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The Minimum Wage, Restaurant Prices, and Labor Market Structure

Daniel Aaronson, Eric French, and James MacDonald

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Abstract

Using store-level and aggregated Consumer Price Index data, we show that restaurant prices rise in response to minimum wage increases under several sources of identifying variation. We introduce a general model of employment determination that implies minimum wage hikes cause prices to rise in competitive labor markets but potentially fall in monopsonistic environments. Furthermore, the model implies employment and prices are always negatively related. Therefore, our empirical results provide evidence against the importance of monopsony power for understanding small observed employment responses to minimum wage changes. Our estimated price responses challenge other explanations of the small employment response too.

*Comments welcome at `efrench@frbchi.org` or `daaronson@frbchi.org`. Author affiliations are Federal Reserve Bank of Chicago, Federal Reserve Bank of Chicago, and Economic Research Service, U.S. Department of Agriculture respectively. Work on the store-level price data was performed under a memorandum of understanding between the Economic Research Service at the Department of Agriculture and the Bureau of Labor Statistics (BLS), which permitted onsite access to the confidential BLS data used in this paper. We thank Bill Cook, Mark Bowman, and Scott Pinkerton of the BLS for their advice and help. We also thank Gadi Barlevy, Jeff Campbell, William Evans (the editor), Bob LaLonde, Derek Neal, Dan Sullivan, the referees, and seminar participants at the Federal Reserve Bank of Chicago, the University of Illinois-Urbana, the Econometric Society, and SOLE meetings for helpful comments and Tina Lam for excellent research assistance. The views expressed herein are not necessarily those of the Federal Reserve Bank of Chicago, Federal Reserve System, Bureau of Labor Statistics, or U.S. Department of Agriculture. Author correspondence to Daniel Aaronson or Eric French, Federal Reserve Bank of Chicago, 230 S. LaSalle St., Chicago, IL 60604. Telephone (312)322-6831, Fax (312)322-2357.

1 Introduction

This paper utilizes unique data to test whether restaurant prices respond to minimum wage changes. We find that restaurant prices unambiguously rise after minimum wage increases are enacted.[1] Furthermore, these price increases are larger for establishments that are more likely to pay the minimum wage. These results are derived from a panel of store-level restaurant prices that are the basis for the food away from home component of the Consumer Price Index (CPI) during a three-year period with two Federal minimum wage increases, and are corroborated using a longer panel of city-level food away from home pricing from the CPI.

Because of the breadth of our price data, we can take advantage of several sources of variation. First, some states set their minimum wage above the Federal level. Second, we can distinguish restaurants that tend to pay the minimum wage from those that do not. Third, the fraction of workers paid at or near the minimum wage varies across geographic areas. All three sources of variation indicate that most, if not all, of the higher labor costs faced by employers are pushed onto customers in the form of higher prices.

As suggested by Brown (1999), the size and sign of these price responses can be used to infer whether monopsony power is important for understanding the employment response to minimum wage hikes. The minimum wage literature has become contentious since Card and Krueger's (e.g. 1995, 2000) research found that an increase in the minimum wage has no, or even a small positive, effect on employment. Therefore, their research contradicts standard models of competitive labor markets, which, prior to their work, most researchers suspected was relevant for industries which primarily employed minimum wage workers.[2]

[1] We are not the first to estimate price pass-through in this context. See, e.g., Converse et al (1981), Card and Krueger (1995), and Aaronson (2001). Card and Krueger use Consumer Price Indices for Food Away from Home in 27 large metropolitan areas over a three year period, finding larger price increases in those cities with higher proportions of low-wage workers. Although their estimates are consistent with full pass-through, their standard errors are extremely large. They cannot reject zero price pass-through in many of their specifications. Moreover, additional evidence from specific state increases in Texas and New Jersey suggests close to no price response. As a result, they conclude that their estimates are "too imprecise to reach a more confident assessment about the effects of the minimum wage on restaurant prices." The size of the panel that we use in this study allows us to estimate price effects much more precisely.

[2] See, for example, Brown et al (1982). The standard competitive model's predictions are generally consistent with recent views reported in a survey of leading labor economists (Fuchs, Krueger, and Poterba 1998) as well. However, a full quarter of respondents believe there is no teenage disemployment effect from a 10 percent increase in the minimum wage.

However, their results are consistent with monopsony power in the labor market, as Stigler (1946) discussed many years ago. The diverse findings reported in the flurry of replies to their work (e.g. Neumark and Wascher (1996, 2000), Deere et al. (1995), Kim and Taylor (1995), Burkhauser et al. (2000), Dickens et al (1999)) led one observer to note that "[Card and Krueger's] lasting contribution may well be to show that we just don't know how many jobs would be lost if the minimum wage were increased...and that we are unlikely to find out by using more sophisticated methods of inference on the existing body of data. What is needed is more sophisticated data" (Kennan 1995).

Restaurant prices complement the existing evidence on employment responses because, as we show below, output prices and employment are unambiguously negatively related in response to an exogenous change in wage rates. In order to show this relationship, we introduce a general model of employment determination that allows for a range of output and input market structures. Part of the reasoning behind the negative relationship between output prices and employment is based on the negative relationship between prices and output. We also add some weak assumptions to the model to show that output and labor input are positively related. Therefore, if the output price rises in response to a minimum wage hike, both output and labor input have fallen. This will be the case if labor markets are competitive. Conversely, if the output price falls in response to a minimum wage hike, total output and labor input have increased. This will potentially be the case if firms have monopsony power in the labor market.

