation, when we reestimate our baseline specification including an interaction between concentration and industry-level coverage of sectoral collective bargaining, the point estimate on this interaction turns out statistically insignificant both in Germany and Portugal (Table 8). This suggests that sectoral collective bargaining plays little role in mitigating the effect of concentration on job security.

Overall, our results suggest that when firms have some monopsony power, the cost for workers materializes in various dimensions of job quality—not only in the form of lower wages, but also in terms of poorer job security. Hence, considering only the wage effects of labor market concentration is likely to underestimate its true cost for workers. For France, Germany, and Portugal, we have estimates of the effect of labor market concentration on both wages and the probability to be hired on a permanent contract. Using information on the willingness

44. This is not due to the fact that, in contrast with Equation 1, Equation 2 is estimated on the subsample of new hires only and does not include individual fixed effects. When reestimating Equation 1 on new hires only and without individual fixed effects, we find that the magnitude of the negative impact of concentration on wages decreases in all countries.

Table 8
Labor Market Concentration and Probability of Being Hired on a Permanent Contract as a Function of Collective Bargaining Coverage—IV Estimates

Dep. Var. Permanent Contract	(1)	(2)
Panel A: Germany		
Log HHI	–0.0680***	–0.0663***
	(0.0123)	(0.0123)
Log HHI*Coverage share, Sector CBA	0.0074	
	(0.0057)	
Log HHI*High coverage dummy, Sector CBA		–0.0017
		(0.0018)
KP *F*-test	85.1	85.1
Observations	4,167,918	4,167,918
Panel B: Portugal		
Log HHI	–0.0514*	–0.0523*
	(0.0303)	(0.0309)
Log HHI*Coverage share, Sector CBA	0.0014	
	(0.0061)	
Log HHI*High coverage dummy, Sector CBA		0.0030
		(0.0023)
KP *F*-test	6.7	6.8
Observations	1,039,792	1,039,792

Notes: 2SLS estimates. The dependent variable is a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract. Local labor markets are defined based on four-digit occupations and functional areas (FAs). "Coverage share, Sector CBA" is the proportion of employees covered by a sector-level collective agreement in the industry where the individual is employed. "High coverage dummy, Sector CBA" is a dummy variable equal to one if the individual is employed in an industry with a proportion of employees covered by a sector-level collective agreement above the median, and zero otherwise. Control variables include gender, education (four categories), yearly dummies for workers' age, whether the individual was employed the year before, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. In Germany, control variables also include the share of employees covered by a works council in the industry interacted with log(*HHI*). Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Log(*HHI*) interacted with a collective bargaining indicator (respectively the works council variable) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation interacted with the collective bargaining indicator (respectively the works council variable). Standard errors are clustered at the labor-market-by-year level. KP *F*-test: Kleibergen–Paap Wald *F*-test. *$p < 0.1$, **$p < 0.05$, ***$p < 0.01$. Observations include singletons.

to pay for being hired on a permanent contract from the literature, we can roughly assess by how much the true cost of labor market concentration is underestimated when considering only its wage effects. Although the literature has estimated the willingness to pay for various types of working conditions and work arrangements (Eriksson and Kristensen 2014; He, Neumark, and Weng 2021; Maestas et al. 2023; Mas and Pallais 2017; Ameriks et al. 2020), few papers specifically look at the value for new hires of holding a permanent contract. Two exceptions are Albanese and Gallo (2020) and Godard and Le Bihan (2023). Using French data, the latter find that workers are indifferent between being hired on a permanent or temporary contract if the latter pays a wage that is at least 14.4 percent higher. A very similar value (11.3 percent) is found by Albanese and Gallo (2020) on Italian data. Using the French estimate, a back-of-the-envelope calculation suggests that when considering the reduction in job security triggered by labor market concentration in addition to its wage cost, the true cost for workers increases by 14 percent in France, 27 percent in Germany, and 29 percent in Portugal,[45] which is far from negligible. These estimates may also be a lower bound since, if labor market concentration degrades other dimensions of job quality, its overall cost for workers will be higher.

