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Imperfect Competition in Labor Markets: A Discrete Choice Approach to Wage-Setting Power with Evidence from Registered Nurse Employment in California

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Nikolaos Papadatos

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by Nikolaos Papadatos

Faculty Advisor: Martin Hackmann

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1 Abstract

Labor economics often assumes a fully competitive labor market where wages (w) equal the marginal revenue product of labor (MRPL). However, recent literature has shown that firms exhibit considerable market power which allows them to pay wages substantially below marginal productivity. The markdown, defined as $(\text{MRPL} - w)/w$, is used as the preferred measure of firms' monopsony power and captures the percent wage increase that would occur if monopsony power were eliminated. I derive the markdown for Registered Nurses working in hospitals across California using a discrete choice modeling framework inspired by the Industrial Organization literature. This approach falls within the class of job differentiation models, where monopsony power arises from workers' heterogeneous preferences over jobs that differ in wages and amenities.

I use detailed hospital-level data from California (2014–2023) to construct measures of wage premiums, effort levels, and market shares for nurses across Health Service Areas (HSAs). The labor supply elasticity to each hospital is estimated through a logit demand model for employment, where the outside option is implicitly captured by the market structure. I account for potential wage endogeneity using two instrumental variable strategies—licensed bed capacity and average rival hospital capacity within the HSA.

Results indicate a statistically significant elasticity of labor supply to wages of approximately 1.82 under OLS, increasing sharply to 17.07 in the IV specification. The latter implies a wage markdown of 6%, suggesting that hospitals pay RNs roughly 6% below their marginal product. While this markdown is lower than values typically reported in the literature (15–50%), it reflects moderate but economically meaningful deviations from perfect competition. Additional findings show that chain-affiliated hospitals are associated with higher effort intensity and lower perceived job utility.

The relatively modest markdown can be interpreted through two lenses: first, as a function of the limitations of hospital-level aggregate data, and second, as evidence of constrained monopsony power in a tightening RN labor market. Recent studies report high turnover and projected shortages in the California nursing workforce, potentially limiting hospitals' ability to suppress wages. Despite these limitations, this study shows the utility of adapting discrete choice models to institutional labor settings and highlights the value of empirically grounded monopsony measures for policy design. In a context where much of the available data is aggregated, such tools can inform the design of regulatory interventions aimed at improving wage-setting conditions and ensuring fair labor standards in imperfectly competitive labor markets.

2 Introduction

Economists have long argued that labor market outcomes exhibit substantial deviations from the prevailing competitive model, in which wages adjust to equate the supply and demand for labor. With the emergence of large-scale micro-level datasets and, in part, due to shifts in ideological norms within the economics profession, research over the past two decades has increasingly recognized that many firms exert some degree of wage-setting power (Card, 2022). This recurring empirical finding has renewed interest in monopsony models, labor market analogs of monopoly models on the demand side, which allow for imperfect competition in labor markets.

A key insight from the literature is that firm-level wage-setting power can arise not only from market concentration, which has traditionally been the focus of monopsony-inspired research, but also from job differentiation and heterogeneous worker preferences over employers. Discrete choice frameworks, widely used in industrial organization (IO), are increasingly adopted in labor economics to model the employment decision as a function of both wage and non-wage job characteristics, offering an insightful way to capture how firms become imperfect substitutes in the eyes of workers.

This study approaches the recent monopsony literature through the lens of IO, and more specifically, follows the modeling framework discussed in Card et al (2018), which provides a microeconomic foundation for imperfect labor market competition by allowing workers to have heterogeneous preferences over the work environment of potential employers. This framework is further developed empirically by Azar, Berry, and Marinescu (2022), who implement a parametric discrete choice model to analyze job application decisions and estimate labor supply elasticities. In this approach, heterogeneous worker preferences form the basis for imperfect competition in labor markets. Workers may accept wages below the competitive level when a job provides higher overall utility relative to their next-best alternative. As a result, a firm that marginally reduces its wage below competitive levels will lose only a fraction of its employees to other employers.

Using yearly data from California's Hospital Disclosure Data (CADD), I model employment, at a hospital level, for registered nurses (RNs) employed full time. As in IO models following Berry (1994), each choice (here, the hospital selected by RNs within a labor market) has both observed and unobserved characteristics. Following evidence on sources of between-hospital competition from Currie et al (2005), hospitals are observably differentiated by factors such as effort intensity, measured as the number of patients per RN, and chain status.

The data has a panel structure, where hospitals are observed across multiple years within local labor markets defined by Health Service Areas (HSAs). The share of RN employment at each hospital in a given HSA-year is interpreted as the outcome of a discrete-choice process, where RNs choose among available employers within their market.

