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Employer concentration and wages for specialized workers

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Abstract

This paper studies how wages respond to a sudden change in employer concentration. It exploits a reform that deregulated the Swedish pharmacy market, which until 2009 was a monopoly. The reform involved a substantial increase in the number of employers on the pharmacy labor market. However, the change in employer concentration was not geographically uniform: certain areas experienced large changes while others were largely unaffected. Exploiting this geographical variation, elasticities of wages with respect to labor market concentration are estimated to be between -0.02 and -0.05. The empirical approach relies only on the variation in concentration controlled by the policymaker to remedy the concern that actual labor market concentration is endogenous. The positive wage effects from reduced labor market concentration are found to be most prevalent for stayers, rather than new hires, as well as those with more industry experience and longer tenure. Overall, the paper adds to a growing literature that finds that market concentration matters for workers' wages, in a context where labor is highly industry-specific.

Keywords: Wages; Competition; Market concentration

JEL Codes: J31; J42; J45

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1 Introduction

Against a background of stagnant wage growth, rising inequality and a falling labor share, a recent literature has revived interest in the question of whether employers have market power.[1] A canonical source of labor market power is labor market concentration.[2] Concentration relates to the existence of only a small number of employers on a particular labor market. High concentration can give employers the ability to depress wages below competitive levels, and a negative relationship between labor market concentration and wages has been documented in recent literature (see Azar, Marinescu, and Steinbaum 2020 and Benmelech, Bergman, and Kim 2018). At the same time, the relation between wages and labor market concentration is oftentimes difficult to shed light on, not least because changes to concentration over time tend to be small and arise from a multitude of decisions taken by employers and employees.

In this paper, I focus on a particular market – pharmacies – and use a major policy reform to deduce quasi-experimental evidence on how increasing the number of employers affects workers' wages. Prior to 2009, only the state-run monopoly *Apoteket* could retail pharmaceuticals in Sweden. In 2009, entry barriers were removed and private firms could enter. Two thirds of *Apoteket*'s pharmacies were privatized as part of the deregulation. Since pharmacists have highly industry-specific skills, the deregulation causes changes to the number of employers in an appropriately defined labor market. I exploit this setting to study how wages respond when labor market concentration falls suddenly. I rely only on the variation in concentration induced by the privatization of pre-existing pharmacies in 2009 to remedy the concern that actual labor market concentration is endogenous. This variation in concentration is controlled by the policymaker. I find that wages respond negatively to local labor market concentration, with estimated elasticities ranging from -0.02 to -0.05.

The setting in this paper is unique and interesting for a number of reasons. First, the deregulation resulted in substantial variation in employer concentration caused by a policy-decision. Second, pharmacists have industry-specific skills such that the deregulation meaningfully changes employer concentration for this group of workers. Third, product prices are largely regulated and set by the state both before and after the reform, which make them independent of product and labor market concentration. The setting allows me to isolate the effects of changes to labor market concentration without confounding it with those related to product market concentration. Fourth, wage-setting in

[1] See Berger, Herkenhoff, and Mongey (2019); Jarosch, Nimczik, and Sorkin (2019); Lamadon, Mogstad, and Setzler (2019); Card, Cardoso, Heining, and Kline (2018); Autor, Dorn, Katz, Patterson, and Van Reenen (2020); De Loecker, Eeckhout, and Unger (2020); and Barkai (2020), among others.

[2] There are other sources of labor market power that arise with a continuum of firms, see Section 2.1. An important distinction is also between market power on the product side and on the labor side.

the industry is decentralized and wages are set in individual negotiations both before and after the deregulation. Fifth, the study uses high quality matched employee-employer data. I can follow workers over time, and the effect of changes to concentration on wages can be estimated holding worker skill constant.

The paper consists of four main parts. In the first part of the paper, I present a simple model that relates wages to labor market concentration. The overall effect of the *deregulation* on wages is ambiguous. When isolating the effects related to reduced *labor market concentration*, however, that model predicts that wages increase post-deregulation. The framework highlights that the overall deregulation effect depends on two channels: changes to labor market power and changes to product market power. A decrease in product market power is likely to decrease wages. This can, for example, arise if profits fall under rent-sharing. The increase in the number of employers can have a counteracting, and positive, effect on wages as monopsony power is reduced. Originally coined by Robinson (1933), monopsony in the strictest sense is a situation with only one buyer. In labor markets, it has more generally come to refer to a situation where individual firms face upward-sloping labor supply (Boal and Ransom 1997).[3] In contrast to competitive labor markets, firms are able to extract rents by setting wages below the marginal revenue product of labor.

The second part of the paper uses rich employee-employer data between 2004 and 2016 to characterize aggregate changes in the pharmacy industry, which is female dominated and highly skilled. The reform led to a large aggregate reduction in labor market concentration, measured by the Herfindahl-Hirschman Index (HHI) in employment shares, from 1 to 0.25. Wages for pharmacy employees increased on aggregate by 2 to 4 percentage points upon deregulation, relative to comparable workers in other industries. The number of pharmacies increased by around 50% and employment by 10%. Overall the market became more fragmented as employees split into more pharmacies operated by more firms.

The third and fourth parts of the paper include the main results. The third part estimates how labor market concentration affects wages. The change in labor market concentration induced by the deregulation differs across local markets. Making use of this geographical variation, elasticities of wages with respect to labor market concentration (or precisely, HHI) are estimated to be between -0.02 and -0.05. These results are consistent with previous studies (see for example Rinz 2020 or Hershbein, Macaluso, and Yeh 2019).

[3]For a recent review of monopsony, see Manning (2020). Given the theoretical link between labor supply elasticities to the firm and labor market power, several papers have estimated the elasticity of labor supply to the firm in particular markets and found these to be far from perfectly elastic (see Sullivan (1989), Staiger, Spetz, and Phibbs (2010) and Matsudaira (2014) for nurses; Falch (2010) and Ransom and Sims (2010) for teachers; and Dube, Jacobs, Naidu, and Suri (2018) for the gig economy). This has been taken as support of that at least specific labor markets are imperfectly competitive, while being agnostic as to the actual source of labor market power.

The effects materialize within two years of the deregulation and are relatively stable until the last period of observation in 2016.

The causal interpretation is supported by a battery of checks. Log wages evolve in parallel in local markets prior to deregulation and effects are not found in a related but unaffected industry. Only variation in labor market concentration from the sale of pre-existing pharmacies in 2009, the year the market was deregulated, is used to remedy the concern that actual labor market concentration is endogenous. This depends on the privatization of pre-existing pharmacies, a process controlled by the policymaker, and is neither affected by firms' decisions to open up new pharmacies nor to mobility decisions that workers make post-deregulation.

The fourth part of the paper studies which employees in the pharmacy industry benefit most from reduced labor market concentration. Stayers rather than new hires benefit with higher wages. That is, conditional on joining a new employer, the results suggest that there is no additional return to making the transition in a labor market with relatively low labor market concentration. This result should be interpreted with caution, however, as the likelihood of moving is itself affected by the deregulation. The positive wage effects are also estimated to be prevalent primarily for those with more industry experience and longer tenure. Individual characteristics do not appear to matter for the wage returns. Instead, similar positive wage returns from reduced labor market concentration are estimated for pharmacy employees of different age, country of birth, gender and education.

The paper is related to three main strands of literature. First, it contributes to the literature that studies the effect of labor market concentration on wages. Whether there exists such a relationship is an old question, reviewed in Boal and Ransom (1997). More generally, this concerns imperfect competition in labor markets under the assumption that labor market concentration captures labor market power (see Manning (2011) for an overview). The question of whether labor market concentration affects wages, which according to Manning (2020) was originally studied by Bunting (1962), was revived recently by Azar, Marinescu, and Steinbaum (2020) and Benmelech, Bergman, and Kim (2018), who estimate a negative relationship between these variables in a U.S. context.[4] Empirically, these papers tend to exploit broad changes in concentration at the region by industry or occupation level. A key concern is that many factors affect both market concentration and wages (Berry, Gaynor, and Scott Morton 2019), and it is hard in data to identify exogenous shifts in concentration. This paper contributes to the literature by focusing on a particular industry where there is a policy change that affects the number

[4]This negative relationship has been confirmed in many other studies including Hershbein, Macaluso, and Yeh (2019), Qiu and Sojourner (2019), Lipsius (2018), Rinz (2020) and Schubert, Stansbury, and Taska (2020) using U.S. data; in Martins (2018) using data from Portugal; and in Marinescu, Ouss, and Pape (2020) using data for France.

of employers.

A second yet scarce strand of literature uses variation from mergers to study how labor market concentration affects workers.[5] Arnold (2020) exploits merger-induced changes in concentration in the U.S. to find negative effects of increased labor market concentration on earnings when the change in concentration is large. A related approach is taken by Prager and Schmitt (2021) who focus on hospital consolidation only. They find a negative effect on wages when the change in concentration is large and worker skills are industry-specific.[6] While compelling, these papers rely on variation from firms that choose whether or not to merge. If firms merge for labor cost-saving reasons, then that could result in a spurious correlation between labor market concentration and wages. I take a complementary approach by focusing on changes in concentration induced by a policy change rather than a firm decision.

Third, the paper contributes to the literature on the wage effects of privatization and deregulation. In contrast to these papers, I isolate the wage effect of the regulatory change to that associated with changes in labor market concentration. The literature on how workers are affected by privatization is scarce. In two recent papers, Olsson and Tåg (2018) find increased unemployment incidence and duration while annual labor income and labor force participation remain unchanged for privatization events in Sweden, while Arnold (2019) finds that privatization lowered incumbents' wages substantially in Brazil. Regarding deregulation, this is often analyzed as a shock to profitability in a specific industry, which under rent-sharing would put downward pressure on wages. Peoples (1998) provides on overview of the wage effects of reducing entry barriers to specific industries. The literature predominantly finds negative industry wage effects of deregulation, including in the U.S. airline (Card 1998, Hirsch and Macpherson 2000), trucking (Rose 1987, Hirsch 1988) and banking industries (Black and Strahan 2001).[7] This literature is often based on cross-sectional data and is, unlike this paper, unable to control for unobserved worker characteristics. An exception is Lergetporer, Ruhose, and Simon (2018), who find negative wage effects for incumbent workers after lifting entry barriers in the German crafts sector.

Overall, the results in this paper underpin that an increase in the number of employers in an industry with specialized labor increases wages. The paper lends support to the literature that finds that labor market concentration can matter for workers' wages. The

[5] In a related paper Hensvik (2012) studies the relation between school competition and teacher wages in Sweden. Hensvik (ibid.) focuses on how public-school hiring and wages are affected by private entry, finding that wages respond positively to the increased competition.

[6] Currie, Farsi, and Macleod (2005) consider labor market effects of mergers for hospitals without linking this explicitly to changes in market concentration. They find increases in nurse effort but no wage effects when studying hospital consolidation.

[7] The deregulation literature also finds that reducing entry barriers leads to employment growth. Bertrand and Kramarz (2002) study entry deterrence in the French retail industry.

rest of the paper proceeds as follows. A theoretical framework is included in Section 2. Key definitions and data are described in Section 3. Institutional details and descriptive patterns are provided in Section 4. Section 5 outlines the empirical strategy, Section 6 provides estimates of the wage effects of reduced labor market concentration and Section 7 considers which employees benefit most from this reduction. Finally, Section 8 concludes.

2 Theoretical framework

2.1 Sources of labor market power

In imperfectly competitive labor markets, monopsony powers can stem from many sources. The quasi-experiment studied in this paper naturally lends itself to studying the effect of labor market concentration on wages. Market concentration relates to the existence of only a small number of potential employers; that is, labor markets are thin. Due to regulatory barriers, only one firm was allowed to operate in the pharmacy industry prior to the reform. Building on a tradition in industrial organization and in antitrust policy, the intuition is that firms are interdependent and take actions of other firms operating in the same labor market into account when making employment decisions. In such a world, it may be profitable for firms to hire fewer workers and thereby set lower wages than in a perfectly competitive world. This type of argument is emphasized by classic models such as Cournot oligopsony and in empirical work by Azar, Marinescu, and Steinbaum (2020) and Benmelech, Bergman, and Kim (2018), among others. Recent theoretical work by Berger, Herkenhoff, and Mongey (2019) provides a micro-foundation to the relationship between market power and market structure. Their model allows for a large but finite number of employers, and market power arises from the ability of firms to exploit the market-share dependent upward-sloping labor supply curves to the firm.[8]

While I focus on a particular source of labor market power – labor market concentration – there are sources of monopsony power that arise even with a continuum of firms. These sources include differentiation and preference heterogeneity (see Bhaskar, Manning, and To 2002 and Card, Cardoso, Heining, and Kline 2018), as well as search frictions (Burdett and Mortensen 1998). Search frictions are key to the seminal dynamic monopsony models by Manning (2003).

[8]Jarosch, Nimczik, and Sorkin (2019) develop a model where a different mechanism gives rise to a relation between market concentration and wages. Market power stems from employer size where each employer recognizes it can exploit its power by eliminating its own vacancies from the worker's outside option and thereby not compete with itself. In the model, employment is not directly affected by market structure. It therefore deviates from the Robinson-style models.