C. Heterogeneity by Gender, Age, Nationality, and Education

The evidence presented so far suggests a strong effect of labor market concentration on job quality. This is consistent with the idea that employers who have market power take advantage of it to reduce labor demand in order to lower the cost associated with wages and nonwage attributes. Employers depressing their labor demand may also be expected to become more selective in hiring in a context of asymmetric information on the labor market. In particular, they may prefer job candidates with long work experience since their resume provides a more accurate signal of their productivity. If this is the case, we may expect labor market concentration to be particularly harmful to the wages and job security of workers with shorter work experience and, in particular, youth, women, and foreign workers.

To test this hypothesis, we interact log(*HHI*) separately with dummies for being: (i) a man or a woman, (ii) younger than 25 or aged 25 and above, and (iii) a domestic or foreign citizen, in Equations 1 and 2. The results are presented in Figures 1, 2, and 3.

As evidenced on Figure 1, Panel A, the effect of labor market concentration on wages appears to be very similar for men and women in all countries. It is slightly larger for women in Denmark, France, and Portugal, but slightly smaller in Germany, and in both cases confidence intervals at the 90 percent level largely overlap. The

45. The estimates of the effect on the overall cost are obtained by multiplying the point estimate of the effect of labor market concentration on the probability of being hired on a permanent contract (Table 6) by 0.144, and then adding the wage elasticity with respect to labor market concentration (Table 2). The percentage increase of the effect is then obtained by dividing this sum by the wage elasticity, subtracting one and multiplying the overall result by 100.

