Minimum Wage Employment Effects and Labor Market Concentration*

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Abstract

This paper shows that more highly concentrated labor markets experience more positive employment effects of the minimum wage. In the most concentrated labor markets, employment rises following a minimum wage increase. The paper establishes its main findings studying the effects of local minimum wage increases on a key low-wage retail sector, and using data on labor market concentration that covers the entirety of the United States with fine spatial variation at the occupation level. The results carry over to the fast-food sector and the entire low-wage labor market, and are robust to using proxies of labor market concentration available for a broader range of industries, such as the number of establishments and population density. A model of oligosponistic competition can explain these effects: there is more room to increase wages in high concentration areas where wages tend to be further below marginal productivity. These findings provide evidence supporting monopsonistic wage setting as an explanation for the near-zero minimum wage employment effect documented in prior work.

Keywords: Minimum Wage, Monopsony, Oligopsony, Labor Markets

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1 Introduction

Many papers document the employment effect of the minimum wage (Neumark and Wascher (1992); Card and Krueger (1994); Dube, Lester and Reich (2010); Meer and West (2016); Jardim et al. (2018); Clemens and Wither (2019); Cengiz et al. (2019), to name only a few). Despite the volume of work there is still debate about whether there is, in fact, an appreciable disemployment effect of the minimum wage, with many studies finding null results while others show a negative employment effect of varying intensity.

A common explanation for the many null findings is the existence of substantial labor market monopsony or oligopsony (Manning, 2011; Naidu, Posner and Weyl, 2018). Under perfect competition, the expectation is a clear and unambiguous reduction in employment caused by a binding increase in the minimum wage. While firms' ability to adjust in other ways, e.g., by increasing prices for consumers, may buffer the employment decline, under perfect competition employment never *increases* as a result of the minimum wage. However, when one departs from the assumption of a perfectly competitive labor market, the minimum wage can increase employment, as Stigler (1946) noted three quarters of a century ago. This is particularly true in the case of labor market monopsony (Robinson, 1969). More generally, employment under oligopsony may increase with the minimum wage, and, when employment does decrease, it can be shown to fall by less under oligosponistic competition than under perfect competition (Bhaskar, Manning and To, 2002).

However, to date there has been no explicit empirical assessment of the monopsony explanation for the limited negative employment effects of the minimum wage. In this paper, we provide the first direct test for the mediating role of labor market concentration - a key source of monopsony and oligopsony power - on the minimum wage employment effect.[1] The prediction we test empirically is that the employment elasticity of the minimum wage is more

[1]See Azar, Marinescu and Steinbaum (2019) on the link between labor market concentration and monopsony power.

positive in cases of greater labor market concentration.

Such a direct empirical test of the monopsony minimum wage story has been partly hindered by lack of fine-grained data on local labor market concentration.[2] Traditional data sets on occupational labor markets (e.g., JOLTS, HWOL) are limited in their geographic granularity, making comparisons by degree of local concentration difficult. To overcome such difficulties, we employ several strategies and data sources to proxy for local labor market concentration in the sectors most affected by minimum wage increases. We first exploit an exceptionally rich data set from Burning Glass Technologies that contains the near universe of U.S firms' online job vacancy announcements from some 40,000 websites at the level of the county, month, and occupation (defined at a six digit standard occupational code level).[3] Based on these data, we can measure occupational labor market concentration using the standard Herfindahl-Hirschman index (HHI) defined over these job vacancies (Azar, Marinescu and Steinbaum, 2019; Azar et al., 2020; Azar, Marinescu and Steinbaum, 2022). We use this measure to study one of the two largest sectors hiring minimum-wage workers in the U.S., General Merchandise Stores (NAICS 452), where we show hiring regularly occurs online.[4] Leading firms in this sector, the largest sector in retail, include Walmart and Macy's.[5] The large fraction of very low-wage workers in the General Merchandise industry, the tendency of industry hiring managers to post jobs for these minimum wage-affected occupations online, and, the significant varia-

[2]Neumark and Wascher (2002) shows that the impact of the minimum wage on employment ranges from positive to negative across states and suggests that this variability may be related to different levels of monopsony power. Okudaira, Takizawa and Yamanouchi (2019) show using Japanese manufacturing data that the minimum wage employment effect varies with the gap between structural estimates of the marginal product of labor and the wage rate in a way consistent with the predictions of the monopsony model.

[3]An occupation-based definition of the local labor market has the advantage of comparing reasonably homogeneous jobs (Azar, Berry and Marinescu, 2019) while worker mobility across occupations is known to be low - 77% of workers when switching jobs stay in their six digit SOC (Schubert, Stansbury and Taska, 2022).

[4]Together, the accommodation/food service and the retail sectors are known to employ 50% of US minimum wage workers (Dube, Lester and Reich, 2010), and, as such have been the subject of much of the prior minimum wage research in the US. Restaurants are contained in the accommodation/food service industry and general merchandise stores are the largest employer in the retail industry.

[5]Typical general merchandise stores are discount stores like Target and Walmart, and department stores like Macy's and Kohl's. More specifically, general merchandise stores are defined as firms that sell a wide range of general merchandise except fresh, perishable foods, and, which have central customer checkout areas, generally in the front of the store, and may have additional cash registers located in one or more individual departments.

tion we find in the measured local labor market concentration for these occupations, make the setting an ideal one to test the predictions of presence of monopsonistic wage setting.

For the other large sector hiring minimum wage workers, food services, a good deal of the hiring is done offline, making the Burning Glass data potentially less suited for measuring labor market concentration.[6] Hence, to study the role of labor market concentration for the food services sector, and, then later for the entire low-wage labor market, we introduce two other strategies for measuring labor market concentration, based on the local number of establishments in the sector and density of population.

These multiple measures of labor market concentration paired with employment and earnings data for the universe of employees in the United States from the Quarterly Census of Employment and Wages (QCEW), available quarterly at the county and industry level, allow us to test whether the employment effects of the minimum wage are more positive in more concentrated local labor markets, respectively. To set the stage for the empirical analysis, the paper introduces a partial-equilibrium model of oligopsonistic competition in the labor market with strategic interactions and solves it numerically to lay out the central hypotheses relating to the impact of the minimum wage in different settings. In contrast to the textbook model of perfect monopsony, the model makes predictions for the impact of the minimum wage across labor markets with continously varying labor market concentration. The model's key prediction is that the employment effect of the minimum wage becomes more positive with increasing labor market concentration, because more concentrated markets have lower wages relative to marginal productivity. The positive labor supply response to higher wages explains why the minimum wage can increase employment in the most concentrated labor markets.[7] Hence, the

[6]In Section 3.1, we provide evidence from the data grounding the common intuition that restaurants are places where the canonical "Help Wanted" sign in the window is still a typical form of job advertisement (and/or that other forms of non-internet-based job posting, e.g. word of mouth search through current employees or worker drop-ins with paper resumes, are common).

[7]This holds as long as the minimum wage is not set too high relative to the market's productivity. Local minimum wages are assumed to be set at the same percentage of the prevailing wage in all labor markets: this implies that the minimum wage effect on earnings is the same in all markets. This is consistent with our empirical findings.

intuition from the perfect monopsony textbook case carries over to oligopsonistic competition.[8]

Our empirical results indicate that, consistent with the theory, more concentrated labor markets exhibit more positive employment effects from the minimum wage. This is true for a variety of specifications, in both the General Merchandise Stores and the Limited Service Restaurant ("fast-food") industries and more broadly, and for different proxies for the degree of labor market concentration. Our baseline results focus on the General Merchandise Stores industry using Burning Glass data to measure concentration via job vacancy HHI for the large, low-wage occupations in that sector. To give a sense of scale of our findings, a standard deviation increase in the measure of labor market concentration is associated with a significant 0.2 *increase* in the employment elasticity of the minimum wage in the General Merchandise Sector. The predictions of our model are obtained in both traditional two-way fixed effects regressions and event-study models that show parallel pre-existing trends.

We extend these baseline findings to the fast-food industry, and, then, to all low-wage workers, using two strategies. First, we replicate our general merchandise analysis using the inverse number of establishments in the county and industry (from the QCEW) as a proxy for labor market concentration. Using this proxy yields nearly identical findings for the General Merchandise Stores industry as when we use the aforementioned Burning Glass HHI, validating the approach. We then apply it to the fast-food industry, where we find the predictions of our model also hold. Second, we replicated the bunching research design introduced by (Cengiz et al., 2019), utilizing Current Population Survey (CPS) data. The authors use shifts in the density of workers' wage bins around the new minimum wage level in order to estimate the employment effects of the policy on all low-wage workers. We compare these employment effects of the minimum wage in areas with high vs. low average occupational HHI, and, also, separately in more vs. less densely populated areas (metro vs. non-metro). We show that our key result holds for the entire low-wage labor market, including all industries at the same time.

[8]A more elaborate general equilibrium version of our setup is introduced by Berger, Herkenhoff and Mongey (2022) to study the optimal minimum wage.

We find that in labor markets that are more concentrated or less densely populated, minimum wage increases lead to overall positive employment effects. Applying this method to either retail or restaurant sectors alone also yields consistent findings with the earlier analysis.[9] An additional advantage of Cengiz et al. (2019)'s approach is that one can see that, in more concentrated labor markets, employment increases are comparatively larger for jobs right above the minimum wage, supporting a causal interpretation of our findings.

The results are robust to a large number of sensitivity tests. One important concern is omitted variable bias: there may be variables correlated with concentration that causally modulate the impact of the minimum wage on employment. Two potential candidates arising from the existing literature are population density and productivity, which are both correlated with HHI and could also, in principle, modulate minimum wage effects directly. We find that results are robust to allowing for a different effect of the minimum wage by both population density and a productivity proxy (average weekly earnings in an area across all sectors). The analysis of the interaction of minimum wages with population density is interesting in its own right, since our results indicate heterogeneity in minimum wage employment effects by population density absent HHI considerations. Thus, like establishment density, this measure could be used as another proxy for labor market power in situations when more direct measures of labor market concentration, such as the HHI of vacancies, are not available. However, when HHI and these proxies are together allowed to modulate minimum wage employment effects, the interaction with the proxies becomes smaller and usually insignificant while the HHI interaction remains significant in the expected direction. This suggests that the HHI measure we used is a preferable measure of labor market power if available. Overall, this affirms our main finding that concentration itself has a significant modulating effect on the employment elasticity of the minimum wage, even after we allow other key variables to have their own modulating effects. However, it does not prove it beyond any doubt, since there could be some other omit-

[9] This analysis uses broader industry categories (taken from the Cengiz et al. (2019) work) than our main analysis. Retail is the broader industry "general merchandise" belongs to, and "restaurants" is the broader industry fast-food belongs to.

ted variables that play a role. Yet, even if this were the case, at a minimum our HHI measure and other proxies of concentration allow policy makers to predict the heterogeneous effects of a minimum wage increase.

Overall, then, our results indicate the size and sign of the minimum wage's employment effects vary substantially on the basis of underlying labor market concentration. There is no such thing as "the" employment effect of the minimum wage applicable everywhere. In our baseline estimates for the general merchandise industry, employment elasticities of the minimum wage are significantly negative with point estimates of approximately -0.2 for the 33% of least occupationally concentrated labor markets, approximately zero for the middle concentration group, and significantly *positive* with point estimates of approximately 0.35 for the 33% of most concentrated of labor markets (which are very concentrated by typical measures).

The paper makes several contributions to the literature. First and foremost, the paper assesses the hypothesis that the minimum wage has differential impacts in markets with varying degree of labor market concentration. While monopsonistic wage setting has been often put forth as a potential explanation for findings of zero or positive effects of minimum wages on employment (e.g Card and Krueger, 1994; Neumark and Wascher, 2002; Giuliano, 2013), this paper is the first to bring to bear data on labor market concentration and quasi-experimental changes in minimum wages to address this question.

The paper's findings shed a new perspective on existing findings in the substantial prior literature on employment effects of minimum wages. For our baseline results, the related own-wage elasticity of labor demand, which we also derive, spans the range of comparable elasticities estimated in the minimum wage literature (see Figure 7), with the 33% of least concentrated markets on the outer left (i.e. the most negative) of the prior work's estimate range, the 33% of most concentrated markets forming the outer rightward bound (most positive) of this range, and, the middle third's estimates being close to zero. This indicates that differences in the degree of monopsony power could play an important role in accounting for the divergent elasticity estimates in prior studies, drawn from a wide range of labor markets that, themselves, have differences in underlying market concentration.

The paper also contributes to several papers assessing heterogeneity in the impact of the minimum wage on employment. While some papers find heterogeneity by skill or labor market experience (Giuliano, 2013; Horton, 2018), others (e.g. Cengiz et al., 2019) do not. Cengiz et al. (2019) show more negative employment effects in tradable industries, consistent with Harasztosi and Lindner (2019). There is also prior literature using data-driven methods to show that minimum wage employment effects systematically vary across geographical labor markets in the US (e.g. Neumark and Wascher, 2002; Wang, Phillips and Su, 2019), but these papers do not empirically document the mechanism that leads to different effects across locales. Showing that the various empirical proxies of concentration in local labor markets studied here – including occupational HHIs, population density, and the number of establishments in a low-wage sector – can separately predict heterogeneous minimum wage employment effects is of interest in its own right.

Our paper is also related to an increasing empirical literature assessing the degree of monopsony or wage setting power in the labor market (e.g. Azar, Berry and Marinescu, 2019; Dube et al., 2020; **?**; Lamadon, Mogstad and Setzler, 2022). With respect to that literature, this paper provides compelling evidence that responses to a key labor market institution (the minimum wage) are influenced by the structure of the labor market. As such, the findings also help to further underscore the role of employer concentration in the labor market.

Last but not least, the paper has shown that a range of empirical proxies can be successfully used to study the differential impact of minimum wages by the degree of monopsony power in local labor market. Given not all sources of data may be available in each setting, these findings will be valuable to other researchers that seek to replicate our findings and study monopsony power in different countries, sectors, or time periods. Our approach cannot distinguish between leading sources of monopsony power, such as labor market concentration, job differentiation, and search frictions. However, Azar, Marinescu and Steinbaum (2019) show that labor market concentration and the elasticity of labor supply (which reflects job differentiation and search frictions) are strongly correlated across markets.[10] Hence, our different proxies are likely to be correlated with several underlying sources of monopsony power, and allow researchers to study how the effect of the minimum wage varies with the structure of the labor market.

The remainder of the paper proceeds as follows. Section 2 provides a brief review of the theoretical considerations motivating the analysis. Section 3 reviews our data and discusses the design considerations motivating our empirical strategy. Section 4 presents our main results and robustness checks. Section 5 concludes.

2 Theory

How does the impact of a minimum wage vary with the degree of concentration among employers in the labor market? To answer this question, we develop and numerically solve a model of Cournot competition among employers. Employers decide on their employment level, anticipating that it will affect the equilibrium wage in a market with a finite number of strategic employers. The purpose of the model is to determine how an increase in the minimum wage affects employment in markets with different levels of concentration. Therefore, in our model, labor markets differ only by the number of employers. Fewer employers yield a higher concentration (Herfindahl-Hirschman index, HHI), lower competition for workers, and lower wages. Consistent with our focus on local minimum wage changes, we assume that the minimum wage is set as a proportion of the equilibrium wage prevailing in each labor market before the minimum wage increase. In this model, the employment effect of the minimum wage becomes more *positive* with increasing labor market concentration. The intuition is simple: the more concentrated a market is, the lower prevailing wages are relative to marginal productivity, and the more room there is to increase wages without reducing employment.

The goal of our model is to generate testable empirical predictions, not to determine the optimum level of the minimum wage. Interestingly, a recent working paper by Berger, Herken-

[10] Job differentiation and search frictions, imply reductions in the wage elasticity of labor supply (e.g., they result in wage decreases having a limited effect on workers' quitting), as discussed in Marinescu and Posner (2019).

hoff and Mongey (2022) uses a model with similar assumptions to ours. Their more elaborate general equilibrium model has the purpose of determining the optimal level of the minimum wage. They also discuss the optimal minimum wage given the empirical results from this paper.

We now provide more details on the setup for our simple model. There are J firms competing in a labor market. Within the market, firms are of equal size because we assume they all use the same production technology; thus the firms are symmetric and concentration (HHI) is simply $1/J$. Firm j's revenues are $Y_j = L_j^{\alpha}$, where L_j is the amount of labor used by firm j and the price of the product that the firms sell is normalized to 1. We assume decreasing returns to scale $\alpha < 1$. All workers in the market are paid the same wage. There is a market-level inverse labor supply curve $w(L)$, which depends only on the aggregate level of labor employed by firms in the market. The inverse labor supply curve has constant elasticity η, so that $w(L) = L^{\frac{1}{\eta}}$. The market level wage $W(L)$ is given by:

$$W(L) = \max(w(L), w_{min}) \tag{2.1}$$

where w_{min} is the minimum wage. The market wage is thus the level determined by aggregate labor demand in the market, if this level is above the minimum wage, and otherwise the market wage is the minimum wage. The firm chooses its level of employment L_j to maximize profits (here revenue minus wage bill). We denote the total employment at all other firms in the market as L_{-j}. The firm's profit maximization problem is given by:

$$\max_{L_j} \pi(L_j) = L_j^{\alpha} - W(L_j + L_{-j})L_j \tag{2.2}$$

The first order condition is $\frac{\partial \pi(L_j)}{\partial L_j} = 0$. Note that this first order condition depends on the minimum wage because the market wage $W(L)$ depends on the minimum wage (see equation (2.1)). To solve for L_j numerically in Julia, we simulate $N = 50$ markets, so the most competitive market has 50 firms, the next most competitive has 49 firms, and so on. We set baseline parameters in the following way. Consistent with our assumption of decreasing returns to scale in the production function, we set $\alpha = 0.6 < 1$. We assume that $\eta = 1.6$, consistent with Azar, Berry and Marinescu (2019).[11]

We arrive at our solution in each oligopsonistic market by first setting the minimum wage to zero, and, solving initially for the level of employment in each of these 50 markets when assuming that each firm sets employment *as if* the market were perfectly competitive, i.e. without taking into account the effect of its own employment L_j on the equilibrium wage in the market W. Under this assumption, when there is no minimum wage, the solution L_j^* to the firm's optimization problem is:

$$W = \alpha (L_j^*)^{\alpha - 1} \tag{2.3}$$

This is the classic result that, under perfect competition, the wage is equal to the marginal revenue productivity of labor (MRPL). We can calculate total market employment by using equations (2.3) and (2.1). When the wage is equal to the MRPL, total employment in a market with J firms is thus:

$$L^* = \left(\alpha J^{1-\alpha}\right)^{\frac{1}{1-\alpha+\frac{1}{\eta}}} \tag{2.4}$$

As this depends on number of firms J, the total employment derived in the equation above is different in each market, as is $W(L^*)$.