2.2 A Cournot oligopsony model

In this section, I outline a simple static Cournot oligopsony model that relates wages and labor market concentration. The framework follows Arnold (2020), who exploits mergers and acquisitions to study how labor market concentration affects workers.[9] There are F firms in a market m, indexed $f = 1, \ldots, F$ and total employment on the market is $L_m = \sum_{f=1}^{F} l_f$, where l_f is firm f's level of employment. For simplicity, labor is assumed to be the only input into production. Each firm maximizes its objective function by choosing its employment, l_f, taking the labor demand of other firms on the same market as given. The market wage $w_m(L_m)$ depends on total employment in the market. This is a posted wage; there is no wage bargaining in this model.

$$\max_{l_f} R_f(l_f) - w_m(L_m) l_f \tag{1}$$

$R_f(l_f)$ is the firm's revenue function. This will depend on product market factors, like price and quantity of goods sold, as well as productivity parameters. More generally, $R_f(l_f)$ can be interpreted as any concave function that is increasing in l_f. This covers the case of the public monopsonist that may have additional objectives beyond maximizing revenue minus cost. This leads to the following first-order condition:

$$\underbrace{\frac{\partial R_f(l_f)}{\partial l_f}}_{\Omega_f \equiv MRPL_f} - \underbrace{\left[w_m(L_m) + \frac{\partial w_m(L_m)}{\partial l_f} l_f \right]}_{MLC_f} = 0 \tag{2}$$

Notice that, in the absence of labor market power, the firm's labor decision would not affect wages and wages would be set to equal the marginal revenue product of labor ($MRPL$). The first order conditions can be re-written as follows:

$$\Omega_f - w_m(L_m)\left[1 + \frac{s_f}{\varepsilon_m}\right] = 0 \tag{3}$$

where $s_f = \frac{l_f}{L_m}$ is firm f's employment share in market m and ε_m is the market-level labor supply elasticity, $\varepsilon_m = \frac{\partial L_m}{\partial w_m(L_m)} \frac{w_m(L_m)}{L_m}$. Multiplying each side by s_f and then summing the first order conditions across all firms, we can rearrange to find an expression for market wages:

$$w_m = \underbrace{\left[\frac{\varepsilon_m}{HHI_m + \varepsilon_m}\right]}_{\sigma_m} \Omega_m \tag{4}$$

where $HHI_m = \sum_f (s_f)^2$ is the Herfindahl-Hirschman Index and $\Omega_m = \sum_f s_f \Omega_f$ is the

[9] Arnold (2020) decomposes the effects of mergers into three components: monopsony effects, product market power effects and productivity effects. It draws on classic Cournot oligopsony results, among others outlined in Boal and Ransom (1997) and Naidu and Posner (2019).

employment-weighted average of the marginal revenue product of labor. HHI_m can take values in the interval $(0, 1]$ where values approaching 0 represent perfect competition and 1 represents only one employer in the market. Unlike in the competitive model, workers will only get a fraction (denoted σ_m) of the average marginal revenue product. The model implies that higher concentration is negatively associated with wages, holding all else constant. Moreover, if market power only stems from market concentration, then the wage approaches the competitive wage as the number of firms increases.

2.2.1 Channels through which deregulation affects wages

In the simple model outlined above, the deregulation may impact wages through changes in labor market power (operating through σ_m) and through changes in productivity, product market power or objectives as ownership shifts from public to private (operating through Ω_m). Taking logs, the average treatment effect of the deregulation on log wages $\widetilde{w}_m$ can be written as follows. *post* refers to post-deregulation and *pre* to pre-deregulation:

$$\mathbb{E}[\widetilde{w}_{m,post} - \widetilde{w}_{m,pre}] = \mathbb{E}[\widetilde{\sigma}_{m,post} - \widetilde{\sigma}_{m,pre}] + \mathbb{E}[\widetilde{\Omega}_{m,post} - \widetilde{\Omega}_{m,pre}] \tag{5}$$

While decreases to labor market concentration will increase wages, decreases to Ω_m are, on the other hand, likely to put downward pressure on wages. To the extent that these two channels are correlated, I am likely to estimate lower bound effects of reduced labor market concentration on wages.

Focusing on the Ω_m component, increased product market competition will in general put downward pressure on prices, and under rent-sharing, also on wages.[10] While the price of prescription drugs, the dominant product category, is regulated, this is not the case for non-prescription drugs and retail items which represents around 25% of revenue. Indeed, earlier deregulation studies (Black and Strahan 2001) have used deregulation as a shock to profitability and find support for the rent-sharing channel as wages fall post-deregulation. Peoples (1998) highlights how labor earnings may fall after deregulation as the bargaining power of workers falls. Recalling that the reform considered in this paper also involves privatization and that the revenue function can be interpreted as any function that is increasing in l_f, Arnold (2019) finds that state-owned enterprises pay significant wage premiums over private firms, also suggesting that changes to Ω_m could put downward pressure on wages. This is consistent with state-owned enterprises having wider objectives than only maximizing profits (Haskel and Szymanski 1993). Indeed, trade unions often fear privatization will lower wages and much academic literature has

[10]The model in Section 2.2 implies that, if decreases to product market power increase employment, then wages also increase. This is not the case with wage bargaining, where lower revenue per worker results in lower wages.

assumed privatization has negative effects on wages.[11]

To isolate labor market power effects (captured by σ_m) from productivity or product market aspects (captured by Ω_m), a sufficient condition is that changes to Ω_m are independent of changes to σ_m, conditional on included controls. For example, if changes to labor market power are correlated with changes to product market power, and if higher product market power has a positive effect on wages, then I am likely to underestimate the effect of labor market concentration on wages. Similarly, if public monopsonists overpay compared to private monopsonists, then the state-to-market quasi-experiment that this paper relies on will also underestimate the effects of reduced labor market concentration. In support of this assumption, firstly note that decisions about product pricing and campaigns are normally taken nationally and the product ranges at pharmacies are relatively homogeneous (Swedish Competition Authority 2013). The empirical strategy only exploits within-industry changes. Moreover, while pharmacies do sell products with unregulated prices, around three quarters of pharmacies' revenues are from products with nationally regulated prices. That product prices are regulated both before and after the reform make them independent of product and labor market concentration.

3 Definitions, data and sample

3.1 Definitions

In order to calculate concentration measures, it is necessary to define what a market is. The definition of the labor market should capture the set of potential employers for a worker. Because workers are tied to their workplace, labor markets tend to be local. In this paper, a local labor market (LLM) m is defined by the interaction of the industry for dispensing chemists and commuting zones (CZ). The industry for dispensing chemists is identified by workplace industry codes included in the data.[12] Commuting zones encompass all industries in a geographic area and are taken from Statistics Sweden who define CZs using commuting patterns.

In line with the theoretical framework, labor market concentration is measured using the Herfindahl–Hirschman Index (HHI) in employment shares. This measure captures concentration among pharmacy firms. For all workplaces that are operating as dispensing

[11] Even so, there is no general theory of how deregulation and privatization will affect wages, and certain product market channels could push wages up. Many papers find that efficiency and profits increase once state owned enterprises are privatized (see Megginson and Netter (2001) for a review), which could put upward pressure on wages if private firms share rents at least as much as public firms do. Earle and Shpak (2019) summarize why wages may rise or fall as a result of privatization.

[12] Industry (SNI) code 52.310 is used until 2007 and 47.730 from 2008 onward. There is a one-to-one mapping between these codes. To remedy potential miss-classifications of workplaces that may occur especially around the time of the deregulation, I (iteratively) assume that a workplace is a dispensing chemist if it was classified as a dispensing chemist in the previous year.

chemists, a unique workplace identifier is assumed to be a pharmacy. A firm is defined using firm identifiers provided in the data as a collection of pharmacies. HHI is calculated separately by year t and LLM m as the sum of squared employment shares s_f across firms in each local pharmacy market:

$$HHI_{mt} = \sum_{f=1}^{F} s_{fmt}^2 \tag{6}$$

An HHI approaching 0 corresponds to perfect competition while an HHI equal to 1 corresponds to a single employer.[13] A higher value means higher concentration and thus lower competition. HHI is a canonical measure of labor market concentration, used among others in Benmelech, Bergman, and Kim (2018), Lipsius (2018) and Rinz (2020), who calculate HHI using employment shares, and in Azar, Marinescu, and Steinbaum (2020), who calculate HHI using vacancy shares. In addition, HHI is widely used in merger control (see the U.S. DoJ and FTC's as well as the European Commission's horizontal merger guidelines) as a measure of market power.

A relevant question is whether industries capture reasonable employment opportunities for workers, and whether workers travel across commuting zone borders to work. In support of that CZs are a reasonable geographical denomination, 92.5% of employees in the pharmacy industry work and reside in the same CZ. A commonly used alternative to industry is occupation.[14] In this paper industry is used instead of occupation both because the deregulation took place at the industry level, and because the data lacks complete and consistent occupation information. Out of the employees that worked at a pharmacy in 2004, only one third have worked in a different five-digit industry after 2004. To the extent that industry is too narrow to represent the employment opportunities for workers, HHI in both the pre-deregulation and post-deregulation period will be too high; the identification relies on differences in HHI over time.

3.2 Data

The data is drawn from a panel of annual register data collected by Statistics Sweden. I have access to full-population data, meaning that I can identify educated pharmacists

[13]The measure assumes that pharmacies operated by the same firm in the same LLM do not compete for workers. In support of this, the estimated returns to being a new hire in a pharmacy in the same LLM between 2004 and 2008 are marginal. For the full sample, the wage returns to joining are not statistically different from 0. Restricting to those who join from another pharmacy, the returns are estimated to be 1.46%. Restricting further to those who move from another pharmacy but are not managers, the return is 1.01%.

[14]Previous work has defined markets both using geography–industry (Berger, Herkenhoff, and Mongey 2019, Lipsius 2018 Benmelech, Bergman, and Kim 2018 and Rinz 2020) and geography–occupation (Azar, Marinescu, and Steinbaum 2020, Azar, Marinescu, Steinbaum, and Taska 2020 and Qiu and Sojourner 2019).

and non-pharmacists working both inside and outside the pharmacy industry. The main dataset is a matched employer–employee register (*RAMS*) that includes firm, workplace and person identifiers as well as gross labor earnings and the months worked at each workplace for all gainfully employed individuals. For those employees with more than one workplace, only one workplace per employee and year is kept, defined as the workplace in November where the individual has the highest annual earnings.[15] Using year, firm and workplace identifiers, firm and workplace characteristics such as workplace industry and ownership structure are matched in. Wages and occupations are taken from structural earnings statistics. Using person and year identifiers, the employer-employee data is linked to demographic registers (*Louise* and *Födelseland*) that include variables such as year and country of birth, gender, education level and field of educational specialization. Commuting zones (*lokala arbetsmarknader*) are taken from Statistics Sweden and are matched in based on the municipality of the workplace.[16] Financial data, only available until 2015, is also matched in at the firm and year level.

As outlined in Section 3.1, the pharmacy market is delineated from the full-population data using the workplace industry code for dispensing chemists. Pharmacists are identified using information on education. The demographic registers not only has information on the level of education that individuals hold but also detailed information on which field the education is in, based on the Swedish educational nomenclature (*SUN*). Consequently pharmacists can be identified using the individual's educational specialization in pharmacy together with the level of education the individual has.[17]

The main wage measure is monthly full-time adjusted wages in Swedish crowns (SEK), measured between September and November each year. Wages are available for all individuals in the public (non-market) sector and for an annual random stratified sample in the private (market) sector.[18] The main analysis uses log full-time equivalent wages as the outcome. These data are accurate but, due to the sampling design, incomplete (see Table 1 for summary statistics). To ensure the results are not driven by extreme outliers, the data is trimmed so that wages 50% below the 1st percentile or 50% above the 99th percentile of the monthly full-time adjusted national wage distribution every year are excluded from the regressions. Sensitivity analyses, included in the appendix, instead use gross monthly earnings (defined as gross annual labor earnings divided by the number

[15]Using November is in line with Sweden's official statistics, in turn based on ILO's methodology. Before identifying the main workplace, I drop employees that cannot be linked to a physical workplace and therefore obtain a false workplace code. I also exclude self-employed, who have a different employment relationship to employees.

[16]The boundaries of commuting zones are revised periodically. To maintain a consistent measure of commuting zones throughout the time period considered, commuting zones from 2013 are used.

[17]Pharmacists are defined using education rather than occupation to have complete coverage in the data. Occupation in the registry data is available for around 50% of workers.

[18]Approximately 50% of private sector employees are included in the sample. The sample is stratified by industry and firm size, with an oversampling of larger firms.

of months worked at the primary employer) from the matched employer-employee data. The earnings measure has complete coverage in the data.[19]

3.3 Main sample

The main sample consists of all employees (pharmacists, pharmacy technicians and non-pharmacists) who are employed at a pharmacy as their main place of work. The sample period is restricted to the years 2004 to 2016, which means pre-trends can be analyzed alongside post-deregulation effects. As the employment relationship is different for self-employed, self-employed are excluded from the sample.[20]

Table 1 includes summary statistics for the full sample period (2004 to 2016) as well as pre-deregulation and post-deregulation. The pharmacy industry is highly skilled and female dominated: 88% are women and 67% have at least post-secondary education. Half the share of employees are educated pharmacists. The statistics also highlight that there have been compositional changes over time that will be important to control for in the empirical analysis. Workers are on average younger and have slightly less tenure and experience in the post-period. The share with post-secondary education has increased over time and the share of foreign born, defined as being born outside of Sweden, has increased sharply from 12% pre-deregulation to 23% post-deregulation.