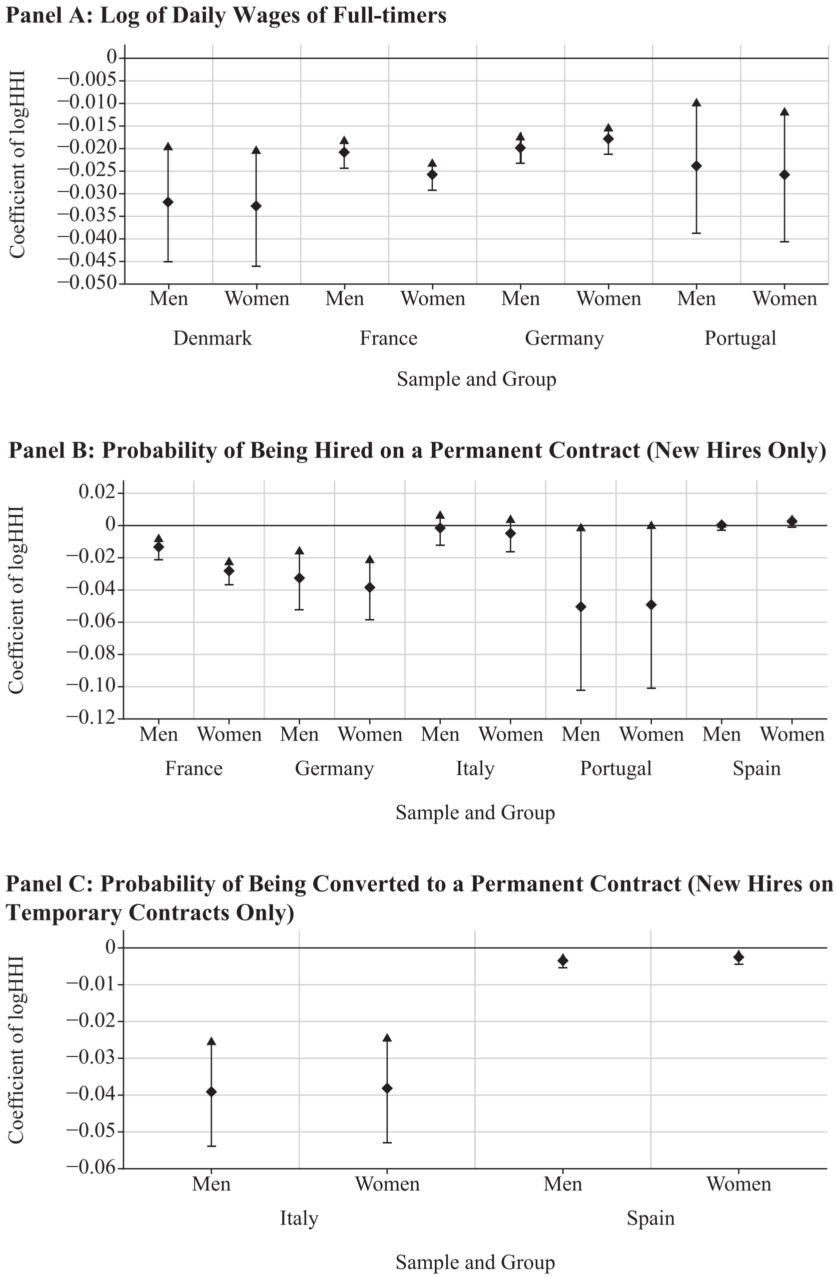

Figure 1
Effect of Labor Market Concentration by Gender

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same holds when considering the probability of being hired on a permanent contract. The point estimate on labor market concentration is slightly more negative for women in France, Germany, and Italy, but this is not the case in Portugal and Spain, and here again confidence intervals overlap (Figure 1, Panel B). Finally, in the case of conversions too, no systematic gender difference is uncovered by our analysis (Figure 1, Panel C).

Regarding youth, results are very similar. Panel A of Figure 2 points to a more negative effect of labor market concentration on youth's wages in Denmark and France and, to a lower extent in Portugal, but the opposite holds in Germany, and confidence intervals largely overlap in most countries. The same goes for the probability of being hired on a permanent contract; see Panel B—the point estimates on labor market concentration tend to be more negative for youth in France, Germany, Italy, and Spain, but this is not the case in Portugal, and confidence intervals overlap in all countries. For conversions, differences in the effect by age are small and go in opposite directions for Spain and Italy (Figure 2, Panel C).

These findings suggest that there is no systematic difference in the effect of labor market concentration on wages and job security across gender or age. However, this does not mean that women and youth are exposed to the same degree of monopsony power as men and older adults. It has been shown that women tend to search for jobs closer to their home and are ready to accept a significant wage penalty for a closer job (Le Barbanchon, Rathelot, and Roulet 2020; Jacob et al. 2019). For the same reason, they may also be willing to accept lower job security. Youth may also search closer to their home if they live with their parents and have not obtained yet their driving license. Thus, a given level of concentration will imply fewer acceptable outside options for women and young workers, and therefore lower wages and possibly lower job

Notes: 2SLS estimates. Panels report point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with gender dummies. Local labor markets are defined based on four-digit occupations and Functional Areas (FAs). Panel A: The dependent variable is log(wage). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel B: The dependent variable is a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract. Control variables include gender, education and yearly dummies for workers' age, whether the individual was employed the year before, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel C reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with gender dummies. The dependent variable is a dummy variable equal to one when the individual was hired on a temporary contract at year *t* and had/started a permanent contract with the same employer in the following calendar year. Control variables include gender, education and yearly dummies for workers' age, whether the individual was employed the year before hiring, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Each interaction between log(*HHI*) and a gender dummy is instrumented by that gender dummy interacted with the instrument for log(*HHI*). Standard errors are clustered at the labor-market-by-year level.