The estimation proceeds in two steps. First, the mean utilities for each hospital are recovered using a logit demand inversion, computed as the log ratio of each hospital's market share to that of the outside option, which is defined as a randomly chosen hospital within an HSA and held constant over time. In the second step, the recovered utility terms are regressed on wage premiums, effort, and hospital characteristics. Wage premiums are calculated as the log difference between a hospital's wage and the wage of the outside option. All wage variables are expressed in real terms, adjusted to 2022 dollars. To address the potential endogeneity of wages, different instruments are employed, with the preferred one being a BLP-style instrument: the average licensed bed capacity of competing hospitals within the market. The estimated wage premium coefficient is then used to compute labor supply elasticities and corresponding wage markdowns, providing a measure of monopsony power in local hospital labor markets.

3 Literature Review and Conceptual Framework

3.1 From Perfect Competition to Monopsony

The notion that labor markets should be analyzed through the lens of imperfect competition has gained increasing attention in the literature over the past 15 years, starting with Manning (2011), who argued that wage-setting power is a pervasive feature of real-world labor markets and contended that adopting this perspective as a standard framework would fundamentally reshape how economists understand key policy topics such as wage inequality, employer-provided training, and the role of labor market regulation.

In the standard perfect competition model found in most textbooks, the labor market is governed by essentially the same forces of supply and demand as any other traded product, meaning that firms treat wages as an external constraint rather than a choice to be optimized. In a perfectly competitive market, a worker's elasticity of labor supply to a single firm is perfectly elastic at the market wage for that type of worker. Any attempt to pay a lower wage results in a complete inability to recruit workers, while any higher wage simply reduces profits. As a result, all employers hiring this type of worker pay the same wage, irrespective of each worker's reservation wage—naturally leading to an employment relationship with zero rents for both sides.

On a general note, imperfect competition implies that the employer, the worker, or both receive some rents from the existing employment relationship. If the employer receives rents, they are worse off when a worker leaves, that is, the worker's marginal product exceeds their wage, and replacement is costly. If the worker receives rents, then losing the current job makes them worse off because an identical job cannot be found at zero cost.

Even though rents can derive from both sides of the employment relationship, economic theory and evidence suggest that they accrue disproportionately to employers. Robinson (1933) laid out a model of a monopsonistic firm facing an upward-sloping labor supply curve, showing that the result is a markdown of wages relative to workers' marginal product of labor. Empirically, economists have long observed violations of the law of one wage (Lester, 1946; Slichter, 1950) and documented persistent industry wage premia (Krueger and Summers, 1988; Katz and Summers, 1989). More recent studies show that displaced workers suffer large and lasting earnings losses (Jacobson et al., 1993; von Wachter et al., 2009), while firms' hiring costs appear comparatively modest.

This view connects naturally to the industrial organization literature, which typically models markets as imperfectly competitive (Tirole, 1988; Pakes, 2016). As noted byCard et al (2018), "although economists widely accept that much of the variation in the prices of cars or breakfast cereal reflects factors beyond marginal cost, the idea that wages can differ substantially among equally skilled workers has remained surprisingly controversial".

For many years, the dominant explanations for such wage dispersion came from two theoretical traditions: search models and differentiated demand models. However, until the advent of large-scale matched employer–employee datasets, it was difficult to empirically separate worker-driven from firm-driven sources of wage variation. As a result, while the idea that firms could systematically pay different wages to similar workers was acknowledged, it remained largely theoretical.

Fueled by the availability of large-scale linked datasets, recent empirical literature has shown that firm heterogeneity plays an important role in wage determination (Kline, 2024). At the same time, there is consistent evidence that firms adjust wages in response to idiosyncratic productivity shocks (Lamadon, Mogstad, and Setzler, 2022; Garin and Silvério, 2023). As a result of the intensified effort to formalize and quantify the origins and magnitude of wage-setting power, Robinson's (1933) theory of monopsony has experienced a notable revival in both theoretical and empirical research (Kline, 2025).

3.2 Monopsony and Job Differentiation in the Nursing Market

This study focuses on an important theoretical framework within the monopsony literature: job differentiation. As discussed by Azar and Marinescu (2024), this framework suggests that monopsony power can arise from workers' heterogeneous preferences over jobs that offer different combinations of wages and non-wage amenities. Assuming that firms do not observe workers' idiosyncratic tastes for non-wage amenities, they cannot price discriminate based on individual reservation wages. Instead, firms post a single wage for each employee level, which attracts workers who find the combination of the posted wage and offered amenities appealing. If the firm wishes to hire more workers, it must raise wages for all employees at that level. Job differentiation thus provides a microeconomic foundation for a finite labor supply elasticity, which in turn allows for the estimation of wage markdowns.