[19]There are two main differences between wages and earnings: (i) due to the sampling design of the official statistics, wages are predominantly available for larger firms, and (ii) earnings are not full-time equivalent. Like for wages, the earnings measure is restricted to drop earnings more than 50% below the 1st percentile or 50% above the 99th percentile of the monthly national wage distribution every year.

[20]Self-employment among pharmacists today is low. Goldin and Katz (2016) find that the fraction of pharmacists in the U.S. who are self-employed has decreased from 40% in 1966 to just under 5% in 2011. Figure A.1 shows the share of self-employment among all educated pharmacists in Sweden between 2004 and 2016. Under 4% of pharmacists are self-employed. There is an increase in self-employment in conjunction with the deregulation. This is not explored further in this paper.

Table 1: Summary statistics, pharmacy industry

	2004–2016		2004–2008		2009–2012		2013–2016	
	Mean	sd	Mean	sd	Mean	sd	Mean	sd
Female	0.88	(0.32)	0.91	(0.29)	0.88	(0.32)	0.86	(0.35)
Age (years)	43.79	(13.96)	45.60	(13.20)	43.32	(14.32)	42.13	(14.23)
Age < 30	0.20	(0.40)	0.16	(0.37)	0.22	(0.41)	0.23	(0.42)
Age ≥ 50	0.40	(0.49)	0.47	(0.50)	0.38	(0.49)	0.33	(0.47)
Foreign born	0.17	(0.38)	0.12	(0.32)	0.17	(0.37)	0.23	(0.42)
Post-secondary	0.67	(0.47)	0.64	(0.48)	0.66	(0.47)	0.71	(0.45)
Pharmacist	0.49	(0.50)	0.51	(0.50)	0.48	(0.50)	0.47	(0.50)
Tenured	0.43	(0.50)	0.50	(0.50)	0.41	(0.49)	0.37	(0.48)
Industry experience	0.73	(0.44)	0.77	(0.42)	0.73	(0.44)	0.69	(0.46)
Non-missing wage	0.69		0.84		0.54		0.67	
Monthly wage (2004 SEK)	26,193	(8,166)	24,246	(7,404)	26,151	(8,096)	29,083	(8,429)
Monthly earnings (2004 SEK)	23,806	(10,353)	22,097	(9,198)	23,707	(9,955)	26,059	(11,614)
Number of employee-year obs.	159,863		59,392		49,277		51,194	

Note: The table shows summary statistics for all employees in the pharmacy industry for the full period (2004–2016), the pre-period (2004–2008) and the post-period, split into two parts (2009–2012 and 2013–2016). Foreign born are those born in a country other than Sweden. Tenured hold at least five years of tenure at a pharmacy. Industry experienced hold at least five years of experience from the pharmacy industry.

4 Institutional setting and descriptive patterns

4.1 Introducing competition on the pharmacy market

Between 1971 and 2009, *Apoteket* (the National Corporation of Swedish Pharmacies) had the exclusive right to retail prescription and non-prescription pharmaceuticals in Sweden. *Apoteket* was a state-owned enterprise whose primary purpose was to ensure a nationwide pharmaceutical supply system. In 2009, a regulatory reform ended the monopoly. The main implication was that *Apoteket* lost its monopoly rights and private firms could enter.

The timeline of the deregulation is included in Figure 1. A special inquiry was commissioned in December 2006 to submit proposals regarding deregulating the market. *Apoteket*'s exclusive rights were abolished through the implementation of three bills between September 2008 and November 2009. Only the second bill (Government Bill 2009b Prop 2008/09:145) is pivotal to the experiment that this paper relies on. The bill, which passed in parliament in April 2009 and into law in July 2009, made it possible for private firms to operate pharmacies. Following this law change, *Apoteket* sold the majority of its pharmacies to private owners but remained in the market. The first private pharmacies opened to customers in January 2010.[21]

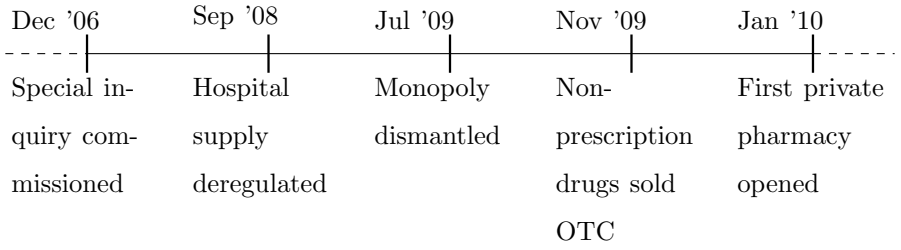


Figure 1: Timeline of regulatory changes

Entry into the market took place through two channels. First, two thirds of *Apoteket*'s pharmacies were privatized. Second, firms could open new pharmacies, subject to obtaining a permit from the Swedish Medical Products Agency. Only variation from the first channel is used in the empirical strategy. Based on public records, 466 out of *Apoteket*'s 946 pharmacies were sold to private firms and a further 150 pharmacies were transferred to a separate state-run company, *Apoteksgruppen*, where entrepreneurs could enter as majority owners (SOU 2017:15 2017). The privatization involved sales of pharmacies

[21]The other two bills are not directly relevant to retail pharmacies. The first bill (Government Bill 2008 Prop 2007/08:142) involves a separate product market – the supply of drugs to hospitals. Prior to 2008, only *Apoteket* or the caregiver could supply inpatient pharmaceuticals. From September 2008, inpatient drugs can be publicly procured from other suppliers. The third bill, implemented in November 2009, implied that certain non-prescription drugs could retail in new locations, such as supermarkets (Government Bill 2009a Prop 2008/09:190). Price competition for non-prescription drugs therefore intensified by breaking up the monopoly and by retailing non-prescription drugs in new locations.

in eight clusters during 2009.[22] The number of clusters and composition of pharmacies within each cluster was formed by the policymaker. The stated aim was to promote competition in and after the bidding process and to achieve competitive neutrality between public and private owners (National Audit Office 2012).

The share of employment at a public owner in the pharmacy industry fell from 100% to 30% and the number of pharmacies increased by over 400 following the deregulation. Likewise, the number of firms increased substantially, reaching over 150 within two years of the deregulation. Labor market concentration also fell substantially. Average HHI in the pharmacy market, weighted by employment in each local labor market, fell from 1 in pre-deregulation to just over 0.25 in 2016.

The variation in HHI that I use in the empirical strategy is based on observed ownership changes in the data in 2009.[23] In this year, most of the privatization occurred but no additional firms had yet entered and no new pharmacies had been opened. Figure 2 maps HHI by local labor market in 2009. Recall that the industry was a monopoly pre-2009. Thus light-colored areas with relatively low levels of HHI in 2009 have experienced the largest reductions in employer concentration, or similarly the largest increases in competition. In line with earlier literature (see Rinz 2020), urban areas are least concentrated post-deregulation and rural areas are most concentrated. Nevertheless, there is substantial variation in HHI across local labor markets.[24]

[22] There were also bids for individual pharmacies in *Apoteksgruppen*. They were first received in 2010 (see National Audit Office 2012).

[23] I observe firm and workplace identifiers over time, as well as which employees work at each workplace and firm. I neither have access to data on the pharmacies included in each cluster nor the winners of the bids.

[24] The reform variation is also illustrated in Figure A.3, which shows the distribution of HHI in 2009 across LLMs and employees. Table A.1 includes pre-reform summary statistics separately for local markets where the change in HHI is high, medium or low. The groups are similar along many dimensions, like the gender and age composition of the workers, but they also differ along some dimensions, like share of foreign born. LLM fixed effects control for permanent differences between local markets in the empirical estimation.

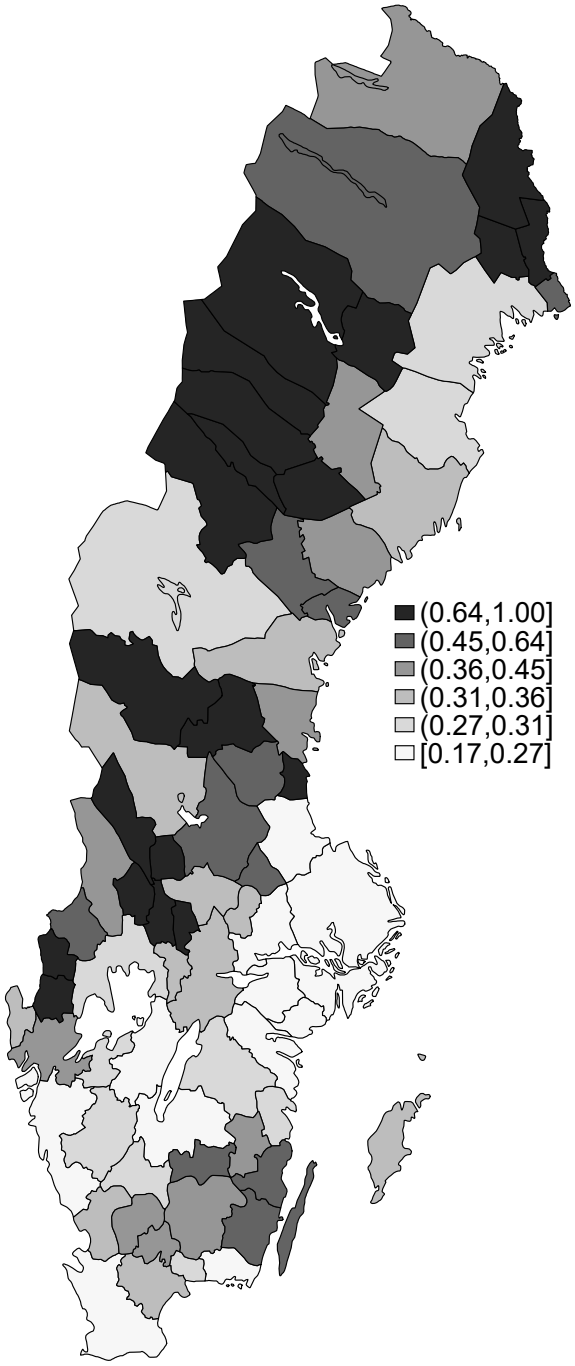


Figure 2: HHI by local labor market (2009)

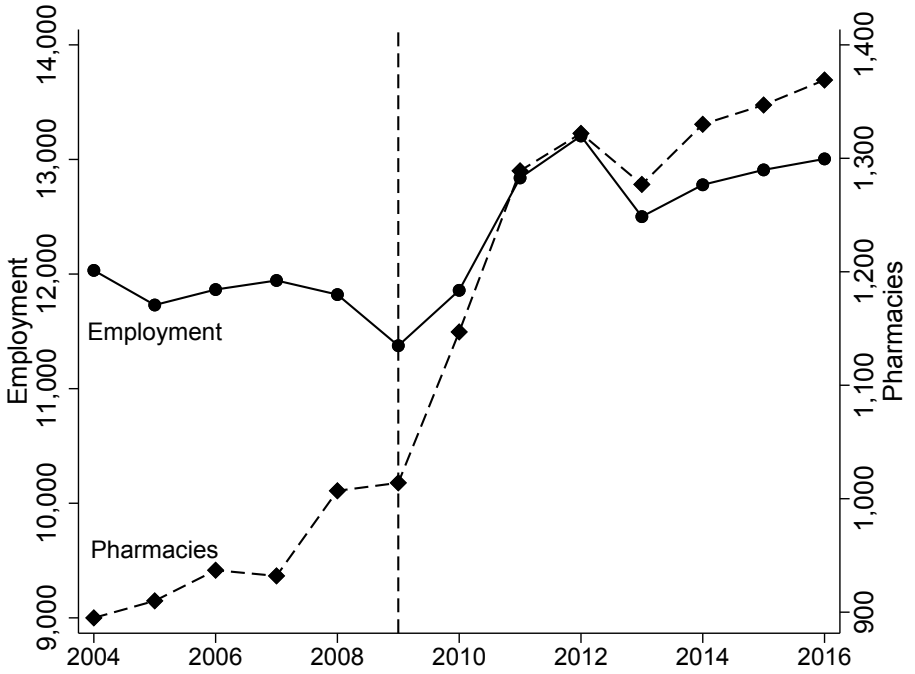


Figure 3: Employment and number of pharmacies in pharmacy industry

4.2 Employees

Figure 3 shows the number of employees and the number of pharmacies in the pharmacy industry annually between 2004 and 2016, based on data in the main sample. Prior to the deregulation, there were 12,000 individuals working in the industry. This has increased by just under 10%, to 13,000, post-deregulation. The number of pharmacies has increased by around 50%, from 900 to nearly 1,400.[25] It follows that the number of employees per pharmacy has decreased and that the market has become fragmented: there are many more pharmacies operated by many more firms, but not equally more employees.

A simple monopsony model predicts that employment should respond positively to reduced monopsony power. This is in line with what we see in the data. Similarly, if the market experienced a market-wide labor demand shock once deregulated, a competitive model would also predict that employment increases. The latter, however, is unlikely. Labor demand is derived from product demand, and the demand for pharmaceuticals is unlikely to jump discontinuously in 2009. I return to this in Section 6.3 below.