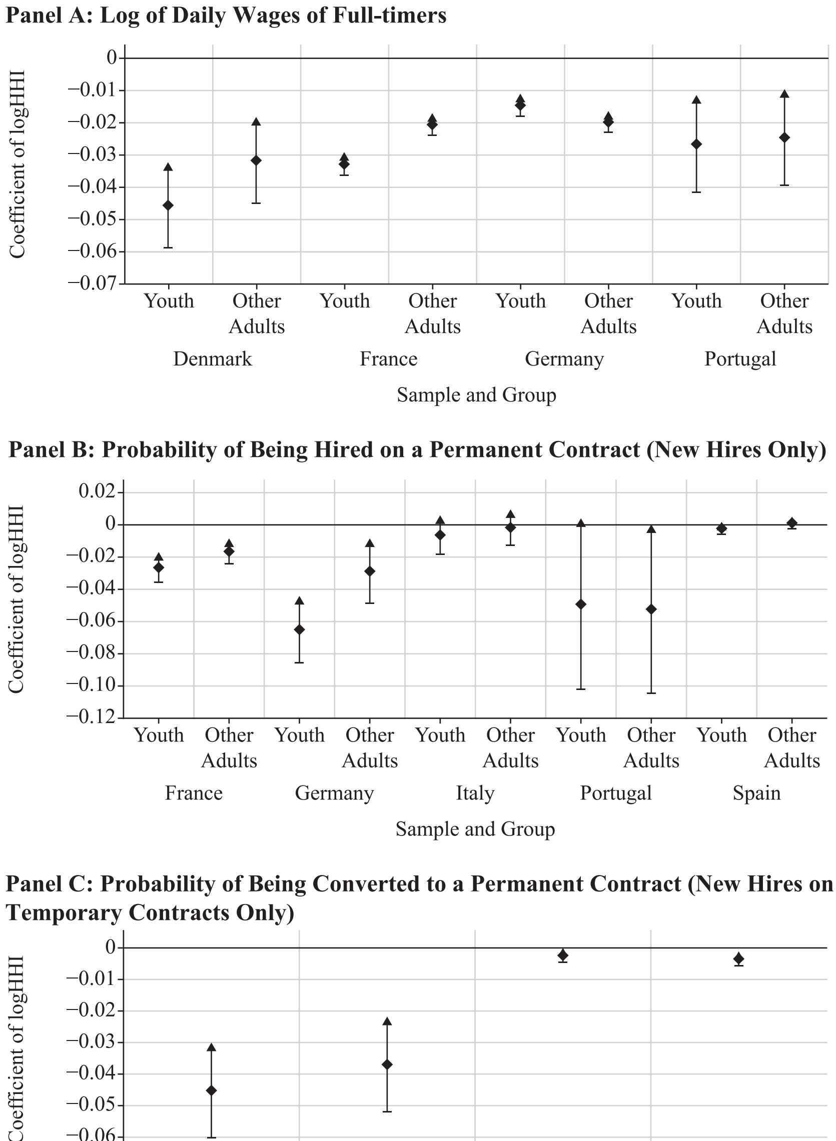

Figure 2
Effect of Labor Market Concentration by Age Group

(legend continued on next page)

security.[46] Despite this difference, an increase in labor market concentration by, say, 10 percent may still have a similar percentage effect on the rarefaction of available alternatives for both men and women on the one hand and older workers and youth on the other hand, consistent with the pattern of results shown in Figures 1 and 2.

As regards nationality, the effect of labor market concentration on wages does not seem to be much different across domestic and foreign citizens (Figure 3, Panel A). It is slightly larger for the former in Denmark, France, and Germany, but confidence intervals largely overlap, and there is virtually no difference across both groups in Portugal. The pattern is very similar for the probability of being hired on a permanent contract. The point estimates are more negative for domestic citizens in Germany and Portugal, but the opposite holds in France and Italy, and the effect is virtually zero in Spain. As in most previous cases, confidence intervals overlap in all countries (Figure 3, Panel B). Finally, concerning conversions, foreign citizens are slightly more affected by labor market concentration than domestic citizens, but the point estimates are within one standard deviation of each other (Figure 3, Panel C). These findings suggest that there is no clear systematic difference in the effect of labor market concentration on wages and job security across nationality.