Next we derive the oligopsonistic equilibrium employment in each market. The procedure we employ is the same whether the minimum wage is set to zero or a positive value. Specifically, we start with setting the employment level at the level we just derived in equation (2.4), $L_j = L_j^* = L^*/J$. We then iteratively let each firm adjust employment L_j to maximize profits (using the first order condition, which depends on the minimum wage), assuming other firms' employment is fixed; we do this until the value of L_j converges.

[11] In appendix Table A1, we show that the impact of the minimum wage on employment as a function of HHI is qualitatively similar when taking a higher or lower level of the labor supply elasticity η or of the production function exponent α: the employment impact of the minimum wage always increases with HHI.

In the absence of a minimum wage, we find that wages and total employment are lower the higher the HHI. The firm's markdown, or Arthur Pigou's rate of "exploitation" (marginal revenue product minus wage, relative to the wage; Boal and Ransom (1997)) is $HHI \cdot \eta^{-1} = 1/(\eta J)$, which is clearly increasing in HHI.[12] However, the rate of exploitation does not pin down wages, since the marginal revenue product of labor also changes with HHI. In Figure 1, panel A, the oligopsonistic level of market-level employment (in dash dot dot purple) decreases with HHI. High HHI markets have fewer firms, each with higher employment (though not high enough to jointly offset the lost employment from the missing firms). Since we assume decreasing marginal productivity of labor ($\alpha = 0.6 < 1$), higher employment at a firm implies lower marginal productivity. Because of lower marginal product, wages in high HHI markets would be lower *even if* they were set equal to the marginal revenue product of labor. And, since firms paying lower wages induces lower labor supply, market employment still decreases with HHI even when wages are set to MRPL (see the brown dotted line in Figure 1, panel A), though less so than under the oligopsonistic level of employment that takes into account the further markdown in wage (the dash dot dot purple line in Figure 1, panel A).

Now, considering the role of minimum wages, we assume that the local minimum wage is set as a fixed percent $x\%$ above the equilibrium oligopsonistic wage *in each market*: therefore, the minimum wage is more likely to be above marginal productivity in lower HHI markets where the equilibrium oligopsonistic wage is closer to marginal productivity. This assumption is plausible to the extent that policymakers look to local wages to set a reasonable minimum wage for the area. By assumption, the local minimum wage increases wages by the same percent, $x\%$, in all markets, regardless of HHI; this assumption is therefore testable in the data.

In our model, we simulate a minimum wage set first at a low level: at 3% higher than the

[12]To see why this is the case, start with the first-order condition of firm j: $\alpha L_j^{\alpha-1} - W(L) - W'(L)L_j = 0$. Dividing by the wage, we obtain $\frac{\alpha L_j^{\alpha-1} - W(L)}{W(L)} = \frac{W'(L)L_j}{W(L)}$. The left-hand side is the firm's markdown. Multiplying and dividing by the aggregate labor supply L on the right-hand side, we obtain $\frac{W'(L)L}{W(L)}\frac{L_j}{L}$, which is the inverse elasticity of labor supply times the labor market share of firm j. With equal sized firms, this is equal to $1/(\eta J)$, which is the same as HHI/η, because $HHI = 1/J$.

prevailing oligopsonistic wage in each market. Then the minimum wage is increased by a larger but still modest 12% and finally the minimum wage is increased by 40% (closer to the large increases advanced in recent years). The minimum wage effect on employment always (weakly) increases with HHI (i.e. as HHI goes up, fewer jobs are lost/more jobs are gained for a given minimum wage increase). There are two distinct cases worth exploring: a low or medium minimum wage (3% over oligopsony wages, then 12% higher than the prior minimum wage) and a high minimum wage relative to prevailing wages (40% higher than the medium level of the minimum wage).

In Figure 1, panel A, we show the equilibrium level of employment without a minimum wage, and at the different levels of minimum wage. In Figure 1, panel B, we show the change in log employment when going from the oligopsony wage (no minimum wage) to a low minimum wage (MW), then from this low minimum wage to a medium minimum wage, and finally from that medium minimum wage to a high minimum wage. For the first two increases, the minimum wage is low enough that it induces employment gains in all but the lowest HHI (most competitive) markets: we have a positive employment effect, the size of which is determined by the labor supply elasticity (a higher labor supply elasticity generally induces a stronger employment increase for a given minimum wage increase, see Appendix Table A1). Overall, in the low or medium minimum wage case, the employment effect increases with HHI, and then is constant with HHI (Figure 1, panel B). In the high minimum wage case, the minimum wage is substantially above marginal productivity in the lowest HHI/most competitive markets, and we observe negative employment effects there. In this case, employment effects become less negative the higher the HHI, i.e. they increase with HHI (green dash dot line in Figure 1, panel B).

To summarize, if the minimum wage is not too much above the equilibrium oligopsonistic wage (i.e., when x% is not *too* high), employment effects become positive in the highest HHI markets; for very large minimum wage increases, employment effects are negative across the board, but closer to zero in high HHI markets. Overall, employment effects of the local minimum wage always increase in HHI, because local minimum wages are further below marginal productivity in high HHI markets. Appendix Figure A1 shows that a higher minimum wage reduces the equilibrium markdown at all levels of the HHI, and this reduction in the markdown is (weakly) higher in concentrated markets than in less concentrated markets.

In Appendix A.1, we discuss the case where the minimum wage is set at the same level in all markets, rather than being proportional to prevailing wages in each market. This case does not correspond to the local minimum wage variation we use in this paper, but it is relevant for a national minimum wage. We find that, as long as the minimum wage is not too high relative to marginal productivity in *high* HHI markets, the employment effects of the minimum wage are more positive in high HHI (less competitive) labor markets.

The key takeaway from our model is that, when local minimum wages are set proportionally to prevailing wages, the minimum wage has the same effect on wages in all markets, and has a more positive employment effect in higher HHI (that is, less competitive) labor markets. In the next section, we discuss the data we use to test these predictions.

3 Data and Design

3.1 Data

Our outcome measure for our main findings – industry-level employment in the general merchandise store sector – comes from the Quarterly Census of Employment and Wages (QCEW), a widely used data set in the minimum wage literature. From the QCEW we also get industry-level earnings in the sector. The publicly available QCEW used by most researchers contains a near census of quarterly, county-level payroll data by detailed industrial classification, including employment counts and average weekly wages.[13]

[13] The underlying micro data in the QCEW is at the establishment level and based on a mandatory survey of all employers participating in the unemployment insurance system. The U.S. Bureau of Labor Statistics publishes a county-level file available on the world wide web at https://www.bls.gov/cew/.

Additional control variables, including the log of county total population, log of total average weekly earnings (across all sectors) in the county, and the log of total employment (across all sectors) in the county, also come from the QCEW, while the log of the county unemployment rate is taken directly from the Bureau of Labor Statistics. Minimum wage data includes all federal, state, and county-level minimum wage changes. In our empirical strategy, federal minimum wage variation would be absorbed by fixed effects, but, as it stands, there is no change in minimum wage at the federal level during our time period of study, Thus, we only use state and county-level variation in the minimum wage.

Job postings data, used to measure the degree of labor market concentration, comes from Burning Glass Technologies and covers the near universe of online US job vacancy postings (culled from some 40,000 websites). This data has recently been used in Azar et al. (2020), Hershbein and Kahn (2018), Deming and Kahn (2018), Modestino, Shoag and Ballance (2016). Importantly, BGT data is fairly similar in terms of industry composition to all vacancies recorded in the nationally representative Job Openings and Labor Turnover Survey (JOLTS) (Hershbein and Kahn, 2018). Furthermore, the occupational distribution in BGT data is similar to the one found in the Occupational Employment Statistics (Hershbein and Kahn, 2018).

The job postings data is cleaned by Burning Glass to remove vacancy duplicates and extract key characteristics for each vacancy. Of interest to our work are the location of the vacancy (county), the time of the initial posting for the job, the name of the employer, and the occupation (categorized by a six digit standard occupation code (SOC) identifier). The name of the employer is normalized by BGT so that similar employer names are grouped together into a single employer: for example, "Bausch and Lomb", "Bausch Lomb", and "Bausch & Lomb" would be grouped together. Still, 35.9% of employer names are missing, partly due to staffing companies not disclosing on whose behalf they are posting a given job. To calculate concentration, we will assume that all these missing employer names are different, thus providing a lower bound for labor market concentration. We utilize uninterrupted data from the first quarter of 2010 until the last quarter of 2016, as there is a gap in the data in 2009.

3.2 Design

When seeking to test for differential employment effects of a minimum wage policy in more vis-a-vis less concentrated labor markets, an ideal setting is one where the following criteria are satisfied about the occupation and the industry:

1. the occupation earns a low median wage (i.e. the minimum wage "bites" from the workers' perspective)

2. the occupation is a sizable fraction of the associated industry employment, implying the associated industry is a low median wage industry (i.e. a minimum wage increase is meaningful, or "bites", from the firm's perspective so that any firm-level employment effects are detectable)

3. the industry primarily recruits online (so that the Burning Glass data can yield a valid measure of HHI)

4. the labor market has enough natural variation across the country in HHI to include both concentrated and competitive occupational labor markets (for meaningful differences in underlying market concentration)

This set of criteria can be summarized in the following statement: an ideal design will study a (near) minimum-wage-earning occupation in a low-wage industry that primarily uses online advertisements to fill jobs and that has a range of high and low labor market concentration levels.

Previous work serves as a guide in choosing the ideal setting. It is well documented that the industries that most intensively use a near-minimum wage workforce are the accommodation/food service and the retail sectors, accounting for 50% of all employees in the US who are paid within 10% of the minimum wage (Dube, Lester and Reich, 2010). The largest (by number of employees) 3-digit NAICS categories within the food & accommodation and the retail sectors are the Food Service and Drinking Places sector (NAICS code 722000) and the General Merchandise Stores sector (NAICS code 452000), respectively.

The two sectors do not, however, use the internet to recruit for low wage jobs at the same rates. As a benchmark for what is typical usage of online recruitment, we take the ratio of Burning Glass Technologies (BGT) job postings for an occupation over the total stock of jobs in the occupation (available from the Occupational Employment Statistics, OES). For the economy as a whole we have a median ratio of 0.09. For the key minimum wage earning occupations in the food service sector (contained in SOC 35-3xxx) the ratio is about half the size (0.05), while, for the key minimum wage earning occupations in the General Merchandise sector (contained in SOCs 43-5xxx and 41-2xxx) this ratio is 0.08 and 0.09, respectively, suggesting more typical rates of online hiring, and, making the Burning Glass data more suitable for study of the latter sector.

As such, we focus on the main minimum wage-heavy occupations in the General Merchandise sector - stock clerks and order fillers (SOC 43-5081), retail salespersons (SOC 41-2031), and cashiers (SOC 41-2011) - which each represent between roughly 20-30% of the General Merchandise work force, and, cumulatively, 65% of all employees.[14] The Burning Glass data, as we later show, also reveals that these occupations exhibit significant spatial variation in labor market concentration levels, with many high and many low concentration areas, an essential feature for answering the question at hand.

To demonstrate this final point, we require greater precision regarding what constitutes high vs. low labor market concentration. This in turn necessitates a measure of concentration, and, for this we rely on the standard definition of the Herfindahl-Hirschman Index (HHI). We calculate concentration at the level of the occupation (SOC-6) by county, which means that we include in our concentration measure job postings for a given occupation from all posters (of whatever industry). Our HHI, is thus:

[14] This is based on the 2016 National Industry-Specific Occupational Employment and Wage Estimates from the Bureau of Labor Statistics for the 452000 "General Merchandise Store" NAICS. Within the industry, stock clerks and order fillers, retail salespersons, and cashiers comprise 17%, 27%, and 21% of total employees, respectively.

$$HHI_{m,t} = \sum_{j=1}^{J} s_{j,m,t}^2$$

where for any firm j in a given occupational labor market m (i.e. occupation by county) and at quarter t, its market share s of job postings for the occupational labor market is defined as the sum of Burning Glass vacancies posted by the firm in the market and quarter divided by the total number of such vacancies posted in that market and quarter.[15] While all labor market definitions will be incomplete in some fashion (either overly broad or overly restrictive), Schubert, Stansbury and Taska (2022) find that the average probability of a worker staying in her 6-digit occupation when she leaves her job is 77%, making the 6 digit SOC a reasonable boundary.

By construction, HHI measures run from 0 to 1, with larger values indicating greater concentration. By Department of Justice/Federal Trade Commission 2010 horizontal merger guidelines, an HHI above 0.15 is considered "moderately concentrated" and an HHI above 0.25 is considered "highly concentrated." As mentioned, an ideal design would include a significant fraction of labor markets above and below the highly concentrated 0.25 threshold. In fact, the three SOCs we work with all have distributions for our HHI measure with coverage throughout the HHI scale's range, so that they are not too skewed in the direction of perfect competition or monopsony and allow for considerable natural variation in labor market concentration. The full distribution of these SOCs' HHIs is presented and discussed in the following section.

In our analysis, we use the average HHI across all available quarters of data in the sample period to give a broadly representative characterization of the underlying labor market concentration. A natural concern this raises, however, is that this may introduce endogeneity if minimum wage changes affect HHI. However, the data suggest this is not the case, as there is no significant relationship at all between minimum wage and HHI level (either at the quarterly HHI level or for the average of quarterly HHIs we use).[16] Furthermore, if we use the first two

[15] Azar et al. (2020) discuss in more detail the construction of this measure and its interpretation.

[16] See Appendix Table A2, discussed further in the next section.

years (2010-2011) to construct an alternative, baseline average HHI measure for each county and then run the equivalent specifications presented below using this alternative HHI measure, but, with only 2012-2016 data for all other variables (thus, eliminating the concern that the analyzed period's minimum wage changes may affect this HHI, since it is defined over the pre-analysis period) we get very similar results to our main results, both qualitatively and in terms of statistical significance.

4 Results

4.1 Main Results

Before turning to the main results on employment, we begin by exploring the "first-stage" result of the minimum wage's effect on earnings. A positive and significant earnings elasticity (with respect to the minimum wage) for workers in the general merchandise store sector is to be expected given the large fraction of workers in the industry in very low wage occupations who are likely to have their wages increased as the minimum wage goes up (in particular, the 65% of the workforce referenced in the previous section who work as stock clerks and order fillers, retail salespersons, and cashiers). However, confirming this empirically is important for the interpretation of later estimates of the minimum wage employment elasticities, as we are interested in explaining real employment responses to *binding* minimum wages (a zero employment elasticity of the minimum wage means something entirely different in an industry where there is no detectable earnings elasticity to begin with vis-a-vis one with a significant earnings elasticity). [17]

Table 1 reports these minimum wage earnings elasticity estimates for the general merchandise sector. The three columns all have the same outcome: average monthly earnings in the general merchandise sector. Each column reflects one of the three alternative baseline samples

[17] A similar need for a research setting where the studied minimum wage increases have bite (are binding) has largely motivated the focus in prior work on populations of restaurant workers, retail workers, or teen workers.

subsequently used (each constructed using one of our three distinct SOCs in the industry to compute the HHI measure, thus, yielding somewhat different sample sizes, as explained further in the discussion of Table 2 that immediately follows). The general finding in Table 1 is a positive and significant (at the 1% level) earnings elasticity with respect to the minimum wage, with a point estimate between 0.09 and 0.10; minimum wage increases are found to yield significant earnings increases in the sector, indicating significant bite of the minimum wage for general merchandise stores. Thus, as expected, the sector is confirmed to be an appropriate setting to study potential employment responses by firms to (actual) minimum-wage-induced increases in labor costs. Additionally, we find no significant difference in the minimum wage earnings elasticity estimates across high and low levels of the market concentration measure introduced in Table 2, as can be seen in Appendix Table A3. The fact that the elasticity of earnings with the minimum wage is roughly the same across HHI is consistent with our theoretical assumption that the minimum wage is set as a fraction of the prevailing wage in each market, with this fraction being the same across markets with different HHI (see section 2).

Turning to the study of the minimum wage's employment effects, we outline our baseline specifications in greater detail. As above, all regressions are performed at the county-quarter level of observation. Our left hand side variable is now the log of general merchandise store employment in the county-quarter. While the log of the governing minimum wage level remains a primary regressor of interest, given our research question, we are especially interested in the interaction of this variable with our measure of labor market concentration, HHI. We construct HHIs for each of the three occupational labor markets identified in the previous section as being large (in their share of total sectoral employment) and minimum-wage heavy occupations: stock clerks and order fillers, retail salespersons, and cashiers. The distribution of county average HHI measures (averaged for each county across the 2010-2016 sample period) is presented for each of the occupational labor market in Figures 2 - 4. As each of these labor markets have distinct variation in HHI - while still each being large relative to the sector's total employment, so that any resulting changes in occupational employment induced by the minimum wage should yield visible changes in total employment - we separately estimate regressions with log minimum wage interacted with each occupational HHI. The basic linear specification takes the form

$$ln(Y_{it}) = \alpha + \beta ln(MW_{it}) + \psi HHI_i + \delta ln(MW_{it}) * HHI_i + \phi ln(TotEmployed_{it}) +$$

$$\zeta ln(TotEarnings_{it}) + \eta ln(Pop_{it}) + \tau_t * ln(Unemp_{it}) + \gamma_i + \tau_{ct} + \chi I_s * t$$

where the outcome $ln(Y_{it})$ is the log of general merchandise store employment in county i and quarter t, $ln(MW_{it})$ is the log of the governing minimum wage, HHI_i is the labor market's average HHI over the sample period, and, the interaction term $ln(MW_{it}) * HHI_i$ is the main variable of interest. Additional control variables include the log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate, and county fixed effects γ_i, census division specific quarter fixed effects τ_{ct}, and state-specific linear time trends $\chi I_s * t$. Standard errors are clustered at the state level.

As an alternative specification we also include a binary HHI variable that separates high and low concentration labor markets using the Department of Justice/Federal Trade Commission 2010 horizontal merger guideline threshold of 0.25 for highly concentrated markets.

$$ln(Y_{it}) = \alpha + \beta ln(MW_{it}) + \psi \mathbb{1}(HHI_i \geq 0.25) +$$

$$+ \delta ln(MW_{it}) * \mathbb{1}(HHI_i \geq 0.25) + \phi ln(TotEmployed_{it}) + \zeta ln(TotEarnings_{it})$$

$$\eta ln(Pop_{it}) + \tau_t * ln(Unemp_{it}) + \gamma_i + \tau_{ct} + \chi I_s * t$$

Table 2 presents the results using both the linear (odd numbered columns) and binary specifications (even numbered columns) for the HHI. In Columns 1 and 2, we present estimates using the HHI measure defined over the stock clerks and order fillers labor market (SOC 43-5081). Columns 3-4 and columns 5-6, respectively, present estimates using the HHI measure defined over the retail sales labor market (SOC 41-2031) and the cashiers labor market (SOC 41-2011).

