Laboratoire d'Economie de Dauphine

DAUPHINE UNIVERSITÉ PARIS | LEDa LEGOS EA 4404

WP n°6/2019

Document de travail

Labor Market Concentration and Stayers'Wages: Evidence from France

Andrea Bassanini
Cyprien Batut
Eve Caroli

Pôle Laboratoire d'Economie et de Gestion des Organisations de Santé (LEGOS)
Place du Maréchal de Lattre de Tassigny 75775 Paris Cedex 16
Tél (33) 01 44 05 44 46 Fax (33) 01 44 05 40 67
Site : www.legos.daupine.fr

Labor Market Concentration and Stayers' Wages: Evidence from France*

July 15, 2020

Andrea Bassanini
(OECD and IZA)

Cyprien Batut
(EHESS and PSE)

Eve Caroli
(Paris-Dauphine Université, LEDa, PSL and IZA)

Abstract

We investigate the impact of labor market concentration on stayers' wages, where stayers are defined as individuals who remain employed in the same firm for at least two consecutive years. Using administrative data for France, we show that the elasticity of stayers' wages to labor market concentration is negative but small (about -0.014) once controlling for firm productivity, product market competition and match-specific heterogeneity. Given the strong wage rigidities characterizing the French labor market, this estimate can be seen as a lower bound of the effect of labor market concentration on stayers' wages in an international perspective.

JEL Codes: J31, J42, L41

Keywords: *labor market concentration, monopsony, wages, stayers, match-specific heterogeneity*

*The views expressed here do not represent those of the OECD or its member countries. Cyprien Batut acknowledges support from the PSE Labor Chair. This research has also been supported by a public grant overseen by the French National Research Agency (ANR) as part of the 'Investissements d'avenir' program (reference: ANR-10-EQPX-17, Centre d'accès sécurisé aux données, CASD). Usual disclaimers apply.

1 Introduction

How labor market concentration affects wages has been the subject of a burgeoning literature in recent years. Many papers have shown that a substantial proportion of individuals are employed in labor markets that are at least moderately concentrated according to the thresholds defined by the US Horizontal Merger Guidelines and that this has a depressing effect on average wages, consistent with a monopsony model (Azar et al., 2018; Martins, 2018; Abel et al., 2018; Rinz, 2018; Benmelech et al., 2018). Beyond average wages, there is evidence that concentration affects posted wages for new vacancies (Azar et al., 2020) and actual wages of new hires (Marinescu et al., 2019).

In this paper we investigate the impact of labor market concentration on stayers' wages, where stayers are defined as individuals who remain employed in the same firm for at least two consecutive years. Looking at stayers is important since they represent a large share of all employees in any given year in all OECD countries (OECD, 2010). Stayers' wages have also been shown to be the largest contributor to aggregate wage growth, at least in recent years (Hahn et al., 2017, 2018). Understanding the impact of monopsony on this group of workers is therefore important to understand how monopsony may explain wage trends. However, so far, no systematic evidence has been provided on how labor market concentration affects stayers' wages. Aggregate effects found in the literature could in principle be driven by new hires only or by both new hires and stayers. In the former case though, this would imply that labor market concentration only affects wages at the margin, i.e. for individuals who change jobs. In this paper, we use French data and show that the elasticity of stayers' wages to labor market concentration is negative but small: about -0.014 in our preferred specification. France is interesting in this respect since, due to automatic extension of collective agreements, almost all employees are covered by them (Babecký et al., 2010) so that wage rigidities are strong. Our results suggest that, even in this context, labor market concentration has a small depressing effect on the earnings of the vast majority of workers, i.e. those who do not change employer from one year to the other.

Estimating the elasticity of wages to labor market concentration on stayers only also has the advantage of cancelling out any potential confounding effects of changes in the composition of the workforce and/or assortative matching between workers and firms. Qiu and Sojourner (2019) find that labor market concentration tends to reduce the average level of education of employees, so that part of the decrease in wages associated with higher concentration could be due to a reduction in the quality of the workforce. Moreover, Macaluso et al. (2019) show that the greater the labor market concentration, the higher the skill requirements imposed by firms, conditional on workers' education. This suggests that firms in concentrated labor

markets are more selective in choosing workers who best fit their specific needs. The effects we estimate in the current paper are net of any composition and/or sorting effect since not only do we account for time-invariant workers' characteristics but also for match-specific heterogeneity.

To our knowledge, the only other paper estimating the impact of labor market concentration on stayers' wages is Arnold (2019). It shows that mergers and acquisitions that lead to higher concentration have a negative effect on the wages of employees who stay in the firms that have merged. We identify the effect of labor market concentration on non-merged companies and show that when labor market concentration changes independently of changes in the firms boundaries due to mergers, stayers' wages are moderately affected. This suggests that employers may be able to take advantage of the reduction in outside options induced by greater concentration to reduce stayers' wages. However, this effect is limited in magnitude presumably because of generalized coverage by collective agreements in France.

The remainder of the paper is structured as follows. Section 2 lays out our empirical strategy. Section 3 describes the data that we use and presents summary statistics. Section 4 presents the results and Section 5 concludes.

2 Empirical specification

2.1 Labor market concentration

As is standard in the literature (Azar et al., 2020; Martins, 2018; Marinescu et al., 2019), we measure employer concentration using the Herfindhal-Hirschman Index (HHI) computed on hirings:

$$HHI_{o,z,t} = \sum_{f=1}^{N_{o,z,t}} s_{f,o,z,t}^2 \tag{1}$$

where $HHI_{o,z,t}$ is the HHI for occupation o in commuting zone z (which define the local labor market $l = (o, z)$) at year t. $N_{o,z,t}$ is the number of firms that have positive hirings in local labor market l at time t and $s_{f,o,z,t}$ is the share of firm f in hirings in local labor market l at time t. With this definition, $HHI_{o,z,t}$ ranges from 0 (no concentration) to 1 (one firm hiring in the market).

We use an HHI based on hirings to measure labor market concentration since, as emphasized by Marinescu et al. (2019), this is the best way to capture job opportunities available to workers searching for a job. To substantiate this claim, consider a search and matching model with granular search where concentration affects wages by affecting workers' outside options - see Jarosch et al. (2019). In this model, the index of labor market concentration that is relevant for wage determination may be measured indifferently with an HHI based on hirings or on the employment stock,

as long as the environment is stationary. However, when the environment in non-stationary, an HHI based on hirings is much more relevant since downsizing firms may have a positive share of the stock of employment in a local labor market, while their hirings are zero so that they do not contribute to creating outside options for workers in that labor market. Nevertheless, since prior literature has used employment stocks to measure labor market concentration (Abel et al., 2018; Rinz, 2018; Benmelech et al., 2018), and since in a standard Cournot model of oligopsony, wages are inversely related to the HHI measured in terms of employment (Boal and Ransom, 1997), we also use the latter as a robustness check.