Labor market concentration could also be particularly harmful to workers with low levels of education, since they tend to have fewer outside options (Caldwell and Danieli 2024). The results obtained when interacting log(*HHI*) with a dummy variable for

46. In fact, the literature on own-firm-labor-supply elasticity systematically finds lower elasticities for women than for men, thereby suggesting that the former are more exposed to monopsony power than the latter (Manning 2021).

Notes: 2SLS estimates. Local labor markets are defined based on four-digit occupations and functional areas (FAs). Panel A reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with dummy variables for being younger than 25 (respectively aged 25 and above). The dependent variable is log(wage). Local labor markets are defined based on four-digit occupations and functional areas (FAs). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel B figure reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with dummy variables for being younger than 25 (respectively aged 25 and above). The dependent variable is a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract. Control variables include gender, education and yearly dummies for workers' age, whether the individual was employed the year before, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel C reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with dummy variables for being younger than 25 (respectively aged 25 and above). The dependent variable is a dummy variable equal to one when the individual was hired on a temporary contract at year *t* and had/started a permanent contract with the same employer in the following calendar year. Control variables include gender, education and yearly dummies for workers' age, whether the individual was employed the year before hiring, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Each interaction between log(*HHI*) and an age category is instrumented by that age category interacted with the instrument for log(*HHI*). Monthly wages instead of daily wages are used for Portugal. Standard errors are clustered at the labor-market-by-year level.