Across specifications and across each of the HHIs defined over the three alternative occupational labor markets, Table 2 shows negative and significant point estimates on the main effect of the minimum wage employment elasticity (see the first row), which corresponds to a low level of labor market concentration. In odd numbered columns, this estimate characterizes the employment effects of the minimum wage in extremely *un*concentrated labor markets (i.e. those approaching the competitive ideal of an infinite number of small firms posting advertisements in equal number for stock clerk jobs, for instance, to take the example of column 1). In even numbered columns, this estimate characterizes the average employment elasticity for those occupational labor markets below the DOJ-FTC-threshold for concentrated markets (HHI below 0.25). This negative point estimate is consistent with the standard prediction, based on a competitive labor market setting, of employment reductions resulting from minimum wage increases. Quantitatively, we find that in these more competitive labor markets, a 10% increase in the minimum wage reduces employment in general merchandise stores by just under 2%.

Across the second and third rows of Table 2, however, there is a consistent finding of a statistically significant and positive log minimum wage-HHI interaction term. This indicates that the more concentrated the occupational labor market is, the more positive the employment elasticity of the minimum wage becomes. To give a sense of scale for the estimates in the second row, taking column 1 as an example, a standard deviation increase in the stock clerk labor market HHI makes the employment elasticity increase by 0.21 (for columns 3 and 5, the equivalent increase is 0.22 and 0.20). The third row estimates indicate the increase in the employment elasticity for those occupational labor markets with HHIs above the DOJ-FTC concentrated markets threshold: the elasticity is higher by about 0.3 or 0.4. In all cases, these estimates indicate a common finding: as labor markets become more concentrated, the disemployment effect from the minimum wage is significantly reduced.

Since we find positive minimum wage employment effects in some markets and negative effects in others, one might wonder about the overall employment effect in our sample. In Appendix Table A4, we show that the overall employment effect is not significantly different from zero, whether or not we weigh counties by their initial employment level. Thus, accounting for heterogeneity in the minimum wage's impact by labor market concentration uncovers effects of the minimum wage (in some places) that would otherwise be masked in aggregate.

Figure 5 presents the estimated employment elasticities for the high and low concentration areas (again, using the 0.25 HHI threshold) that result from the binary specifications in Table 2. As can be seen across occupations, when moving from low concentration areas to high concentration areas, the employment elasticity changes dramatically, going from approximately -0.2 to a *positive* point estimate, though it is only statistically significant at the 5% level for the retail sales occupation.

Table 3 and Figure 6 make the pattern even clearer. Here we separate labor markets into HHI terciles, rather than simply binary bins (otherwise following the specifications of Table 2). As can be seen across the columns (again, corresponding to the alternative occupational labor markets) of Table 3, we estimate a negative point estimate for the minimum wage employment elasticity in the lowest HHI tercile, with the elasticity significantly increasing (becoming more positive) for upper terciles, and the size of that increase being considerably larger in magnitude for the upper HHI tercile-log minimum wage interaction.[18]

Figure 6 plots the point estimates and 95% confidence intervals for the minimum wage employment elasticities in each HHI tercile group (based on the specifications of Table 3). For the bottom HHI tercile group, the employment elasticity is significantly negative (at or very near the 5% level) across each panel of the figure, at about -0.2. For the middle HHI tercile group, the employment elasticity is not significantly different from zero (with a point estimate that is essentially zero across all panels). For the top HHI tercile group, the employment elasticity increases further, with positive point estimates of about 0.35 that are significantly different from

[18] For stock clerks, the terciles run from 0 to 0.346, 0.346 to 0.713, and, 0.713 to 1. For retail sales, the terciles run from 0 to 0.144, 0.144 to 0.507, and, 0.507 to 1. For cashiers, the terciles run from 0 to 0.362, 0.362 to 0.690, and, 0.690 to 1.

0 at the 5% level in all panels.

Figure 7 presents a comparison of our estimates, taking labor market concentration into consideration, and estimates from the existing literature, that do not explicitly account for the level of labor market concentration. For the survey of the existing literature, we make use of estimates provided by Harasztosi and Lindner (2019) (see their Web Appendix), who report the employment elasticity with respect to the wage (own-wage elasticity of labor demand) derived from minimum wage-based research designs published in peer-reviewed academic journals. This elasticity is distinct from the canonical employment elasticity with respect to the minimum wage focused on so far, but, it can easily be computed using both the employment elasticity with respect to the minimum wage and the earnings elasticity with respect to the minimum wage (it is effectively, the ratio of the two). In the figure, we present the own-wage employment elasticity for each HHI tercile and for each occupation that results from our above estimates.[19] A general pattern emerges. In comparison to the estimates from the broader literature, our low concentration estimates are on the outer left (most negative) of the range of prior work's point estimates, and in two of the three cases include only prior estimates with negative elasticites in their 95% confidence intervals. Our middle-range concentration estimates are centered around zero and are not significant. Our upper-range concentration estimates, representing very highly concentrated labor markets, are to the far right of the existing range of estimates from the prior literature and have 95% confidence intervals that include mostly positive values. Such a pattern is to be expected if indeed labor market concentration is an important, previously unobserved and unaccounted for, factor in determining the employment effects of the minimum wage.[20]

[19]To construct our HHI-tercile-specific own-wage employment elasticities for Figure 7 we utilize the minimum wage earnings elasticity estimates from Table 1, which are averaged for the entire HHI distribution, because, as with the binary HHI measure-log minimum wage interactions presented in Appendix Table A3, there is no significant difference in the minimum wage earnings elasticity estimates across HHI terciles (Appendix Table A5)

[20]An increasing number of cities have set their own minimum wages. While many of these changes do not appear to have noticeable employment effects (Dube and Lindner, 2021), some do (Karabarbounis, Lise and Nath, 2022). An interesting avenue for future work would be to assess whether heterogeneity in the impact of the minimum wage at the city level is related to the degree of competitiveness in relevant low-wage industries.

4.2 Robustness

We now consider the robustness of the above-reported findings. We start by considering extensions of the prior specifications. We then discuss the possible interpretations of the findings and consider alternatives to the hypothesized channel. We also test for the robustness of results to alternative event-study specifications. Next, we extend the analysis to the fast food industry using an alternative to the Burning Glass HHI as a proxy for concentration. Finally, we use the bunching specification from Cengiz et al. (2019) to extend our analysis to the low-wage labor market across all industries.

4.2.1 Baseline Specification Extensions

We first consider a simplified version of the specifications in Table 2. In recent years, there has been debate in the minimum wage literature over the appropriateness of including census-division-by-period fixed effects and linear state-specific time trends à la Dube, Lester and Reich (2010) (Neumark, Salas and Wascher (2014) Allegretto et al. (2013) Allegretto et al. (2017)). Neumark, Salas and Wascher (2014) have been particularly critical of this approach, and, have instead argued for the basic two-way fixed-effect estimator commonly used in the literature in the 2000s and before. In Table A6 in the Appendix, we report a two-way fixed-effect version of Table 2. Results are similar across the two tables, and reveal the same pattern of significantly more positive employment effects of the minimum wage as labor market concentration increases. This convergence across specifications is consistent with the recent work of Cengiz et al. (2019), who find that differences across specifications disappear when focusing on minimum wage changes in recent years (as we do given our reliance on the Burning Glass data).

In Table A7 of the Appendix we explore the robustness of our main results to the exclusion of counties where Burning Glass data is more limited. Specifically, Appendix Table A7 uses in the analysis only the counties for which an occupational average HHI can be formed from at least 8 quarters of BGT job advertisement data (otherwise repeating the specifications of of Table 2). This adjustment can be thought of as a way of getting an HHI average that is potentially more reflective of the overall time period. The downside is that the exclusion of counties with limited Burning Glass data reduces the sample size somewhat. However, the change has little effect on the results.

Next, as a minimal test to reveal a possible spurious correlation we perform a placebo test in a relatively high wage industry that has virtually no minimum wage jobs. To guide the selection of industry and occupation, we start by focusing on the highest paid of the 10 largest (by employment) occupations in the US, registered nurses (SOC 291141)[21] . BLS Occupational Employment Statistics data on the distribution of this occupation's wages indicates that under no circumstances should we expect these workers to be affected by minimum wage increases, either marginally or inframarginally. OES occupation-industry matrices further indicate that the largest industry to employ this occupation while also employing essentially no minimum wage-level workers[22] is the offices of physicians (NAICS code 621100), which is also the second-largest employer of the occupation overall. This setting is ideal because we can focus on the non-minimum wage occupation in the sector without any likelihood that firm adjustments to pay for other minimum wage-level occupations will have ripple effects on the former group's employment. Our placebo test explores whether variation in labor market concentration for this non-minimum wage occupation mediates the way sectoral employment correlates with the minimum wage, looking to see if we find the same relationship observed in the general merchandise store sector. If so, our prior results would be called into question as evidence of labor market concentration's effect on the minimum wage employment elasticity.

Table A9 in the Appendix presents equivalent specifications to those in Table 2 using the physicians' office employment numbers as the outcome and registered nurse HHI in columns 1 and 2. In the the next four columns, we also include two other HHIs for two related medical

[21]This occupation has high prevalence of online recruitment (above the rate for the median occupation as assessed by the measure discussed in Section 3.2), as do the other occupations we consider in the placebo tests in this section, making use of the Burning Glass HHI appropriate for the placebo exercise we undertake here.

[22]No minimum-wage level occupation makes up more than 1 percent of the workforce. Therefore, as expected, Appendix Table A8 demonstrates a null minimum wage earnings effect for the industry.

occupations in the sector that also have moderate, above-minimum-wage earnings. These are medical assistants (SOC 319092) and Licensed Practical and Licensed Vocational Nurses (SOC 292061). The former has the advantage of being the largest detailed occupational category in the physician's office industrial sector, comprising 15% of the workforce (compared to 8% for registered nurses), making it non-trivial to firms in the industry and another attractive placebo setting; on the other hand, its median wage of $15 an hour in 2016 is closer (than the $30 an hour for registered nurses) to the range of concern for possible minimum wage ripple effects from upward wage pressure in the broader economy. The LPN/LVN occupation enjoys an intermediate wage ($19 an hour) but is a smaller fraction of the industry workforce (4%). In the last two columns of the table we report similar specifications for another placebo test with an occupation in a different industry that has closer co-movement in employment with the occupations studied in the general merchandise sector (given the symbiotic relationship between the primary sectors employing them). Following the same logic in the selection of registered nurses and the physicians' office NAICS, we focus on: the accountant and auditor occupation (SOC 132011), the next largest (in terms of employment) non-management occupation with comparably high earnings to those of registered nurses (the median wage for accountants is within fifty cents of that for registered nurses), and, the industry which most employs it and essentially no minimum wage workers, Accounting, Tax Preparation, Bookkeeping, and Payroll Services (NAICS code 541200).[23] These last two columns thus report NAICS 541200 employment as the outcome and take the HHI for SOC 132011.

In all four cases, across all specifications of Table A9 of the Appendix, the results do not yield a positive, significant interaction term between HHI and the minimum wage - just as we would expect in a setting where the underlying theoretical considerations governing monopsony and minimum wage do not apply. These results provide additional confidence in the validity of the

[23] 541200 is the largest industry to employ this occupation, and, no minimum-wage level occupations makes up more than 1 percent of the industry's workforce (and thus there is therefore no increase in average earnings in the industry following a minimum wage increase - see Appendix Table A8). One quarter of all workers in the accounting and auditor SOC are in the industry and nearly one third of the industry workforce is in the SOC.

positive interaction term seen in the main results of Table 2 and related specifications.

In Appendix Table A10, we use a single HHI measure that is a weighted average of the HHI in the three occupations we focus on in general merchandise. Not surprisingly, the results are very similar to those seen in Table 2.

In Table A11, we replicate our results for general merchandise (Table 2) using an HHI that takes into account workers' transitions across occupations using the transition matrix from Schubert, Stansbury and Taska (2022). In particular, we use the methodology of Arnold (2021) to include vacancies from occupations other than the focal occupation in the calculation of the market shares. The method gives lower weight to the vacancies that are in occupations that are inferred to be less valuable to workers in the focal occupation based on their relative transition rates and market shares. Predictably, since 77% of workers stay in their occupation when they change jobs (see Schubert, Stansbury and Taska (2022)), accounting for across-occupation transitions does not materially affect the results.

4.2.2 Consideration of Alternative Channels

Throughout the paper, we are estimating minimum wage effects on employment as a function of labor market concentration. The variation in labor market concentration is cross-sectional, and, unlike the variation in the minimum wage, this variation cannot be easily mapped to sudden policy shocks. Therefore, the question arises as to whether HHI is correlated with some other omitted variable in the market that is the true reason behind the variation in the impact of the minimum wage on employment. We address this concern below, but, first note that, from a policy perspective, even if the differential employment effects of the minimum wage are ultimately caused by something other than concentration, concentration measures still allow policy makers to assess which markets are likely to have zero or even positive employment effects when the minimum wage is increased.

One possible area of concern for an omitted variable bias arises from the fact that HHIs tend to be higher in more rural areas.[24] And, rural areas are plausibly less productive. Independent of labor market concentration measures, then, this productivity difference might affect employment responses to the minimum wage. Our expectation, however, would be that the minimum wage depresses employment more in less productive areas because increases in the minimum wage above the federal level are more likely to result in local minimum wages above workers' marginal productivity. This kind of bias goes against our finding that the minimum wage tends to increase employment in the most concentrated areas.

However, this is something that can be assessed empirically. In Appendix Table A13, we attempt to control for the main factors we believe to determine wages and to be correlated with HHI: productivity and population density. We control for population density directly and approximate productivity through the log of total average weekly earnings in the county across all sectors.[25] We find that there are no marked changes to the estimated coefficients, as can be seen in Appendix Table A13 which yields results that are very similar in magnitude and significance to our main results in Table 2.

While including these controls in the main results allows the employment level to be different in lower density or lower productivity counties, and partially addresses the omitted variable bias channel, there is an additional question of whether the minimum wage-HHI interactions remain significant and positive when also controlling for the way these additional characteristics moderate the minimum wage employment response. Appendix Table A14 allows for interactions of the minimum wage with log population density as well as with average weekly earnings in the county (across all industries). Our main finding remains: a higher HHI is associated with a more positive minimum wage effect on employment. In additional analysis, we drop HHI and its minimum wage interaction and look at the predictive power of these other characteristics: log population density has a similar mediating role for the employment

[24]See Azar et al. (2020). For reference, in our sample, the bivariate correlation of HHI for cashiers, for example, with log population density is -0.68.

[25]Following convention, the log of total average weekly earnings in the county across all sectors is already included in the set of regressors for our baseline results in Table 2 as well (excluding this regressor doesn't change results, see Appendix Table A12).

effect of the minimum wage as HHI (see Appendix Table A15).[26] However, as Appendix Table A14 makes clear, HHI is the better predictor of heterogeneity in the minimum wage employment effect. These results lend further credence to the idea that labor market concentration is not merely a proxy for other factors like productivity or population density, but can by itself modulate the employment effect of the minimum wage.

It is also possible, in principle, that the minimum wage may have a direct effect on labor market concentration, perhaps because it forces some firms to close. This could complicate the interpretation of the employment effect of the minimum wage by concentration. However, empirically, there is no significant relationship between the minimum wage level and the level of labor market concentration (Appendix Table A2).

A final identification concern involves differential price pass through responses to the minimum wage. When the minimum wage increases, firms may raise prices, and this may be especially the case in high concentration areas. This alternative mechanism would thus dampen a negative effect of the minimum wage on employment (shifting the policy's cost to consumers rather than workers), but should not by itself yield a positive minimum wage employment effect. When examining the impact of the minimum wage on prices (without an interaction with concentration), the prior literature has found mostly positive price effects (Allegretto and Reich, 2018; Aaronson, French and MacDonald, 2008; Renkin, Montialoux and Siegenthaler, 2020; Leung, 2021; Ashenfelter and Jurajda, 2021) while some studies find smaller effects (MacDonald and Nilsson, 2016; Ganapati and Weaver, 2017). Leung (2021) documents significant geographic heterogeneity in the pass through effect, which is potentially consistent with a role of labor market concentration. Empirically, we check for a differential impact of the minimum wage on the Consumer Price Index (overall, and for commodities) in higher concentration areas. We do not find statistically significant differences. However, the power of this analysis is limited because temporarily and geographically rich CPI data is scarce and there are only ten

[26]Our productivity proxy, on the other hand, does not have a significant negative interaction term with log minimum wage, though the coefficient is negative, as would be expected if more competitive markets pay higher wages. See Appendix Tables A15 and A16.

metropolitan areas with usable CPI data (see more details in appendix section A.3). Overall, we conclude that price effects are unlikely to explain away our key result that minimum wage employment effects increase with labor market concentration.

4.2.3 Event-Study Specification

As an alternative identification strategy, we utilize an event-study specification. Besides providing a robustness test on the above findings, the event-study design graphically displays the counterfactual and its identifying assumption can easily be assessed visually. As shown below, the findings here are consistent with the previous sections' results.

We use large-bite, first-time minimum wage reforms as the "events" in our analysis. Specifically, we use all instances where the minimum wage increase exceeded 50 cents and where this increase was not preceded by a minimum wage increase in the previous two years. This effectively ensures the event time of zero actually indicates a first-time minimum wage event within our entire sample period, and, that this minimum wage increase, and resulting additional firm labor costs for minimum wage workers, are sizeable. All other locations that only had smaller (typically, absent) minimum wage changes in our sample period are treated as control groups in the initial analysis. As a robustness check, we also perform the analysis using as a control only those places that did not ever have a minimum wage increase (even a small one) during our sample period.[27] The results are quite similar regardless of which control group is chosen.

The basic event-study specification is:

$$ln(Y_{it}) = \sum_{j=-4}^{-2} \alpha_j D_i \mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j) +$$

$$\sum_{j=0}^{3} \rho_j D_i \mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j) + \gamma_i + \tau_{ct} + \chi I_s * t +$$

$$\phi ln(TotEmployed_{it}) + \zeta ln(TotEarnings_{it}) + \eta ln(Pop_{it}) + \tau_t * ln(Unemp_{it})$$

[27] The corresponding results for this alternative control group are summarized at the end of this section with associated figures available in the Appendix (Appendix Figures A9, A10, A11).

where, as before, $ln(Y_{it})$ is the log of general merchandise store employment in county i and quarter t, and additional control variables include the log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate, county fixed effects γ_i, census division specific quarter fixed effects τ_{ct}, and state-specific linear time trends $\chi I_s * t$, with standard errors clustered at the state level. New in the specification is the D_i term, an indicator of treatment that is equal to 1 if the county ever experiences a large ($>$ 50 cent) minimum wage increase that was not preceded by a minimum wage increase in the previous two years.