2.2 Labor market concentration and wages

We estimate the impact of labor market concentration on individual wages. Our baseline specification is as follows:

$$log(w_{i,f,o,z,s,t}) = \beta log(HHI_{o,z,t}) + \boldsymbol{X_{i,f,o,z,s,t}}\gamma + \mu_i + \mu_{f,t} + \mu_{o,z} + \epsilon_{i,f,o,z,t} \quad (2)$$

where i indexes the individual, f the firm and s the sector in which the firm is active. w denotes the individual wage, $\boldsymbol{X}$ is a vector of age dummies - one for each year of age - and μ are fixed effects. In this baseline specification, we control for individual and local-labor-market - $l = (o, z)$ - fixed effects. We also control for firm-by-time fixed effects to capture firm productivity. Standard errors are clustered at the commuting zone level.

We estimate equation (2) on stayers only, i.e. on individuals who remain employed in the same firm for at least two consecutive years. Focusing on stayers as opposed to movers is interesting *per se* since they represent the vast majority of the labor force (more than 80% in France - see Section 3). Moreover, as emphasized by Qiu and Sojourner (2019), labor market concentration and wages may be affected by workforce composition. In turn, this may generate assortative matching between workers and firms. For example, firms with buyer power on the labor market may become more selective and retain only those workers that better match their idiosyncratic needs. Focusing on stayers' wages has another advantage in this respect: it allows controlling for match-specific heterogeneity by augmenting equation (2) with a spell fixed effect ($\mu_{i,S,f}$), where S denotes the spell of individual i in firm f - see equation (5) below. When doing so, $\hat{\beta}$ is identified only by variations within the same match, defined as a consecutive spell of employment of a worker within a firm.

Arguably, when estimating the impact of labor market concentration on wages, product market competition is a potential confounder. In our baseline specification, we control for firm-by-time fixed effects. If local firms produce for the national or international market - and not only for the local one -, product market competition is firm specific and firm-by-year fixed effects would control for it. However, if firms in

a given geographical area produce for the local market, a better way to control for product market competition is to include sector-by-commuting-zone-by-year fixed effects. This is what we do in our extended specifications - see equation (5) below.

A key threat to identification in this set-up is that an omitted time-varying variable could be correlated with the HHI and determine wages. This is the case, for example, if a negative shock on the supply of labor takes place in a local labor market $l = (o, z)$. This shock is likely to raise wages. If productivity stays unchanged, unit labor costs go up, thereby likely reducing the number of local firms which find it profitable to employ this type of labor. As a consequence, labor market concentration would increase thus giving rise to a positive correlation between HHI and wages that would, in fact, be due to reverse causality. To deal with this endogeneity problem, the literature suggests instrumenting $log(HHI_{o,z,t})$ with the average of $log(1/N_{o,z',t})$ - where $N_{o,z',t}$ is the number of firms with positive hirings in all other commuting zones z' for the same occupation and time period (Azar et al., 2020; Martins, 2018; Qiu and Sojourner, 2019). $1/N_{o,z',t}$ corresponds to the value of the HHI in local labor market $l' = (o, z')$ when all firms have the same hiring share in that market. This instrument provides a source of variation in labor market concentration relying on national rather than local changes in the occupation we consider.[1]

However, since individuals living close to the border of a commuting zone may be working either in this zone or in the bordering one, any shock on the local labor supply in a given occupation taking place in the periphery of a commuting zone is likely to affect the bordering commuting zones too. To deal with this issue, when building our instrument, we improve on the existing literature by not only removing the commuting zone we consider but also all the zones that have a common border with it. Thus doing, we considerably reduce the risk that spillovers across local labor markets may threaten the orthogonality of our instrument.

Another problem raised by this instrument is that since its variation essentially relies on national changes in occupations (i.e. variations in the o and t dimensions only), it does not allow including occupation-by-time fixed effects in the model. Now, if there is a national demand or supply shock in occupation o at time t, the instrument may capture this shock, which will affect both labor market concentration and wages in $l' = (o, z')$ and in $l = (o, z)$. In this case, the instrument is no longer exogenous. To allow controlling for aggregate occupational shocks (by including occupation-by-time fixed effects in our regressions), we develop a related instrument exploiting the industry composition within occupations. Namely, we instrument $log(HHI_{o,z,t})$ with the weighted sum of $log(1/N_{o,z',t})$, where the weights are indexed on the proximity in the industry composition of each occupation across commuting

[1] Instrumenting a variable in one zone using the average of this variable in other zones (Hausman instruments) is standard in international economics and industrial organization - see e.g. Hausman et al. (1994), Autor et al. (2013), Bai et al. (2017) and Azar et al. (2019a)

zones z and z'.

More specifically, for a given occupation, we take each commuting zone in which this occupation exists and we consider the employment share of each industry s within that local labor market ($e_{s,o,z,t}$). We then define the proximity of labor markets $l = (o, z)$ and $l' = (o, z')$ as regards industry s as:

$$p_{s,o,z,z',t} = \sqrt{1 - (e_{s,o,z,t} - e_{s,o,z',t})^2} \tag{3}$$

The intuition behind this measure of proximity is that whenever two local labor markets have the same industry composition, $e_{s,o,z,t} = e_{s,o,z',t}$, so that proximity p is equal to 1.

In order to compute the proximity in industry composition across labor markets $l = (o, z)$ and $l' = (o, z')$, one could be tempted to average $p_{s,o,z,z',t}$ across all industries. However, thus doing, proximity would be equal to 1 whenever an industry is absent from both labor markets. This would give too large a weight to industries in which both labor markets are not specialized. More generally this would be a problem for all industries whose employment share is small in both labor markets.

To overcome this problem we define an adjusted measure of proximity in industry composition across labor markets $l = (o, z)$ and $l' = (o, z')$:

$$adj\_p_{o,z,z',t} = \frac{1}{n_s} \sqrt{\sum_{s=1}^{n_s} (e_{s,o,z,t} \times p_{s,o,z,z',t})(e_{s,o,z',t} \times p_{s,o,z,z',t})} \tag{4}$$

where n_s is the number of industries.[2] By multiplying p by the employment share of industry s in each local labor market, we ensure that when these employment share are small, adjusted proximity is small as well.