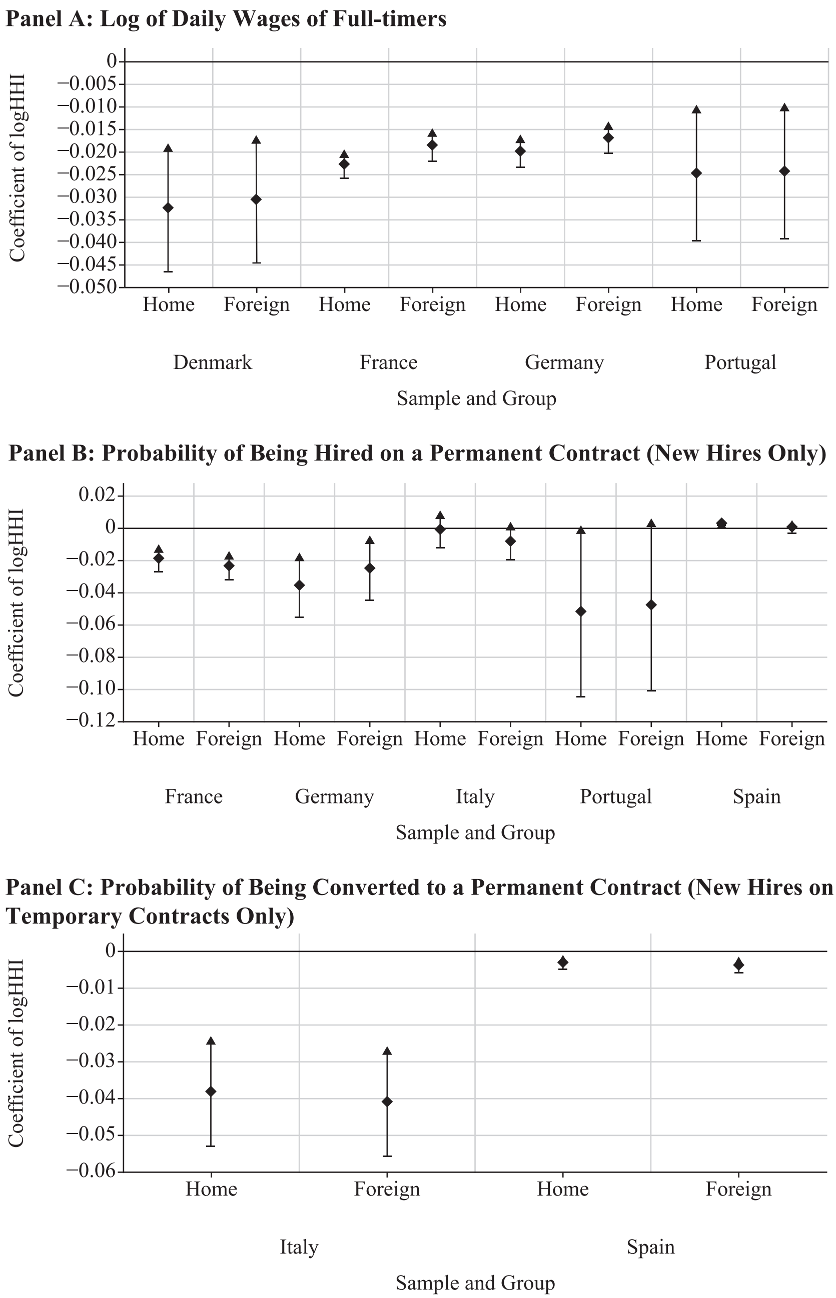

Figure 3

Effect of Labor Market Concentration by Nationality

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having upper secondary education or above, rather than below, are presented in Figure 4. As regards the effect on wages, in all countries except France, the point estimates are more negative for high- than for low-educated workers, although confidence intervals overlap in all countries (Figure 4, Panel A). This may seem surprising considering that, as mentioned above, high-education workers tend to have more outside options. However, it may be explained by the fact that more educated workers tend to work in highly skilled occupations, where pay levels are more individualized and hence more exposed to downward wage pressures when employers' monopsony power increases. Wages of lower-educated workers, by contrast, are more often set up by minimum wage legislation or sectoral collective agreements and therefore tend to be more protected against such pressures. Our findings regarding the probability of being hired on a permanent contract are consistent with this interpretation. In this case, in all countries, the negative effect of labor market concentration tends to be stronger for low-educated workers,[47] although, as before, confidence intervals overlap across both groups (Figure 4, Panel B).[48] This overall pattern of results is consistent with the fact that, as mentioned in Section IV.B above, contract type is largely unregulated by sectoral collective agreements.

47. In Italy, the point estimates are significant for both low- and high-educated workers, whereas our estimate for the whole population (Table 6) is insignificant at conventional levels. This is due to differences in the number of observations across samples since the information on education is missing for 21 percent of the observations in our original sample.

48. We find virtually no significant difference across education levels as regards conversions (Figure 4, Panel C).

Notes: 2SLS estimates. Local labor markets are defined based on four-digit occupations and functional areas (FAs). Panel A reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with nationality dummies. The dependent variable is log(wage). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel B reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with nationality dummies. The dependent variable is a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract. Control variables include gender, nationality, education and yearly dummies for workers' age, whether the individual was employed the year before, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel C reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with nationality dummies. The dependent variable is a dummy variable equal to one when the individual was hired on a temporary contract at year t and had/started a permanent contract with the same employer in the following calendar year. Control variables include gender, nationality, education and yearly dummies for workers' age, whether the individual was employed the year before hiring, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Each interaction between log(*HHI*) and a nationality dummy is instrumented that by that nationality dummy interacted with the instrument for log(*HHI*). Monthly wages instead of daily wages are used for Portugal. In France, *Foreign* corresponds to foreign born. Standard errors are clustered at the labor-market-by-year level.