Thus, the estimates characterizing the employment effects of a minimum wage increase are the coefficients on the interactions of D_i with the event-time dummies, $\mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j)$ that characterize time (in six month increments) from the minimum wage increase event. These dummies are equal to 1 when the observation is $j = -4, ..., 0, ..., 3$ bi-annual time periods from T_i^*, the date when the minimum wage was increased in the county ($j = -1$ is the omitted category). Observations more than 2 years before or after the minimum wage increase event are captured by dummies, $\mathbb{1}(t - T_i^* + 5 \leq -4)$ and $\mathbb{1}(t - T_i^* - 5 \geq 3)$. The point estimates α_j characterize the evolution of general merchandise log employment in the eventually treated counties before the minimum wage increase *net* of changes in untreated counties after adjusting for the model covariates, and, as such, allows for a direct evaluation of the assumption that the location and timing of the minimum wage events analyzed is unrelated to pre-event changes in general merchandise employment. The point estimates ρ_j then describe the divergence of outcomes j (bi-annual) periods *after* the minimum wage increase *net* of changes in untreated counties after adjusting for the model covariates (relative to the six months before the minimum wage increased, $j = -1$).

We first present the results from this specification separately for high and low labor market concentration areas (HHI above or below 0.25). Figure 8 shows the estimates for the low concentration areas and Figure 9 for high concentration areas using cashier HHI (equivalent figures using stock clerk and retail sales HHIs are presented in the Appendix Figures A3, A4, A6 and A7; while we focus on the cashier figures in this section, as can be seen in the results that follow, the findings are broadly similar across occupations).[28]

For both high and low concentration areas, we see no evidence of a differential trend in employment in treated locations before the minimum wage increase event (Figures 8 and 9). The pre-minimum wage effects α are all statistically insignificant and close to zero in magnitude. After the minimum wage increase, however, the employment effects in the low concentration areas trend negative. As can be seen in Figure 8, they become significantly different from zero after six months. The results underlying Figure 8 indicate that the average effect across the first two years post-event is a statistically significant -0.0176 reduction in log employment (p-value of 0.015). In high concentration areas, on the other hand, following the minimum wage increase the employment effects are immediately significant and *positive*. The results underlying Figure 9 indicate that across the first two years the average effect is a statistically significant .0163 increase in log employment (p-value of 0.013).

The difference between the post-event effect sizes in high and low concentration areas can be assessed more formally with the following specification

$$
\begin{gathered}
ln(Y_{it}) = \textstyle\sum_{j=-4}^{-1} \nu_j D_i \mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j) + \\
\textstyle\sum_{j=0}^{3} \mu_j D_i \mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j) + \\
\textstyle\sum_{j=-4}^{-2} \lambda_j D_i \mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j) \mathbb{1}(HHI_i \geq 0.25) + \\
\textstyle\sum_{j=0}^{3} \sigma_j D_i \mathbb{1}(t - T_i^* - j = j | t - T_i^* - j - 1 = j) \mathbb{1}(HHI_i \geq 0.25) + \\
\psi \mathbb{1}(HHI_i \geq 0.25) + \beta D_i \mathbb{1}(HHI_i \geq 0.25) + (1 + \mathbb{1}(HHI_i \geq 0.25))(\gamma_i + \tau_{ct} + \chi I_s t + \xi_t + \\
\phi ln(TotEmployed_{it}) + \zeta ln(TotEarnings_{it}) + \eta ln(Pop_{it}) + \tau_t * ln(Unemp_{it}))
\end{gathered}
$$

which includes the same regressors as the previous specification plus an indicator for labor markets being above the DOJ-FTC "high concentration" threshold, $\mathbb{1}(HHI_i \geq 0.25)$, and its

[28] Appendix Figures A12-A14 report the corresponding earnings effects. As with the previous findings, there is no significant difference in the size of the post-event earnings effect across high and low concentration places.

interaction with the other main regressors. The key coefficients of interest for our hypothesis are the λ_j and σ_j terms which characterize the evolution of the difference in log employment between eventually treated counties and controls in high concentration areas *net* of the equivalent difference in low concentration areas (relative to the difference in this difference in the six months before the minimum wage increased, $j = -1$, all after adjusting for model covariates).

The point estimates λ_j characterize the evolution of the difference in this difference in the pre-event period, and, as can be seen in Figure 10 these estimates are all insignificant and close to zero, providing validation for the event-study identifying assumption (the same can be said of the specifications using stock clerk or retail sale HHI instead, whose corresponding figures are Appendix Tables A5 and A8). As a summary measure that effectively averages across λ_j terms from $j = -2$ to $j = -4$, row 1 of Table 4 reports the event-study estimates of the difference in this difference in the pre-minimum wage window of 2 years to six months prior to the minimum wage (again, relative to the difference in this difference in the six months before the minimum wage increased). For all 3 occupation HHIs, the coefficient magnitude is very small and not statistically different than zero.

The point estimates ρ_j describe the divergence of this difference-in-the-difference *after* the minimum wage increase. As can be seen in Figure 10 these differences are positive and significant (at the 10% level initially and then the 5% level), with the average effect across the first two years post-minimum wage increase being a .031 increase in log employment (significant at the 1% level), as reported in row 2, column 3 of Table 4. Similar effect sizes are reported for the other occupation HHIs as well (significant at the 1% level for stock clerks and at the 10% level for retail sales). Rows 3 and 4 of Table 4 also report equivalent estimates to rows 1 and 2 (finding very similar results across the board) using the alternative control group defined above, locations that do not experience any change in the minimum wage during the sample period (in the Appendix we also present equivalent figures to Figure 10 for each of the occupation HHIs using this alternative control group; see Appendix Figures A9, A10, A11).

As a final robustness check on the event-study analysis, we also estimate a stacked event-study specification to alleviate negative weighting concerns, following the methods of (Cengiz et al., 2019).[29] The stacked specification analogues to Figures 8 - 10 and Appendix Figures A3 - A8 can be found in new Appendix Figures A15-A23. As can be seen, the results are robust to this alternative event-study specification.

Taken together, the event-study analysis is quite consistent with our hypothesis and the two-way-fixed effect findings in the previous section. The increasing effect size over time seems to be largely driven by the slight delay in the disemployment effect in the low concentration areas seen in Figure 8, itself possibly due to either, or both, a) competitive firms attempting to initially absorb increased labor costs in the first six months, and, after failing to do so resorting to worker cuts, and, b) the fact that in virtually all of the minimum wage increase events analyzed additional increases were mandated within two years of the event (usually after a year), making these labor costs even larger in the time periods after $t = 0$, and, thus, even harder to absorb without job cuts.[30]

Overall, the event study shows that there are no pre-trends in employment and that a large minimum wage increase leads to decreases in employment in low concentration markets and increases in employment in high concentration markets. Critically, there is a significantly more positive employment effect in high concentration labor markets than in low concentration ones. The event-study results thus confirm our main estimates.

[29] The procedure for the stacked event-study specification is as follows. For each of our minimum wage events, we construct an event-specific data set that includes the counties treated by that particular large minimum wage increase (included only for the periods within the ± 8 quarter event window around the event's onset) and any counties without a large minimum wage increase happening during the ± 8 quarter event window around the event's onset (included only for the periods within this same time span as well). Each of these event-specific data sets are then stacked on top of each other to form one stacked data set. On this stacked data set we run the event-study specifications in this subsection with the modification that the calendar time and county fixed effects are interacted with indicators for the event-specific data set from which the observation comes (with standard errors now clustered at the state by event-specific-data set level). The resulting estimates on the event-time dummies provide us with an average effect across all the events.

[30] This feature of the minimum wage policy implementation makes it difficult to disentangle from the current specifications the extent to which the $t + 2$ and $t + 3$ employment estimates are larger because of delayed onsets from the minimum wage increase we observe in $t = 0$, or, because of even larger cumulative minimum wage increases that occur in the post-event window. We leave this assessment for further analysis elsewhere.

4.2.4 Extension to Fast Food Industry with Alternative Concentration Measure

In this section, we show how our results generalize to the other key minimum wage sector in the United States, the limited service restaurant, or, "fast-food", industry (NAICS 722513).[31] Since the Burning Glass data, as previously noted, is not representative for restaurant hiring, we utilize the QCEW data alone to create a modified concentration measure that enables this extension. While we otherwise follow the methodology in the previous subsection, for this additional analysis we instead approximate the concentration in the fast-food industry with the inverse number of establishments in the limited service restaurant NAICS at the county level. This is an imperfect measure of concentration, since some firms will indeed have multiple establishments in a county. At the same time, generally speaking, fewer establishments in a place means fewer employers and greater labor market concentration for the occupations they employ. Notably, this establishment-based concentration measure is also very close to the one we use in our theoretical model in section 2, where we compute concentration as the inverse number of (identical) firms.

We first validate this alternative concentration measure by observing that the Burning Glass-derived HHIs in the general merchandise analysis are strongly negatively correlated with the number of establishments in the general merchandise industry taken from the QCEW data (correlation of -0.47), with a t-stat on the bivariate specification of these two measures well over 10. Moreover, when using the number of establishments in the general merchandise industry as a concentration measure, we get very similar results to our prior analysis where concentration was measured per our Burning Glass-generated HHI (compare Figure A24 in the Appendix to Figure 10).

This allows us to move to a QCEW-based analysis with confidence in the use of the number of establishments as a proxy for Burning Glass HHI estimates of labor market concentration, opening up analysis to the restaurant industry, where the Burning Glass job vacancy data was

[31] In 2012 the limited service restaurant code was renamed from 722211 to 722513. We refer to it under the new code.

not precise enough. The upshot is that, when using this alternative measure of concentration for the fast-food industry, the results echo the previous findings: more positive minimum wage employment effects are present in more concentrated markets. We now describe the results for the fast-food analysis in greater detail.

We define the cutoff for high concentration in the limited-service restaurant industry so that it is substantively comparable to the cutoff used for the general merchandise industry. Specifically, we define high concentration fast-food markets as those in the bottom decile of the number of county-level establishments. In this bottom decile, the median number of fast-food establishments is 4, which is equivalent to the implied number of equally-sized firms for an HHI of 0.25 - the DOJ-FTC cutoff used in the corresponding general merchandise analysis.

Figures 11-13 take the sample of locations used in Figures 8-10 of the general merchandise analysis and apply the same event-study specifications - but now with limited-service restaurant sector employment as the outcome instead, and, the number of establishments in the fast-food sector used as the labor market concentration proxy to partition the sample into newly defined low and high concentration areas for this sector. Specifically, the coefficients in Figure 11 are estimates of α and ρ derived from the first event-study specification in Section 4.2.3 when applied to the sample of low concentration (many fast-food establishment) locations. Similarly, Figure 12 presents an equivalent figure to Figure 9 for the high concentration fast-food locations (those counties inside the bottom decile of the fast-food establishment distribution). Figure 13 presents an equivalent figure to Figure 10 and compares the resulting difference in the fast-food employment effects of the minimum wage in high vs. low labor market concentration counties (the coefficients are event-study estimates of λ and σ using the second event-study specification in Section 4.2.3). The results echo the general merchandise findings, with significantly more positive employment effects from the minimum wage in more concentrated labor markets.[32]

[32] Appendix Figure A25 reports fast-food minimum wage earnings effects in the high and low concentration areas, respectively. Consistent with our model and prior results, there is a not a significant difference between the size of these earnings effects.

4.2.5 Bunching Specification (Cengiz et al., 2019)

Having focused thus far on the two key industries employing minimum wage workers, in this section, we further generalize our results to all minimum wage-level workers (of whatever industry) by utilizing the data and bunching design of Cengiz et al. (2019) applied separately to low and high concentration markets. We also separately zoom in on retail and restaurant workers and show that results are consistent with our above findings when using this alternative bunching methodology.

We start with the publicly available replication package from Cengiz et al. (2019). The Current Population Survey dataset used in Cengiz et al. (2019) has limited geographic information for each person-level observation. In particular, it records the state in which a person is located and whether they are in a metropolitan or non-metropolitan location.[33] To assign an HHI to each person in the dataset, we first calculate a vacancy-weighted average of the SOC $\times$ county HHIs for each geographic area available (that is, for each State x Metropolitan/Non-Metropolitan indicator pair). For example, if we know that a person is located in a metropolitan area in the state of Arkansas, we assign the vacancy-weighted average of the SOC $\times$ county HHIs in metropolitan areas in Arkansas in the BGT dataset.[34] For this average, we use data between 2010 and 2016, as above. This effectively gives us a summary measure of how concentrated metropolitan Arkansas is across all occupations and (county-level) locations that constitute it, i.e. across its various occupational labor markets. We then classify the geographic areas in Cengiz et al. (2019) (that is, each State x Metropolitan/Non-Metropolitan pair) into low and high average concentration based on the DOJ-FTC HHI threshold of 0.25 used in the rest of this

[33] The specific metropolitan or micropolitan area in which a person is located is available for only a subset of the observations, and only since May 2004.

[34] To clarify, we do the following: for a given geographic area G (e.g., metropolitan Arkansas), we calculate its average HHI by taking a vacancy-weighted average of the HHIs in all the SOC$\times$county$\times$year-quarter labor markets (indexed by m for notational simplicity) in that geographic area:

$$\overline{HHI}_G = \sum_{m \in G} w_m HHI_m,$$

where w_m is the share of vacancies in geographic area G that are in market m.

paper. We estimate the Cengiz et al. (2019) specification (specifically the specification in their Table 3 that underlies Figure 2) separately using workers in low vs. high concentration areas.

Figure 14 shows the results. Panel A shows the results using workers from all industries. Subfigure (a) shows the results for workers in areas with average HHIs below 0.25, that is, in less concentrated areas. Increases in the minimum wage have a negative effect on employment for workers in areas with average HHIs below 0.25. The blue bars in the figure show for each dollar bin (relative to the minimum wage) the estimated average employment changes in that bin during the five years post treatment relative to the total employment in the state one year before the treatment. There is an expected decline in employment in the bins right below the new minimum wage, and an increase in the five bins above the new minimum wage. However, the increase in jobs in the bins above the minimum wage is smaller than the decrease in jobs in the bins below the minimum wage, as indicated by the dashed red line which shows the running sum of employment changes up to the wage bin it corresponds to. However, although the implied elasticity of employment with respect to the affected wage and with respect to the minimum wage is negative, it is not statistically significant, as can be seen in Appendix Table A17, column 1.

Subfigure (b) in panel A of Figure 14 shows the results for workers in areas with average HHIs above 0.25, that is, in more concentrated areas. Increases in the minimum wage have an overall positive effect on employment for workers in these geographic areas. There is again a decline in employment in the bins right below the new minimum wage, and an increase in the five bins above the new minimum wage, but now, the increase in jobs in the bins above is substantially larger than the decrease in jobs in the bins below the minimum wage. Moreover, as shown in Appendix Table A17, column 2, this is a statistically significant result, with an implied employment elasticity of the minimum wage with respect to the affected wage of 1.82 (1% significance level) and an employment elasticity with respect to the minimum wage of 0.26 (10% significance level).

Figure 14 panels B and C show analogous results separately for Retail and Restaurants, following the industry definitions in Cengiz et al. (2019). For the weighted average HHIs by geographic area, we now use weights w_m (defined in footnote 34) that reflect only the total vacancies in Retail (44-45) and Restaurants (722 NAICS), respectively. The worker counts are of course also restricted to the corresponding industry. For retail, the minimum wage effect on employment in areas with low average concentration is close to zero, while the effect on employment in areas with high average concentration is positive and statistically significant: for the later case, Table A17, column 6 reports the employment elasticity with respect to the affected wage of 3.18 (1% significance level) and an employment elasticity with respect to the minimum wage of 0.69 (5% significance level). Interestingly, the former is in the range of our high HHI estimate in Figure 7, and, also to the right of the prior literature's estimates (that do not disaggregate by labor market concentration). In the case of restaurants, the employment effect is negative in low-concentration areas, and positive in high-concentration areas, but the magnitude of the effect is small, and in both cases the implied elasticity is statistically insignificant. This is perhaps not surprising given that, as discussed earlier, the BGT dataset is not well suited to accurately measure concentration in the restaurant sector.

With this in mind, in Appendix Figure A26 and Appendix Table A18 we also show results using metropolitan vs. non-metropolitan areas instead of low vs. high concentration areas. The pattern is in fact quite similar. This coherence is consistent with our above finding for the general merchandise industry where population density – when included on its own – has a significant mediating impact on the minimum wage employment effect (see again Appendix Table A15). These results speak to the possibility of using population density as a proxy for HHI when data limitations require it, though, as we previously noted, HHI is a better predictor for employment effects (see again Appendix Table A14).

4.2.6 Summarizing Heterogeneity by HHI vs. Alternatives

Regarding this final point, we summarize (in Table 5) the differences in heterogeneity by HHI vs. the alternative proxies considered. The specifications underlying the table return to our baseline specification for general merchandise in Section 4.1 and are mostly taken from results reported elsewhere in the paper (aggregated here for summary purposes).[35] Column 1 of Panel A reports the heterogeneity in the employment elasticity with respect to the minimum wage by HHI (from Table 2). Column 2 leaves out HHI and instead reports the heterogeneity in the employment elasticity with respect to the minimum wage by population density (from Table A15). Column 3 leaves out both the log minimum wage interaction with HHI and log population density and instead interacts log minimum wage with the log of the number of establishments in the sector. In all three cases there is clear and statistically significant heterogeneity in the minimum wage employment elasticity in the theoretically expected direction: the employment elasticity goes up (becomes significantly less negative or more positive) in the more concentrated (higher HHI) occupational labor markets, in the less densely populated places, and in the places with fewer establishments in the sector. Finally, in Column 4 we report the heterogeneity (from Table A15) in the employment elasticity with respect to the minimum wage by productivity (also considered on its own, and, again, measured by log of county average earnings), which, as we have seen does not yield any significant heterogeneous effect.

However, though HHI, population density, and, number of low-wage sector establishments all yield significant heterogenous effects, of the three variables occupational HHI is the better predictor. Panel B in Table 5 illustrates this. In it we report the estimates on the interaction of log minimum wage with each of the the mediating variables explored in Panel A - when all interactions are included in the specification together. Only the interaction with HHI remains statistically significant. Its magnitude is essentially unchanged from Panel A. Occupational labor market HHI thus has separate empirical content from the other mediators that is not easily explained by them (and their mediating effect is no longer statistically discernible when HHI is accounted for). This suggests that while some more readily available proxies (population or low-wage firm density) may be useful measures of heterogeneous minimum wage employ-

[35]We present here the results for cashiers, but, the conclusions are the same when looking at either the stock clerks or retail sales results instead.

ment effects, the HHI measure is a preferable if available.

5 Conclusion

Economic theory predicts that the minimum wage reduces employment in a competitive labor market. However, empirical evidence has often failed to find a negative impact. In contrast, it is well known that minimum wages could have zero or positive effects if firms have monopsony power. In this paper, we present the first direct empirical evidence showing measures of labor market concentration – a key determinant of monopsony power – modulate the impact of the minimum wage on employment. This robust finding holds for the two main sectors employing minimum wage workers in the United States, and for different proxies of labor market concentration.