Our instrument is then computed as the weighted sum of $log(1/N_{o,z',t})$, where the weights are given by the vector of $adj\_p_{o,z,z',t}$ normalized so that $\sum_{z'} adj\_p_{o,z,z',t} = 1$. It provides a source of variation in labor market concentration based on national industry shocks that affect the number of firms active in specific local labor markets depending on their exposure to that particular industry. To the extent that it varies with the industry composition of local labor markets, this instrument has a significant variation in the o, z and t dimensions. This allows controlling for both occupation-by-time and occupation-by-sector fixed effects.

Our most complete specification therefore writes:

$$log(w_{i,f,o,z,s,t}) = \beta log(HHI_{o,z,t}) + \boldsymbol{X_{i,f,o,z,s,t}}\gamma + \mu_{iSf} + \mu_{f,t} + \mu_{o,z} + \mu_{s,z,t} + \mu_{o,t} + \mu_{o,s} + \epsilon_{i,f,o,z,t} \tag{5}$$

where μ_{iSf} is the spell fixed effect.

[2] We compute this proximity as of 2009 using 17 aggregate sectors from NAF Rev.2, which is approximately equivalent to the 1-digit letter of the NACE Rev.2 classification.

3 Data

We use two datasets extracted from the French Social Security records (DADS). The first dataset (DADS-Postes) covers the universe of workers and establishments in all industries except agriculture, part of the food-processing industry, rural financial institutions (e.g. Crédit Agricole) and public administrations. This contains information on establishment location (municipality) and the firm to which the establishment belongs. Moreover it provides information on hours worked, gross wages (constructed as gross annual wages divided by the number of hours worked), workers' age, gender and 4-digit occupation for all employees with non-zero hours worked in a given year. Establishments and firms have a unique identifier which is invariant over time, except when sold out to another company, in which case they are assigned a new identifier. By contrast, for the sake of anonymity, workers' identifiers are changed every year. However, for any given year, we know in which establishments employees were working the year before. We use data starting in 2009 since information on occupations was not systematically reported before that date. We match each municipality contained in the DADS-postes with the 2010 commuting zones using a mapping provided by the French Statistical Institute (INSEE).

For the subset of workers in the DADS-Postes who are born in October of each year, there exists a panel which maintains the same identifier over time for each worker and hence allows following workers across various employers and years. This panel (DADS-Panel) is currently available until 2015. For this reason, we limit our analysis to 2009-2015.

We use the whole DADS-Postes to construct HHIs based on hirings (and, as a robustness check, on employment). We only consider business companies, and exclude workers on training contracts or on occasional jobs.[3] Employment is defined in full-time equivalent terms. A new hire in a given year is defined as a worker who did not work for any establishment of the firm the year before. We only keep local labor markets with at least 10 employees in each year of our time window.

Descriptive statistics of concentration in French local labor markets are reported in Appendix Table A1 and Figures A1 and A2. When measured with reference to hirings, mean concentration weighted by employment is relatively stable over time around 0.11, which is below the threshold for moderate concentration (0.15) defined by the US antitrust authorities. Unsurprisingly, it is even lower when measured on the basis of employment: about 0.09. However, mean values of HHIs turn out to be much larger than median values, suggesting that a number of local labor markets are highly concentrated. As a matter of fact, although 79.2% of workers are employed in a market where the HHI based on hirings is lower than 0.15 (resp. 83.8% for the HHI based on employment), 11.4% (resp. 9.1%) are employed in local

[3] The so-called *emplois annexes*.

labor markets where the HHI based on hirings (resp. employment) is higher than 0.25, which corresponds to high concentration - see Appendix Table A2. Moreover, the (unweighted) proportion of local labor markets with an HHI above 0.25 is non negligible - see Figures A3 and A4 -, at least when labor market concentration is defined on the basis of hirings. In this case, almost 6% of labor markets even have $HHI = 1$ in 2009. This is consistent with the rather low share of individuals facing highly concentrated labor markets since large markets tend to be less concentrated than smaller ones.

As in most countries - see Abel et al. (2018), Rinz (2018) and Azar et al. (2019b) - local labor markets are more concentrated in mostly rural than in mostly urban commuting zones in France - see Figures A5 and A6. This is accounted for in our regressions by including occupation-by-commuting zone fixed effects.

We estimate our wage regressions on the subset of stayers employed in business companies and for whom we dispose of a panel. We keep workers aged 15 to 74. As is standard when using the DADS and to eliminate implausible values of hourly wages due to misreporting of either annual wages or hours worked, we drop the lowest and highest percentiles of the hourly wage distribution each year. Descriptive statistics for this panel are presented in Appendix Table A3. Our observations are individual-by-firm-by-year triples. Stayers represent 83.5% of this sample. Their average age is 40, and men represent about 51.9% of this group, i.e. about the same proportion as in the whole sample.

4 Results

We first estimate the impact of labor market concentration on individual wages using a measure of the HHI based on hirings. As evidenced in Table 1, OLS estimates are negative although insignificant at conventional levels, no matter which set of fixed effects we include. These results could suggest that labor market concentration has no impact on stayers' wages, as one could expect in a country with high wage rigidity. However, this could also be due to endogneity if local labor supply shocks simultaneously drive wages and the number of firms in the local labor market - see Section 2.2.

In order to disentangle between these explanations, we run IV estimates in which $Log(HHI_{o,z,t})$ is instrumented by the weighted sum of $log(1/N_{o,z',t})$, where $N_{o,z',t}$ is the number of firms with positive hirings in commuting zones z' excluding z and all commuting zones that have a border with z. The weights are proportional to the adjusted proximity in the industry composition within occupation o across commuting zones - see equation (4). This instrument is strongly correlated with labor market concentration, as evidenced by the first-stage F-statistics reported at the bottom of Table 2. When estimated in this way, the impact of the HHI on

individual wages turns out to be negative and significant, whatever the specification we consider - except in col (3). Controlling for individual and firm-by-time fixed effects or for spell fixed effects yield very similar results - see cols (1) and (2). When controlling for product market competition and aggregate occupational shocks, by adding sector-by commuting zone-by-year fixed effects and occupation-by-time fixed effects, our estimates become larger - see col. (4). A similar point estimate is found when adding sector-by-occupation fixed effects - see col (5).

As a robustness check, we estimate the impact of labor market concentration on individual wages using a measure of the HHI based on employment. Point estimates are close in magnitude to those obtained when using an HHI based on hirings - see Table 3. Whatever measure of HHI we use, our results suggest that, on average, a 10% increase in labor market concentration decreases stayers' wages by 0.14 to 0.15% - see col (5) of Tables 2 and 3 -, corresponding to an elasticity of -0.014 (-0.015 respectively). This suggests that labor market concentration has a limited depressing effect on stayers' wages in France.