[25]The number of pharmacies in the data is similar but generally slightly higher than that reported by the Dental and Pharmaceutical Benefits Agency (*TLV*), see Dental and Pharmaceutical Benefits Agency (2018) Figure 2. This could arise if the industry code for dispensing chemists is wider than that used by *TLV*, and/or if some pharmacies organize the pharmacy under more than one workplace identifier. Similarly, the number of employees is higher than that reported by the industry organization *Sveriges Apoteksförening*. In addition to the above, this could for example arise if fewer employee categories are included in the industry organization's reporting.

4.2.1 Type of pharmacy employees

Pharmacists make up 50% to 60% of the workforce in pharmacies, pharmacy technicians around 25% and non-pharmacists, including managers, the remainder. Pharmacies must have a pharmacist on duty during opening hours. They have an occupational license, issued by the National Board of Health and Welfare.[26] In 2016, 64% of pharmacists worked in a pharmacy. The second largest industry for pharmacists at the five-digit level was manufacture of pharmaceutical preparations (6%) and the third largest wholesale of pharmaceutical goods (5%).[27] Pharmacy technicians primarily work with sales and advice on non-prescription drugs and retail items, and can assist with dispensing medicines. They are tied to pharmacies – over 95% of pharmacy technicians work there – but are harder to identify in data as they lack unifying educational backgrounds or occupational codes.[28] Prior to the deregulation, *Apoteket* internally trained pharmacy technicians. Post-deregulation, there are vocational degrees. In 2016, 62% had upper secondary schooling or less.

4.3 Wages

Despite being a state-owned enterprise, employees negotiate wages individually both pre-deregulation and post-deregulation. The pharmacy market is fully covered by collective agreements. Wages are not specified in the collective agreements but are set flexibly in annual wage negotiations between the employee and their manager.

4.3.1 Aggregate changes in wages

The deregulation may on aggregate result in either increases or decreases in wages, as outlined in Section 2. To descriptively gauge aggregate changes in wages for pharmacy employees over time, workers in the pharmacy industry are compared either to the school sector (preschools and compulsory schools) or to the labor market as a whole. The school sector is used because spillovers to this sector are unlikely, and because it shares important features with the pharmacy industry:[29]

[26] Formally, there are two types of pharmacists: those who hold at least a Master's degree in Pharmacy (*apotekare*) and those who hold a Bachelor's degree in Pharmacy (*receptarie*). The legal requirements apply to either type of pharmacist, and there are only small differences in tasks performed. Consequently no distinction is made between the two categories in this paper.

[27] Table A.2 includes the top 5 industries by employment in 2016 for educated pharmacists.

[28] From 2014, there is an occupational code for pharmacy technicians. The occupational code prior to 2014 is broader and does not only encompass pharmacy technicians.

[29] Both public and private organizations run schools. Similar to pharmacies, it is highly skilled and female dominated (see Figure A.2, which compares workers in the pharmacy industry to workers in schools and the full labor market). Teachers are certified and, from 2011, there is an occupational license.

$$\ln(w_{it}) = \alpha_1 Pharmacy_{it} + \alpha_2 Pharmacy_{it} \times Post_t + \lambda_t + \beta X_{it} + \varepsilon_{it} \quad (7)$$

The coefficient of interest is α_2. This summarizes the aggregate change in wages in pharmacies in the post- compared to pre-period, relative to changes in wages on the whole labor market and in schools. Individual-level controls, X_{it}, for age, gender, foreign-born and education are included in the specification to make the comparison between similar workers. In certain specifications, CZ fixed effects are included to control for permanent regional wage differences, or individual by CZ fixed effects to control for individual heterogeneity in wages. Estimates of α_2 in Table 2 indicate that wages on aggregate increased by 2 to 4 percentage points more for workers in pharmacies in the post-period (from 2009 onward) than they did in the rest of the labor market or for individuals working in schools.[30]

Table 2: Aggregate changes in ln(*wage*)

	(1)	(2)	(3)
Panel A: Full labor market			
Pharmacy × Post	0.030***	0.034***	0.023***
	(0.008)	(0.009)	(0.002)
R^2	0.394	0.419	0.936
N	26,587,246	26,587,246	25,613,134
Panel B: Pharmacies and schools			
Pharmacy × Post	0.040***	0.042***	0.020***
	(0.006)	(0.006)	(0.002)
R^2	0.646	0.663	0.958
N	3,256,605	3,256,605	3,113,725
Year FE	Yes	Yes	Yes
Controls	Yes	Yes	Yes
CZ FE		Yes	
Person × CZ FE			Yes

Note: This table provides the estimates of α_2 from estimating equation (7). Controls are age (in five categories), gender, foreign born and level of education (in five categories). Standard errors are clustered by commuting zone (CZ) and reported in parentheses.

[30] Comparing changes in mean (deflated) wages over time gives a similar picture. Wages have increased 2 (3.7) percentage points more in pharmacies than in the full labor market (schools).

5 Empirical strategy

The deregulation provides sudden and stark variation in the number of employers. While the national pharmacy market was deregulated, the change in local labor market concentration varied over the country. The empirical strategy exploits the fact that the deregulation gave rise to differences in the size of changes in labor market concentration across local labor markets within the pharmacy industry. Recognizing that actual HHI is potentially endogenous, it only makes use of the variation in concentration from the privatization of pre-existing pharmacies. The empirical strategy is outlined in detail below.[31]

I would like to estimate regressions of the following form to understand the effect that labor market concentration has on wages:

$$\ln(w_{imt}) = \alpha \ln(HHI_{mt}) + \lambda_m + \lambda_t + \beta X_{imt} + \varepsilon_{imt} \tag{8}$$

λ_m are local labor market (LLM) fixed effects, λ_t are year fixed effects and X_{imt} are additional controls. The concern is that actual HHI_{mt} is potentially endogenous, leading to biased estimates of the coefficient of interest, α. For example, market concentration depends on firms' location choices. If unobserved factors affect both the choice of where to locate pharmacies as well as wages in those locations in ways which are not controlled for, $\hat{\alpha}$ will be biased. Another potential concern is that HHI is based on employment shares, which relate to individuals' decisions of where to work. The reform coincides with increased aggregate mobility on the labor market (see Figure 7) and wages are a key component in the labor supply decision.[32]

To address these concerns, I adopt an approach where, instead of using actual changes to concentration, only the change that arises from the privatization of pharmacies that exist when the market is deregulated is used. This is the part of the variation that was within the policymakers' control. Reduced form regressions of the following form are estimated for everyone who is employed at a pharmacy:

$$\ln(w_{imt}) = \gamma[\ln(HHI_{m,2009}) \times Post_t] + \lambda_m + \lambda_t + \beta X_{imt} + \varepsilon_{imt} \tag{9}$$

As above, λ_m are LLM fixed effects and λ_t are year fixed effects. X_{imt} are additional controls for age, gender, foreign-born, level of education, pharmacist, tenure and industry

[31] The empirical models are estimated in Stata using the Multi-Way Fixed Effects estimator (Correia 2017).

[32] Schubert, Stansbury, and Taska (2020) stress that wage–HHI regressions are likely to be biased if workers' outside options are ignored. This is likely to be less of a concern in this setting for at least two reasons. First, pharmacists and pharmacy technicians are closely tied to pharmacies, so outside options that require their skills are limited. Second, permanent differences in outside options across local labor markets are controlled for in the specifications.

experience. $\ln(HHI_{m,2009})$ is log HHI in 2009 and $Post_t$ is a dummy variable equal to one from 2009 onward. Each local labor market receives a constant value of $\ln(HHI_{m,2009})$, such that $\ln(HHI_{m,2009}) \times Post_t$ captures treatment intensity at the local market level from the year of deregulation onward. The specification exploits the full variation in market concentration based on the sale of pre-existing pharmacies as a predictor of the actual change in concentration. $HHI_{m,2009}$ is neither affected by firms' decisions to open up new pharmacies nor to any mobility decisions that workers make post-deregulation, but is instead a function of the deregulatory design (see Section 4 for details).[33]

The local labor market fixed effects control for permanent differences across local markets and year fixed effects for general time trends in industry wages, including those due to inflation. Importantly, national effects of deregulating the market will be absorbed in the time fixed effects. The empirical strategy allows me to capture the effects of changes to labor market concentration without confounding it with effects related to product market concentration, which are likely to be national.[34] The local labor market fixed effects are important both because wages are likely to differ throughout the country for reasons unrelated to the reform, and because firms generally want to locate pharmacies in highly populated areas. Finally, as wages are particularly driven by individual characteristics, certain specifications use person by local labor market fixed effects instead of local labor market fixed effects. This keeps composition constant. The identifying variation then comes from workers who have stayed in their local labor market over time and consequently experienced different levels of concentration pre- and post deregulation.

Equation (9) is a type of difference-in-difference specification with different treatment intensities that, in addition to correct functional form, relies on parallel trends in log wages across local markets in the absence of the deregulation. To support this assumption, I also estimate event versions of equation (9):

$$\ln(w_{imt}) = \sum_{t \neq 2008} \gamma_t \ln(HHI_{m,2009}) \mathbb{1}[year = t] + \lambda_m + \lambda_t + \beta X_{imt} + \varepsilon_{imt} \tag{10}$$

Estimating responses by year serve a dual purpose. This both shows whether there are pre-treatment trends in log wages relative to 2008 (the omitted year) and the time trajectories of post-treatment effects. In Section 6.3 I include extensive additional robustness checks.

[33] We may worry that there is some measurement error in employment in 2009 as employment seems to dip slightly in that year, see Figure 3. The results are robust to instead calculating HHI based on pharmacies and employment from 2004 but ownership structures from 2009, see Figure B.2.

[34] Qiu and Sojourner (2019) highlight the importance of distinguishing between concentration in product and in labor markets. Kroft, Luo, Mogstad, and Setzler (2020) study the construction industry to find that firms in this industry enjoy rents both due to markups of prices and markdowns of wages.

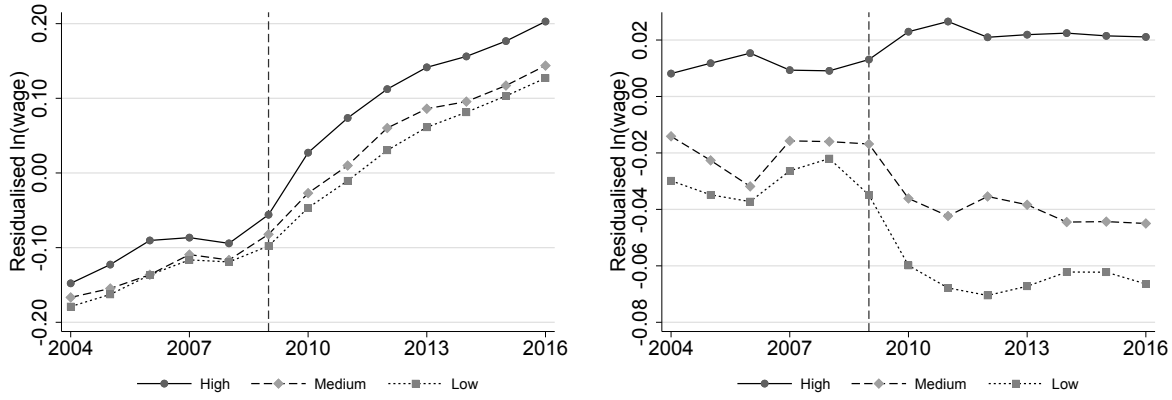
6 Results

6.1 Wage effects

6.1.1 Descriptive patterns

Before formally estimating the effect of labor market concentration on wages, I show descriptively how wages have evolved over time in local markets that experienced different changes in employer concentration. Figure 4 plots residualized log wages separately by change in HHI. Pharmacy workers are split into three categories: those that work in local markets where the change in HHI due to the deregulation is high, medium or low. Wages are residualized by age, gender, foreign born and level of education in Panel A. To control for general time trends in wages, year fixed effects are also included in Panel B.

The figure suggests that wages evolve in parallel prior to deregulation in markets that ended up experiencing different changes in concentration. Notice that mean wages are not the same even though HHI is the same everywhere in the pre-period. Wages are set flexibly, and differences in wages are driven by factors other than employer concentration. Upon deregulation, wages grow faster in markets that experience larger reductions in labor market concentration.



(a) Without year FE

(b) With year FE

Figure 4: Residualized log wages by change in labor market concentration

Note: This plots mean residualized log wages separately by three groups of changes in labor market concentration: "High" (where $HHI_{m,2009}$ is less than or equal to the 25th percentile of the distribution of $HHI_{m,2009}$ by LLM), "Medium" (where $HHI_{m,2009}$ is above the 25th but less than or equal to the 50th percentile), and "Low" (where $HHI_{m,2009}$ above the 50th percentile). In Panel A, log wages are residualized by age, gender, foreign born and level of education. In Panel B, year fixed effects are additionally included.