The paper first establishes its main findings studying a key low-wage retail sector (General Merchandise) and using data on labor market concentration that covers the entirety of the United States with fine spatial variation at the occupation-level. It then shows the results carry over to the fast-food sector and the entire low-wage labor market and to alternate proxies of labor market concentration available for a broader range of industries, such as the number of low-wage sector establishments and population density.

Specifically, the paper finds that the employment elasticity of the minimum wage robustly and significantly increases (becomes more positive) with the concentration of occupational labor markets. In the most concentrated third of markets in the General Merchandise Sector, the minimum wage employment elasticity is even estimated to be significantly positive. Compared to the existing literature, these results yield related own-wage elasticity of labor demand estimates for the lowest tercile of labor market concentration on par with the lowest estimates from the minimum wage literature, effects close to zero and statistically insignificant for the middle tercile, and positive estimates that are larger than most in the literature for the highest tercile. The findings, thus, offer an empirically-founded candidate explanation as to why null employment effects from the minimum wage abound in the literature (due to the averaging of real underlying positive and negative effects), as well as insight into why we may see significant variation in the employment effect of the minimum wage across studies (due, in part, to differing underlying levels of labor market concentration in a study's settings).

In total, these results provide compelling evidence that the structure of the labor market influences the effects of a key labor market institution, the minimum wage. The findings imply the degree of local labor market concentration is helpful for understanding the employment effects of a minimum wage increase. Additionally, the results bolster the evidence for monopsony power in the labor market by demonstrating that key policy effects conform to the predictions of the monopsony model. Finally, our findings imply that multiple proxies for labor market concentration can be used to study the differential effects of minimum wages, a result useful for future research on the minimum wage.

Data Availability Statement The data underlying this study were provided by Burning Glass Technologies (Lightcast) by permission. The code is provided in our online repository: https://doi.org/10.5281/zenodo.8212833

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Table 1. Minimum Wage Effect on Earnings

	Dependent Variable: Log Earnings		
	(1)	(2)	(3)
Log Min Wage	0.091***	0.093***	0.095***
	(0.031)	(0.031)	(0.033)
County Fixed Effects	Y	Y	Y
Cen. Div. Period Fixed Effects	Y	Y	Y
State-Specific Time Trends	Y	Y	Y
Additional Controls	Y	Y	Y
adj. R^2	0.839	0.840	0.841
N	56536	57280	56592

Notes: The table presents estimates of the earnings elasticity with respect to the minimum wage in the General Merchandise Store industrial sector (NAICS industry code 452). All specifications in the table take the log of county-quarter general merchandise store average monthly earnings as the outcome. The three columns report estimates using the alternative baseline samples subsequently used throughout the paper (see Table 2) to compute HHI. Each sample is constructed using information from a different low-wage occupational labor market in the industry, resulting in different sample sizes. Column 1 corresponds to the sample for stock clerks and order fillers (SOC 435081); Column 2 to retail salespersons (SOC 412031); Column 3 to cashiers (SOC 412011). In addition to the log of the governing minimum wage as a regressor, all specifications include county fixed effects, census division specific period fixed effects, state-specific linear time trends, and, the following additional control variables: log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate. Standard errors (in parenthesis) are clustered at the state level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 2. Minimum Wage Effect on Employment by Concentration in Occupational Labor Market

	Dependent Variable: Log Employment					
	Stock Clerks		Retail Sales		Cashiers	
	(1)	(2)	(3)	(4)	(5)	(6)
Log Min Wage	-0.312***	-0.178**	-0.219***	-0.168**	-0.310***	-0.183**
	(0.082)	(0.074)	(0.063)	(0.071)	(0.093)	(0.075)
Log Min Wage * Avg HHI	0.702***		0.802***		0.728***	
	(0.152)		(0.220)		(0.194)	
Log Min Wage × High HHI		0.291***		0.432***		0.308***
		(0.060)		(0.082)		(0.067)
County Fixed Effects	Y	Y	Y	Y	Y	Y
Cen. Div. Period Fixed Effects	Y	Y	Y	Y	Y	Y
State-Specific Time Trends	Y	Y	Y	Y	Y	Y
Additional Controls	Y	Y	Y	Y	Y	Y
adj. R^2	0.994	0.994	0.994	0.994	0.994	0.994
N	56536	56536	57280	57280	56592	56592

Notes: The table presents estimates of the employment elasticity with respect to the minimum wage, and, variation in this estimate across places with differing levels of labor market concentration for key low-wage occupations. All specifications in the table take the log of county-quarter general merchandise store employment as the outcome. In addition to the log of the governing minimum wage as a regressor, all specifications include the interaction of this variable with labor market concentration (HHI). We construct HHI for each of the three key low-wage occupational labor markets in the industry: stock clerks and order fillers (SOC 435081) in columns 1-2; retail salespersons (SOC 412031) in columns 3-4; cashiers (SOC 412011) in columns 5-6. In odd-numbered columns, this concentration measure is the county's Herfindahl-Hirschman Index (HHI) for the relevant labor market (averaged across all quarters of the 2010-2016 sample period), and, we report its interaction with the log minimum wage. In even-numbered columns, we instead report the interaction of log minimum wage with a binary concentration measure that separates high and low concentration labor markets based on whether their labor market HHI is above or below 0.25 (the Department of Justice/ Federal Trade Commission threshold for highly concentrated markets). All specifications further include county fixed effects, census division specific period fixed effects, state-specific linear time trends, and, the following additional control variables: log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate. Standard errors (in parenthesis) are clustered at the state level. For the corresponding average minimum wage earnings effect in the full sample, and, for the high and low concentration sample minimum wage earnings effects see Table 1 and Table A3, respectively. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 3. Minimum Wage Effect on Employment by HHI Terciles

	Dependent Variable: Log Employment		
	(1) Stock Clerks	(2) Retail Sales	(3) Cashiers
Log Min Wage	-0.200**	-0.189**	-0.153*
	(0.078)	(0.072)	(0.078)
Log Min Wage * Mid HHI	0.161**	0.231**	0.176*
	(0.074)	(0.095)	(0.094)
Log Min Wage $\times$ High HHI	0.586***	0.553***	0.475***
	(0.138)	(0.139)	(0.150)
County Fixed Effects	Y	Y	Y
Cen. Div. Period Fixed Effects	Y	Y	Y
State-Specific Time Trends	Y	Y	Y
Additional Controls	Y	Y	Y
adj. R^2	0.994	0.994	0.994
N	56536	57280	56592

Notes: The table presents estimates of the employment elasticity with respect to the minimum wage by labor market concentration (HHI) terciles. HHI is constructed for each of the three key low-wage occupational labor markets in the general merchandise store industry: stock clerks and order fillers (SOC 435081) in column 1; retail salespersons (SOC 412031) in column 2; cashiers (SOC 412011) in column 3. All specifications in the table take the log of county-quarter general merchandise store employment as the outcome. Specifications 1-3 are identical to specifications 2,4,6, respectively, in Table 2 with the exception that log minimum wage is here interacted with two indicator variables identifying high and medium HHI terciles (rather than simply a binary measure as in Table 2). All specifications include county fixed effects, census division specific period fixed effects, state-specific linear time trends, and, the following additional control variables: log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate. Standard errors (in parenthesis) are clustered at the state level. For the corresponding low, mid, and high concentration minimum wage earnings effects see Table A5. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4. Event-Study Estimates of Minimum Wage Employment Effects in High Concentration Labor Markets Net of Effects in Low Concentration Labor Markets

	Dependent Variable: Log Employment		
	(1) Stock Clerks	(2) Retail Sales	(3) Cashiers
Inclusive Control Group			
Relative to $t = -1$			
2 Years to 6 Months Before Minimum Wage Change	0.004	-0.006	0.002
	(0.007)	(0.010)	(0.006)
Minimum Wage Change to 2 Years After	0.031***	0.032*	0.031***
	(0.009)	(0.016)	(0.008)
Restrictive Control Group			
Relative to $t = -1$			
2 Years to 6 Months Before Minimum Wage Change	0.005	-0.002	0.005
	(0.007)	(0.010)	(0.006)
Minimum Wage Change to 2 Years After	0.037***	0.033*	0.034***
	(0.009)	(0.018)	(0.007)
County Fixed Effects	Y	Y	Y
Cen. Div. Period Fixed Effects	Y	Y	Y
State-Specific Time Trends	Y	Y	Y
Additional Controls	Y	Y	Y

Notes: The table reports the event-study estimates derived from the second specification in Section 4.2.3 when using (for rows 1 and 3) a pre-event event-time category of 2 years to 6 months before the minimum wage increase, or, (for rows 2 and 4) a post-event event-time category of 0 months to 2 years after the minimum wage increase. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase. Zero represents the onset of the minimum wage. Estimates measure, for the respective event-time window, the difference in log general merchandise store employment between eventually treated counties and control counties in high labor market concentration areas ($HHI \geq 0.25$) net of this difference in low concentration areas (again, relative to the difference in this difference in the six months before the minimum wage increased) after adjusting for model covariates. High and low labor market concentration areas are defined alternatively over each of the three key low-wage occupational labor markets in the general merchandise store industry: stock clerks and order fillers (SOC 435081) in column 1; retail salespersons (SOC 412031) in column 2; cashiers (SOC 412011) in column 3. Results are presented for two separate control groups, as described in the text: in the *inclusive control group* panel, locations that only had small (50 cents or less) or no minimum wage change in our sample period are treated as the control group, while the *restrictive control group* panel only takes as a control group those places that did not ever have a minimum wage increase (even a small one) during our sample period. All specifications include county fixed effects, census division specific period fixed effects, state-specific linear time trends, and, the following additional control variables: log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate. Standard errors (in parenthesis) are clustered at the state level. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

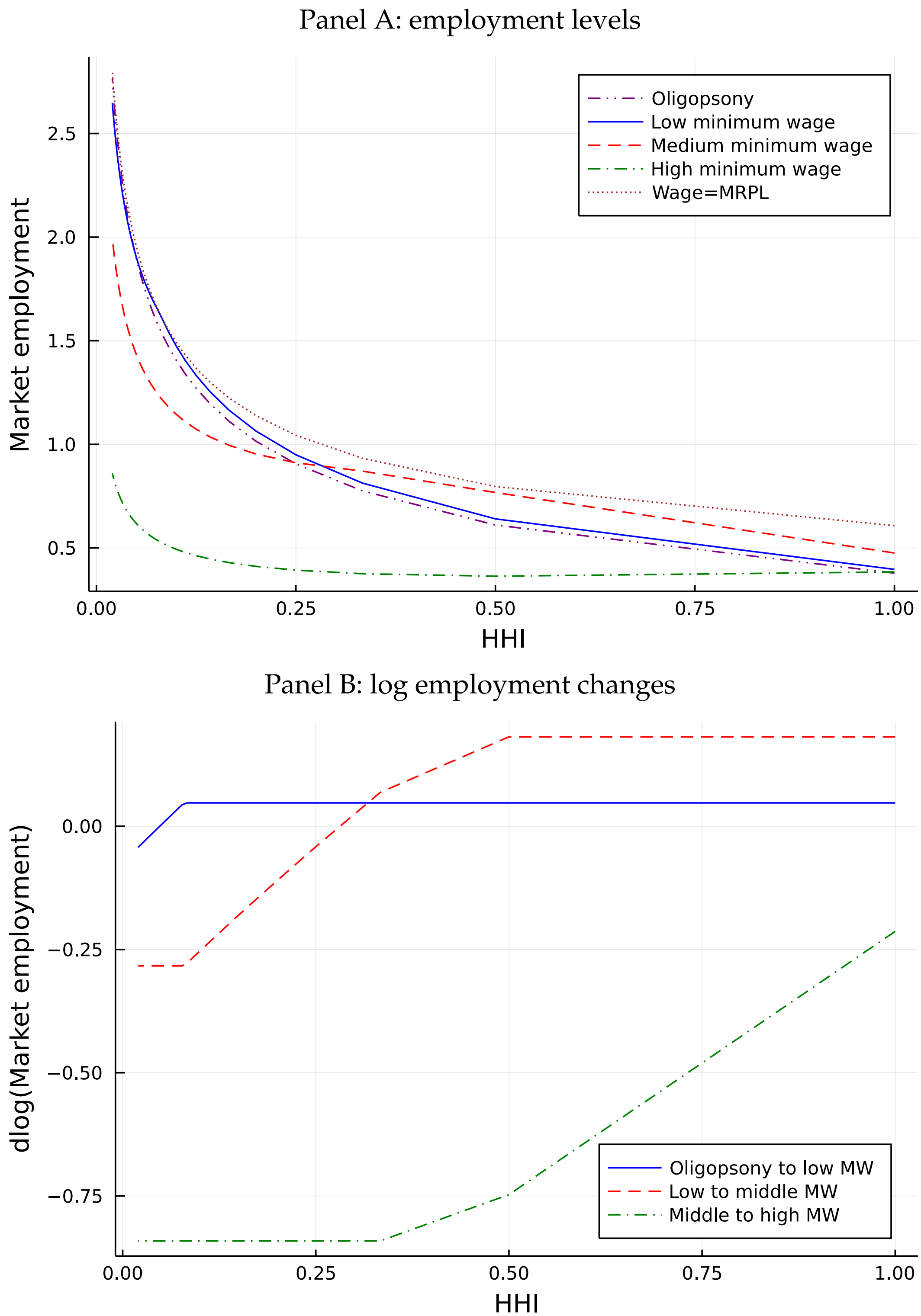

Figure 1. Theoretical effect of a minimum wage increase by labor market concentration (HHI)
Notes: We simulate market-level employment effects as a function of HHI for a minimum wage ("MW") set first 3% higher than the market-specific oligopsony wage (low minimum wage). Then the minimum wage is increased by 12% (medium minimum wage) and finally the minimum wage is increased by 40% (high minimum wage). In panel A, for the brown dotted line, the wage is set equal to the marginal revenue product of labor ("MRPL"). See more details in section 2.

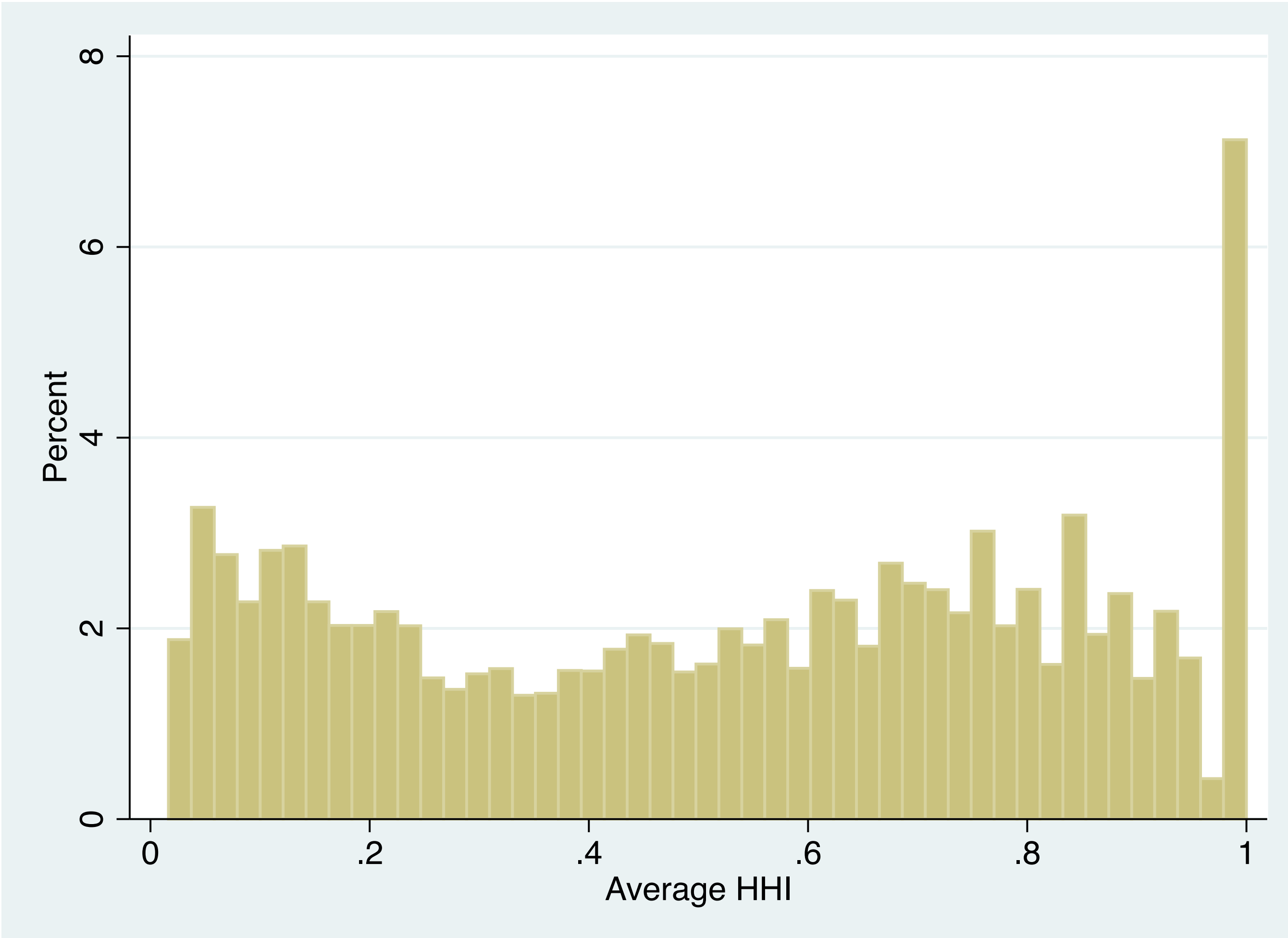

Figure 2. Distribution of County Avg. HHI Measure for Stock Clerks Occupational Labor Market
Notes: The figure plots the distribution across counties of the average HHI measure used in the "Stock Clerks" regressions (Columns 1 and 2 of Table 2 and equivalent samples in other tables).