We also re-estimate our IV model with the standard instrument used in the literature. Namely, we instrument $Log(HHI_{o,z,t})$ with the average of $log(1/N_{o,z',t})$, computed over all commuting zones z' that have no border with z. As mentioned in Section 2.2, when doing so, we cannot include occupation-by-time fixed effects since this is essentially the dimension of the instrument. The results are presented in Table 4. The point estimates are of the same order of magnitude as those obtained with our instrument, although slightly larger than in the corresponding specifications of Tables 2 and 3 - cols (1) to (3). Results in column (3) of Table 4 can be roughly compared to what Marinescu et al. (2019) obtain for new hires in France insofar as they control for firm-level labor productivity and concentration in the product market, along with individual fixed-effects. Their preferred estimate corresponds to an elasticity of -0.09, which suggests that the impact of labor market concentration on stayers' wages is 15 times lower than that estimated on new hires. This is not surprising given the importance of collective wage setting regulations on the French labor market. However, this suggests that for the vast majority of workers, who do not change employer from one year to the other, labor market concentration is unlikely to have a massive depressing effect on wages.

5 Conclusion

Using French administrative data, we have investigated the impact of labor market concentration on stayers' wages. By focusing on stayers and using a new instrument that exploits industry composition within occupations, we are able to control, not only for labor productivity and product market competition, as standard in the literature, but also for match-specific heterogeneity as well as any occupational shock

at the aggregate level. Our findings show that when labor market concentration increases by 10%, stayers' wages decrease by 0.14%. Due to the automatic extension of collective agreements, wage rigidities are strong in France. This probably explains why the effect we find on stayers' wages is small, although negative. As such, it can be considered as a lower bound in an international perspective.

Our results complement Marinescu et al. (2019) who find that labor market concentration reduces the wages of new hires in France with an elasticity of -0.09. The effect that we find on stayers is about 15 times smaller than on new hires. Our findings also complement those of Arnold (2019) who finds that mergers that increase labor market concentration reduce stayers' wages in merged companies. We find that concentration affects stayers' wages in non-merged companies too, although to a limited extent.

To the extent that stayers represent the vast majority of the labor force - more than 80% in France -, our results suggest that in countries where coverage by collective agreements is the norm, labor market concentration is unlikely to put strong downward pressure on wages. This may be one of the reasons why the labor share remained quite stable in France since the end of the 1980s, while it substantially decreased, e.g. in the USA (Cette et al., 2019). Replicating this analysis on US data - or data from other countries - would allow investigating the role of the concentration-wage channel as a determinant of changes in the labor share. This is a promising avenue for further research.

References

W. Abel, S. Tenreyro, and G. Thwaites. Monopsony in the UK. *CEPR Discussion Paper*, No. DP13265, 2018.

D. Arnold. Merger and acquisitions, labor market concentration and worker outcomes. *Mimeo*, Princeton University: https://scholar.princeton.edu/sites/default/files/dharnold/files/jmp.pdf, 2019.

D. H. Autor, D. Dorn, and G. H. Hanson. The China syndrome: Local labor market effects of import competition in the United States. *American Economic Review*, 103(6):2121–68, 2013.

J. Azar, S. Berry, and I. E. Marinescu. Estimating labor market power. *Available at SSRN 3456277*, 2019a.

J. Azar, I. Marinescu, and M. Steinbaum. Measuring labor market power two ways. *AEA Papers and Proceedings*, 109(5):317–21, 2019b.

J. Azar, I. Marinescu, and M. Steinbaum. Labor market concentration. *Journal of Human Resources*, pages 1218–9914R1, 2020.

J. A. Azar, I. Marinescu, M. I. Steinbaum, and B. Taska. Concentration in US labor markets: Evidence from online vacancy data. *NBER Working Paper*, No. 24395, 2018.

J. Babecký, P. Du Caju, T. Kosma, M. Lawless, J. Messina, and T. Rõõm. Downward nominal and real wage rigidity: Survey evidence from European firms. *The Scandinavian Journal of Economics*, 4(112):884–910, 2010.

J. Bai, S. Jayachandran, E. J. Malesky, and B. A. Olken. Firm growth and corruption: Empirical evidence from Vietnam. *The Economic Journal*, 129(618): 651–677, 2017.

E. Benmelech, N. Bergman, and H. Kim. Strong employers and weak employees: How does employer concentration affect wages? *NBER Working Paper*, No. 24307, 2018.

W. M. Boal and M. R. Ransom. Monopsony in the labor market. *Journal of economic literature*, 35(1):86–112, 1997.

G. Cette, L. Koehl, and T. Philippon. Labor shares in some advanced economies. Technical report, National Bureau of Economic Research, 2019.

J. K. Hahn, H. R. Hyatt, H. P. Janicki, and S. R. Tibbets. Job-to-job flows and earnings growth. *AEA Papers and Proceedings*, 107(5):358–63, 2017.

J. K. Hahn, H. R. Hyatt, and H. P. Janicki. Job ladders and growth in earnings, hours, and wages. *Society for Economic Dynamics Meeting Papers*, No. 908, 2018.

J. Hausman, G. Leonard, and J. D. Zona. Competitive analysis with differenciated products. *Annales d'Economie et de Statistique*, pages 159–180, 1994.

G. Jarosch, J. S. Nimczik, and I. Sorkin. Granular search, market structure, and wages. *NBER Working Paper*, No. 26239, 2019.

C. Macaluso, B. Hershbein, and C. Yeh. Concentration in US local labor markets: Evidence from vacancy and employment data. *Society for Economic Dynamics Meeting Papers*, No. 1336, 2019.

I. E. Marinescu, I. Ouss, and L.-D. Pape. Wages, hires, and labor market concentration. *SSRN*, No. 3453855, 2019.

P. Martins. Making their own weather? Estimating employer labour-market power and its wage effects. *QMUL Working Papers*, No. 95, 2018.

OECD. *Employment Outlook*. Organisation for Economic Co-operation and Development, 2010.

Y. Qiu and A. Sojourner. Labor-market concentration and labor compensation. *SSRN*, No. 3312197, 2019.

K. Rinz. Labor market concentration, earnings inequality, and earnings mobility. *Center for Administrative Records Research and Applications Working Paper*, 10, 2018.