6.1.2 Estimations

Columns (1) to (3) in Table 3 report the results of estimating equation (8) in Panel A, and equation (9) in Panel C. Columns (4) and (5) include the first stage and IV results of instead using $\ln(HHI_{m,2009}) \times Post_t$ as an instrument for $\ln(HHI_{mt})$. Focusing first on the OLS results, the point estimates are negative and statistically significant at the 1% level. Taken at face value, they imply elasticities of wages with respect to labor market concentration between -0.016 (see 0.003) and -0.048 (se 0.007). The reduced form estimates in Panel C are between -0.022 (se 0.006) and -0.048 (se 0.011), and the elasticities implied by the IV estimation are between -0.020 (se 0.006) and -0.046 (se 0.008).[35] The similarity in estimates across estimation techniques arises because the change in market concentration in the year of deregulation is highly predictive of actual changes in market concentration on the pharmacy market. The first stage coefficient is between 1.023 and 1.065, and the F-statistic between 80.45 and 107.21, depending on whether person by LLM fixed effects are used or not. Figure A.4 in the appendix plots actual HHI against $\ln(HHI_{m,2009}) \times Post_t$. The two measures are highly correlated: the R^2 of regressing $\ln(HHI_{m,t})$ on $\ln(HHI_{m,2009}) \times Post_t$ is 0.91.

To put the estimated effects into perspective, a local market that ended up at the 25th percentile of market concentration instead of the 75th percentile following the deregulation (an HHI of 0.28 instead of an HHI of 1 in the LLM-year distribution) would have wages that are 2.5 to 6 percent higher. The negative effect that market concentration has on wages echoes a growing literature in labor economics.[36] The results are very similar to those found by Hershbein, Macaluso, and Yeh (2019) and Rinz (2020), who report elasticities between -0.01 and -0.05, and somewhat smaller than those found by Azar, Marinescu, and Steinbaum (2020) who report elasticities between -0.03 (OLS) and -0.14 (IV). The somewhat smaller magnitude could reflect that the public monopsonist did not depress wages as much as its market power would allow it to. It could also be that labor market institutions, such as the high degree of collective bargaining and the fact that this is a highly specialized group of labor, limit the firm's bargaining power pre-deregulation. In line with this, Jarosch, Nimczik, and Sorkin (2019)'s model implies that the elasticity of wages with respect to concentration becomes smaller when worker bargaining power increases, i.e. wages are less sensitive to concentration when workers have more bargaining power. Moreover, the estimates are likely lower-bound estimates of the effect of labor market concentration on wages, as discussed in Section 2.2.1.

[35] Similar but somewhat more imprecise results are found when using log earnings as the outcome, see Appendix Table A.3.

[36] See Arnold (2020); Azar, Marinescu, and Steinbaum (2020); Benmelech, Bergman, and Kim (2018); Hershbein, Macaluso, and Yeh (2019); Jarosch, Nimczik, and Sorkin (2019); Lipsius (2018); Martins (2018); Prager and Schmitt (2021); Qiu and Sojourner (2019); Rinz (2020); Schubert, Stansbury, and Taska (2020).

Table 3: Effect of labor market concentration on ln($wage$) – OLS, RF and IV results

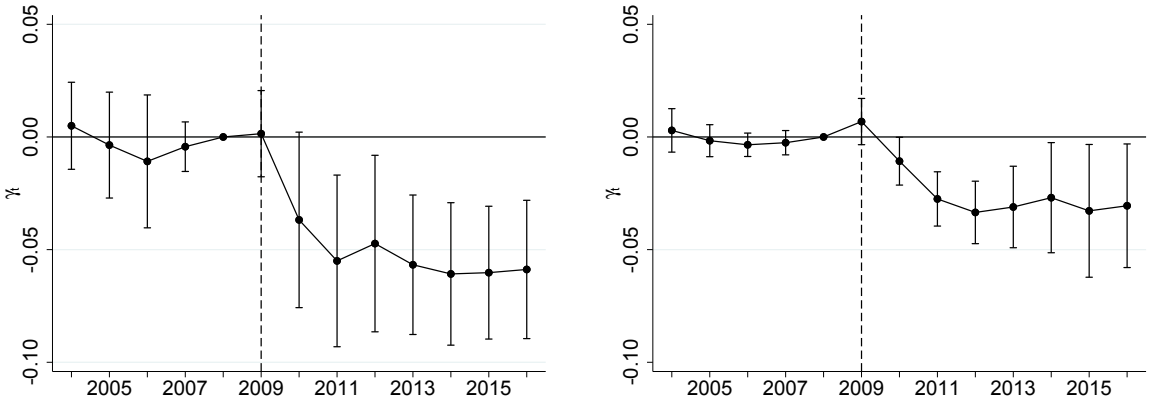
	(1)	(2)	(3)	(4)	(5)
	OLS			IV	
Panel A: OLS & IV					
$ln(HHI_{mt})$	-0.045***	-0.048***	-0.016***	-0.046***	-0.020***
	(0.009)	(0.007)	(0.003)	(0.008)	(0.006)
R^2	0.217	0.552	0.924		
Panel B: First stage					
$ln(HHI_{m,2009}) \times Post_t$				1.023***	1.065***
				(0.114)	(0.103)
F-statistic				80.45	107.21
Panel C: Reduced form					
$ln(HHI_{m,2009}) \times Post_t$	-0.033**	-0.048***	-0.022***		
	(0.013)	(0.011)	(0.006)		
R^2	0.216	0.551	0.924		
N	110,825	110,722	104,968	110,722	104,968
Year FE	Yes	Yes	Yes	Yes	Yes
LLM FE	Yes	Yes		Yes	
Person × LLM FE			Yes		Yes
Controls		Yes		Yes	

Note: This provides the results of estimating equations (8) and (9) for log wages. OLS and IV results are presented in Panel A. The first stage in Panel B show the results of regressing ln(HHI_{mt}) on the instrument ln($HHI_{m,2009}$)×$Post_t$ and exogenous regressors. Controls are age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), tenure (in three categories) and industry experience (in three categories). Standard errors are clustered by LLM and reported in parentheses. The F-statistic is the Kleibergen-Paap Wald rk F-statistic.

In the remainder of the discussion, I focus on the reduced form results from estimating equation (9). To check pre-reform parallel trends in log wages as well as post-reform differences in effects, Figure 5 plots estimates of γ_t from equation (10). The results support that wages evolve in parallel in local labor markets prior to deregulation, responding positively to the lower concentration post-deregulation.[37] As expected, there is no response to the deregulation in 2009: while the deregulation came into force in 2009, the new firms formally began trading in 2010. The positive wage effects rise between 2009 and 2010, and again between 2010 and 2011, and at least at the 5% level they are not equal

[37]Figure 4 suggests that the evolution of mean wages slows down in 2008. The parallel trends are robust to instead using 2004 as the base year.

between 2009 and 2011. From 2011 onward, the estimated effects are relatively stable. This stability over time is also supported by regressions of equation (9) when the pre-2009 period is grouped and the post-2009 period is split into two (2010–2012 and 2013–2016). The coefficients on the two sub-periods post-deregulation are not found to be statistically significantly different from each other.



(a) LLM fixed effects

(b) Person by LLM fixed effects

Figure 5: The effect of $\ln(HHI_{m,2009})$ on $\ln(wage)$ over time

Note: This plots estimates of γ_t from equation (10) with 95% confidence intervals. Panel A also controls for age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), workplace tenure (in three categories) and industry experience (in three categories).

6.2 Composition effects

The results in Table 3 show that there are differences in point estimates across models. The estimates of γ that are only identified from differences in concentration across LLMs over time, without any controls, are -0.033. When adding controls, the point estimates increase to nearly -0.048. This suggests that the composition of workers bears at least partial importance for the estimated wage effects.[38] The estimates that use within-individual variation over time and therefore fully control for composition are around -0.02. That these estimates are slightly smaller suggests that high-wage workers become employed in markets where concentration changed more. Put differently, the lower concentration that materializes post-deregulation both results in higher wages for workers conditional on worker quality (seen by the regressions that include individual by local labor market fixed effects), and changes the composition of workers toward higher-paid workers.[39]

[38]That labor market concentration can affect the types of workers hired is highlighted by Hershbein, Macaluso, and Yeh (2019), who find that local labor market concentration is negatively correlated with wages but positively correlated with skill demand.

[39]This result is also supported by the fact that when the sample is restricted to incumbents, the estimated wage effects from reduced labor market concentration are -0.035 (compared to -0.048 for the full sample), in a model that includes year fixed effects, local labor market fixed effects and controls.

To gauge how composition changes following the deregulation, Figure 6 plots the estimated γ-coefficients from estimating equation (9) without controls for five indicator outcomes: female, foreign born, post-secondary education, pharmacist and young workers (defined as being below age 40). The results indicate that there is no statistically significant effect at the 5% level on the share with post-secondary education. Local markets where where concentration was reduced the most have experienced faster growth in the share of foreign born and the share of young employees. On the contrary, the share of women and pharmacists has declined more in these markets, though the effect is only statistically significant at the 5% level for women, not pharmacists.

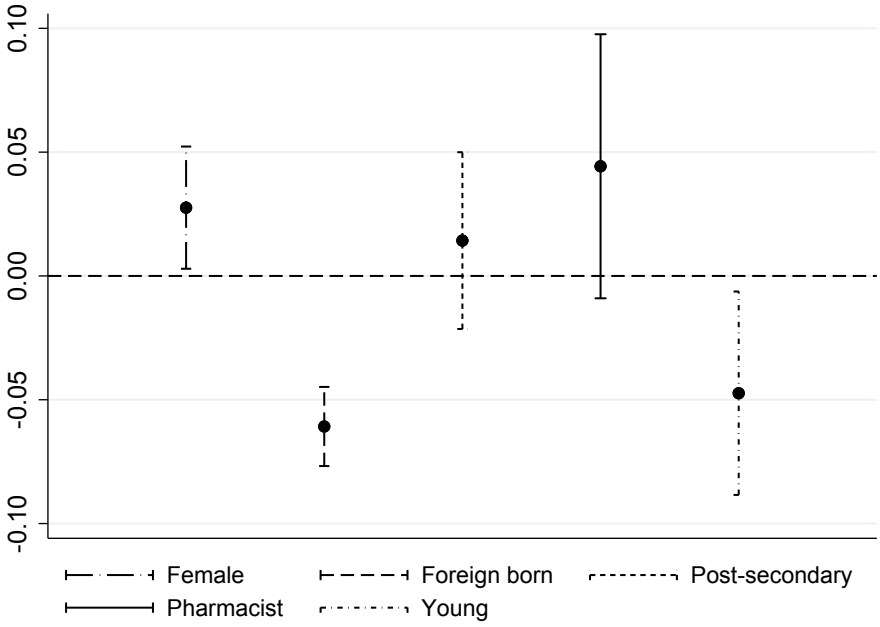


Figure 6: Compositional changes

Note: The figure plots estimated γ-coefficients with 95% confidence intervals from the following model for five indicator outcomes: $Y_{imt} = \gamma[\ln(HHI_{m,2009}) \times Post_t] + \lambda_m + \lambda_t + \varepsilon_{imt}$. Young is defined as being below age 40.

6.3 Robustness checks

Robustness checks are included in Appendix B. To assess the stability of the estimates, a set of robustness checks estimates equation (9) using different samples or controls. The γ-estimates are included in Appendix Table B.1. In column (1) local labor market fixed effects and controls are included and in column (2) individual by local labor market fixed effects are included. Baseline results are shown in Panel A.

We may be worried that the relationship between labor market concentration and wages is biased if labor demand changes in conjunction with concentration. Boal and Ransom (1997) highlight that a negative correlation between labor market concentration

and wages cannot be taken as evidence against competitive labor markets unless total labor demand is fixed. If the deregulation coincides with a market-level increase in labor demand, we may see increases in wages as well as the number of firms who enter, leading to lower levels of market concentration. This would bias $\hat{\gamma}$ away from 0.[40] To alleviate this concern, I include controls for log number of pharmacy employees at the LLM-year level.[41] The results of this estimation are included in Panel B of Table B.1. The point estimates at between -0.020 and -0.046 are very similar to the baseline results of -0.022 to -0.048, indicating that the results are not driven by labor demand effects. We also expect demand effects, to the extent that they exist, to be moderated over time. The event analysis presented in Figure 5 shows that the wage effects from reduced labor market concentration persist also over the longer term, which supports that the estimated effects do not reflect a demand shock.

In Panel C of Table B.1, controls for log value added per employee at the firm-level are included to remedy the concern that productivity changes coupled with rent-sharing may drive observed wage effects. In Panel D, urban areas (precisely, local markets that encompass the three largest cities in Sweden) are excluded to ensure that the effects are not driven by these markets alone. The share of managers employed in the pharmacy industry increases from 2% to between 6% and 8% post-deregulation. In Panel E, managers are excluded from the sample to ensure that the results are not driven by the increase in the number of managers. In Panel F only the public sector is included. This is to remedy the concern that the effects are driven by the entry of private firms only, that may have different objectives to public firms. Finally, in Panel G controls are included for mean log wages at the CZ and year level, to remedy the concern that wages may in general be growing faster in certain commuting zones. The estimates are remarkably stable across the empirical models, and the result that the reduction in market concentration positively influences wages is robust to these checks. The most conservative estimates that hold composition constant find estimates between -0.013 and -0.023.

While the analysis above displays reassuring pre-trends (see Figure 5) and is robust to controlling for log wages in the CZ (see Panel G of Table B.1), we may nevertheless worry that the regressions are picking up spurious changes within these local labor markets rather than effects related to the competitive changes. To alleviate this concern, a placebo analysis is performed where event specification (10) is estimated for the school sector. Figure B.1 plots γ_t separately for the school sector and the pharmacy market.