FollowingCard et al (2018), who argue that it is plausible for firms to have more wage-setting power in some labor markets than others, this study focuses on hospital nursing staff in California, and particularly on registered nurses (RNs).

While early contributions theorized hospitals as classic monopsonists (Yett, 1970), empirical studies found limited evidence that hospital market power significantly depresses nurse wages (Sullivan, 1989; Hirsch & Schumacher, 1995). In their review of the literature, Boal and Ransom (1997) concluded that any monopsony power in the nurse labor market appeared modest.

More recent work has produced mixed results. Using data from the same 1989–1999 hospital disclosure files, Currie et al (2005) find that hospital takeovers and increasing market concentration were associated with greater nurse workloads but had little effect on wages, suggesting limited direct wage-setting power. Mukherjee (2011), using semi-parametric methods, finds limited support for monopsonistic exploitation, suggesting that limited worker mobility—rather than employer concentration—may better explain wage and workload outcomes. In contrast, Staiger, Spetz, and Phibbs (2010) exploit an exogenous wage change at VA hospitals and estimat a labor supply elasticity for RNs of just 0.1, implying significant monopsony power, at least in the short run.

This study empirically examines wage dispersion in the labor market for RNs using a static differentiated products modeling framework developed byCard et al (2018). In contrast to much of the rent-sharing literature, this framework assumes that employers set wages to maximize profits, subject to constraints on the relationship between wages and labor supply. The firm's ability to set wages arises from job differentiation.

4 Data

4.1 Hospital Employment Data and Variable Construction

The data used in this study are sourced from California's Hospital Disclosure Data (CADD) for fiscal years 2014–2022. CADD consists of information from non-federal hospital financial reports (disclosure reports), which are submitted annually to the Office of Statewide Health Planning and Development (OSHPD). These reports include information on hospital ownership, for-profit or non-profit status, number of beds, costs, revenues, and personnel. Hospitals also report productive and non-productive hours, as well as hourly wages, across several occupational categories, including Registered Nurses (RNs), Licensed Vocational Nurses (LVNs), aides and orderlies, supervisors, technical staff, clerical workers, and support service personnel. A final category includes all other roles, such as salaried physicians and non-physician medical staff. This study focuses specifically on the labor market for RNs, both due to the relevance of existing literature and the higher quality and consistency of RN-specific data in the CADD dataset.

To reliably estimate the value derived from working in a particular hospital, the sample is restricted to hospital-year observations reporting at least 2,080 productive hours, which is the equivalent of a full-time (FTE) RN, following the methodology of the Bureau of Labor Statistics (BLS, 2023). Observations with reported hourly wages below the California minimum wage for full-time employees are also excluded. All wage measures used in the analysis are expressed in real terms, deflated using annual Consumer Price Index (CPI) values from the BLS with 2022 as the base year.

Table 1 presents summary statistics for key hospital-level variables averaged across years. On average, the dataset includes 369 hospitals per year, with the number ranging from 352 to 399. Approximately 27% of hospitals are part of a larger chain, though this share varies from 23.3% to 31.8% across the sample. Following Currie et al (2005), I classify a hospital as part of a chain if its owner appears in the dataset as the owner of more than 10 hospitals.

The average real hourly wage for RNs, adjusted to 2022 dollars, is $56.78, with wages ranging from $23.42 to $101.58. The average volume of RN labor input, measured by total productive hours, is 386,448 hours per hospital-year. Dividing by the full-time equivalency threshold of 2,080 hours, this corresponds to an average of 186 FTE RNs per hospital. This figure ranges widely across institutions—from as few as 1 to over 1,600 FTE nurses—indicating substantial heterogeneity in scale.

Table 1: Summary Statistics of Key Variables (Averaged Across Years)

Metric	Mean	Sd	Min	25%	50%	75%	Max
Hospitals	369	20	352	353	357	393	399
% Chain	26.60	2.60	23.30	25.15	26.00	26.10	31.80
Hourly RN Wage	56.78	14.86	23.42	46.88	53.87	64.38	101.58
Productive Hours	386,448	466,428	2,694	78,145	217,101	548,766	3,383,190
Equivalent FTE RNs	186	224	1	38	104	264	1,627
Effort	6.02	12.04	0.38	2.33	2.87	4.88	143.73
Licensed Beds	214	176	10	78	165	316	1,023

In addition to wages, this study incorporates effort as a measure of job differentiation, reflecting the intensity of work that nurses experience across hospitals. This approach builds on the insights of Currie et al (2005), who highlight that within monopsonistic and contracting frameworks, hospitals compete not only on wages but also on working intensity conditions. They argue that effort, unlike wages, is rarely contractible and can be more readily adjusted in response to financial or organizational pressures, especially in the context of hospital mergers and takeovers. Their analysis shows that such institutional changes may lead to higher patient loads per nurse (their measure of effort), even when wages remain largely unaffected, underscoring the importance of modeling non-wage aspects of job quality in labor market analyses. Empirically, they define effort as the number of patients each nurse is responsible for during a 24-hour period. In the data employed for this study, we observe considerable variation in nurse workload across hospitals. As shown in Table 1, the effort measure ranges from 0.38 to 143.73, with a mean of 6.02, meaning that on average nurses are responsible for approximately six patients per 24-hour period.