Table 1: HHI based on Hirings - OLS

	(1) Log(Wage)	(2) Log(Wage)	(3) Log(Wage)	(4) Log(Wage)	(5) Log(Wage)
$Log(HHI_{o,z,t})$	-.00004	-.00004	-.00000	-.00012	-.00008
	(.00025)	(.00024)	(.00023)	(.00019)	(.00019)
Individual FE	Yes	No	No	No	No
Spell FE	No	Yes	Yes	Yes	Yes
Firm×Year FE	Yes	Yes	Yes	Yes	Yes
Commuting zone×Occupation FE	Yes	Yes	Yes	Yes	Yes
Sector×Commuting zone×Year FE	No	No	Yes	Yes	Yes
Occupation×Year FE	No	No	No	Yes	Yes
Occupation×Sector FE	No	No	No	No	Yes
Age Dummies	Yes	Yes	Yes	Yes	Yes
Observations	10,965,520	10,965,520	10,965,520	10,965,520	10,965,520
R^2	0.928	0.941	0.942	0.943	0.943

Note: *** p<0.01, ** p<0.05, * p<0.1. Standard errors are clustered at the commuting zone level. Age dummies include one dummy variable for each year of age of the individual.

Table 2: HHI based on Hirings - IV

	(1) Log(Wage)	(2) Log(Wage)	(3) Log(Wage)	(4) Log(Wage)	(5) Log(Wage)
$Log(HHI_{o,z,t})$	-.00490**	-.00509*	-.00299	-.01421**	-.01369**
	(.00246)	(.00256)	(.00236)	(.00683)	(.00696)
Individual FE	Yes	No	No	No	No
Spell FE	No	Yes	Yes	Yes	Yes
Firm×Year FE	Yes	Yes	Yes	Yes	Yes
Commuting zone×Occupation FE	Yes	Yes	Yes	Yes	Yes
Sector×Commuting zone×Year FE	No	No	Yes	Yes	Yes
Occupation×Year FE	No	No	No	Yes	Yes
Occupation×Sector FE	No	No	No	No	Yes
Age Dummies	Yes	Yes	Yes	Yes	Yes
F-test on instrument	144.1	135.38	141.27	130.80	131.38
Observations	10,965,520	10,965,520	10,965,520	10,965,520	10,965,520

Note: *** p<0.01, ** p<0.05, * p<0.1. Standard errors are clustered at the commuting zone level. Age dummies include one dummy variable for each year of age of the individual.

Table 3: HHI based on Employment - IV

	(1) Log(Wage)	(2) Log(Wage)	(3) Log(Wage)	(4) Log(Wage)	(5) Log(Wage)
$Log(HHI_{o,z,t})$	-.00475**	-.00497*	-.00282	-.01519**	-.01479*
	(.00242)	(.00252)	(.00230)	(.00759)	(.00775)
Individual FE	Yes	No	No	No	No
Spell FE	No	Yes	Yes	Yes	Yes
Firm×Year FE	Yes	Yes	Yes	Yes	Yes
Commuting zone×Occupation FE	Yes	Yes	Yes	Yes	Yes
Sector×Commuting zone×Year FE	No	No	Yes	Yes	Yes
Occupation×Year FE	No	No	No	Yes	Yes
Occupation×Sector FE	No	No	No	No	Yes
Age Dummies	Yes	Yes	Yes	Yes	Yes
F-test on instrument	281.6	265.5	298.5	56.19	55.27
Observations	10,965,520	10,965,520	10,965,520	10,965,520	10,965,520

Note: *** p<0.01, ** p<0.05, * p<0.1. Standard errors are clustered at the commuting zone level. Age dummies include one dummy variable for each year of age of the individual.

Table 4: HHI based on Employment and Hirings - Alternative IV

	(1) Log(Wage)	(2) Log(Wage)	(3) Log(Wage)	(1) Log(Wage)	(2) Log(Wage)	(3) Log(Wage)
HHI based on	Hirings	Hirings	Hirings	Employment	Employment	Employment
$Log(HHI_{o,z,t})$	-.00631***	-.00663***	-.00474**	-.00655**	-.00691**	-.00483*
	(.00245)	(.00257)	(.00240)	(.00257)	(.00270)	(.00247)
Individual FE	Yes	No	No	Yes	No	No
Spell FE	No	Yes	Yes	No	Yes	Yes
Firm×Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Commuting zone×Occupation FE	Yes	Yes	Yes	Yes	Yes	Yes
Sector×Commuting zone×Year FE	No	No	Yes	No	No	Yes
Age Dummies	Yes	Yes	Yes	Yes	Yes	Yes
F-test on instrument	168.1	157.7	163.5	283.4	269.9	299.0
Observations	10,965,520	10,965,520	10,965,520	10,965,520	10,965,520	10,965,520

Note: *** p<0.01, ** p<0.05, * p<0.1. Standard errors are clustered at the commuting zone level. Age dummies include one dummy variable for each year of age of the individual.

A Appendix (Not for publication *or* for online publication only)

Table A1: Descriptive statistics - Local labor markets

(1) Year	(2) Mean	(3) SD	(4) Median	(5) Obs.	(6) % HHI=1
HHI based on Hirings					
2009	0.1091	0.1635	0.0474	55,818	5.7
2010	0.1082	0.1611	0.048	55,818	4.5
2011	0.1049	0.1541	0.0480	55,818	4.2
2012	0.1093	0.1594	0.0493	55,818	4.8
2013	0.1071	0.1578	0.0487	55,818	4.9
2014	0.1073	0.1567	0.0487	55,818	4.8
2015	0.1064	0.1559	0.0488	55,818	4.9
HHI based on Employment					
2009	0.0844	0.1415	0.0317	55,818	0.4
2010	0.0898	0.1486	0.0334	55,818	0.5
2011	0.0856	0.1407	0.0333	55,818	0.4
2012	0.0888	0.1456	0.0343	55,818	0.5
2013	0.0890	0.1466	0.0350	55,818	0.5
2014	0.0889	0.1453	0.0350	55,818	0.5
2015	0.0887	0.1452	0.0347	55,818	0.5

Note: The mean, standard deviation and median value of HHIs reported in cols (2) to (4) are weighted by employment in each local labor market.