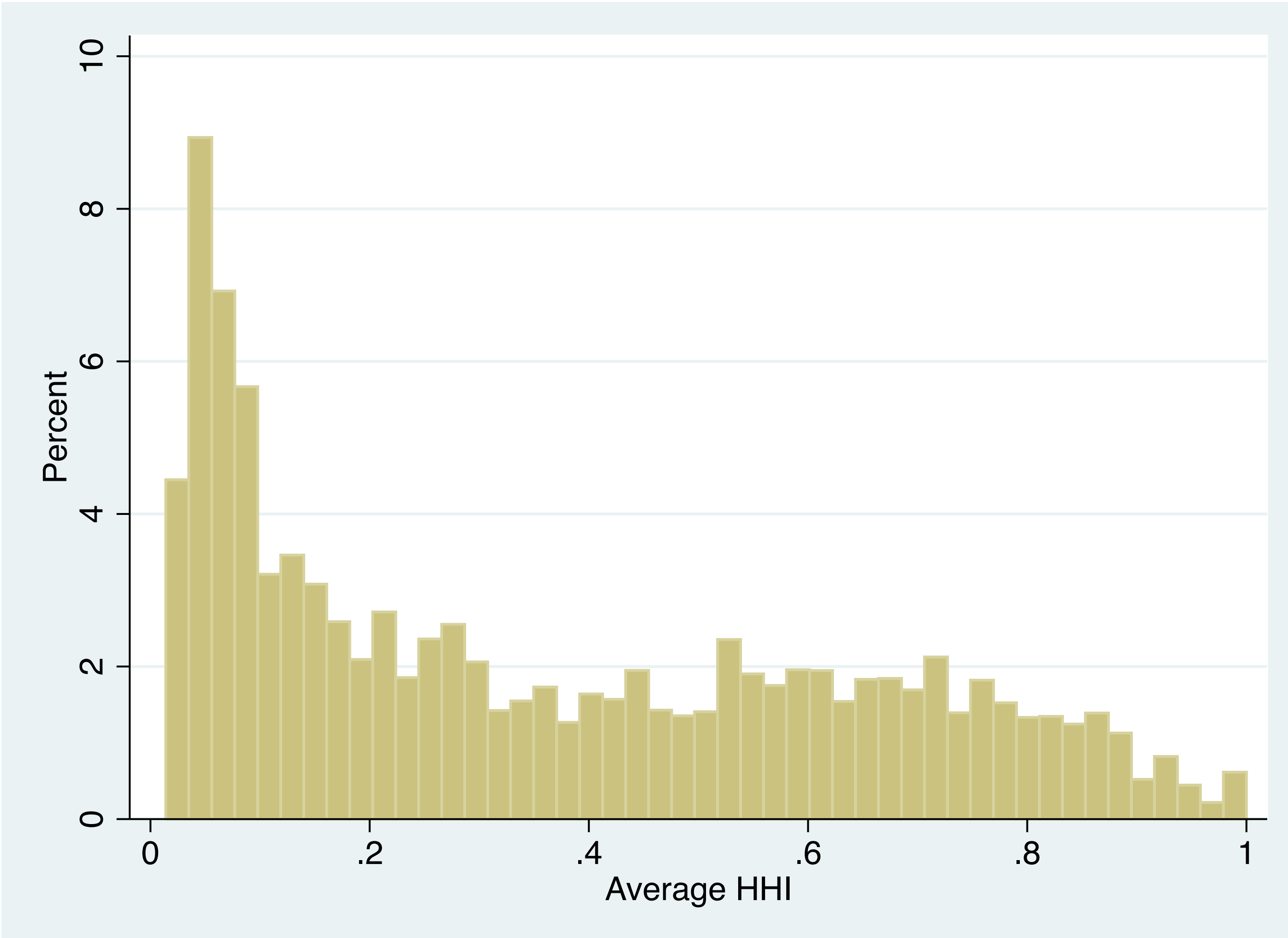

Figure 3. Distribution of County Avg. HHI Measure for Retail Sales Occupational Labor Market
Notes: The figure plots the distribution across counties of the average HHI measure used in the "Retail Sales" regressions (Columns 3 and 4 of Table 2 and equivalent samples in other tables).

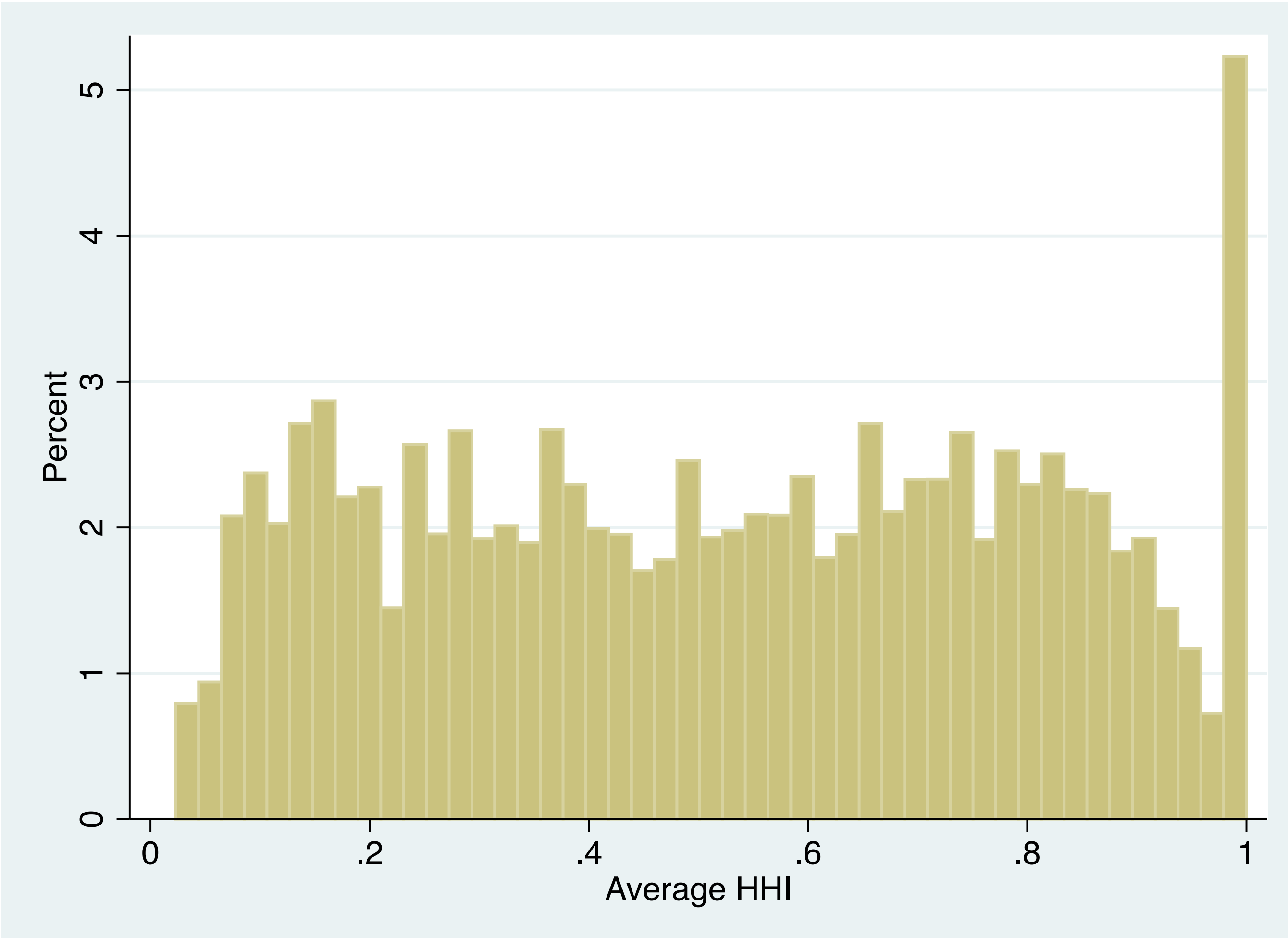

Figure 4. Distribution of County Avg. HHI Measure for Cashiers Occupational Labor Market
Notes: The figure plots the distribution across counties of the average HHI measure used in the "Cashiers" regressions (Columns 5 and 6 of Table 2 and equivalent samples in other tables).

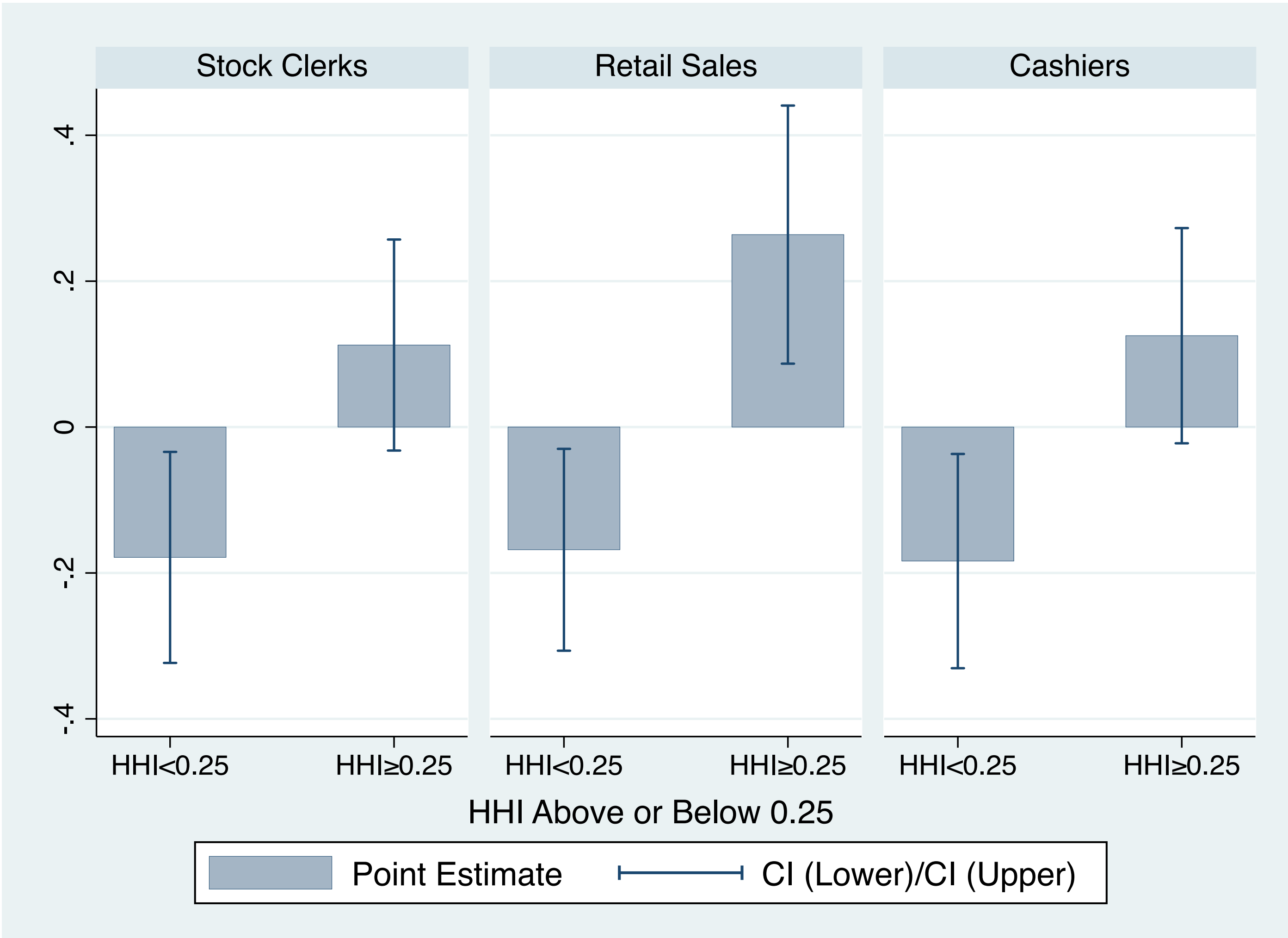

Figure 5. Employment Elasticities by High/Low Occupational Labor Market Concentration
Notes: The figure reports the estimated employment elasticity with respect to the minimum wage for occupational labor markets having high (HHI above 0.25) and low (HHI below 0.25) concentration. Each panel represents estimates using a different low-earning occupational labor market, with estimates corresponding to specifications in columns 2, 4, and, 6 of Table 2. As a reminder, Table 2 presents estimates of the employment elasticity with respect to the minimum wage, and, variation in this estimate across places with differing levels of labor market concentration for key low-wage occupations. All specifications in the table take the log of county-quarter general merchandise store employment as the outcome. In addition to the log of the governing minimum wage as a regressor, all specifications include the interaction of this variable with labor market concentration (HHI). We construct HHI for each of the three key low-wage occupational labor markets in the industry: stock clerks and order fillers (SOC 435081) in columns 1-2; retail salespersons (SOC 412031) in columns 3-4; cashiers (SOC 412011) in columns 5-6. In odd-numbered columns, this concentration measure is the county's Herfindahl-Hirschman Index (HHI) for the relevant labor market (averaged across all quarters of the 2010-2016 sample period), and, we report its interaction with the log minimum wage. In even-numbered columns, we instead report the interaction of log minimum wage with a binary concentration measure that separates high and low concentration labor markets based on whether their labor market HHI is above or below 0.25 (the Department of Justice/ Federal Trade Commission threshold for highly concentrated markets). All specifications further include county fixed effects, census division specific period fixed effects, state-specific linear time trends, and, the following additional control variables: log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate. Standard errors are clustered at the state level. 95% confidence intervals are shown.

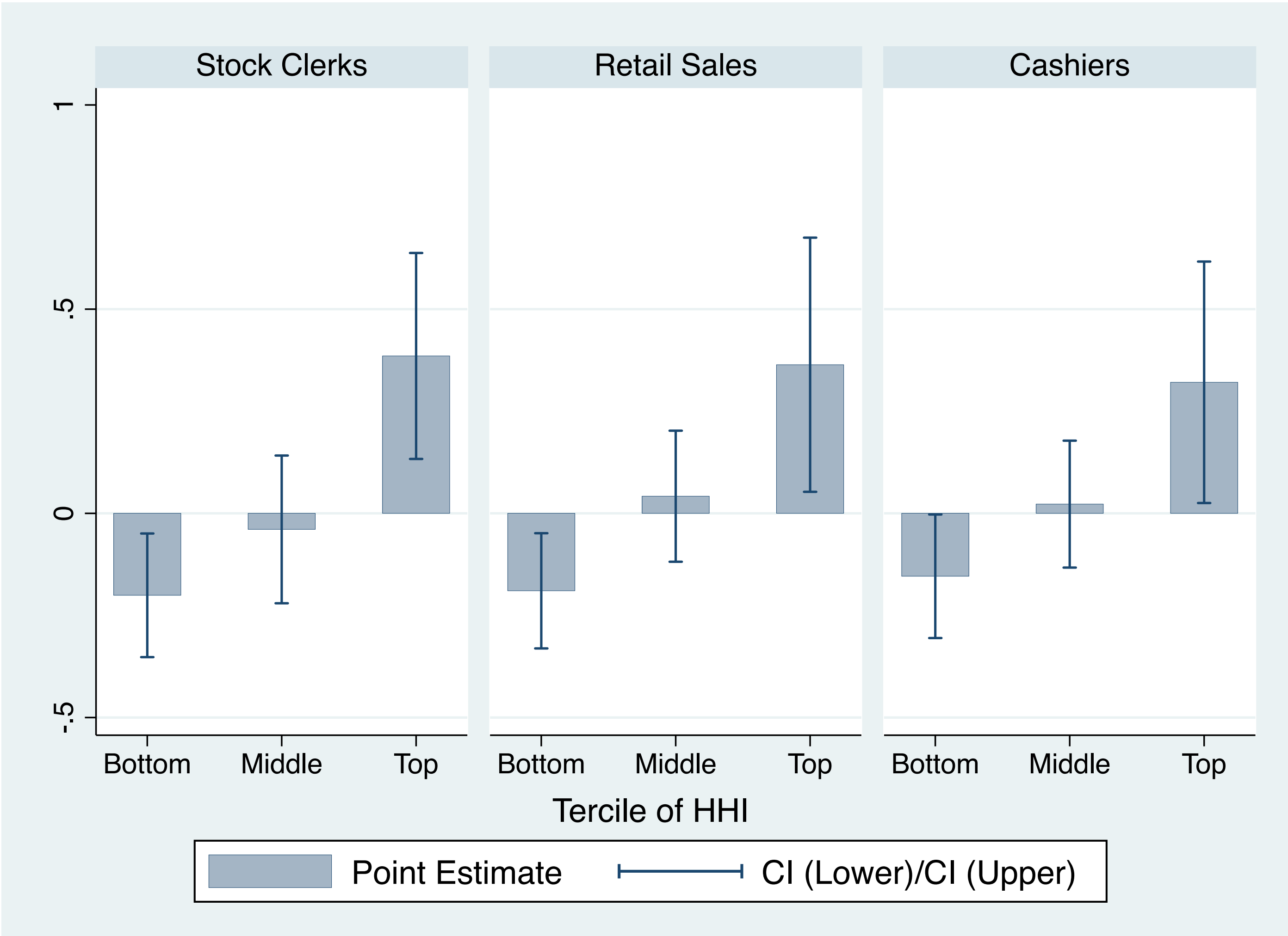

Figure 6. Employment Elasticities by Terciles of Occupational Labor Market Concentration
Notes: The figure reports the estimated employment elasticity with respect to the minimum wage for occupational labor markets having high, medium, and low concentration, as indicated by HHI tercile. Each panel represents estimates using a different low-earning occupational labor market, with estimates corresponding to specifications in columns 1, 2, and, 3 of Table 3. As a reminder, Table 3 presents estimates of the employment elasticity with respect to the minimum wage by labor market concentration (HHI) terciles. HHI is constructed for each of the three key low-wage occupational labor markets in the general merchandise store industry: stock clerks and order fillers (SOC 435081) in column 1; retail salespersons (SOC 412031) in column 2; cashiers (SOC 412011) in column 3. All specifications in the table take the log of county-quarter general merchandise store employment as the outcome. Specifications 1-3 are identical to specifications 2,4,6, respectively, in Table 2 with the exception that log minimum wage is here interacted with two indicator variables identifying high and medium HHI terciles (rather than simply a binary measure as in Table 2). All specifications include county fixed effects, census division specific period fixed effects, state-specific linear time trends, and, the following additional control variables: log of county total population, the log of total average weekly earnings (across all sectors) in the county, the log of total employment (across all sectors) in the county, the log of the county unemployment rate. 95% confidence intervals are shown.