As established in the literature, firm productivity and wages are strongly correlated, raising concerns about endogeneity in wage regressions. To address this issue, we instrument for wages using hospital bed capacity. The number of beds is plausibly correlated with wages, as larger hospitals tend to offer higher compensation, but conditional on observed measures of effort, chain status and year fixed effects, bed capacity is unlikely to directly influence a nurse's utility from the job. This satisfies the exclusion restriction and makes beds a credible instrument.

In our main specification, we use the number of licensed beds as the instrument, following Currie et al (2005), though we also explore specifications with alternative bed definitions, including staffed and available beds. In addition to using a hospital's own beds, we also construct a rival beds instrument, defined as the average number of licensed beds in competing hospitals within the same market and year. This BLP-style instrument captures variation in local market structure that influences wage-setting but is plausibly exogenous to any one hospital's employment outcomes. We report results using both instruments.

4.2 Defining Local Labor Markets and Geographic Variation

Defining the boundaries of local labor markets is a central challenge in empirical studies of labor market power. As emphasized by Staiger, Spetz, and Phibbs (2010) and Azar, Marinescu, and Steinbaum (2022), estimates of labor supply elasticity and wage-setting power are highly sensitive to how markets are geographically delineated. An overly narrow market definition may exaggerate monopsony effects by understating worker mobility, while an overly broad one may mask genuine frictions that restrict job switching across employers. Hence, the validity of monopsony models depends critically on choosing market boundaries that reflect actual search behavior and employer competition.

Table 2: HSA-Level Averages of Key Variables Across All Years

HSA	Hospitals	% Chain	Hourly RN Wage	FTE RNs	Effort	Licensed Beds
1	34	28.52	50.08	59	8.22	78
2	27	63.98	64.17	239	4.11	202
3	16	25.93	70.49	121	3.67	128
4	23	28.93	80.13	245	5.33	310
5	27	39.04	77.79	195	4.66	218
6	20	32.32	59.47	141	3.65	159
7	13	11.90	79.74	355	4.58	341
8	11	37.63	71.66	115	4.25	150
9	33	26.69	49.03	158	7.38	198
10	14	28.70	55.39	168	4.34	226
11	108	18.36	52.77	210	8.04	254
12	48	27.89	48.09	177	5.71	183
13	32	21.91	50.81	169	3.91	209
14	28	12.90	50.42	258	5.11	291

This study adopts Health Service Areas (HSAs) as the geographic unit of analysis to define local labor markets. HSAs are established by the U.S. National Center for Health Statistics and are designed to capture areas where residents typically receive most of their hospital care. Their construction reflects both patient flow data and health care infrastructure, making them an accepted proxy for hospital labor markets in the health economics literature (Currie, Farsi, & MacLeod, 2005).

Several key variables in this study are computed at an HSA level. First, hospital market shares are calculated within each HSA-year, reflecting the distribution of RN employment across hospitals in a given local labor market. Second, the outside option is defined as a randomly selected hospital within the same HSA, constant over time. The wage premium is then computed as the difference between each hospital's RN wage and the wage of this outside option. Additionally, a BLP-style instrument is constructed by averaging the number of licensed beds at competing hospitals within the same HSA-year.

Table 2 summarizes key hospital characteristics aggregated at the HSA level. Substantial heterogeneity is observed across the 14 HSAs in California. For example, the number of hospitals per HSA ranges from 11 to 108, while average RN wages range from approximately \$48 to \$80 per hour. Variation is also evident in staffing intensity, with average FTE RN counts and effort levels differing significantly across areas. This geographic dispersion supports the use of a market-level framework and motivates the modeling of employer choice as a function of both wage and non-wage attributes, conditional on local market structure.

5 Model and Estimation

5.1 Theoretical Framework

Hospital-level labor demand is modeled using a discrete choice framework in which registered nurses select among hospitals within local labor markets defined by HSAs. This setting captures firm-specific wage-setting power that arises from job differentiation and limited substitutability between employers. The outside option is defined as a reference hospital within each HSA, serving as the baseline for evaluating relative preferences.