Table A2: Monopsony in Local Labor Markets in France

	Unconcentrated (HHI<0.15)	Mildly-concentrated (0.15<HHI<0.25)	Highly-concentrated (HHI>0.25)
HHI based on Hirings	79.2	9.4	11.4
HHI based on Employment	83.8	7.1	9.1

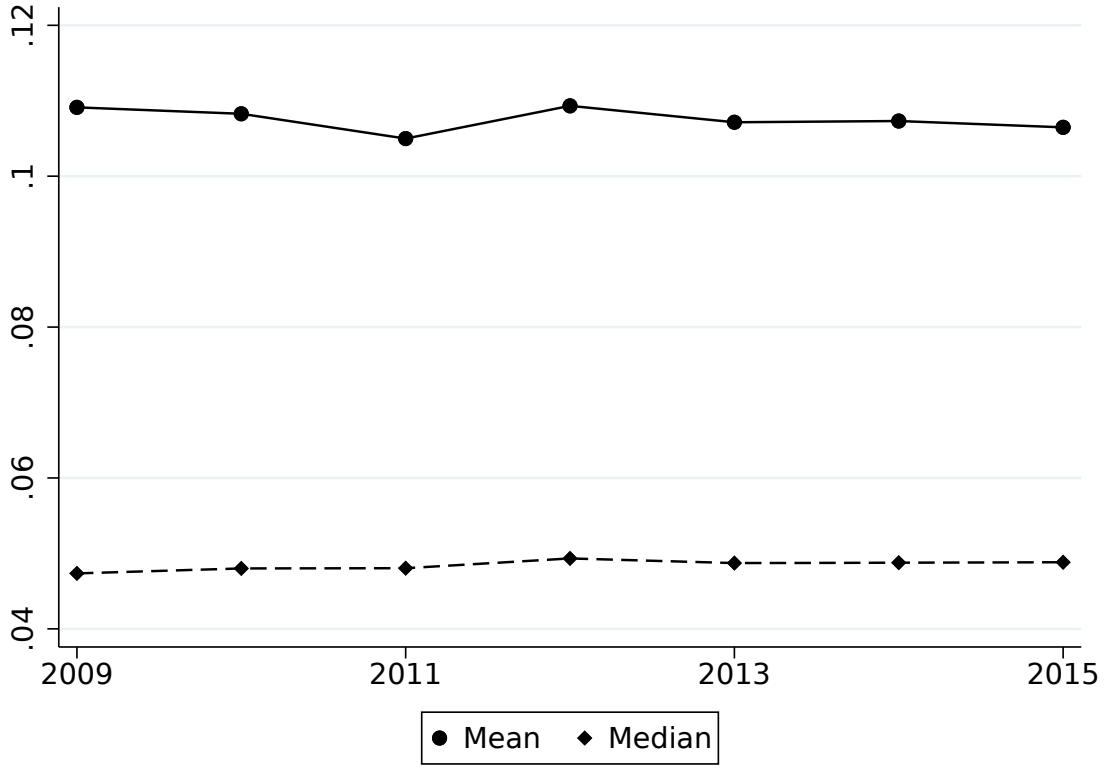
Note: Proportion of local labor markets (weighted by their employment) in the DADS-Postes 2009-2015 according to their level of concentration. Our categorization is similar to the one used by the US Department of Justice.

Table A3: Individual characteristics - Panel DADS, 2009-2015

Mean	Wage	Men	Age	% Stayers	Observations
All	12.13	51.7	39	83.5	13,123,026
Stayers	12.51	51.9	40	-	10,965,520

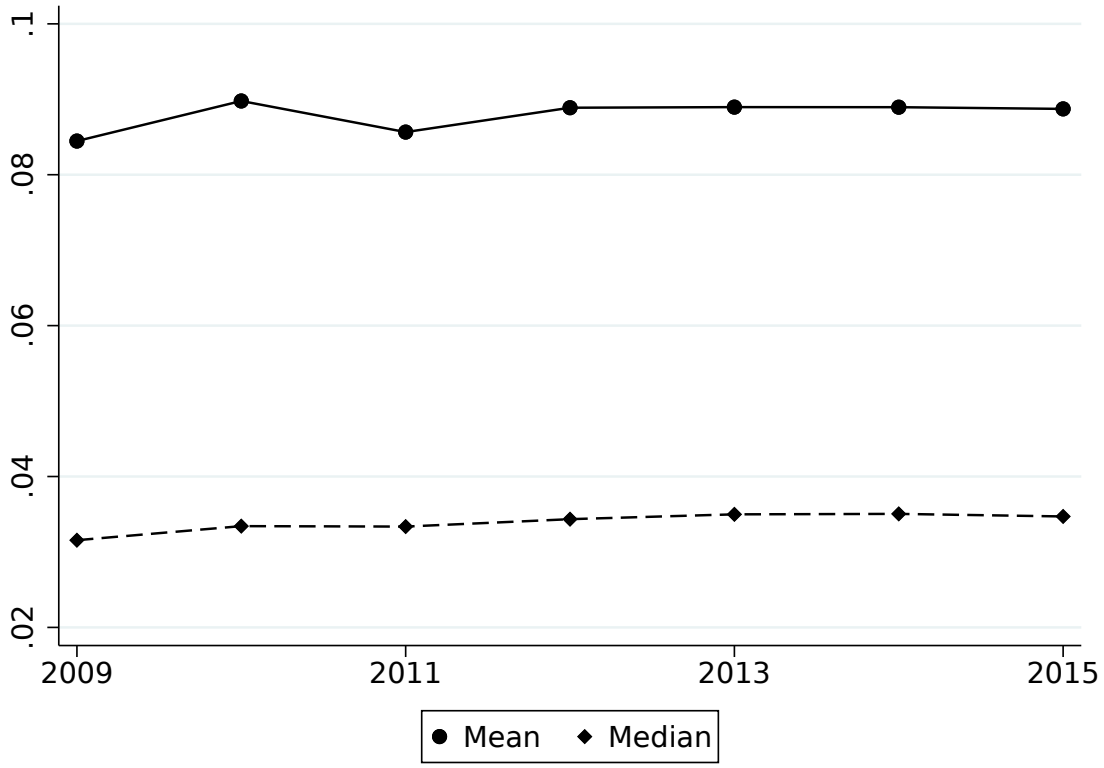
Note: Each observation is an individual-by-firm-by-year triple. A stayer is a triple observed for at least two consecutive years in a given firm.