Research on monopsony power has recently been revitalized by the empirical and theoretical work of Card and Krueger (1995), Burdett and Mortenson (1998), Bashkar and To (1999), Ahn and Arcidiacono (2003), Flinn (2006), Manning (1995), and Rebitzer and Taylor (1995).[3] But our results suggest that competition is likely more relevant than monopsony. Moreover, in Aaronson and French (2007), we show that a computational model of labor demand with a competitive labor market structure predicts price responses that are comparable to those found in this paper. The employment elasticities that are derived from that

[3] Although few believe that low wage labor markets are characterized by pure monopsony, as in Stigler (1946), many models give rise to monopsony-like behavior that corroborate Card and Krueger's findings of small or even positive employment movements after a minimum wage increase. These include models where transportation (Bashkar and To 1999) or employee search (Burdett and Mortenson 1998; Ahn and Arcidiacono 2003; Flinn 2006) is costly and employers are not able to discriminate high and low reservation wage workers. Efficiency wage models such as Manning (1995) and Rebitzer and Taylor (1995) can also cause monopsony-like employment effects. See Boal and Ransom (1997) and Manning (2003) for broader reviews.

calibrated model are within the bounds set by the empirical literature.

To be clear, as Boal and Ransom (1997), among others, point out, our results do not necessarily prove labor markets are competitive. Although the results are clearly consistent with this conclusion, if the minimum wage is set high enough, positive comovement between the minimum wage and prices may be consistent with the monopsony model as well. We discuss this point more formally below.

However, our results provide evidence against the hypothesis that monopsony power is important for understanding the observed small employment responses found in some minimum wage studies. Indeed, our estimated price responses provide evidence against other explanations of the small employment response as well, including the potential substitution of nonwage for wage compensation and the importance of endogenous work effort or efficiency wages. They do, however, provide support for a model of "hungry teenagers," whereby higher income resulting from a minimum wage increase causes low wage workers to buy more minimum wage products, attenuating the disemployment effect of the minimum wage. Although our test answers a fairly narrow question, we believe that the answer to this question is of broad interest. Given that the low observed employment responses to minimum wage changes sparked particular interest in the importance of monopsony power in the labor market, our results should temper this interest.

Finally, it is important to emphasize that our estimates are for the restaurant industry only. This industry is a major employer of low-wage labor and therefore a particularly relevant one to study.[4] However, as a result of different intensities of use of minimum wage labor, substitution possibilities, market structure, or demand for their products, other industries might face different employment responses. See Aaronson and French (2007) for further details.

[4]Eating and drinking places (SIC 641) is the largest employer of workers at or near the minimum, accounting for roughly a fifth of such employees in 1994 and 1995. The next largest employer, retail grocery stores, employs less than 7 percent of minimum or near minimum wage workers. Moreover, the intensity of use of minimum wage workers in the eating and drinking industry is amongst the highest of all sectors, with approximately 23 percent of all workers, encompassing 11 percent of the industry wage bill, within 10 percent of the minimum wage. All calculations in this footnote are based on the Current Population Survey's outgoing rotation groups. Other prominent examples of studies that concentrate on the restaurant industry include Katz and Krueger (1992), Card and Krueger (1995,2000) and Neumark and Wascher (2000).

2 Data

Under an agreement with the Bureau of Labor Statistics (BLS), we were granted access to the store-level data employed to construct the food away from home component of the CPI during 1995 to 1997.[5] While the time frame is short, this three-year period contains an unusual amount of minimum wage activity. A bill signed on August 20, 1996 raised the federal minimum from \$4.25 to \$5.15 per hour, with the increase phased in gradually. An initial increase to \$4.75 (11.8 percent) occurred on October 1, and the final installment (8.4 percent) took effect on September 1, 1997. Moreover, additional variation can be exploited since price responses will vary geographically. This occurs for two reasons. First, market wages may exceed minimum wages in some areas but not in others. Second, some states set minimum wages above the federal level. We capture the latter source of heterogeneity by allowing the effective minimum wage to be the maximum of the state and federal level.[6]

The sample itself is based on nearly 7,500 food items at over 1,000 different establishments in 88 Primary Sampling Units (PSUs).[7] Because restaurants in some geographic areas are surveyed every other month, all numbers are reported as bimonthly (every other month) price changes.[8] Within an establishment, specific items, usually 7 or 8, are selected for pricing with probability proportional to sales. During our time frame, an "item" usually was a meal, as the BLS aimed to price complete meals as typically purchased (for example, a meal item might consist of a hamburger, french fries, and a soft drink). Our dataset codes items broadly, such as breakfast, lunch, dinner, or snacks. Unfortunately, because there are no specific item descriptions, we cannot tie price changes to item-specific measures of input price changes (such as ground beef or chicken price indexes). Nevertheless, the BLS strives to price

[5] Because the BLS introduced a complete outlet and item resampling in January 1998, we only use data through December 1997. Data prior to 1995 are no longer available. Bils and Klenow (2004) use the same 1995 to 1997 period.