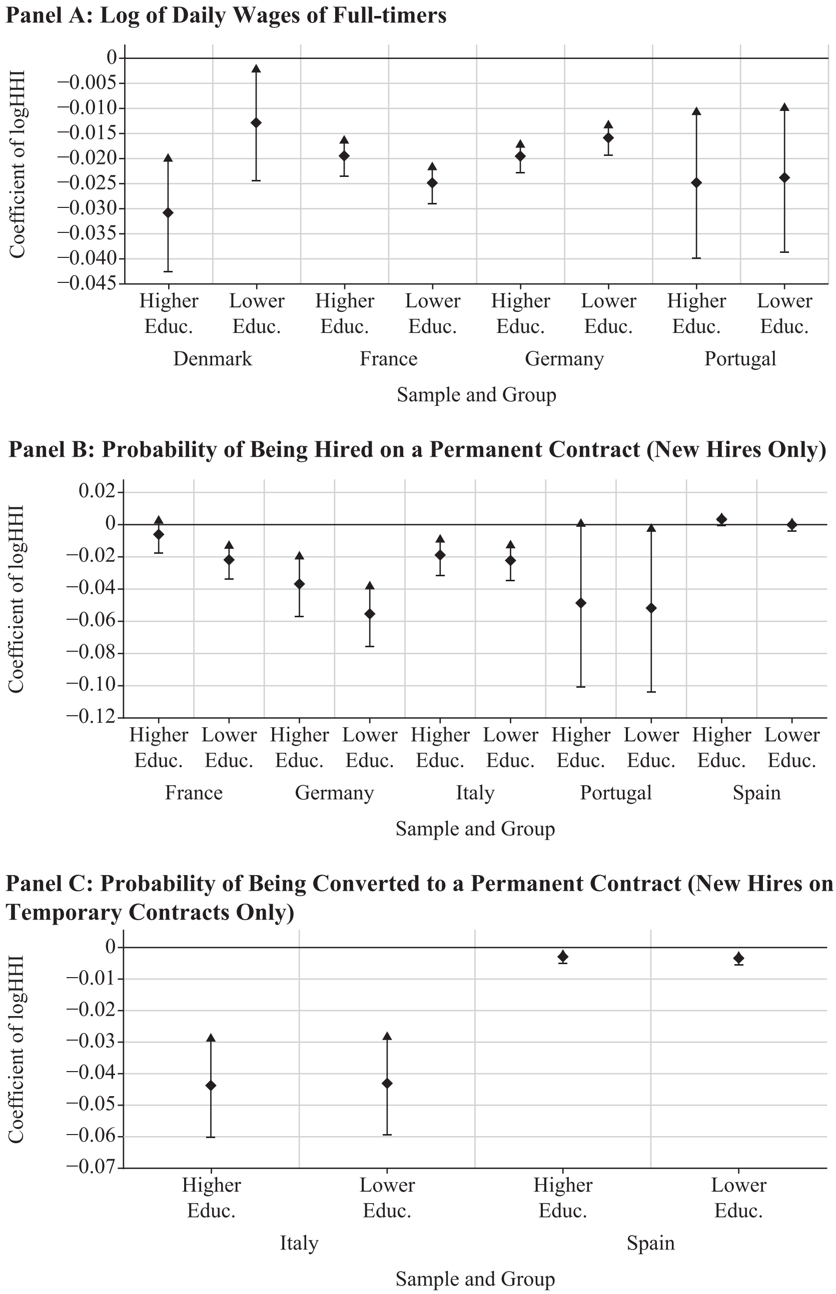

Figure 4

Effect of Labor Market Concentration by Education

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Having fewer outside options, low-educated workers therefore tend to be more exposed to a deterioration of job security when their employers gain monopsony power.

V. Conclusion

This work contributes to the debate on the effects of labor market concentration on job quality. We leverage rich administrative linked employer–employee data from Denmark, France, Germany, Italy, Portugal, and Spain in the 2010s to provide the first comparable cross-country evidence in the literature. First, we show that the distribution of labor market concentration is similar across these six countries. Second, controlling for productivity and local product market concentration, we show that, despite different labor market institutions, the elasticities of wages with respect to labor market concentration are very similar across countries, ranging from −0.019 in Germany to −0.033 in Denmark. We then consider a second dimension of job quality, namely job security. Our results suggest that higher labor market concentration reduces the probability of being hired on a permanent contract in France, Germany, and Portugal, with elasticities as large as −0.046, −0.053, and −0.233, respectively. In Italy and Spain, where most workers are anyway predominantly hired on temporary contracts, we detect no significant effect of labor market concentration on such probability. However, we find that higher concentration significantly reduces the probability of being converted to a permanent contract once hired on a temporary one.