Figure A1: Change in labor market concentration 2009-2015: Hirings



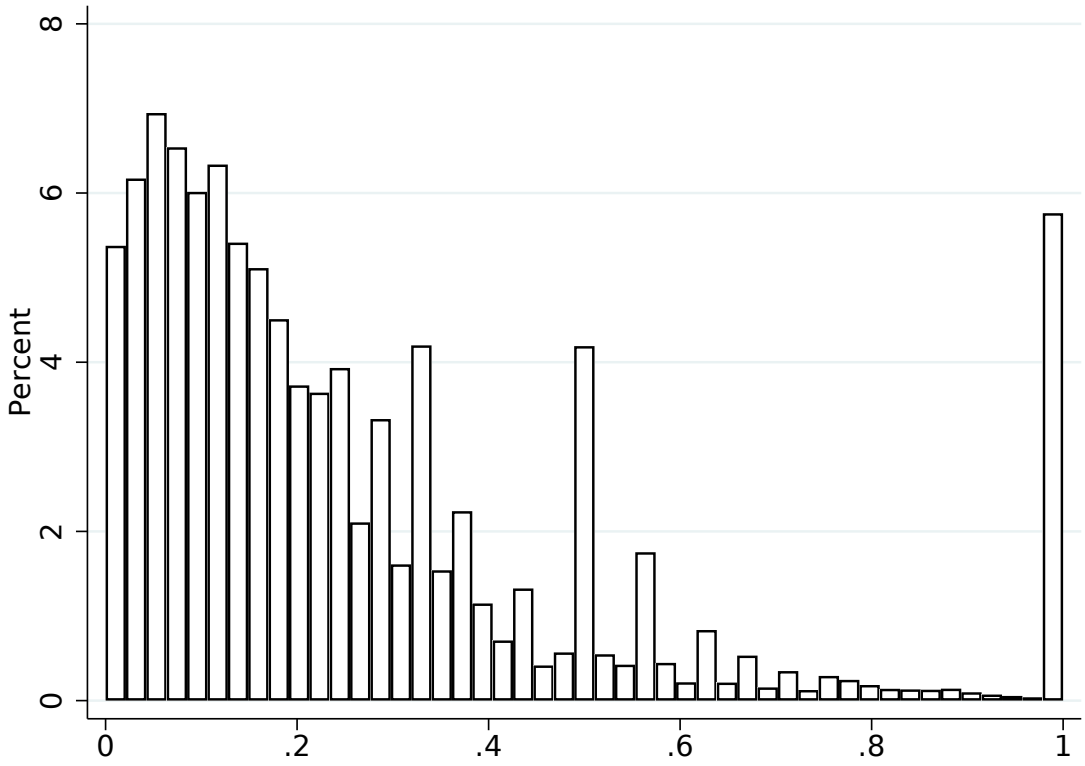
Note: Average HHI (weighted by employment) by year.

Figure A2: Change in labor market concentration 2009-2015: Employment



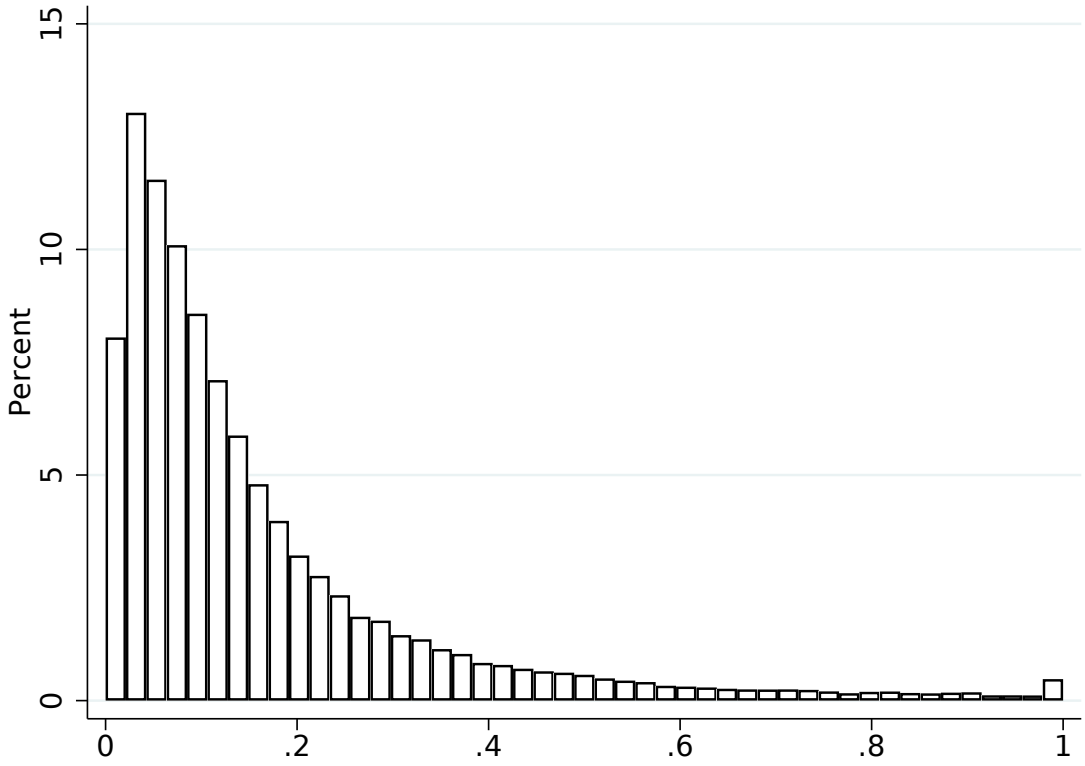
Note: Average HHI (weighted by employment) by year.