[6] This source of variation is especially useful in section 3.3, when we look at city level variation between 1979 and 1995. During 1995 to 1997, 10 states (not including Alaska which is always \$0.50 above the federal level) had minimum wages above the federal level for some part of the period. Six states (HI, MA, NJ, OR, VT, and WA) were at or above the federal minimum of \$4.75 prior to October of 1996. Three states (HI, MA, and OR) were at or above the federal minimum prior to September 1997. Further variation is available from states (CA, CT, NJ, WA) that were between the old \$4.75 minimum wage but below the new \$5.15 minimum wage.

[7] The 88 PSUs cover 76 metropolitan statistical areas and 12 other areas representing the urban non-metro U.S.

[8] PSUs were assigned to one of three reporting cycles: outlets in the five largest were surveyed each month, while others were surveyed in two bimonthly cycles of odd and even numbered months. For sample size and consistency, we randomly assigned outlets in the five largest PSUs to odd or even two month cycles.

identical items over time, and codes in our database describe temporal item substitutions due to discontinuance and alteration. Our analysis focuses on price changes for identical items, and we do not compare prices where the BLS has made an item substitution.[9]

A particular advantage of this data is its depiction of the type of business. Limited service (LS) outlets, which account for roughly 30 percent of the sample, are those stores where meals are served for on- or off-premises consumption and patrons typically place orders and pay at the counter before they eat. Roughly half the sample is comprised of full service (FS) outlets, establishments that provide wait-service, sell food primarily for on-premises consumption, take orders while patrons are seated at a table, booth or counter, and typically ask for payment from patrons after they eat.[10] The minimum wage is likely to increase wages at LS restaurants more than at FS restaurants, for two reasons. First, wages for cashiers and crew members are higher, perhaps by 60%, at FS restaurants than LS restaurants.[11] Thus, a higher fraction of workers are paid the minimum wage at LS restaurants. Second, many FS employees are paid through tips and the federal tipped minimum wage remained $2.13 throughout our sample period. Thus, minimum wage changes have smaller effects on wages

[9]Firms could respond to a minimum wage increase by reducing quality, instead of raising price. While we do not have direct measures of quality, the dataset notes whether an item is the same as the item priced in the previous month. There is no evidence of any increased incidence of item changes or substitutions following minimum wage increases, suggesting that quality changes or item substitution are not a standard means of dealing with a cost shock. There also might be concern that a minimum wage increase changes the composition of items sampled. If revenues are negatively correlated with prices and sampling probability is a function of sales, a change in the minimum wage could result in a shift in the distribution of sampled products towards high priced items (or stores with fewer minimum wage workers). To minimize this concern, we ran everything with sampling weights and found identical results.

[10]The BLS replaced an old ordering with these types of business codes in July 1996, and began to report price indexes for type of business groupings after our data period in January 1998. Businesses surveyed early in our sample period were retroactively assigned the new codes. The remaining fifth of non-LS and non-FS outlets, which we usually exclude from this analysis, include meals consumed at department stores, supermarkets, convenience stores, gas stations, vending machines, and many other outlets.

[11]Assuming that LS wage rates are identical to U.S. McDonald's wage rates collected by McKinsey Global Institute and reported in Ashenfelter and Jurajda (2001), then wage rates among cashier and crew members in FS establishments in the outgoing rotation files of the CPS are about 60 percent higher than wage rates in LS establishments. Ashenfelter and Jurajda report the average U.S. McDonald's wage for crew and cashier workers was $6.00 and $6.50 in December 1998 and August 2000, respectively. We compared these figures to the average wage of $7.81 and $8.52 for CPS workers that report their industry as eating and drinking places and their occupation as food preparation and service occupation, janitors and cleaners, or sales counter clerks during the fourth quarter of 1998 or the third quarter of 2000. Assuming all LS establishments pay the same wage as McDonalds and noting that the 1997 Economic Census of Accommodations and Food Services reports that 48 percent of all employees in FS and LS establishments are employed in the LS sector, we can back out that FS establishments pay roughly 60 percent higher hourly wages than LS establishments within these occupation codes. The Economic Census also reports average weekly wages that are approximately 20 percent higher in FS establishments. But this calculation does not correct for differences in hours worked per week and cannot refine the sample by occupational class.

in FS outlets than LS outlets because tip earnings usually exceed effective minimum wages.[12]

3 Estimates of Price Pass-Through

One problem with price (as well as employment) data is that they are potentially measured with error.[13] In other words, $p_{ijkt} = p^*_{ijkt} + \epsilon_{ijkt}$, where p_{ijkt} is the measured price of item k at outlet j in state i during month t, p^*_{ijkt} is the model predicted value, where the model is described in section 4, and ϵ_{ijkt} is measurement or model misspecification error.[14] We approach the data in three ways.

3.1 Store-level Descriptions of Price Increases and Decreases

Our first approach ignores errors in the price data and simply tabulates price increases and decreases after a minimum wage change. In the model described in section 4, we formally show that price data can be used to infer labor market structure. In particular, price cuts tend to be an outcome unique to monopsonistic labor markets. In the absence of measurement error, observed price cuts allow us to identify individual firms that potentially have monopsony power. On the other hand, if variability in ϵ_{ijkt} is significant, measurement or misspecification error may cause us to erroneously infer monopsony power when in fact none exists.