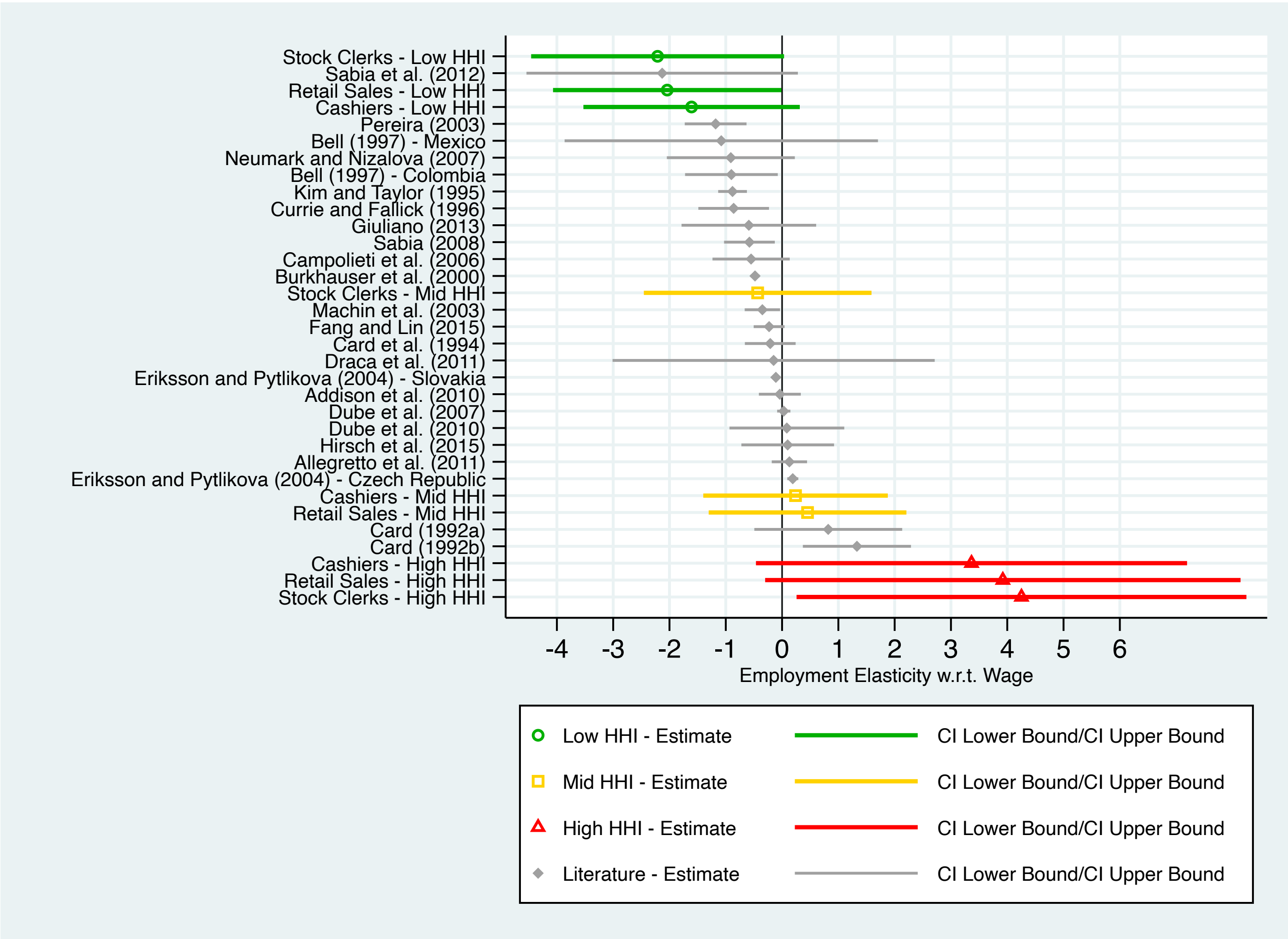

Figure 7. Our Elasticities Compared to Those in the Literature

Notes: The figure presents the employment elasticity with respect to the wage from other studies in the minimum wage literature (own wage elasticity of labor demand), as well as those estimated in this study based on Table 3 for the employment effect and Table 1 for the earnings effect. We utilize the review of the literature by Harasztosi and Lindner (2019). The estimates come from: Sabia, Burkhauser and Hansen (2012); Pereira (2003); Bell (1997); Neumark and Nizalova (2007); Kim and Taylor (1995); Currie and Fallick (1996); Giuliano (2013); Sabia (2009); Campolieti, Gunderson and Riddell (2006); Burkhauser, Couch and Wittenburg (2000); Machin, Manning and Rahman (2003); Fang and Lin (2015); Card, Katz and Krueger (1994); Draca, Machin and Van Reenen (2011); Eriksson and Pytlikova (2004); Addison, Blackburn and Cotti (2012); Dube, Naidu and Reich (2007); Dube, Lester and Reich (2010); Hirsch, Kaufman and Zelenska (2015); Card (1992*b*,*a*).

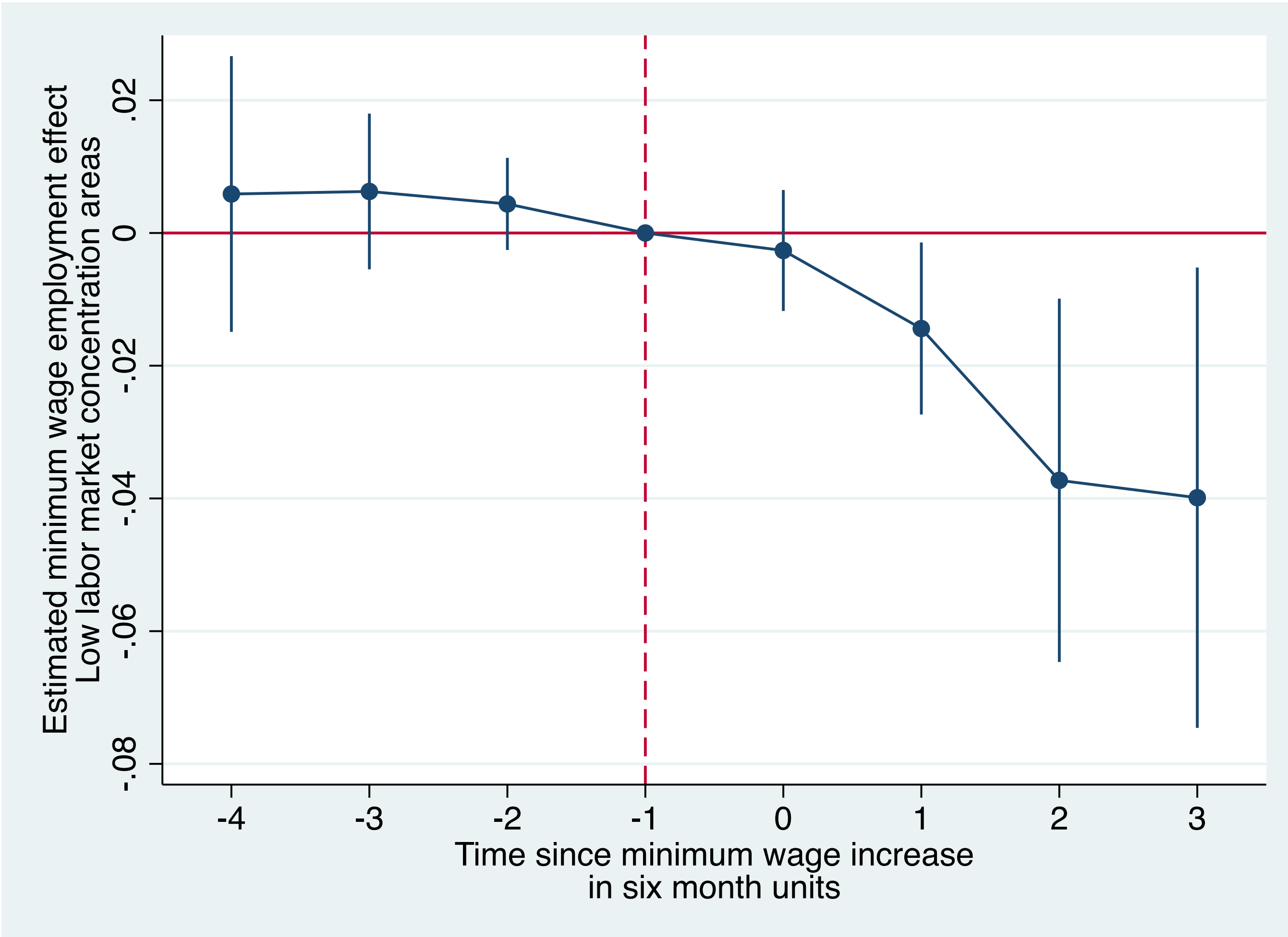

Figure 8. Event-study Estimates of Minimum Wage Employment Effect in Low Concentration Cashier Occupational Labor Markets

Notes: The figure reports the event-study estimates of α and ρ derived from the first event-study specification in Section 4.2.3 when using the sample of locations where cashier occupational labor market concentration levels are low ($HHI < 0.25$). Solid lines are 95% confidence intervals. The dependent variable is log employment in the general merchandise store industry. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase (with this reference time period indicated by the dashed vertical line). See text for further model details.

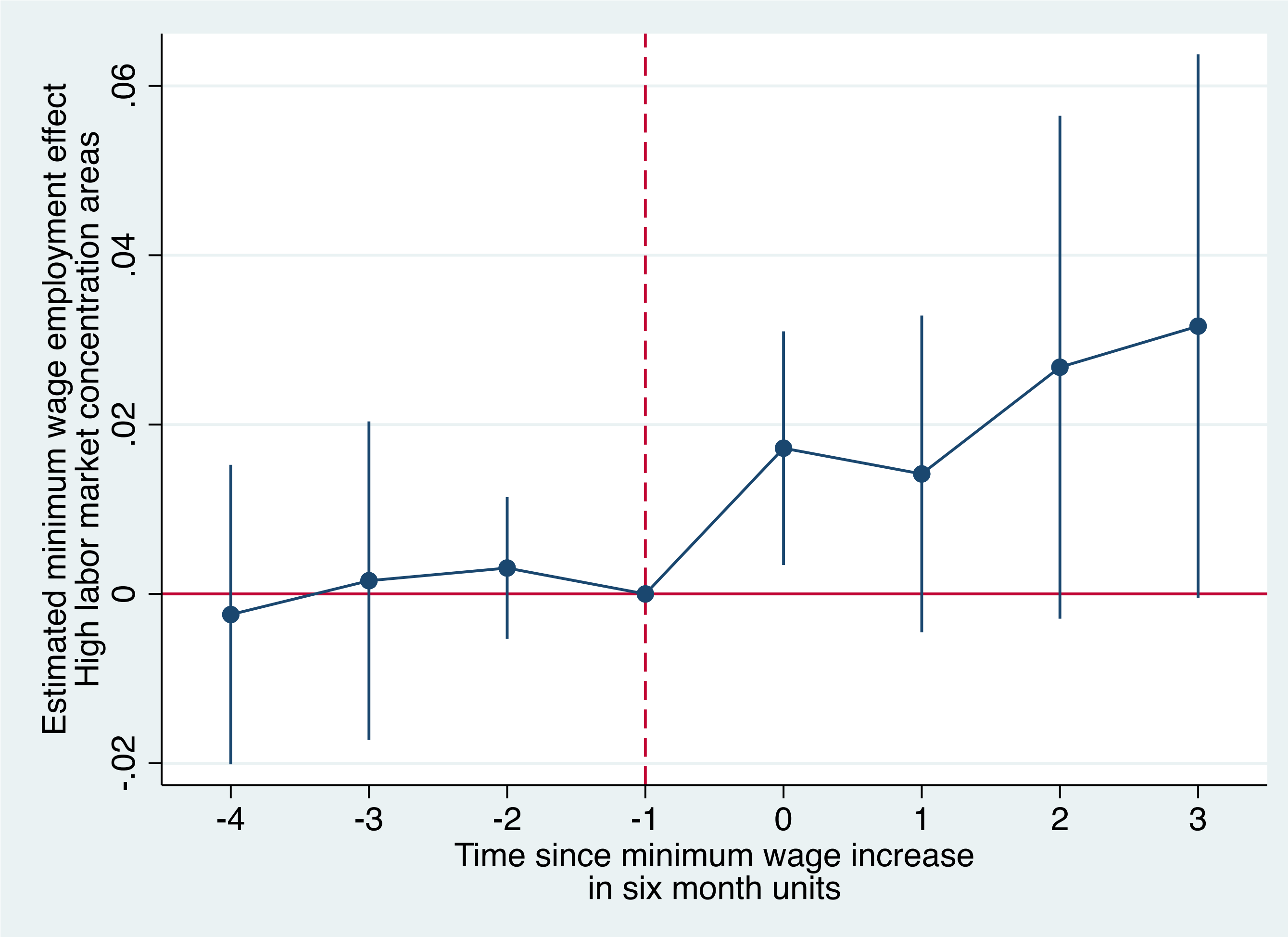

Figure 9. Event-study Estimate of Minimum Wage Employment Effect in High Concentration Cashier Occupational Labor Markets

Notes: The figure reports the event-study estimates of α and ρ derived from the first event-study specification in Section 4.2.3 when using the sample of locations where cashier occupational labor market concentration levels are high ($HHI \geq 0.25$). Solid lines are 95% confidence intervals. The dependent variable is log employment in the general merchandise store industry. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase (with this reference time period indicated by the dashed vertical line). See text for further model details.

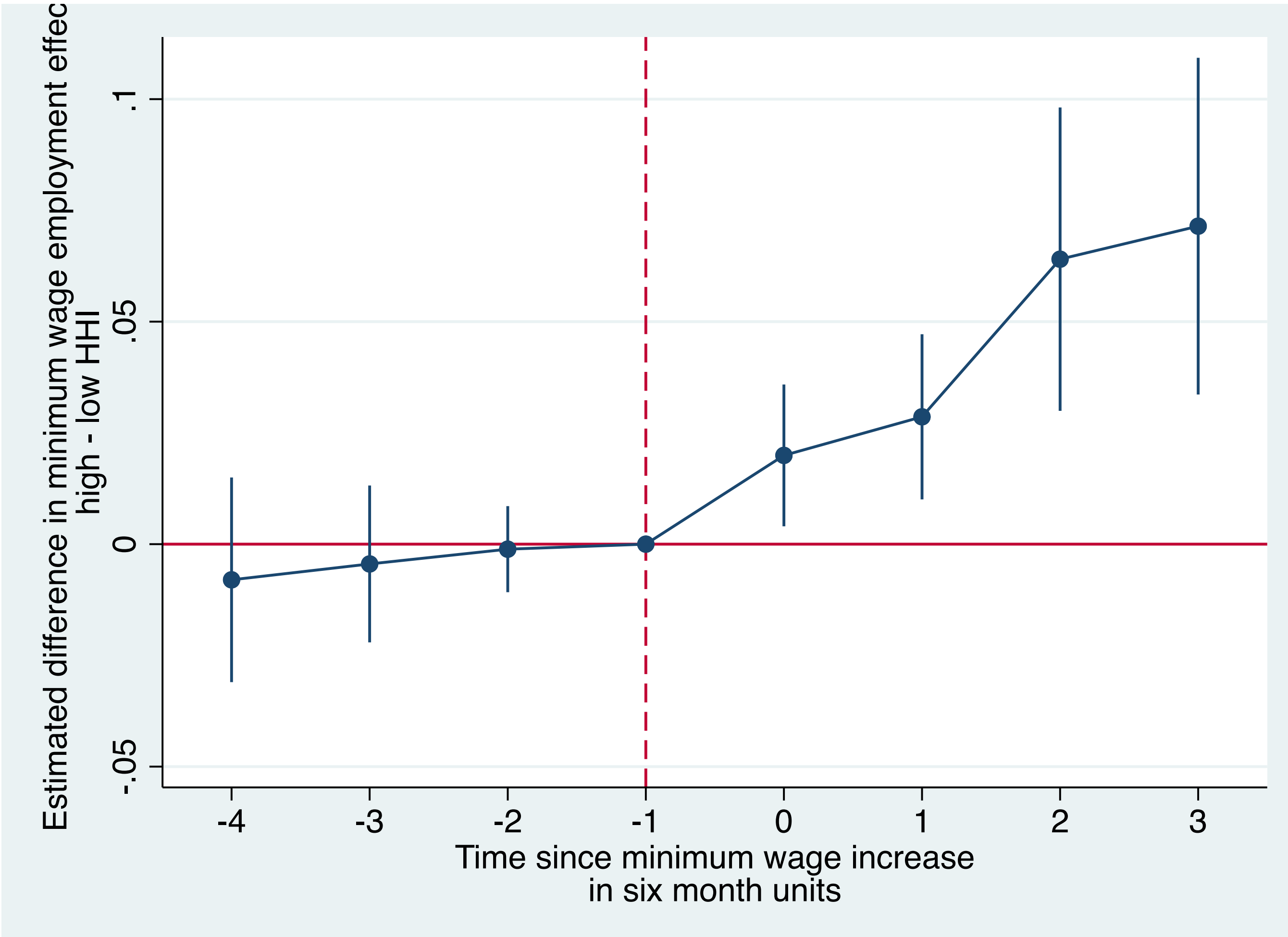

Figure 10. Estimated Difference in Minimum Wage Employment Effect in Areas with High vs. Low Concentration in the Cashier Occupational Labor Market

Notes: The figure reports the event-study estimates of λ and σ derived from the second event-study specification in Section 4.2.3 when defining high and low labor market concentration levels over the cashier occupational labor market. Solid lines are 95% confidence intervals. The dependent variable is log employment in the general merchandise store industry. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase (with this reference time period indicated by the dashed vertical line). Estimates measure the evolution in the difference in log general merchandise store employment between eventually treated counties and control counties in cashier labor markets with high concentration ($HHI \geq 0.25$) net of this difference in low concentration areas (relative to the difference in this difference in the six months before the minimum wage increased) after adjusting for model covariates. See text for further model details.