Let $j \in \mathcal{J}_m$ index hospitals in HSA m, and let $j = 0$ denote the reference hospital. FollowingCard et al (2018), the indirect utility each worker i receives from hospital j in market m and year t is:

$$u_{ijmt} = \alpha \log(w_{jmt} - b_{mt}) + \beta \cdot \text{effort}_{jmt} + X'_{jmt}\gamma + \xi_{jmt} + \varepsilon_{ijmt} \tag{1}$$

where:

- w_{jmt} is the real hourly wage for RNs at hospital j,
- b_{mt} is the reference hospital wage at time t,
- effort_{jmt} is a measure of job intensity,
- X_{jmt} includes observed hospital characteristics,
- ξ_{jmt} is an unobserved job-specific utility component,
- ε_{ijmt} is an idiosyncratic Type I Extreme Value error.

Given posted wages and assuming that RNs are free to choose any hospital within their local market, the mean utility of working at hospital j, denoted by δ_{jmt}, is specified as:

$$u_{ijmt} = \alpha \log(w_{jmt} - b_{mt}) + \beta \cdot \text{effort}_{jmt} + X'_{jmt}\gamma + \xi_{jmt} \tag{2}$$

Assuming that ε_{ijmt} are iid and following McFadden (1973), RNs have multinational logit choice probabilities of the form:

$$\Pr\left(y_{ijmt} = 1 \mid \beta, \gamma, \text{effort}_{jmt'}, x_{jmt'}, \xi_{jmt'}\right) = \frac{\exp(\delta_{jmt})}{\exp(\delta_{0mt}) + \sum_{k \in \mathcal{J}_m} \exp(\delta_{kmt})}, \tag{3}$$

where $y_{ijmt} = 1$ if registered nurse i chooses to work at hospital j in market m at time t, and zero otherwise.

Using aggregated market shares (s_{jmt}), the choice probabilities become:

$$s_{jmt} = \frac{1}{M}[M \cdot \Pr\left(y_{ijmt} = 1 \mid \beta, \gamma, \text{effort}_{jmt'}, x_{jmt'}, \xi_{jmt'}\right)] = \frac{\exp(\delta_{jmt})}{\exp(\delta_{0mt}) + \sum_{k \in \mathcal{J}_m} \exp(\delta_{kmt})} \tag{4}$$

The observed market share of hospital j within HSA m and year t is defined as $\hat{s_{jmt}}$. The share of the reference hospital, $j = 0$, is denoted as $\hat{s_{0mt}}$. The systematic utility component $\hat{\delta_{jmt}}$ is recovered from observed market shares following the inversion approach of Berry (1994), which expresses utility differences relative to the outside option as:

$$\delta_{jmt} - \delta_{0mt} = \log\left(\frac{s_{jmt}}{s_{0mt}}\right) \tag{5}$$

This log-difference transformation linearizes the multinomial logit model in terms of observed shares and enables estimation using aggregate data.

Although the reference hospital is fixed within each HSA, its observed market share s_{0mt} varies across years due to changes in employment patterns and staffing levels. Consequently, the utility δ_{0mt} also varies over time. However, because δ_{0mt} is constant across all j within a given HSA-year, it can be absorbed by HSA-year fixed effects in estimation. Thus, substituting the utility specification into the Berry (1994) inversion yields:

$$\log\left(\frac{s_{jmt}}{s_{0mt}}\right) = \alpha \log(w_{jmt} - b_{mt}) + \beta \cdot \text{effort}_{jmt} + X'_{jmt}\gamma + \xi_{jmt} \tag{6}$$

5.2 Empirical Specification

Given the above framework, I estimate the following regression model:

$$\log\left(\frac{s_{jmt}}{s_{0mt}}\right) = \alpha \log(w_{jmt} - b_{mt}) + \beta \cdot \text{effort}_{jmt} + \gamma \cdot \text{Chain}_{jt} + \gamma_2 \cdot \text{Year} \times \text{HSA} + \varepsilon_{jmt} \tag{7}$$

where:

- The left-hand side is the log market share ratio between hospital j and the reference hospital $j = 0$ within HSA m and year t,
- α captures the semi-elasticity of hospital choice with respect to the wage premium, defined as the difference between hospital j's wage and the reference hospital's wage within the same market.
- β measures the effect of job intensity (effort) on hospital choice,

- γ and γ_2 are indicator variables for chain affiliation and HSA-year, respectively, controlling for time-invariant chain characteristics and aggregate temporal market shifts.