[12] Federal law sets a separate cash minimum for tipped employees (which is \$2.13 throughout our sample period), but requires that tips plus cash wages must at least equal the nontipped employee minimum. For example, in September 1996, \$2.62 (\$4.75-2.13) in tips are allowed to be applied to a tipped employee's wage to reach the minimum wage. In 1996, only Rhode Island and Vermont changed their state-specific tipped minimum. In 1997, Maryland, Michigan, North Dakota, and Vermont did as well. Of these states, only Maryland and Michigan are included in our CPI sample.

[13] Measurement error is unlikely to be important in this data (Bils and Klenow 2004). The BLS has procedures in place to flag and investigate unusual observations. Nevertheless, some measurement error may exist for the following three reasons. First, the price on the menu may differ from the price paid because of discounts and coupons distributed outside of outlets. Prices are collected net of sales and promotions but some, particularly those not run by the outlet itself may be missed. Second, high frequency but short-term price changes may not be captured by our monthly data (Chevalier et al 2003). Third, surveyors may falsely report last month's price instead of going to the restaurant to record the price, a practice known as "curbstoning". There is no reason to think that any of these sources of error are correlated with the presence of a minimum wage change.

[14] There are other interpretations for ϵ_{ijkt}, such as menu costs of switching prices. Moreover, many firms in our sample offer short term sales for reasons that are unrelated to changes in input prices.

Outlet type	Limited Service	Limited Service	Full Service	Full Service
Two month period with minimum wage change	No	Yes	No	Yes
A. Share of price changes, observation is an item				
Percent increases	11.5	22.6**	10.7	12.0**
Percent decreases	2.9	2.5	1.8	1.6
Item Observations	25,815	3,853	44,632	7,045
B. Share of price changes, observation is the average of all items at a store				
Percent increases	24.1	38.3**	19.5	22.4
Percent decreases	8.0	6.7	5.1	4.8
Store Observations	3,799	551	6,809	1,036
C. Size of price changes (in percent)				
Mean item price change—increase	5.3	4.8**	4.8	4.9
Mean item price change—decrease	8.4	8.2	7.5	9.3*

* (**) = Statistically different from months without a minimum wage increase at the 5(1) percent level.

Table 1: The Frequency and Magnitude of Store-Level Price Changes

Table 1 reports descriptive statistics on the frequency and size of price changes for Limited Service (column 2) and Full Service (column 4) outlets in the two months immediately after a minimum wage change. For comparison, all other two month periods are reported in columns 1 and 3. In panel A, the observational unit is a food item. Because multiple items are surveyed for each store, individual stores are in these computations up to 8 times each period. Panel B computes price changes by store. That is, an average price is calculated from each store's sampled items and consequently a store is included, at most, once every two months.

There are several notable features of the data. First and foremost, prices increase in response to a minimum wage change. During the two months after a minimum wage increase, 22.6 percent of LS items increased in price. This is almost double the 11.5 percent share of LS price quotes that are increased in months without a minimum wage increase.[15] Moreover, as expected, the minimum wage effect is substantially smaller, although still statistically significant, for full service outlets. In such stores, the share of quotes that are higher than the previous two months is 12 percent, exceeding months that do not follow a minimum wage increase by 1.3 percentage points. Excluding small price changes (say, those less than 2 percent) that might be driven by measurement error reinforces these differences. Price increases over 2 percent are more than twice as likely in minimum wage months in LS establishments (17.6 versus 8.6 percent) and 22 percent more likely in FS establishments (8.9 versus 7.2 percent).

Conversely, there is little evidence that minimum wage increases cause price declines. The share of prices that decline is stable throughout the three years regardless of whether the minimum wage has been altered. The results are identical when small changes are excluded. Therefore, based on incidence alone, the data suggest that many firms raise their price, but few reduce it, in response to a change in the minimum wage.

These results are robust to looking at price movements at the store-level, as reported in panel B. Here, prices are computed by averaging the price of all items in a store. Again, there is a notable acceleration of LS outlet price increases following a minimum wage increase (38.3 percent versus 24.1 percent) but no unusual increase in price declines during these periods.[16]

[15] The unconditional probability of a price change matches Bils and Klenow (2004), who also use CPI microdata.

[16] We have explored looking at longer intervals but are quite limited by our data. The September 1997 increase is only 4 months before the end of our sample, so we are forced to rely on a single comparison: pre- and post- the 1996 increase. The results are a bit more muted but similar inferences can be drawn. LS price

Finally, increasing the frequency of price changes is not the only avenue for firms to raise or lower prices. The size of price changes could be altered as well.[17] However, panel C shows that, if anything, price increases tend to be slightly smaller in size after minimum wage increases. The size of LS price cuts are unaffected by minimum wage changes. There is some evidence of larger price cuts in FS establishments after minimum wage hikes, but this is a rare event (only 1.6 percent of all FS item observations). It is worth noting that, while price cuts (in FS and LS stores) are rare, they can be large. Over a quarter of all food away from home price cuts exceed 10 percent, and over a tenth exceed 20 percent. By comparison, price increases are more tightly concentrated, with about half under 4 percent and less than a tenth above 10 percent.[18] However, there is little evidence that the size of price cuts change in any meaningful way after a minimum wage increase. Kolmogorov-Smirnov D-tests for differences in price distributions finds no significant shift in the size of price cuts following minimum wage increases.[19] The fact that these large declines exist in roughly the same fashion in non-minimum wage change periods suggests to us that the largest cuts typically reflect temporary sales.