Figure A3: Distribution of labor market concentration in 2009: Hirings



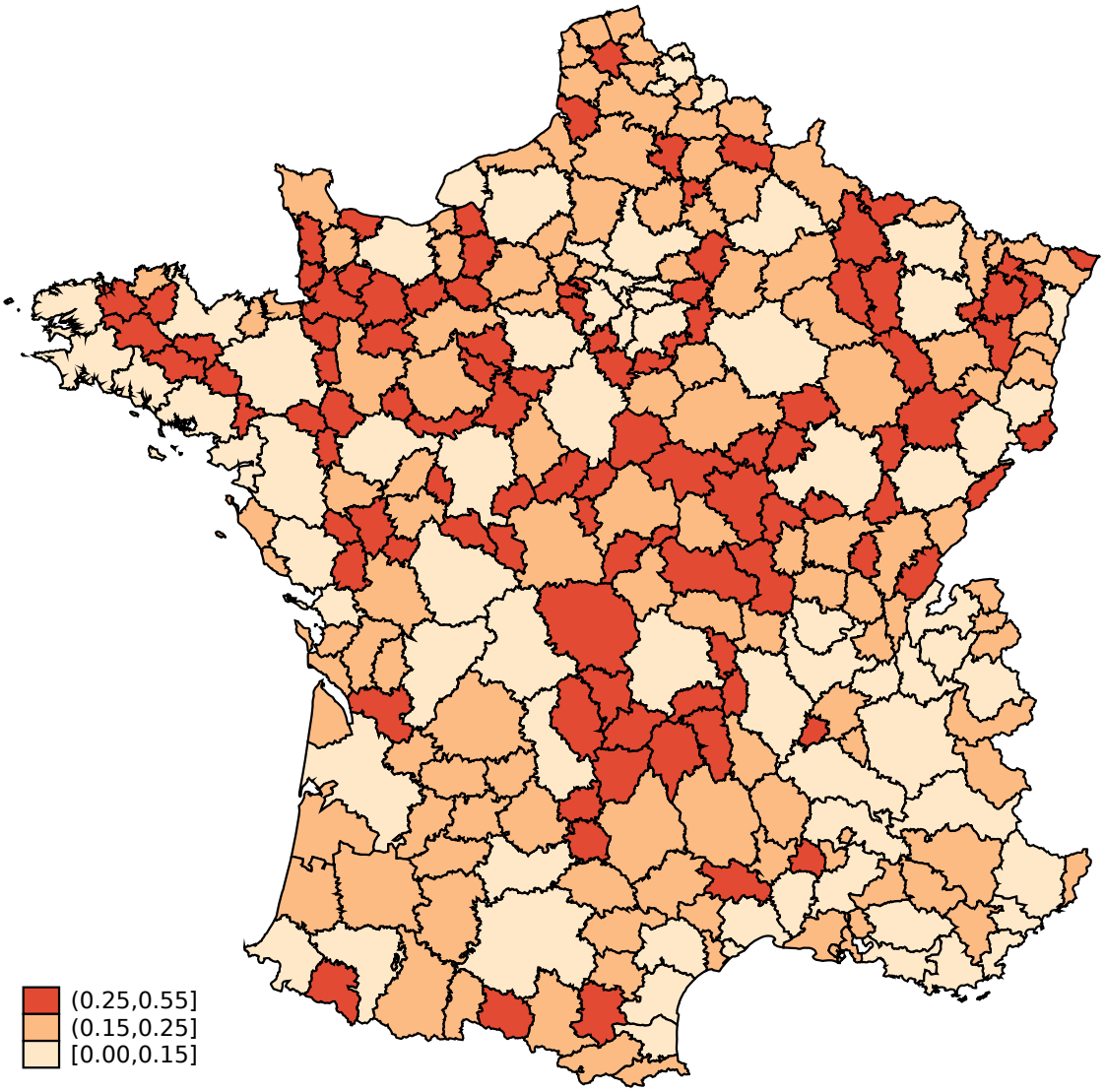
Note: Each observation is a local labor market in 2009. The figure plots the unweighted distribution of hiring-based HHI across these labor markets.

Figure A4: Distribution of labor market concentration in 2009: Employment



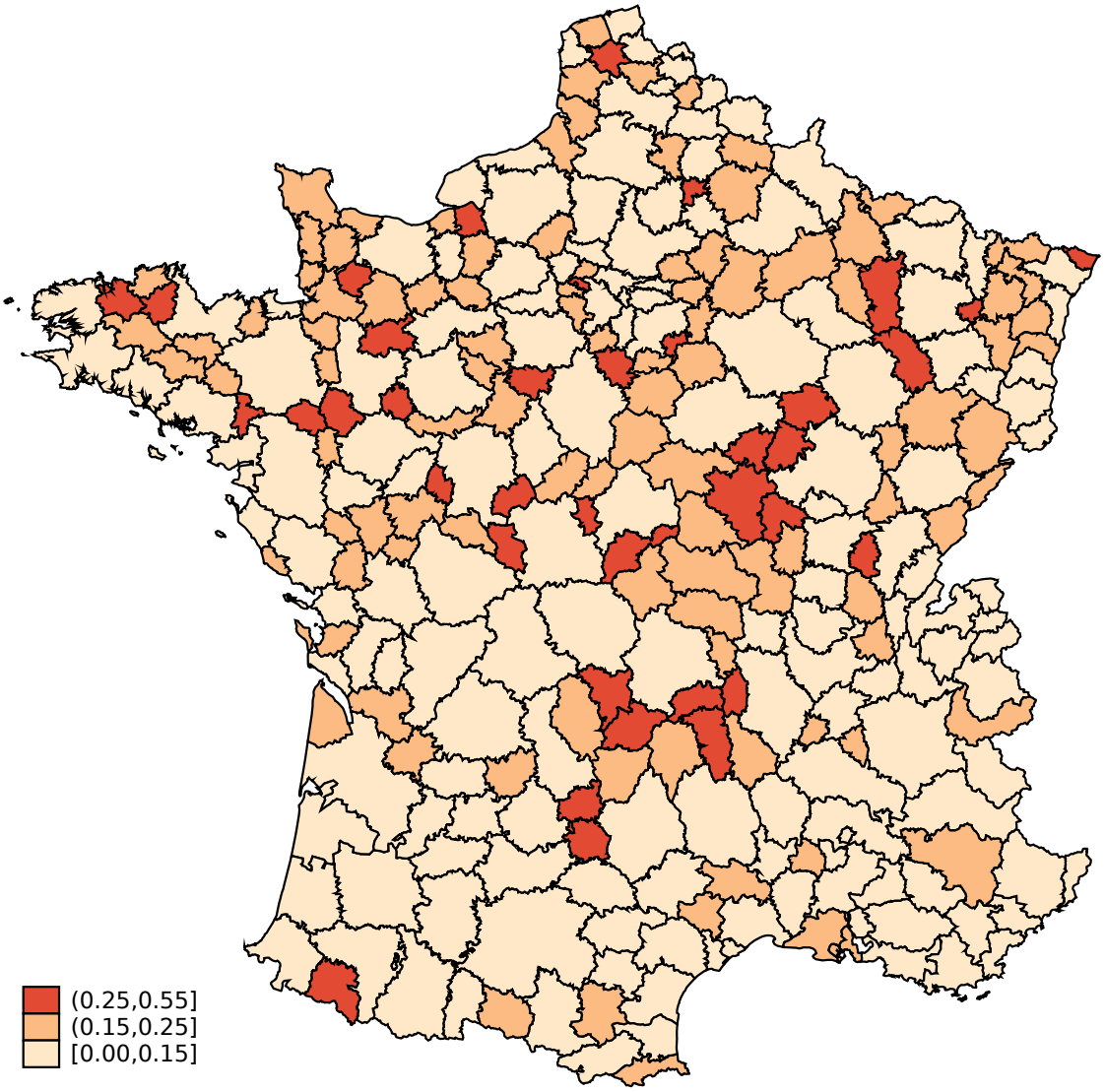
Note: Each observation is a local labor market in 2009. The figure plots the unweighted distribution of employment-based HHI across these labor markets.

Figure A5: Labor market concentration in French commuting zones in 2009: Hirings



Note: Average HHI (weighted by employment) by commuting zone.

Figure A6: Labor market concentration in French commuting zones in 2009: Employment



Note: Average HHI (weighted by employment) by commuting zone.