Market shares are defined as the proportion of total productive RN hours in a given HSA-year that are worked at hospital j:

$$s_{jmt} = \frac{\text{RN hours}_{jmt}}{\sum_{k \in \mathcal{J}_m} \text{RN hours}_{kmt}} \tag{8}$$

Identification comes from variation in wages and effort across hospitals within HSA-year markets, holding the reference hospital fixed. The inclusion of fixed effects ensures that utility shocks common across hospitals are differenced out.

This specification permits a structural interpretation of α as the semi-elasticity of hospital choice with respect to wages. Under standard discrete choice assumptions, this elasticity directly maps into a labor supply elasticity, from which we compute the implied markdown as a measure of monopsony power.

5.3 Wage Markdown and Labor Supply Elasticity

To quantify monopsony power, I follow the approach outlined in Azar, Berry, and Marinescu (2022), which derives the markdown (μ) from a simple model of a monopsonist employer. Let the marginal revenue product of labor be denoted by a, and the wage paid by the firm be w, which depends on the level of employment L. The firm's profits are given by:

$$\pi(L) = (a - w) \cdot L \tag{9}$$

Suppose the firm considers a wage change of Δw, inducing a change in labor supply ΔL. The resulting change in profits is:

$$\Delta\pi = \Delta L \cdot (a - w - \Delta w) - \Delta w \cdot L \tag{10}$$

The wage markdown is profitable if and only if this profit change is positive, i.e.,

$$\Delta L \cdot (a - w - \Delta w) > \Delta w \cdot L \tag{11}$$

Dividing both sides by wL gives:

$$\frac{\Delta L}{L} \cdot \left(\underbrace{\frac{a - w}{w}}_{\mu} - \frac{\Delta w}{w} \right) > \frac{\Delta w}{w} \tag{12}$$

Rearranging terms (and noting the inequality reverses due to the negative wage change), the condition becomes:

$$\frac{\Delta L / L}{\Delta w / w} < \frac{1}{\mu - \Delta w / w} \tag{13}$$

The left-hand side is the labor supply elasticity η. Thus, the critical elasticity for a markdown μ and wage change $\Delta w / w$ is:

$$\eta \approx \frac{1}{\mu - \Delta w / w} \tag{14}$$

This expression provides a benchmark for evaluating whether a firm could profitably exercise monopsony power in a given market. In line with antitrust guidelines, a 5% wage reduction is typically used to define the threshold for a “small but significant non-transitory reduction in wages” (SSNRW). Therefore, the critical elasticity below which such a markdown is profitable is determined using $\Delta w / w = 0.05$.

5.4 Instrumental Variable Strategy

A key concern in estimating equation (6) is the potential endogeneity of hospital wages. Unobserved amenities (e.g., job satisfaction, prestige, work culture) may correlate with both wages and the utility component ξ_{jmt}, biasing OLS estimates of α.

To address this issue, I implement a two-stage least squares (2SLS) strategy using an instrument derived from local labor market structure. Specifically, I instrument $\log(w_{jmt} - b_{mt})$, the wage premium relative to the outside option, using the **average number of licensed beds at competing hospitals** within the same HSA, excluding hospital j.

This approach follows prior work in product markets (e.g., Berry, Levinsohn, and Pakes, 1995) and its adaptation to labor settings by Azar, Berry, and Marinescu (2022). The intuition is that an increase in the licensed bed capacity of competing hospitals raises the local demand for RNs, thereby increasing wage pressure at hospital j, but does not directly influence a nurse's utility from working there beyond its effect on wages. The instrument captures variation in local competition and capacity that shifts relative wages while satisfying the exclusion restriction. Formally, the instrument is constructed as:

$$\text{RivalBeds}_{jmt} = \frac{1}{|\mathcal{J}_m \setminus j|} \sum_{k \in \mathcal{J}_m \setminus j} \text{Beds}_{kmt} \tag{15}$$

where Beds_{kmt} is the number of licensed beds at hospital k in the same HSA m and year t, and $\mathcal{J}_m \setminus j$ denotes all hospitals in the HSA excluding hospital j.

We then estimate the following first-stage equation:

$$\log(w_{jmt} - b_{mt}) = \pi_0 + \pi_1 \text{RivalBeds}_{jmt} + \pi_2 \cdot \text{effort}_{jmt} + X'_{jmt}\gamma + u_{jmt} \tag{16}$$

The second stage substitutes the predicted values $\log(\widehat{w_{jmt} - b_{mt}})$ into equation (6) to consistently estimate α. Standard errors are clustered at the HSA level to account for intra-market correlation.

This IV approach permits a causal interpretation of α as the semi-elasticity of hospital choice with respect to the wage premium, capturing how wage deviations from local norms influence RN employment decisions.