Assuming that all markets are competitive (or, as we point out in Section 4.3, factor markets are competitive and product markets are monopolistically competitive) and firms have a constant returns to scale production function, it is straightforward to show that all cost increases will be passed onto consumers in the form of higher prices. If minimum wage labor's share of total firm costs is s_{min}, then a 10% increase in the minimum wage should increase the product price $s_{min} \times 10\%$.

To get a sense of whether the observed price responses are consistent with competition, we note that minimum wage labor's share of total costs is equal to labor's share of total

hikes in the 10 months after October 1996 are 6 percentage points (53 vs. 47 percent) more common than in the average 10 month period in the year and a half prior to October 1996. Price declines occur slightly more frequently (but not statistically so) in the 10 months after the increase: 6 percent versus 5 percent.

[17] See MacDonald and Aaronson (2006) for a more extensive discussion of the various ways restaurants construct price changes. The price change distributions are available upon request from the authors.

[18] Price changes cluster near the mean, with excess kurtosis of 62.0 and 80.8 for price changes among LS and FS outlets, respectively. Distributions of increases and decreases are also quite peaked compared to normal distributions, with excess kurtosis of 14.2 (LS) and 8.6 (FS) for increases, and 1.6 (LS) and 6.8 (FS) for decreases. Kashyap (1995) also reports positive excess kurtosis in his sample of catalog prices.

[19] In particular, the K-S D-test suggests no significant difference in the distribution of increases among LS outlets. Small changes (less than 2 percent) occur among 22 percent of LS price changes in minimum wage bimonths, compared to 25 percent in all other bimonths. Large changes (greater than 10 percent) occur among 12 percent of LS price changes observed in minimum wage bimonths, compared to 13 percent in all other bimonths. There is a statistically significant difference, driven by the higher incidence of very small increases, among FS outlets.

costs multiplied by minimum wage labor's share of labor costs. 10-K company reports, the Economic Census for Accommodations and Foodservices, and the IRS' Statistics on Income Bulletin all provide an estimate of labor's share of total costs, and in each, the sample median and mean are around 30 to 35 percent.[20] Unfortunately, we are less certain of minimum wage labor's share of total labor costs for the average firm. Using household level data, we know that about a third of all restaurant workers are paid near the minimum wage over this time period, constituting 17% of all payments to labor.[21] Using these values, we make two calculations that bound the competitive response. If there is only one type of labor, all firms have the same employment level, and all firms either pay 0% of their workers or 100% of their workers the minimum wage, depending on the labor market, then 33% of all firms pay the minimum wage. Given this, and the fact that about 33% of total costs are in the form of labor costs, then a 10% increase in the minimum wage raises prices by $10\% \times 33\% \times 33\% = 1.09\%$. Alternatively, if all firms hire above minimum wage labor in equal proportions, then each restaurant must have 17% of its labor costs going to minimum wage labor. Thus, a 10% increase in the minimum wage should raise prices by $10\% \times 33\% \times 17\% = 0.56\%$.

Aaronson and French (2007) use a calibrated model of labor demand that accounts for both firm and worker heterogeneity to show that when these factors are explicitly accounted for, the competitive model predicts prices will increase by roughly 0.7%. Moreover, because

[20] Of the 17 restaurant companies that appear in a search of 1995 reports using the U.S. Security and Exchange Commissions (SEC) Edgar database, the unconditional mean and median of the payroll to total expense ratio from 10-K reports is 30 percent. This search uses five keywords: restaurant, steak, seafood, hamburger, and chicken. Limited service establishments, like McDonalds and Burger King, are at or below the mean. Full service restaurant companies, like Bob Evans and California Pizza Kitchen, lie above. Similarly, the 1997 Economic Census for Accommodations and Foodservices reports payroll as 31 and 25 percent of sales at full and limited service restaurants, respectively. Since 10-Ks from food away from home companies generally show that wages account for 85 percent of compensation, the Economic Census' estimate of labor share based on compensation is roughly 36 and 29 percent at full and limited service restaurants. Another method of calculating lfabor's share is through a sampling of 1995 corporate income tax forms from the Internal Revenue Service's Statistics on Income Bulletin. Because operating costs are broken down by category, it is possible to estimate labor's share. We restrict the sample to partnerships because of IRS concern that labor costs are notoriously difficult to decompose for corporations. Despite the quite different sampling of firms relative to the Edgar Database, labor cost as a share of operating costs for eating place partnerships is of a similar magnitude to the other estimates, roughly 33 percent. Finally, these figures correspond well to a 2002 survey of restaurants by Deloitte and Touche (2003). Among limited service establishments, Deloitte and Touche find that wages and salaries make up 31 percent of total expenses. Benefits add another 2 percent. The payroll to expense ratio is roughly 2 percent higher for full service establishments.

[21] See Aaronson and French (2007) for a description of this calculation. Because wage distributions are not available in company reports, we estimate the share of employees that are paid at or near the minimum wage from the outgoing rotation files of the CPS for the two years prior to the 1996 legislation. We use a survey in Card and Krueger (1995, p. 162) to account for the share of workers paid slightly above the minimum wage that are impacted by new legislation.

limited service restaurants are more likely to pay the minimum wage than full service restaurants, competition would imply larger price increases at limited service restaurants. Given that some restaurants do not increase prices after minimum wage hikes, but restaurants that do raise their prices usually do so by more than 0.7%, it is difficult to compare the observed price response to the competitive prediction. Section 3.2 presents a statistical model to better make this comparison.