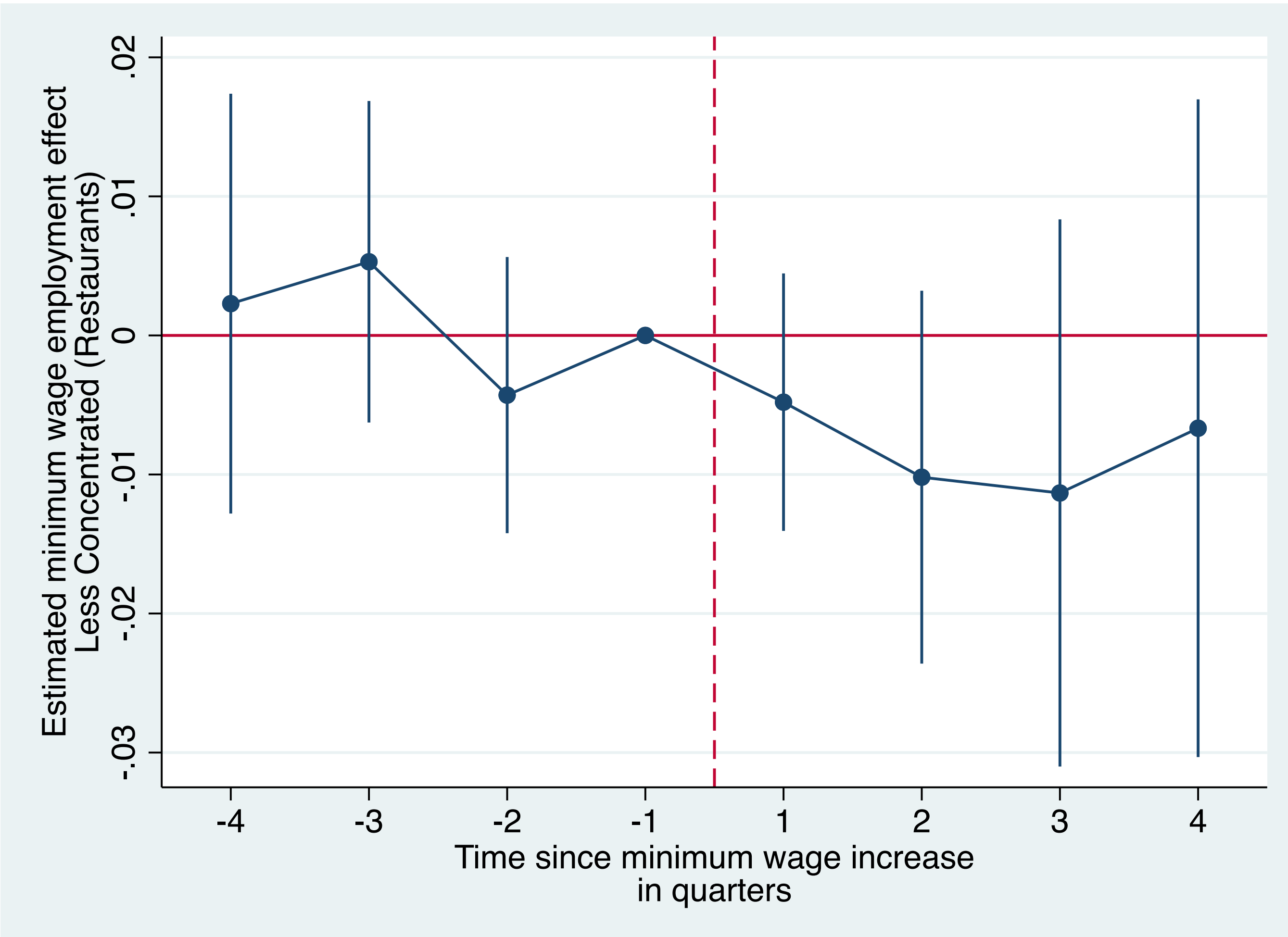

Figure 11. Restaurant Event-study Estimates of Minimum Wage Employment Effect in Low Concentration Labor Markets (no. of establishment proxy)

Notes: The figure presents an equivalent figure to Figure 8 but now for limited service restaurant industry employment. Specifically, the coefficients are event-study estimates of α and ρ derived from the first event-study specification in Section 4.2.3 when applied to the sample of low concentration locations as defined by our alternative proxy for concentration, the number of establishments in the industry. Low concentration locations with this metric are those above the bottom decile (this decile's median size is 4, which corresponds to the number of equally sized firms implied by the HHI cutoff of 0.25 used in the general merchandise analysis). Solid lines are 95% confidence intervals. The dependent variable is log employment in the limited service restaurant industry. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase (with this reference time period indicated by the dashed vertical line). See text for further model details.

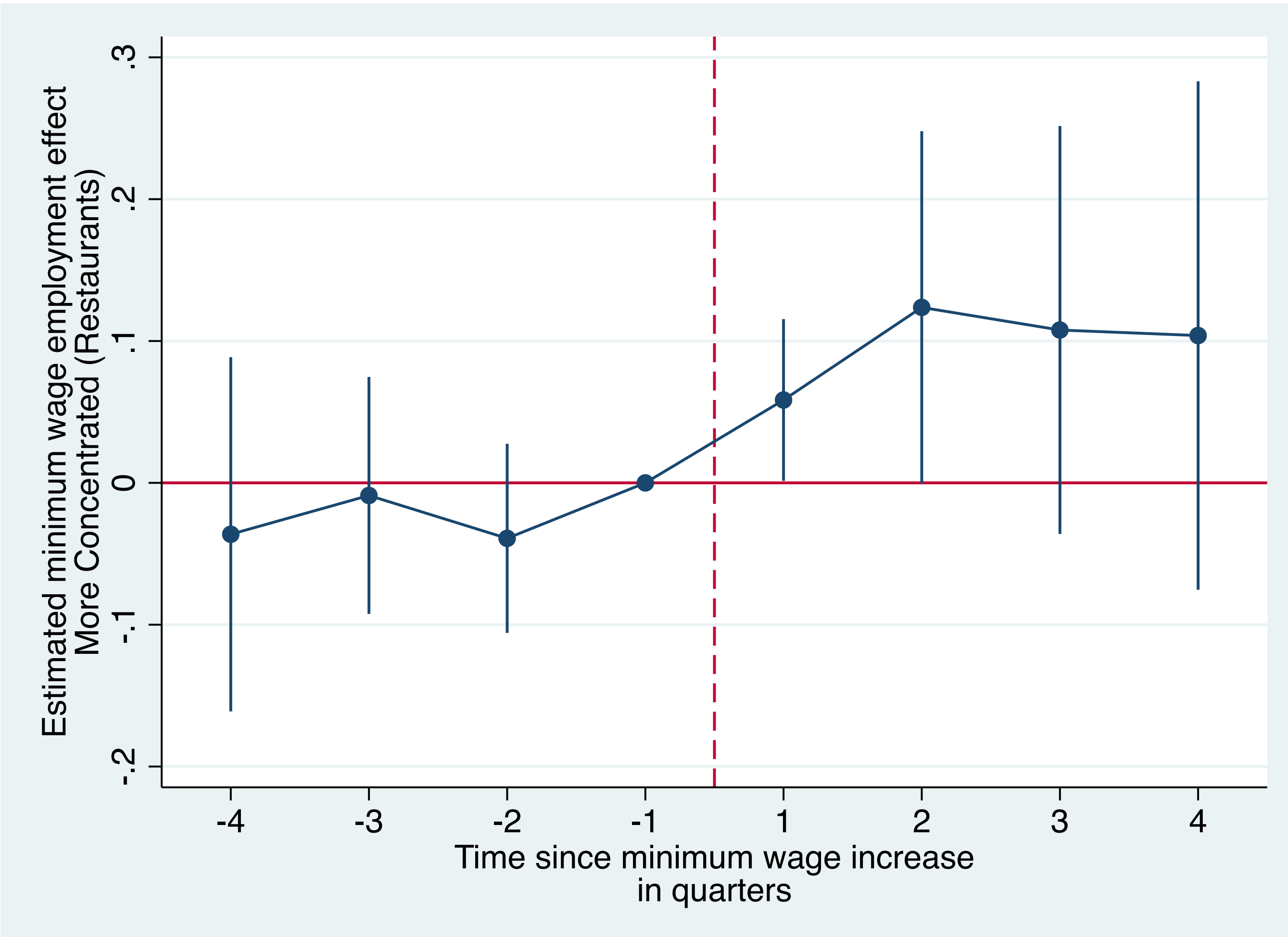

Figure 12. Restaurant Event-study Estimates of Minimum Wage Employment Effect in High Concentration Labor Markets (no. of establishment proxy)

Notes: The figure presents an equivalent figure to Figure 9 but now for limited service restaurant industry employment. Specifically, the coefficients are event-study estimates of α and ρ derived from the first event-study specification in Section 4.2.3 when applied to the sample of high concentration locations as defined by our alternative proxy for concentration, the number of establishments in the industry. High concentration locations are those in the bottom decile (this decile's median size is 4, which corresponds to the number of equally sized firms implied by the HHI cutoff of 0.25 used in the general merchandise analysis). Solid lines are 95% confidence intervals. The dependent variable is log employment in the limited service restaurant industry. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase (with this reference time period indicated by the dashed vertical line). See text for further model details.

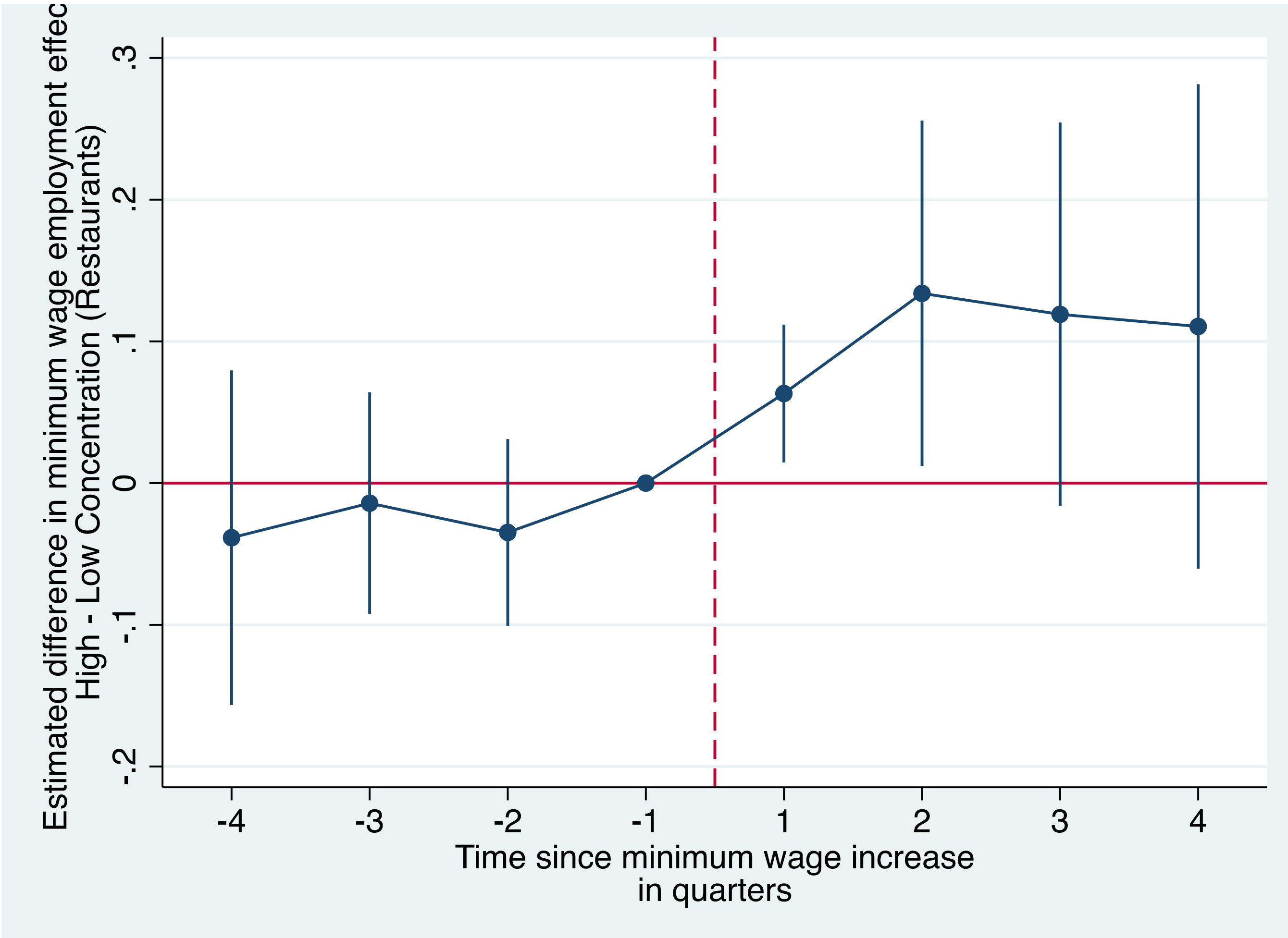

Figure 13. Estimated Difference in Restaurant Minimum Wage Employment Effect in Areas with High vs. Low Concentration (no. of establishments proxy)

Notes: The figure presents an equivalent figure to Figure 10 but for limited service restaurant industry employment instead. Specifically, the coefficients are event-study estimates of λ and σ derived from the second event-study specification in Section 4.2.3 when defining high and low labor market concentration levels using our alternative proxy for concentration, the number of establishments in the industry. High vs. low concentration compares the bottom decile of this distribution to the rest (this decile's median size is 4, which corresponds to the number of equally sized firms implied by the HHI cutoff of 0.25 used in the general merchandise analysis). Solid lines are 95% confidence intervals. The dependent variable is log employment in the limited-service restaurant industry. All effects are relative to a $t-1$ time window of 6 months prior to the minimum wage increase (with this reference time period indicated by the dashed vertical line). Estimates measure the evolution in the difference in log limited service restaurant employment between eventually treated counties and control counties in the high concentration markets net of this difference in low concentration areas (relative to the difference in this difference in the six months before the minimum wage increased) after adjusting for model covariates. See text for further model details.

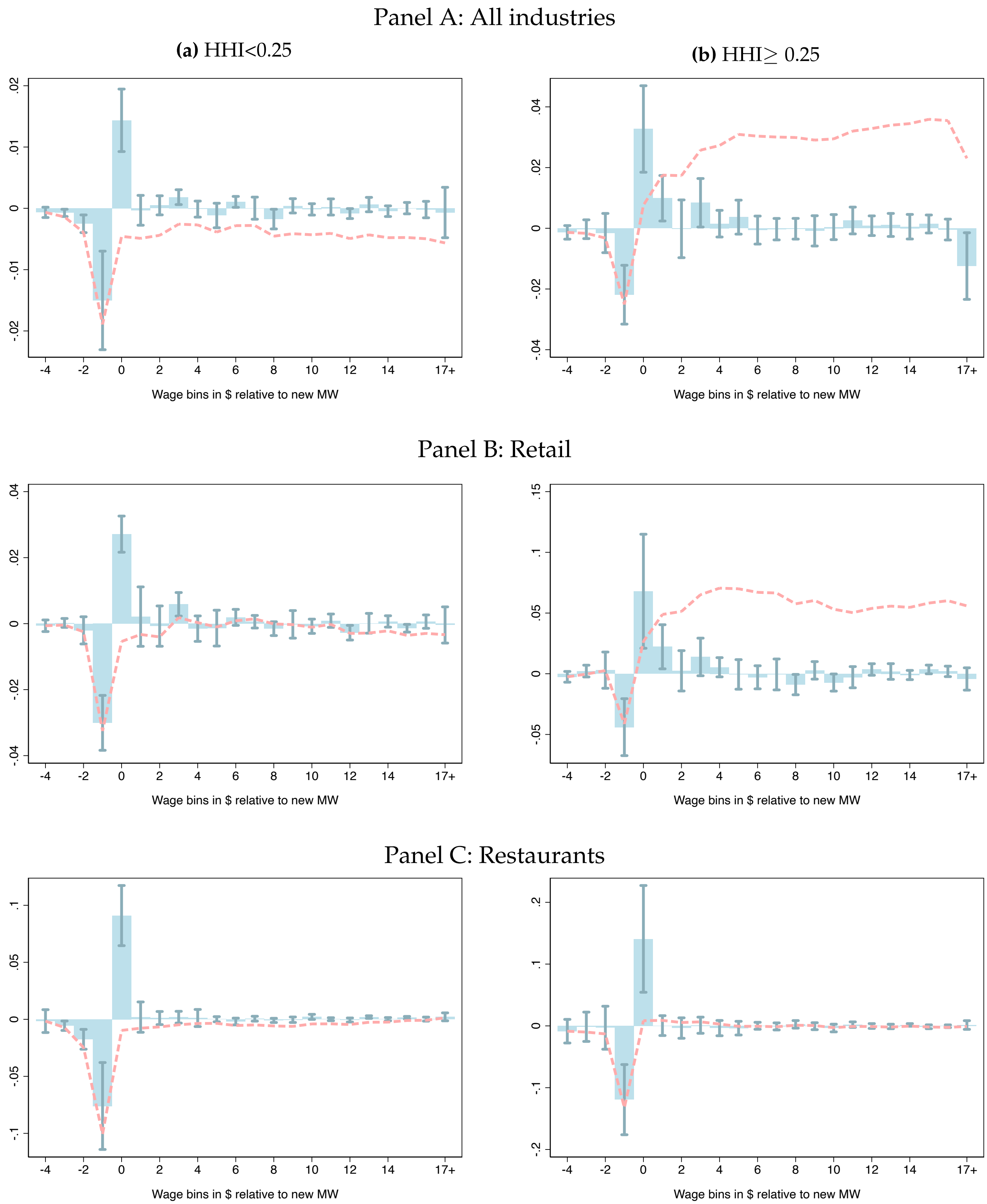

Figure 14. Impact of minimum wage on the wage distribution, by level of HHI
Notes: The Figure shows the bunching estimates of the impact of the minimum wage from Cengiz et al. (2019) for both low (HHI<0.25) and high (HHI≥0.25) concentration labor markets. HHI is the vacancy-weighted average of the county-SOC HHIs by State x Metropolitan/Non-Metropolitan pair. The dashed red line shows the running sum of employment changes up to the wage bin it corresponds to. The blue bars show for each dollar bin (relative to the minimum wage) the estimated average employment changes in that bin during the five-year post treatment relative to the total employment in the state one year before the treatment. The error bars show the 95% confidence interval using standard errors that are clustered at the state level. See Section 4.2.5 and the discussion of Figure 2 in Cengiz et al. (2019) for further details as we follow the methodology they outline using their replication files.

Table 5. Sources of Heterogeneity: HHI vs. Alternatives

	Dependent Variable: Log Employment			
Panel A. Interactions One by One	(1)	(2)	(3)	(4)
. . . where Interaction is with	HHI	Ln(PopDensity)	Ln(No. Estabs.)	Ln(Productivity)
Log Min Wage	-0.310***	0.259**	0.455***	1.332
	(0.092)	(0.128)	(0.133)	(0.925)
Log Min Wage × Interaction	0.728***	-0.0518**	-0.154***	-0.196
	(0.194)	(0.0212)	(0.0356)	(0.140)
adj. R^2	0.994	0.994	0.994	0.994
N	56,592	56,592	56,592	56,592
Panel B. Interactions All Together				
Log Min Wage	-1.178	-1.178	-1.178	-1.178
	(1.316)	(1.316)	(1.316)	(1.316)
Log Min Wage × Interaction	0.803**	0.074	-0.100	0.118
	(0.354)	(0.046)	(0.066)	(0.220)
adj. R^2	0.994	0.994	0.994	0.994
N	56,592	56,592	56,592	56,592

Notes: The table reports heterogeneity in the minimum wage employment elasticity by HHI and other variables considered in the text. All specifications in the table take the log of county-quarter general merchandise store employment as the outcome. Panel A presents coefficients on the log minimum wage base term and its interaction with HHI and each alternative one by one. Panel B presents coefficients on the base terms and the interaction coefficients when all the interactions are included together in the same specification. Column 1 in Panel A interacts log minimum wage with cashiers occupational labor market HHI (it is identical to the specification in Column 5 of Table 2; see the table's notes for more details). Column 2 in Panel A interacts log minimum wage with the log of population density (it is identical to the specification in Column 5 of Table A15; see the table's notes for more details). Column 3 of Panel A replicates the specification in column 1 except we replace HHI and the HHI interaction with the log of the number of general merchandise establishments (averaged throughout the period) and this variable's interaction with log minimum wage. Column 4 in Panel A interacts log minimum wage with a productivity proxy, log of total average earnings across all sectors (it is identical to the specification in Column 6 of Table A15; see the table's notes for more details). In Panel B we report coefficients when taking the baseline specification of Table 2 and including the HHI interaction, the population density interaction, the number of establishments interaction, and, the productivity interaction (and each base term) in the same specification. As the panel reports coefficients on the respective interaction terms across the 4 columns (all drawn from the same regression) the coefficient on the log minimum wage regressor is the same across the panel's 4 columns. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$