6 Results

Table 3 reports the estimated coefficients for different specifications of the empirical model introduced in Section 5. The dependent variable is the log odds of hospital employment, interpreted as the relative utility RNs derive from selecting a given hospital compared to an alternative employment option within the same local labor market. All specifications include year fixed effects and control for chain affiliation. Columns (1) and (2) report OLS estimates, while columns (3) and (4) use an instrumental variables strategy, leveraging variation in licensed and rival bed capacity, respectively.

Table 3: Estimates from Log-Odds Models of Hospital Employment Choice

	OLS		**IV**	
	(1)	(2)	(3) Own Beds	(4) Rival Beds
Wage Premium	1.819***	1.817***	17.066***	17.079***
	(0.229)	(0.275)	(0.845)	(2.501)
Effort	-0.0397***	-0.0456***	0.0032	0.0032
	(0.008)	(0.009)	(0.008)	(0.012)
Chain Status	-0.2181	-0.0642	1.391***	1.392***
	(0.146)	(0.109)	(0.102)	(0.257)
Year FE	✓			
Year × HSA FE		✓	✓	✓
N	3,320	3,320	3,320	3,320
R^2	0.208	0.537	0.829	0.711

Notes: Dependent variable is $\log(s_{jmt}/s_{0mt})$. All models include controls for effort and chain status. Columns (3) and (4) are instrumented using the number of licensed beds and the average number of beds at rival hospitals within the HSA, respectively. Standard errors clustered at the HSA level in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

In the specification presented in Column (1), the coefficient on the wage premium is estimated to be 1.819 (SE = 0.229), indicating that, holding other hospital characteristics constant, hospitals offering higher wages are more likely to attract registered nurses. Even after controlling for the substantial heterogeneity across HSAs in Column (2), the results remain nearly unchanged. These OLS estimates suggest that hospitals face upward-sloping labor supply curves for RNs with respect to wages, a finding consistent with previous work in the field (Staiger, Auerbach, & Buerhaus, 2011). However, market-level conditions may not fully account for variation in unobserved factors that influence perceived job utility.

Columns (3) and (4) present two-stage least squares (2SLS) estimates to address the potential endogeneity of hospital wages. Both instrumental variable (IV) strategies demonstrate strong first-stage relevance, with F-statistics well above the conventional threshold of 10. The validity of the exclusion restrictions is supported by institutional and theoretical considerations: while bed capacity affects hospitals' labor demand, it is unlikely to directly influence RN preferences once wages and other hospital characteristics are accounted for.

In Column (3), the number of licensed beds is used as an instrument, yielding a wage premium estimate of 17.01 (SE = 0.845). Column (4) employs the average licensed bed capacity of competing hospitals within the same HSA as an instrument, resulting in a similar estimate of 17.08 (SE = 2.501). Consistent with findings in the recent literature (Azar, Marinescu, & Steinbaum, 2022), these IV estimates are substantially larger than the corresponding OLS coefficients, suggesting that the OLS estimates are biased downward, likely due to unobserved hospital amenities or job characteristics that are positively correlated with RN utility but not fully captured by observed controls.

The coefficient on effort is negative and statistically significant in the OLS specification, consistent with prior findings that higher job intensity reduces job desirability (Currie et al, 2005). However, once wages are instrumented, the effort coefficient becomes statistically insignificant, suggesting that, conditional on exogenous wage variation, differences in patient load are not a primary determinant of RN hospital choice.

In addition, the coefficient on independent hospital status is positive and statistically significant in the IV models, indicating that independent hospitals provide, on average, greater utility to nurses compared to chain-affiliated hospitals. This result aligns conceptually with the findings of Currie et al. (2005), who show that chain hospitals tend to impose higher effort levels on staff, thereby reducing overall job utility

Monopsony Power

The estimated wage premium coefficient from the econometric model captures the elasticity of hospital employment shares with respect to the wage differential between a given hospital and the outside option within the same local market. By interpreting this coefficient as the labor supply elasticity for registered nurses, we can calculate the implied wage markdown. The markdown provides a structural measure of monopsony power in California hospital labor markets, indicating the degree to which hospitals are able to pay RNs wages below their marginal value of labor.

Using the IV estimates from Columns (3) and (4), which yield an elasticity of $\varepsilon = 17.0$, the implied markdown is:

$$\mu = \frac{17.0}{17.0 - 1} \approx 1.06 \tag{17}$$

implying a 6% wage markdown. These values suggest moderate but economically meaningful deviations from competitive wage-setting, consistent with a labor market characterized by imperfect competition.