3.2 Estimates of Price Pass-Through

The next approach provides a more complete statistical model of the price response to a minimum wage change. In our basic model we regress the log change in prices for item k at outlet j in period t on the percentage change in the minimum wage in state i over the contemporaneous, lag, and lead periods, and a set of controls:

$$\Delta \ln p_{ijkt} = \sum_{h=1}^{H} \alpha_h \Delta \ln p_{ijkt-h} + \sum_{h=0}^{2} \beta_h \Delta \ln PPI_{t-h} + \sum_{h=-1}^{1} \omega_h \Delta \ln w_{min,it-h} + u_{ijkt} \qquad (1)$$

We include $w_{min,it-1}$ and $w_{min,it+1}$ to allow a more flexible response to the legislation, as price responses can play out over time. Consider the timing of the 1996-97 federal increase. When the law was passed, firms knew that minimum wages would be increased on October 1, 1996, and again on September 1, 1997. It is conceivable that firms could react to the expectation of an increase (i.e. at the bill discussion or passage), rather than the enactment dates. However, empirically, we found no evidence of longer leads or lags.[22]

The vector of controls include contemporaneous and lags of changes in the producer price index for processed foods to account for material input price shocks faced by sample outlets ($\ln PPI_t$).[23] To allow for mean-reverting price movements that typically occur after sales or

[22] Businesses knew of the 1996 increase just 2 to 4 months prior to implementation. They knew of the 1997 increase, specified in the 1996 bill, 13 to 15 months before implementation. The 1996 increase could not have been predicted until shortly before the House of Representatives vote on May 23, after a week of legislative maneuvering that almost consigned the bill to defeat (Weisman 1996 and Rubin 1996). Even then, the final timing of the minimum wage increase did not become clear until adoption of the conference report on August 2. Aaronson (2001) shows that longer-run price pass-through estimates are roughly the same size as short-run estimates using aggregated U.S. and Canadian price data.

[23] Aaronson (2001) accounts for the costs of particular food items, such as chicken, beef, bread, cheese, lettuce, tomatoes, and potatoes, and finds similar aggregated results to those reported here. We also controlled for broader labor market pressures using changes in CPS median wages and fixed chain and PSU (i.e. city) effects (not shown). Minimum wage point estimates and standard errors are quite robust to the inclusion of these variables. Furthermore, the fixed effects themselves added almost nothing to the model's fit. Controls for mealtype (breakfast, lunch, or dinner) are also included but have no impact on the results.

price hikes, we also experimented with controls for lags in $\ln p_{ijkt}$. However, the minimum wage estimates barely change whether lagged dependent variables are accounted for or not (the version reported in table 2 includes them).[24]

Table 2 presents the basic results. Since quotes from the same outlet are unlikely to be statistically independent, all standard errors account for quote (i.e. menu items within an establishment) clustering, using Huber-White robust estimation techniques. We also checked for error clustering by city, chain, and outlet. While within-outlet effects were important, within-city and within-chain effects were not. All results are reported as elasticities and multiplied by 10 to gauge the impact of a 10 percent minimum wage increase.

[24]To be clear, the inclusion of a lagged dependent variable potentially leads to inconsistent parameter estimates. In practice, this bias appears to be negligible. But for completeness, we also ran a specification that included one lag in $\ln p_{ijkt}$ that is instrumented by thrice lagged prices and found statistically indistinguishable results. Moreover, to capture any asymmetry in the mean-reverting process, we separately measure percentage increases and decreases in an item's price in the previous periods. Parameter estimates are of the expected sign: past price cuts lead to current period price increases, and past price increases lead to current price cuts, presumably because the original increase reflected temporary cost increases or because rivals didn't match the price increase. However, these responses are dampened substantially. Full reversion to prior prices implies absolute coefficient values of 1, while the estimated effects fall well below 1 and usually below 0.1.

Variable	All	Limited service	Full service	All	Limited service	Limited service	Full service	Full service
Column	1	2	3	4	5	6	7	8
$\Delta \ln w_{min,it-1}$	0.229	0.334	0.234	0.225	0.295	0.202	0.249	0.246
	(0.064)	(0.117)	(0.082)	(0.067)	(0.120)	(0.326)	(0.087)	(0.157)
$\Delta \ln w_{min,it}$	0.407	0.940	0.190	1.444	2.695	2.392	1.039	1.245
	(0.070)	(0.135)	(0.086)	(0.531)	(0.883)	(1.005)	(0.692)	(0.706)
$\Delta \ln w_{min,it+1}$	0.077	0.275	-0.102	0.078	0.243	0.451	-0.082	-0.228
	(0.063)	(0.136)	(0.073)	(0.067)	(0.149)	(0.431)	(0.079)	(0.164)
$\Delta \ln w_{min,it}$*wage20				-0.161	-0.278	-0.242	-0.128	-0.133
				(0.078)	(0.133)	(0.134)	(0.100)	(0.100)
Total effect	0.713	1.549	0.322					
	(0.140)	(0.275)	(0.168)					
At wage20= \$5.00				0.94	1.84	1.84	0.56	0.60
At wage20= \$7.00				0.62	1.29	1.35	0.31	0.33
At wage20= \$9.00				0.30	0.73	0.87	0.05	0.07
month dummies?	no	no	no	no	no	yes	no	yes
include PPI?	yes	yes	yes	yes	yes	no	yes	no
R^2	0.070	0.167	0.017	0.070	0.171	0.175	0.019	0.021
N	71,077	21,883	36,928	63,630	18,691	18,691	33,875	33,875

See text for detail. Controls not shown in table include three lags in $\ln p_{ijkt}$ and mealtype (breakfast, lunch, or dinner).
Huber-White standard errors corrected for clustering at the item and establishment level are in parentheses.
Sample sizes in columns 2 and 3 do not add up to column 1 because some establishments are not categorized as full or limited service restaurants.
Wage20 is the 20th percentile of the MSA's hourly wage distribution, calculated from the 1996 CPS.