[40] Given that labor demand is derived from product demand, and that there is little reason to expect demand for pharmaceuticals to be affected by the deregulation, this alleviates the concern that there is a concurrent labor demand shock.

[41] An alternative is to control for log number of pharmacies. The correlation between log number pharmacies and log employment is 0.98, thus I do not include both at the same time. Instead controlling for the number of pharmacies also yields statistically significant and negative estimates of employer concentration on wages at the 5% level.

$HHI_{m,2009}$ is defined as previously. Reassuringly, the analysis shows that employees in schools have not experienced corresponding wage growth as in the pharmacy industry, supporting that the estimated wage effects are real rather than spurious. The analysis for the school sector suggests that wages have grown slightly faster over time in markets where concentration has changed more, but these effects are much smaller than those estimated in the pharmacy industry. Taken together, the results support that reducing market concentration leads to a positive wage effect not observed in general in local labor markets.

7 Heterogeneity in estimated effects

This section explores who benefits most from reduced labor market concentration. First, I study new hires and stayers in the pharmacy industry. I consider how mobility is affected in Section 7.1.1, and how wage effects of reduced labor market concentration differ for new hires and stayers in Section 7.1.2. Second, I consider whether the estimated wage effects differ by individual characteristics in Section 7.2.

7.1 New hires and stayers

7.1.1 Mobility effects

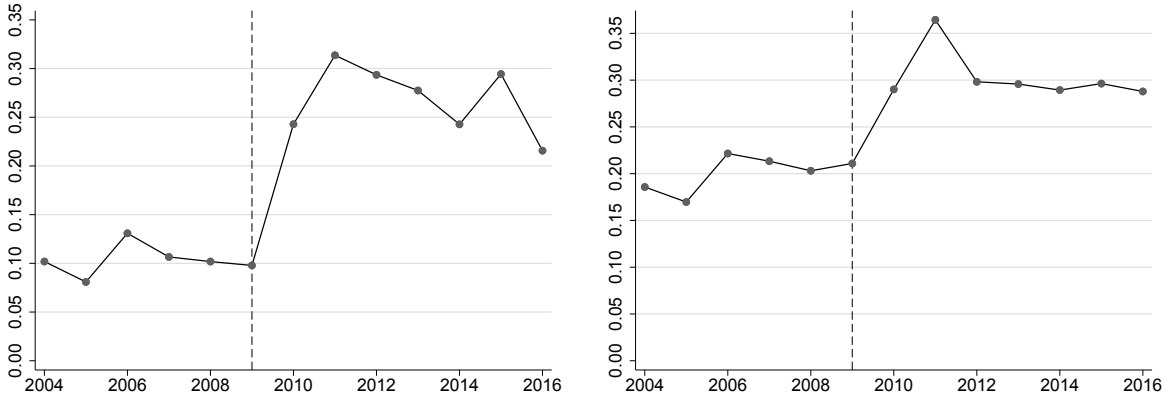
To understand the types of career options that individuals face, it is instructive to consider transitions on the labor market. The rate at which workers change jobs is informative of the extent of competition between employers, as the ability of workers to leave for another employer limits the wage-setting power that employers hold. Mobility can be measured in different ways. Mobility to firms (where one firm can have many workplaces, or in this case, pharmacies) captures the fraction of employees who change firms or transition from non-employment, while mobility to workplaces captures the fraction of employees that change workplace (which can either be within the same firm, across firms or from non-employment). For wages, mobility between firms is likely important: changing firms involves changing employers, and is an opportunity to renegotiate wages. Mobility between workplaces can also be important for wages and represent a career transition.

To gauge whether the opportunity to move has been affected by the deregulation, Figure 7 plots the fraction of new hires to a firm (Panel A) or pharmacy (Panel B) across all employees working at a pharmacy.[42] This shows that mobility in the pharmacy industry as a whole increased following the deregulation. A new hire is defined to be an employee who works at the firm (pharmacy) in year t but did not work there the year before, in $t-1$. To capture genuine recruits rather than moves that happen due to

[42]New hire is used interchangeably with mover, joiner or recruit in the text.

restructuring, the restriction that it is not a move if over 50% of the employees at the firm (pharmacy) in year t come from the same firm (pharmacy) in $t-1$ is applied.

Prior to the deregulation, the fraction of new firm hires was 10%. Because there was only one firm in the pharmacy industry, this represents transitions from firms in other industries and from non-employment. Firm mobility fluctuates between 25% and 30% post-deregulation, spiking at 31% in 2011. Mobility to pharmacies is higher than to firms, at 20% pre-deregulation. Post-deregulation, the share of new hires to pharmacies increases by around ten percentage points, to 30%. Just like firm mobility, pharmacy mobility peaks in 2011 at 36%, supporting that there is a reshuffling of employees in close conjunction with the deregulation. The increase in mobility between pharmacies is not only driven by the opening of new pharmacies but is also present at pre-existing pharmacies (though the changes are less stark; see Appendix Figure A.5).



(a) Share new hires to firm

(b) Share new hires to pharmacy

Figure 7: Mobility in pharmacy industry

Note: Panel A and B show the annual share of employees that join a new firm and join a new pharmacy respectively.

To assess how the fraction of new hires relates to changes in labor market concentration, Table 4 reports estimates from equation (9) where the outcome is joining a firm (Panel A) or joining a pharmacy (Panel B). In column (1) the full sample is included. The results suggest that the likelihood of joining a new firm is positively affected by reduced labor market concentration: the point estimate of -0.046 implies that the likelihood of moving increases by 0.44 percentage points when HHI decreases by 10%. The effects for joining a new pharmacy are not statistically significantly different from zero. This implies that the share who move pharmacy does not increase more in areas where the change in concentration is high compared to low.

New hires can either come from another employer or from non-employment. Manning (2003) explains that a simple measure of monopsony power is the share of recruits

from non-employment. If workers who quit are easily replaced by new hires from non-employment, the threat of quitting is limited and we expect employers' wage-setting powers to be high. In column (2), attention is limited to the sample of new hires and the outcome is a dummy that indicates whether the new hire came from non-employment. The results suggest that the likelihood of joining a firm from non-employment is negatively related to the reduction in labor market concentration, but the effects are only statistically significant at the 10% level. A 10% decrease in HHI is related to a 0.78 percentage point decrease in the likelihood of joining a firm from non-employment, or similarly a 0.78 percentage point increase in the likelihood of making a firm-to-firm transition.

Table 4: Effect of labor market concentration on hiring

	(1) Join	(2) Non-employment
Panel A: Firm hires		
$ln(HHI_{m,2009}) \times Post_t$	-0.046***	0.082*
	(0.008)	(0.042)
R^2	0.367	0.166
N	159,633	31,140
Panel B: Pharmacy hires		
$ln(HHI_{m,2009}) \times Post_t$	0.006	-0.015
	(0.007)	(0.019)
R^2	0.454	0.204
N	159,633	41,083
Sample	Full	New hires
Year FE	Yes	Yes
LLM FE	Yes	Yes
Controls	Yes	Yes

Note: This provides estimates of γ from equation (9) for the likelihood of joining a new firm (Panel A) or a new pharmacy (Panel B). In column (1) the outcome is an indicator variable for joining the pharmacy or firm from anywhere, and in column (2) it is an indicator for joining from non-employment, conditional on being a new hire to the firm or pharmacy. Column (1) uses the full sample and column (2) only new hires. Controls are included for age, gender, foreign born, pharmacist, level of education, tenure and industry experience. Standard errors are clustered by LLM and reported in parentheses.

7.1.2 Wage effects

Next, I consider wage effects in the pharmacy industry for new hires and stayers. The analysis should be interpreted with caution: it is based on comparisons of a selected sample, and we saw above that the likelihood of moving firm is itself related to labor market concentration. Table 5 shows results of estimating equation (9) separately for four sub-groups: stayers, new hires, new hires from another firm (which can be inside or outside the pharmacy industry), and new hires from non-employment. As previously, a new hire is defined to be an employee who works at the firm in year t but does not work there in $t-1$.[43] A stayer is someone who is neither a new hire in year t nor in year $t+1$.

Table 5: Effect of labor market concentration on ln(*wage*) – firm stayers and new hires

	(1)	(2)	(3)	(4)
	Stay	*Join*	*Firm-to-firm*	*Non-empl.*
$ln(HHI_{m,2009}) \times Post_t$	-0.048***	-0.023	0.011	-0.026
	(0.013)	(0.027)	(0.042)	(0.034)
R^2	0.560	0.555	0.539	0.516
N	90,554	14,123	9,388	4,732
Year FE	Yes	Yes	Yes	Yes
LLM FE	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Note: This provides estimates of γ from equation (9) for different sub-samples, specified in the column headings. Columns 3 and 4 break out the new hires into two mutually exclusive types: those coming from another firm or those coming from non-employment. Controls are included for age, gender, foreign born, pharmacist, level of education, tenure and industry experience. Standard errors are clustered by LLM and reported in parentheses.

The results show that stayers (compared to other stayers) experience positive wage gains from the reduced labor market concentration while no statistically significant is found for joiners (compared to other joiners). Put differently, conditional on moving to a new firm, there is no additional return to moving in a local market with relatively low levels of concentration.[44] The effect for new hires is not distinguishable from zero irrespective of whether they are coming directly from another firm or from non-employment (see columns 3 and 4).[45]

[43] The same restriction that it is not a move if over 50% of the employees at the firm (pharmacy) in year t come from the same firm (pharmacy) in $t-1$ is applied.

[44] Similar estimates are found for pharmacy hires and stayers, see Appendix Table A.4.

[45] The results should not be interpreted as saying that there are no wage gains to moving. Table A.5 compares wages for joiners and non-joiners in the same firm or pharmacy before and after deregulation.

Employees' tenure and industry experience is related to the career moves that they make. Figure 8 instead considers the effects of labor market concentration separately for employees with different levels of tenure and industry experience. Tenure is defined as the number of years that the employee has spent at the same pharmacy since 2000. Industry experience is analogously defined as the number of years that the employee has spent working in the pharmacy industry since 2000. The results indicate that positive wage effects of lower labor market concentration materialize predominantly for those with longer tenure and more industry experience. This echoes the results for stayers found above. These workers are also likely to be the hardest to replace, given that labor supply in this market is likely to be relatively inelastic.

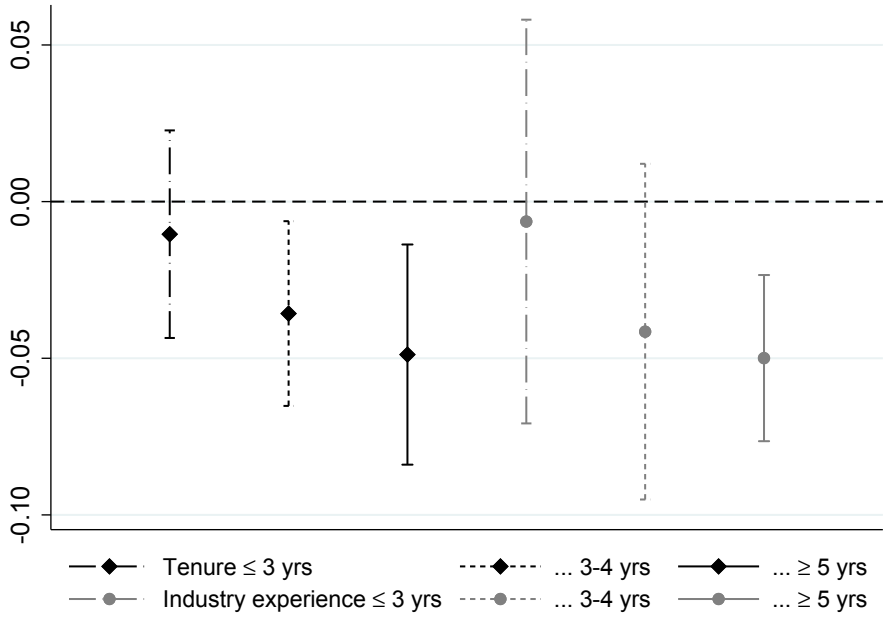


Figure 8: Effect of labor market concentration on ln($wage$) – tenure and industry experience

Note: This plots estimated γ-coefficients with 95% CI from estimating equation (9) for subsamples, as specified by the labels. Full results are included in Table A.6.

7.2 Individual characteristics

Next, I turn to consider how the wage effects from reduced labor market concentration differ by individual characteristics. Just like the analysis for movers, the results should be interpreted with caution since the deregulation has involved compositional changes (see

This abstracts from the role of labor market concentration. Focusing on firm hires in Panel A, the results indicate that new hires to the firm in the pre-period have lower wages than those already working in the firm. In the post-period, the pattern is reversed. Similar results are found when making comparisons in the same pharmacy.

Section 6.2). Figure 9 shows the results of estimating equation (9) by five sub-groups based on gender, country of birth, level of education, being a pharmacist and age. The results are remarkably stable across groups. Overall they suggest that all groups benefit from reduced labor market concentration.