In contrast, the OLS elasticity in Columns (1) and (2), $\varepsilon = 1.82$, yields a markdown of:

$$\mu = \frac{1.82}{1.82 - 1} \approx 1.55 \tag{18}$$

which would imply that wages are approximately 55% below marginal revenue productivity—a figure notably higher than those estimated in the literature, which range between 15% and 50% (Azar and Marinescu, 2024). This result underscores the upward bias introduced by endogeneity in the model and highlights the necessity of using valid instruments when estimating firm-level labor supply elasticities.

7 Discussion of Limitations

Our preferred specification results indicate a modest but non-negligible degree of wage-setting power in the California market for registered nurses. One way to interpret these results is by considering the limitations of our estimated model.

As with many studies in the monopsony literature, model misspecification has meaningful consequences, with more complex model specifications predicting lower markdowns. Naturally, all empirical results are shaped by the underlying assumptions embedded in the model.

One important assumption in our framework is that RNs can, without cost, move between hospitals within an HSA, and that their choice set includes only other acute-care hospitals in the same HSA. While this is econometrically justifiable, it is conceptually limiting. In reality, RNs often have a broader set of employment opportunities even within the same geographic area, including positions in nursing homes, outpatient clinics, schools, or physicians' offices, none of which are captured in our definition of the labor market or the outside option.

In addition, the use of the CADD data likely introduces measurement bias. This data are known to be noisy, and for the purpose of this study, I had to restrict the analysis to RNs in Daily Hospital Services with reliable wage and productive hours data. This excludes other relevant nursing activities such as ambulatory care, ancillary services, and research roles, as well as other types of nursing personnel, which may drive important within-hospital variation in wages and labor supply that our estimates cannot capture.

Finally, the CADD data are aggregated at the hospital level, not the individual level, which limits the extent of variation available for estimation and reduces the credibility of causal inference. In contrast, more recent and influential studies such as those by Azar, Berry, and Marinescu (2022) and Roussille and Scuderi (2023), rely on high-quality, matched employer-employee data, which allow for better modeling of heterogeneous worker preferences and labor supply responses. Richer micro-data would enable more precise estimates of hospital-specific monopsony power and a clearer identification of markdowns.

8 Conclusion

Wage-setting power in labor markets is both theoretically grounded and empirically well documented. The empirical literature has developed multiple approaches to quantify monopsony power, including measures based on labor market concentration, estimates of labor supply elasticities, and production function frameworks. Across the majority of studies, implied markdowns, the difference between what workers earn and what they would earn in a perfectly competitive labor market, range from approximately 15% to 50%. However, the estimated degree of monopsony power is highly sensitive to the definition of the labor market, the occupation under study, and the empirical strategy employed.

This study reports a markdown of 6%, a result significantly lower than the numbers seen above. Following the monopsony literature, this finding suggests that hospitals are able to exert wage-setting power, but to a limited extent. As discussed in Section 7, one way to approach this misalignment is through the limitations inherent in the empirical strategy, the most important of which is the reliance on aggregated hospital-level data rather than matched employer-employee data.

An alternative way to interpret this result is by situating it within the context of past literature. As discussed in Section 3.2, prior studies examining monopsony power in nursing labor markets, particularly through the lens of employer concentration, have produced mixed findings, with many suggesting that hospitals exert limited wage-setting power. Specifically for California, more recent labor market developments may help explain the relatively low markdown. Spetz, Chu, and Blash (2021, 2024) document a persistent and potentially worsening shortage of RNs in the state. Drawing on surveys through 2023, they report rising RN turnover, increasing burnout, and projected labor market tightening despite growth in nursing school graduations. These shortages have likely constrained hospitals' ability to suppress wages in recent years, limiting their monopsony power and contributing to the relatively modest markdown observed in our estimates.

Despite its clear limitations, the decision to examine monopsony through a discrete choice framework using aggregated data is, in my view, well justified. As discussed in both the introduction and the literature review, once the assumption of perfect competition in labor markets is relaxed, important policy implications emerge. Deviations from the competitive model open the door to interventions that can potentially lead to more Pareto-efficient outcomes. Manning (2011) and Azar and Marinescu (2024) offer concrete examples of how monopsony insights can inform policy decisions. It is increasingly evident that models capable of estimating wage-setting power should play a central role in evaluating firm mergers, setting minimum wages, and assessing other labor market regulations. Given that much of the publicly available data—particularly from government sources—are in aggregated form, as in this study, I believe there is strong public value in developing theoretically sound and empirically implementable models that allow researchers to extract meaningful insights from such data. These tools can help inform evidence-based policy aimed at promoting fairer and more efficient labor market outcomes both for firms, which need to design more effective strategies to engage in online labor markets and attract better employees, and for workers, who seek fair compensation. As David Card noted in his 2022 address to the American Economic Association, "if your employer set your wage, it is hard to believe that it is too high" (Card, 2022).

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