Table 2: The Price Response to a 10 Percent Minimum Wage Increase

Column (1) reports the minimum wage effect for the full sample of food away from home establishments. We find that a 10 percent increase in the minimum wage increases prices by roughly 0.7 percent (with a standard error of 0.14), of which over half the response occurs within the first two months after the minimum wage change. However, because this result combines outlets where the minimum wage is binding with those where it might be less important, two particular sources of variation can be used to identify the price response to a minimum wage increase.

First, as in table 1, we can take advantage of variation in the intensity of minimum wage worker use between limited (column 2) and full (column 3) service restaurants. The price increase generated by a 10 percent minimum wage hike is roughly 1.55 percent (standard error of 0.28 percent) for limited service outlets but a fifth that size for full service enterprises.[25]

Second, market wages vary across local labor markets. Where prevailing low-skill wages far exceed minimum wages, minimum wage increases will have little impact on market wages and consequently costs. Where the minimum binds for low-skill workers, changes in the minimum wage will have strong effects on wages.[26] Therefore, we test whether the price response varies with respect to the pay of low-skilled workers. We are able to perform this test because our data include precise outlet locations (addresses and telephone numbers) that we link to MSA hourly wage distributions estimated from the 1996 Current Population Survey (CPS). Columns (4), (5), and (7) interact one version of these measures, the 20th percentile from the MSA's hourly wage (wage20) distribution, with the contemporaneous minimum wage change using the full sample and subsample of limited and full service outlets.[27] We find that minimum wage increases have larger effects on prices in low wage areas, among both limited and full service outlets. An MSA where the 20th percentile of the 1996 hourly wage is $5 leads to a 0.56 percent price increase among full service outlets and a 1.84 percent increase

[25] Aaronson (2001) finds an elasticity of around 1.5 for fast food restaurants in the American Chamber of Commerce price survey, consistent with the findings on limited service establishments.

[26] We can look at this directly using the outgoing rotation files of the CPS. Using a state panel developed from the 1979-2002 files, we regressed log hourly earnings in the restaurant industry on the prevailing minimum wage and state and year fixed effects. The data were too noisy and replete with missing observations at the monthly level. Nevertheless, we find that wages rise by 4.4 percent in the restaurant industry following a 10 percent increase in the minimum wage. Although we cannot distinguish full and limited service establishments, we note that wages rise by 10.7 percent among teens and 7.1 percent among high school dropouts. The results are similar when we restrict the sample to the CPI cities.

[27] The results are also comparable when changing the year used to calculate *wage*20, interacting wage20 with the lag and lead minimum wage change. Wage data for the 12 non-metro PSU's are drawn from the non-metro parts of the outlet's state. CPS codes are unavailable for 9 MSAs, so sample sizes decline when area wage data are included in the analysis.

among limited service outlets. At a wage of \$7 (\$9), this effect drops to 0.31 (0.05) and 1.29 (0.73) percent for full and limited service firms. We have also used the share of minimum wage workers, $prob(w_{it} = w_{min,it})$, as a measure of variation in minimum wage bindiness across local labor market. The LS results are similar to those reported here, although less precisely estimated.[28] The coefficient on $\Delta \ln w_{it} \times prob(w_{it} = w_{min,it})$ is 0.49 (0.39) for LS establishments and 0.17 (0.30) for FS establishments. At the mean value of the share of workers paid at or near the minimum wage (6 percent), a 10 percent increase in the minimum wage increases LS prices by 1.15 percent and FS prices by 0.38 percent, very similar to columns (5) and (7) in table 2.

As a robustness check, columns (6) and (8) report results of a regression with the wage20 interaction that also includes a full set of month dummies.[29] The month dummies eliminate the possibility that the minimum wage changes are confounding other contemporaneous national inflation or economic trends or seasonal factors (since the two federal changes occur in September and October). But as can be seen, this does not appear to be an economically or statistically important concern, either for LS or FS establishments. This is also true when we use the share of minimum wage workers rather than the 20th percentile of the wage distribution.

As a final alternative, we estimated logit models that explore the relationship between minimum wage changes and the probability of a price increase or decrease by outlet type. These regressions use a very similar specification to equation (1), but substitute $\Delta \ln p_{ijkt}$ with an indicator of whether there is a price increase ($p_{ijkt} > 0$) in one specification and a price decrease ($p_{ijkt} < 0$) in another.[30] The first four columns of table 3 report a specification that includes the lag, contemporaneous, and lead minimum wage change measure, along with the controls described in the table. We find that the likelihood of a price increase in LS outlets jumps from 12 to 28 percent if a 10 percent increase in the minimum wage is introduced in a

[28] There is some debate on the extent to which minimum wage increases impact workers paid above the minimum. See Card and Krueger (1995), Abowd et al (2000), and Lee (1999). We have experimented with using those at or below the new minimum wage, as well as allowing for spillovers up to 20 percent above the new minimum. This has little impact on the results.