These results suggest that firm monopsony power not only negatively affects wages but also degrades job security. Policy interventions limiting employer concentration

Notes: Panels reports point estimates and 90 percent confidence intervals of the coefficients of the interactions of log(*HHI*) with education dummies. *Higher education* is a dummy variable for having upper secondary education or above. *Lower education* is a dummy variable for having less than upper secondary education. Local labor markets are defined based on four-digit occupations and Functional Areas (FAs). Panel A: The dependent variable is log(wage). Control variables include yearly dummies for workers' age, whether the individual is a new hire, whether the individual was employed the year before if new hire, as well as individual fixed effects, firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel B: The dependent variable is a dummy variable equal to one if the individual is hired on a permanent contract and zero if hired on a temporary contract. Control variables include gender, education and yearly dummies for workers' age, whether the individual was employed the year before, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Panel C: The dependent variable is a dummy variable equal to one when the individual was hired on a temporary contract at year t and had/started a permanent contract with the same employer in the following calendar year. Control variables include gender, education and yearly dummies for workers' age, whether the individual was employed the year before hiring, whether they work full-time or not, as well as firm-by-municipality-by-year fixed effects, sector and establishment fixed effects (where not collinear with firm-by-municipality fixed effects), and local labor market fixed effects. Log(*HHI*) is instrumented by the average of the log inverse number of firms with positive hiring in other FAs for the same occupation. Each interaction between log(*HHI*) and an education dummy is instrumented by that education dummy interacted with the instrument for log(*HHI*). Monthly wages instead of daily wages are used for Portugal. Standard errors are clustered at the labor-market-by-year level.

and/or its effects are therefore likely to improve labor market outcomes along both dimensions. Potential interventions may include enforcement actions by antitrust authorities, such as taking systematically into account labor market outcomes in merger reviews and cracking down on labor market collusion, including on no-poaching and wage-fixing agreements (for example, Hovenkamp and Marinescu 2019). On the one hand, evidence suggests that mergers increasing concentration do not need to create dominant employers to have a strong negative effect in the labor market (for example, Arnold 2021; Prager and Schmitt 2021). On the other hand, collusion is more likely to occur in concentrated markets, since coordination among fewer actors is typically easier to sustain (for example, Asker and Nocke 2021). More generally, the promotion of entrepreneurship and firm creation, leading to the emergence of new employers, will also reduce the labor market power of incumbent firms.

Other interventions to counteract the effect of labor market concentration are rather in the realm of labor policy. These notably include direct interventions to facilitate collective bargaining.[49] Labor unions and collective bargaining have been shown to help counterbalance the effect of firm market power in case of monopsonistic competition. In line with this literature, we indeed find small and strikingly similar wage elasticities in Denmark, France, Germany, and Portugal and show that the effect of labor market concentration on contract type is much larger than on wages. We provide suggestive evidence that this pattern of results could be explained by the fact that the countries we study have high coverage of collective bargaining, which takes place predominantly at the industry level. Since wages are largely set by sectoral collective agreements, whereas contract types are not, firms likely have more margin of maneuver to reduce job security rather than adjust wages downward when labor market concentration increases. This suggests that enlarging the scope of collective bargaining to contract types may help limit the effect of monopsony power on workers' welfare.

While contract types are an important feature of the employment relationship, exploring the impact of monopsony power on additional dimensions of job quality would be important to fully measure its overall impact on workers' welfare. As discussed in Manning (2003), we would expect a negative effect of monopsony power on other nonwage attributes that are both costly to the employer and valuable to the employee, for example, paid days off, employers' contributions to private pension plans, employer-paid extensions of health insurance coverage, and other fringe benefits. As our administrative data lack information on such aspects of the employment relationship, we were not able to pursue this analysis here. However, exploring the effects of firm monopsony power on further dimensions of job quality remains a promising avenue for future research.

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