Gender and country of birth: If the monopsonist engages in monopsonistic discrimination (Robinson 1933), which arises from the fact that firms can set lower wages to groups of individuals with more inelastic labor supply to the firm, we expect wages for women and foreign born to respond more strongly to the change in concentration.[46] The estimates in Figure 9 do not, however, suggest that this is the case. There are several plausible explanations for this. A public monopsonist may be less likely to discriminate than a private monopsonist and the collective bargaining institutions may dampen potential discrimination across groups. Moreover, the majority of employees are women both before and after the deregulation.[47]

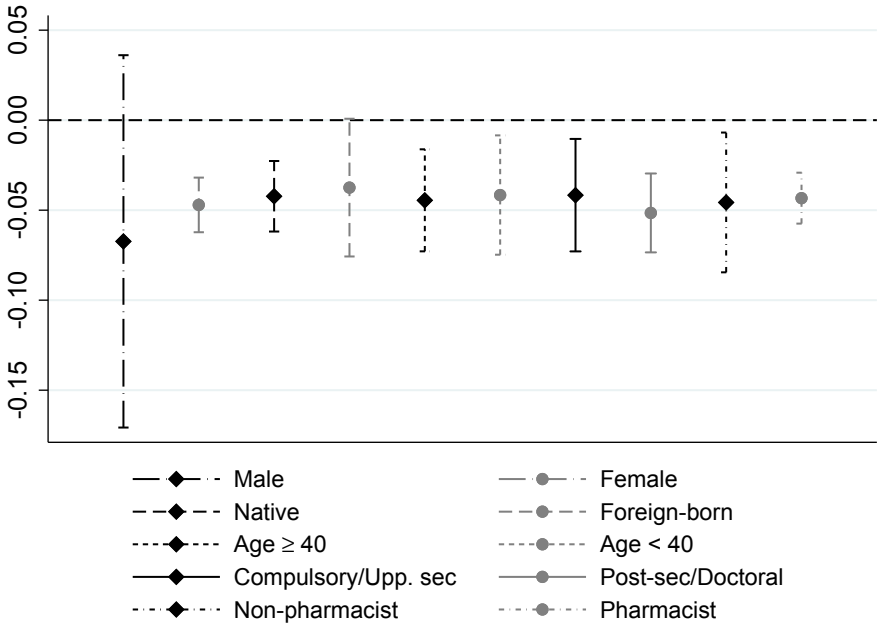


Figure 9: Effect of labor market concentration on ln($wage$) – individual characteristics

Note: This plots estimated γ-coefficients with 95% CI from estimating equation (9) for subsamples, as specified by the labels. Full results are included in Table A.7.

Age: Another dimension along which employers potentially discriminate is age. In the gray-colored bars in Figure 9, the sample is split into employees in the pharmacy industry that are below 40 or 40 and above. The point estimates suggest that both

[46] There are several papers on firm-level monopsony and the gender pay gap (see Hirsch 2009, Barth and Dale-Olsen 2009 and Webber 2016) and the immigrant-native pay gap (see Hirsch and Jahn 2015).

[47] This is also in line with Goldin and Katz (2016) who find that the pharmacy profession in the U.S. is among the most egalitarian profession of all.

younger and older workers experience positive wage effects of similar magnitudes from the reduction in market concentration.

Pharmacists and level of education: We expect the deregulation to change competition for specialized labor but not necessarily non-specialized labor. As Boal and Ransom (1997) point out, concentration of hospitals is not likely to monopsonize the market for hospital housekeepers. Turning to the analysis based on education, the results suggest that both those with at least some post-secondary education and those educated as pharmacists have gained similarly from the reduction in concentration. At the same time, the results also support that those with less education and non-educated pharmacists have gained from the deregulation. While initially surprising, this can arise for several reasons. First, as explained in Section 4, the group of non-pharmacists consists to a large extent of pharmacy technicians. This group is tied to the pharmacy industry and is consequently likely to experience positive wage gains from reduced market concentration. The majority of pharmacy technicians have less than post-secondary education. Second, the similarity in point estimates for pharmacists and non-pharmacists could relate to the bargaining power of pharmacists before the deregulation. If their bargaining power is high – pharmacies are not allowed to operate without a pharmacist on site – this can limit the firm's labor market power in the pre-period and attenuate the estimated gain from reduced labor market concentration for pharmacists. In the absence of this bargaining power, we might expect larger returns to reduced labor market concentration for pharmacists than for less specialized labor.

8 Conclusion

This paper studies a classic question in novel way. It uses a rare natural experiment that provides geographic variation in employer concentration coupled with rich matched employee-employer data to estimate the effect of reduced labor market concentration on wages. In 2009, the Swedish state-run pharmacy monopoly was deregulated. Entry barriers were lifted and private firms entered the market. Because pharmacists and pharmacy technicians primarily work in pharmacies, the reform provides a stark change in employer concentration for employees in the industry. The change in labor market concentration differs across local markets. Making use of this geographic variation in a difference-in-difference design, elasticities of wages with respect to labor market concentration are estimated to be between -0.02 and -0.05. The empirical strategy relies only on the variation in concentration induced by the sale of pre-existing pharmacies in 2009 to remedy the concern that actual labor market concentration is endogenous. This variation in concentration arises directly from the privatization of pharmacies, a process controlled by the policymaker, but does not relate to the opening of new pharmacies post-reform.

The positive wage effects from reduced labor market concentration are robust to extensive checks and are not observed in a similar, but unaffected, industry.

Regarding who benefits from reduced labor market concentration, similar positive wage returns are found for employees of different age, country of birth, gender, educational level and educational specialization. Effects differ for individuals who make different career moves: the positive wage effects primarily arise for stayers, but not new hires, and those with longer tenure and more experience. This does not mean that wages do not respond positively to changing employer. Rather, conditional on changing employer, the results suggest that it does not matter for wages whether the move is made in a local market where the change in labor market concentration is high or low. The result should be interpreted with caution, however, as the likelihood of moving is itself affected by the deregulation.

Overall, the paper lends support to the growing literature that finds that labor market concentration can matter for workers' wages. It finds that wages respond positively to reduced employer concentration in a context with high industry-specificity in skills. This is likely to be true in many other settings where labor is similarly tied, such as teachers or health professionals.

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Appendix A: Additional description and results

Table A.1: Means by change in labor market concentration (2004–2008)

	$HHI_{m,2009}$ interval		
	≤25th (*high*)	>25th & ≤50th (*medium*)	>50th (*low*)
Panel A: Characteristics of LLMs			
Employment	433.44	138.68	34.64
Pharmacies	31.16	12.38	3.59
Population ('000)	341.19	100.70	25.08
$HHI_{m,2009}$	0.25	0.35	0.76
$HHI_{m,2009}$ (empl. weighted)	0.25	0.35	0.55
Number of LLM-year obs.	95	85	185
Number of LLMs	19	17	37
Panel B: Characteristics of employees			
Female	0.91	0.91	0.89
Age (years)	45.64	45.36	45.80
Age < 30	0.16	0.17	0.17
Age ≥ 50	0.47	0.46	0.49
Foreign born	0.14	0.09	0.04
Post-secondary	0.66	0.61	0.58
Pharmacist	0.52	0.50	0.48
Tenured	0.50	0.47	0.52
Industry experience	0.79	0.76	0.71
Monthly wage (2004 SEK)	24,650	23,492	23,102
Monthly earnings (2004 SEK)	22,459	21,449	20,975
Number of employee-year obs.	41,177	11,788	6,408
Number of employees (2008)	8,182	2,366	1,273

Note: The table shows means across LLMs (panel A) or employees (panel B) for the pre-deregulation period, separately by three groups of changes in labor market concentration: "High" (where $HHI_{m,2009}$ is less than or equal to the 25th percentile of the distribution of $HHI_{m,2009}$ by LLM), "Medium" (where $HHI_{m,2009}$ is above the 25th but less than or equal to the 50th percentile), and "Low" (where $HHI_{m,2009}$ above the 50th percentile). The 25th (50th) percentile of $HHI_{m,2009}$ is 0.306 (0.433). Foreign born are born in a country other than Sweden. Tenured hold at least five years of tenure at a pharmacy. Industry experienced hold at least five years of experience from the pharmacy market. $HHI_{m,2009}$ (empl. weighted) is the mean value of HHI 2009, weighting by employment in 2008.

Table A.2: Top 5 industries for pharmacists (2016)

	%
Dispensing chemist	63.90
Manufacture of pharmaceutical preparations	6.16
Wholesale of pharmaceutical goods	4.61
Specialised hospital somatic activities	3.60
Inspection, control, permit & licensing activities of central & local gov't	3.16

Note: The table shows the top 5 industries for pharmacists in 2016 by share of employment. Pharmacists are identified by their educational level and specialisation. Industries are defined by five-digit SNI codes.

Table A.3: The effect of market concentration on ln($earnings$) – OLS, RF and IV results

	(1)	(2)	(3)	(4)	(5)
	OLS			IV	
Panel A: OLS & IV					
$ln(HHI_{mt})$	-0.035**	-0.051***	-0.034***	-0.029**	-0.019*
	(0.014)	(0.012)	(0.012)	(0.013)	(0.010)
R^2	0.082	0.306	0.726		
Panel B: First stage					
$ln(HHI_{m,2009}) \times Post_t$				1.035***	1.063***
				(0.108)	(0.091)
F-statistic				91.66	136.80
Panel C: Reduced form					
$ln(HHI_{m,2009}) \times Post_t$	-0.003	-0.030**	-0.020*		
	(0.012)	(0.015)	(0.012)		
R^2	0.081	0.305	0.726		
N	137,664	137,511	131,149	137,511	131,149
Year FE	Yes	Yes	Yes	Yes	Yes
LLM FE	Yes	Yes		Yes	
Person × LLM FE			Yes		Yes
Controls		Yes		Yes	

Note: This table provides the results of estimating equations (8) and (9) for the outcome log earnings. OLS and IV results are presented in Panel A. The first stage in Panel B show the results of regressing $\ln(HHI_{mt})$ on the instrument $\ln(HHI_{m,2009}) \times Post_t$ and exogenous regressors. Controls are age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), tenure (in three categories) and industry experience (in three categories). Standard errors are clustered by LLM and reported in parentheses. The F-statistic is the Kleibergen-Paap Wald rk F-statistic.

Table A.4: Effect of labor market concentration on ln($wage$) – pharmacy stayers and new hires

	(1)	(2)	(3)	(4)
	Stay	*Join*	*Plant-to-plant*	*Non-empl.*
$ln(HHI_{m,2009}) \times Post_t$	-0.049***	-0.023	-0.022	-0.041
	(0.014)	(0.015)	(0.013)	(0.028)
R^2	0.569	0.540	0.532	0.513
N	79,208	21,139	16,679	4,458
Year FE	Yes	Yes	Yes	Yes
LLM FE	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes

Note: This provides estimates of γ from equation (9) for the outcome ln($wage$) for different sub-samples, specified in the column headings. A joiner is someone who is working at the pharmacy in year t but not in year *t-1*, while a stayer is someone who is neither a joiner in year t or *t+1* (see text for full details). Columns 3 and 4 break out the new hires into two mutually exclusive types: those coming from another pharmacy or those coming from non-employment. Controls are included for age, gender, foreign born, pharmacist, level of education, tenure and industry experience. Standard errors are clustered by LLM and reported in parentheses.

Table A.5: Wages – new hires vs. non-new hires

	(1)	(2)
Panel A: Firm hires		
Join firm	-0.064***	-0.047***
	(0.004)	(0.006)
Join firm × Post	0.070***	0.058***
	(0.008)	(0.004)
R^2	0.545	0.573
N	110,722	93,872
Year × firm FE	Yes	Yes
Panel B: Pharmacy hires		
Join pharmacy	-0.014***	-0.007**
	(0.003)	(0.003)
Join pharmacy × Post	0.026***	0.022***
	(0.008)	(0.007)
R^2	0.674	0.691
N	109,778	92,555
Year × pharmacy FE	Yes	Yes
Sample	Full	Incumbents
Controls	Yes	Yes

Note: The table compares wages for new hires and non-new hires in the same firm or pharmacy in the pharmacy industry. It provides estimates of θ_1 and θ_2 from estimating $\ln(w_{it}) = \theta_1 Join_{ipt} + \theta_2 Join_{ipt} \times Post_t + \lambda_{pt} + \beta X_{it} + \varepsilon_{it}$, where p is either the firm or the pharmacy. A joiner to a firm (pharmacy) is someone who is working at the firm (pharmacy) in year t but not in year *t-1* (see text for full details). In column (1) the full sample is included, while in column (2) only incumbents are included, defined to be employees who are working in the pharmacy industry in the pre-period. Controls are included for age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), tenure (in three categories) and industry experience (in three categories). Standard errors are clustered by LLM and reported in parentheses.

Table A.6: Heterogeneous effects of labor market concentration on ln($wage$) – tenure and experience

	(1)	(2)	(3)
Tenure (years):	*<3*	*3–4*	*≥5*
$ln(HHI_{m,2009}) \times Post_t$	-0.010	-0.036**	-0.049***
	(0.017)	(0.015)	(0.018)
Mean log wage (pre)	10.05	10.08	10.1
R^2	0.544	0.566	0.566
N	34,007	22,190	54,525
Industry experience (years):	*<3*	*3–4*	*≥5*
$ln(HHI_{m,2009}) \times Post_t$	-0.006	-0.042	-0.050***
	(0.032)	(0.027)	(0.013)
Mean log wage (pre)	9.88	9.93	10.12
R^2	0.502	0.542	0.554
N	9,375	12,200	89,145
Year FE	Yes	Yes	Yes
LLM FE	Yes	Yes	Yes
Controls	Yes	Yes	Yes

Note: This provides estimates of γ from equation (9) for different subsamples, specified in the column headings. Controls are age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), tenure (in three categories) (only in the bottom panel) and industry experience (in three categories) (only in the top panel). Standard errors are clustered by LLM and reported in parentheses.

Table A.7: Heterogeneous effects of labor market concentration on ln($wage$) – individual characteristics

	(1)	(2)	(3)	(4)	(5)	(6)
	Female		Foreign born		Young	
	no	*yes*	*no*	*yes*	*no*	*yes*
$ln(HHI_{m,2009}) \times Post_t$	-0.067	-0.047***	-0.042***	-0.037*	-0.045***	-0.042**
	(0.052)	(0.008)	(0.010)	(0.019)	(0.014)	(0.017)
Mean log wage (pre)	10.18	10.08	10.09	10.04	10.14	9.96
R^2	0.528	0.571	0.553	0.593	0.538	0.569
N	11,094	99,626	92,654	18,067	71,109	39,612
	Post secondary		Pharmacist			
	no	*yes*	*no*	*yes*		
$ln(HHI_{m,2009}) \times Post_t$	-0.042***	-0.052***	-0.046**	-0.043***		
	(0.016)	(0.011)	(0.019)	(0.007)		
Mean log wage (pre)	9.88	10.19	9.93	10.22		
R^2	0.480	0.391	0.493	0.400		
N	35,443	75,279	52,975	57,747		
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
LLM FE	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes

Note: This provides estimates of γ from equation (9) for different sub-samples, specified in the column headings. Young is defined to be someone below age 40. Controls are age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), tenure (in three categories) and industry experience (in three categories). Standard errors are clustered by LLM and reported in parentheses.

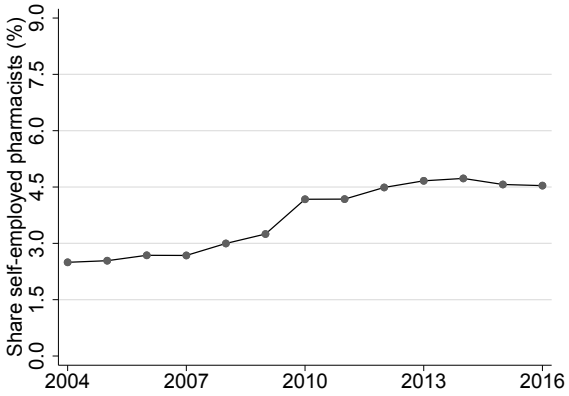
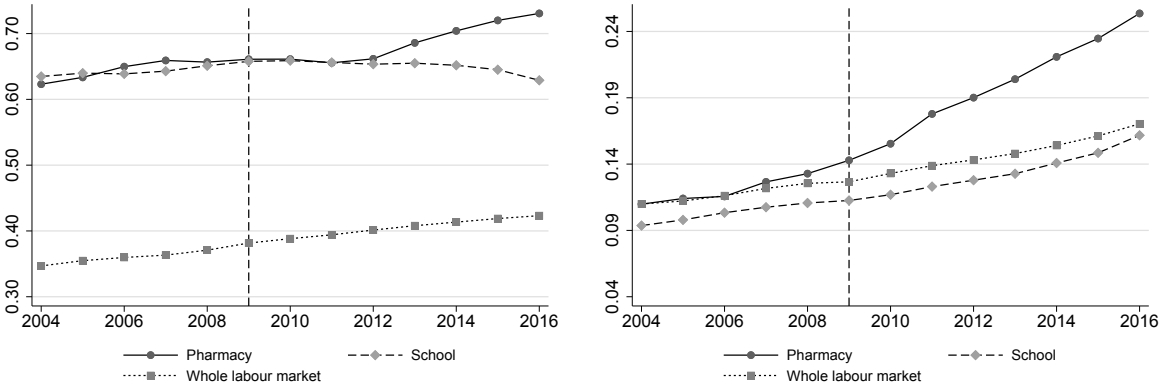


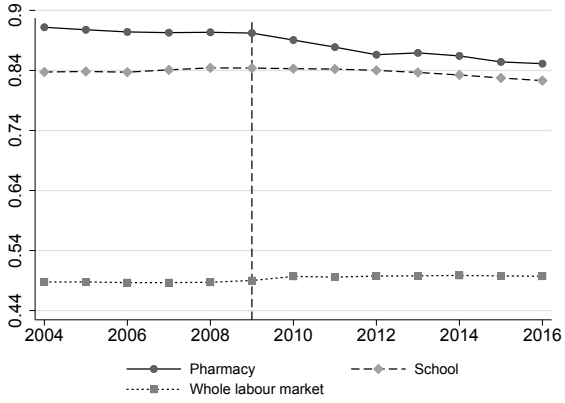
Figure A.1: Self-employment among pharmacists

Note: The figure shows the share of all educated pharmacists that are self-employed.



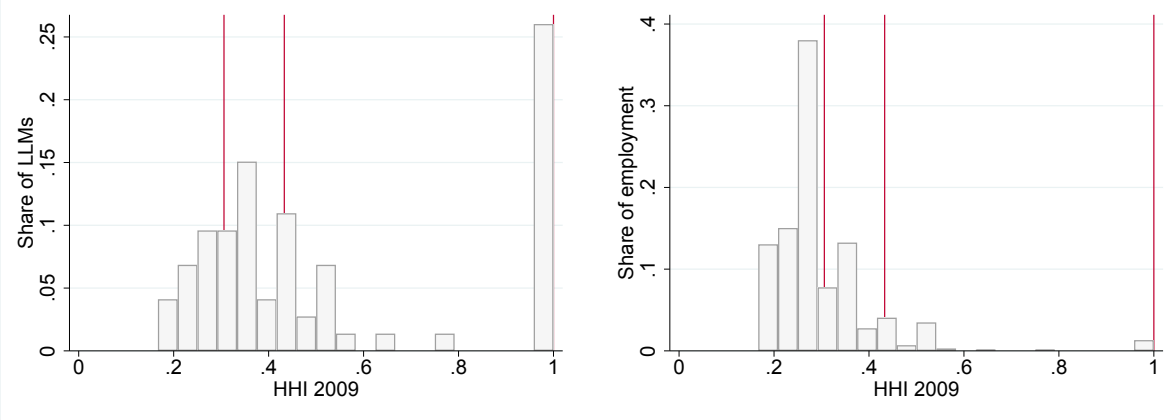
(a) Share post-secondary

(b) Share foreign born



(c) Share female

Figure A.2: Worker composition in pharmacy, school sector and whole labor market



(a) LLMs (b) Employment

Figure A.3: Distribution of $HHI_{m,2009}$

Note: The figure shows the distribution of $HHI_{m,2009}$ across LLMs (Panel A) and employees (Panel B). The three lines mark the 25th, 50th and 75th percentiles of the distribution of $HHI_{m,2009}$ across LLMs. They are 0.306, 0.433 and 1 respectively.

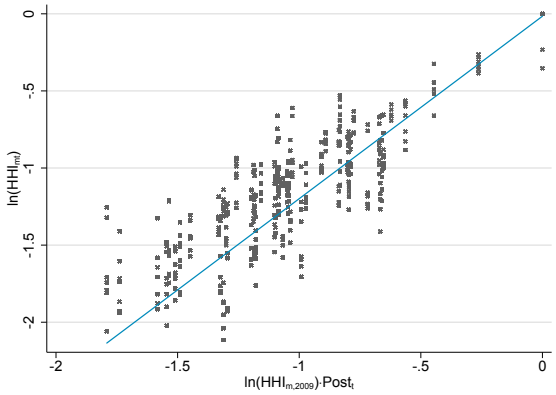


Figure A.4: Scatter of $\ln(HHI_{mt})$ vs. $\ln(HHI_{m,2009}) \times Post_t$

Figure A.5: Share new hires to pharmacy (pre-reform pharmacies)

Appendix B: Robustness checks

Table B.1: Robustness – Effect of labor market concentration on ln($wage$)

	(1)	(2)
Panel A: Baseline		
$ln(HHI_{m,2009}) \times Post_t$	-0.048***	-0.022***
	(0.011)	(0.006)
R^2	0.551	0.924
N	110,722	104,968
Panel B: Control for nr employees		
$ln(HHI_{m,2009}) \times Post_t$	-0.046***	-0.020***
	(0.010)	(0.006)
R^2	0.551	0.924
N	110,722	104,968
Panel C: Control for value added		
$ln(HHI_{m,2009}) \times Post_t$	-0.040***	-0.017***
	(0.011)	(0.006)
R^2	0.544	0.926
N	99,052	93,588
Panel D: Omit urban areas		
$ln(HHI_{m,2009}) \times Post_t$	-0.039***	-0.014*
	(0.011)	(0.008)
R^2	0.607	0.921
N	58,254	55,326
Panel E: No managers		
$ln(HHI_{m,2009}) \times Post_t$	-0.051***	-0.023***
	(0.010)	(0.006)
R^2	0.566	0.913
N	104,041	98,272
Panel F: Public sector only		
$ln(HHI_{m,2009}) \times Post_t$	-0.051***	-0.013***
	(0.019)	(0.004)
R^2	0.541	0.945
N	87,237	82,697

Continued on next page

Table B.1 – continued from previous page

	(1)	(2)
Panel G: Control for mean wage		
$ln(HHI_{m,2009}) \times Post_t$	-0.047***	-0.021***
	(0.011)	(0.006)
R^2	0.551	0.924
N	110,722	104,968
Year FE	Yes	Yes
LLM FE	Yes	
Person × LLM FE		Yes
Controls	Yes	

Note: This provides robustness checks for estimating equation (9). Column (1) controls for age, gender, foreign born, pharmacist, education, tenure and industry experience. Both columns control for log number of employees per LLM (Panel B), log value added per employee (Panel C), and mean log wages on the whole labor market by CZ and year (Panel G). Value added is available until 2015. Urban areas are LLMs that encompass Stockholm, Gothenburg and Malmö (the three largest cities in Sweden). There is some measurement error which reports a too high share in the public sector particularly between 2009 and 2011. Omitting these years leads to somewhat larger point estimates in absolute terms. Standard errors are clustered by LLM and reported in parentheses.

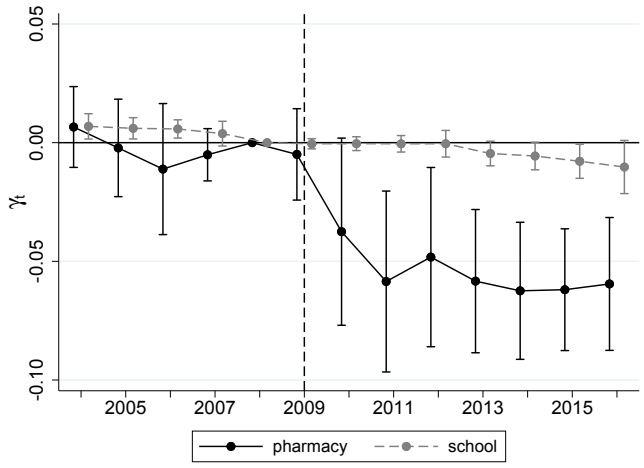


Figure B.1: Placebo analysis of the effect of market concentration on wages

Note: The figure plots γ_t from Equation (10) with 95% confidence intervals for regressions that are run separately for the pharmacy industry and for the school sector (preschools and compulsory schools). The outcome is log wages. Included in X_{imt} are age (in five categories), gender, foreign born and level of education (in five categories). Standard errors are clustered by LLM.

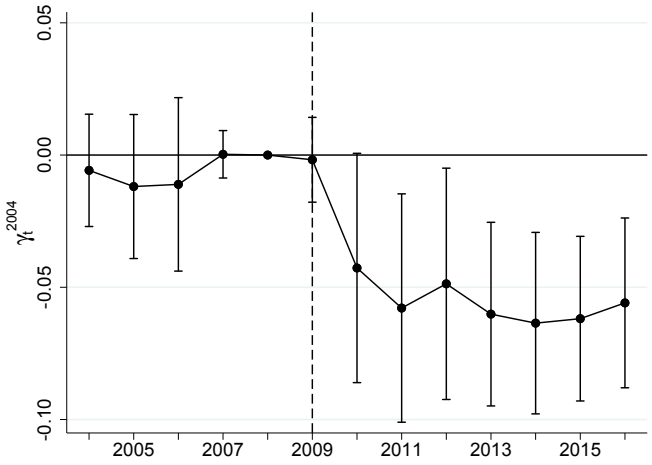


Figure B.2: The effect of $\ln(HHI_{m,2009})$ on $\ln(wage)$ – 2004 pharmacies

Note: The figure plots γ_t from Equation (10) with 95% confidence intervals. HHI is based on employment and pharmacies from 2004 and ownership structures from 2009. Included in X_{imt} are age (in five categories), gender, foreign born, pharmacist, level of education (in five categories), tenure (in three categories) and industry experience (in three categories). Standard errors are clustered by LLM.