[29] For identification, we must drop the PPI controls in this specification.

[30] The main deviation from equation (1) is that we include a dummy for prices ending in 99 cents, to account for the extra stickiness apparent at such price points. See MacDonald and Aaronson (2006) or Kashyap (1995) for further discussion. In the version reported here, we also include three lags in changes in item prices, an indicator of whether any sampled price was changed by the outlet in the previous period, and indicators for meal type. Inferences are not contigent on the inclusion of these additional covariates.

period with otherwise stable prices. The probability of a price increase in FS establishments increases from 10 to 13 percent following a similar sized minimum wage change. However, a 10 percent increase in the minimum wage has no statistically significant impact on the probability of a price decline in either type of establishment. The final two columns show that MSAs with a lower 20th percentile wage are much more likely to see price increases in both LS and FS establishments after a minimum wage increase. There is no such effect among price declines (not shown).

Establishment type	Limited service	Full service	Limited service	Full service	Limited Service	Full Service
Price change	Increase	Increase	Decrease	Decrease	Increase	Increase
Column	1	2	3	4	5	6
$\Delta \ln w_{min,it-1}$	0.010	0.019	-0.007	0.004		
	(0.014)	(0.011)	(0.022)	(0.018)		
$\Delta \ln w_{min,it}$	0.104	0.027	-0.020	0.008	0.253	0.233
	(0.012)	(0.011)	(0.021)	(0.022)	(0.095)	(0.092)
$\Delta \ln w_{min,it+1}$	0.017	0.010	-0.029	0.014		
	(0.015)	(0.012)	(0.029)	(0.021)		
wage20 $* \Delta \ln w_{min,it}$					-0.024	-0.032
					(0.015)	(0.014)
Constant	-1.974	-2.150	-3.732	-4.403	-1.842	-2.485
	(0.074)	(0.063)	(0.137)	(0.136)	(0.444)	(0.381)
Base Probability of price change	0.122	0.104	0.023	0.012	0.137	0.077
After 10% minimum wage increase	0.282	0.132	0.019	0.013		
at wage20=$5					0.349	0.189
at wage20=$7					0.241	0.121
at wage20=$9					0.158	0.076

Coefficients and standard errors are derived from a logit model.
See text for detail. Controls not shown in table include whether the price ended in 99 cents,
three lags in $\ln p_{ijkt}$, an indicator for whether any sampled price item was changed in the previous period,
and mealtype (breakfast, lunch, or dinner).

Table 3: The Probability of a Price Increase or Decrease in Response to a 10 Percent Minimum Wage Increase

3.3 City-level Price Responses

The micro data suggest that prices move higher in response to minimum wage changes that occurred between 1995 and 1997. In this section, we show that the results are robust to looking at a longer earlier period. Here, we use the publicly available city-level price indices of the CPI between 1979 and 1995 to test whether cities with higher fractions of restaurant workers impacted by the minimum wage laws are more likely to change their food prices. Hence, identification is based on the intensity of minimum wage worker usage. Our results are based on a slightly modified form of equation (1):

$$\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}} = \gamma prob(w_{it} = w_{min,it}) + \beta' x_{it} + u_{it} \tag{2}$$

where i denotes city and the coefficient $\gamma = E[\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}} | w_{it} = w_{min,it}, x_{it}]$ is the price response to increases in the minimum wage. If producers have a constant returns to scale production function, competitive theory implies that 100 percent of the higher labor costs are passed on to the consumer in the form of higher prices. As we pointed out earlier, 100% pass through implies that the percent increase in product price equals the percent increase in the minimum wage multiplied by labor's share. Therefore, γ should equal labor's share under perfect competition.

Figure 1 maps each city's price response to a minimum wage hike against the share of minimum wage workers in the city. Each observation, of which there are 82, represents a city around the time of a minimum wage change. The data cover 4 federal minimum wage hikes - in 1980, 1981, 1990, and 1991 – and a small number of state increases between 1979 and 1995.[31] The horizontal axis plots $prob(w_{it} = w_{min,it})$, the share of workers in a city's restaurant industry that are paid the minimum wage. This is computed from the outgoing rotation files of the Current Population Survey (CPS). Because employees paid just above the minimum wage are also affected by the law, we include anyone paid within 120 percent of

[31] The federal minimum wage increased from \$2.90 to \$3.10 per hour in January 1980, to \$3.35 in January 1981, to \$3.80 in April 1990, and to \$4.25 in April 1991. State increases tend not to occur in states represented by CPI survey cities. The 27 CPI cities are New York City, Philadelphia, Boston, Pittsburgh, Buffalo, Chicago, Detroit, St Louis, Cleveland, Minneapolis-St. Paul, Milwaukee, Cincinnati, Kansas City, DC, Dallas, Baltimore, Houston, Atlanta, Miami, Los Angeles, San Francisco, Seattle, San Diego, Portland, Honolulu, Anchorage, and Denver. After 1986, prices for 12 of these cities – Buffalo, Minneapolis-St. Paul, Milwaukee, Cincinnati, Kansas City, Atlanta, San Diego, and Seattle, Portland, Honolulu, Anchorage, and Denver – are reported semiannually. Therefore, we only include pre-1986 observations for these cities.

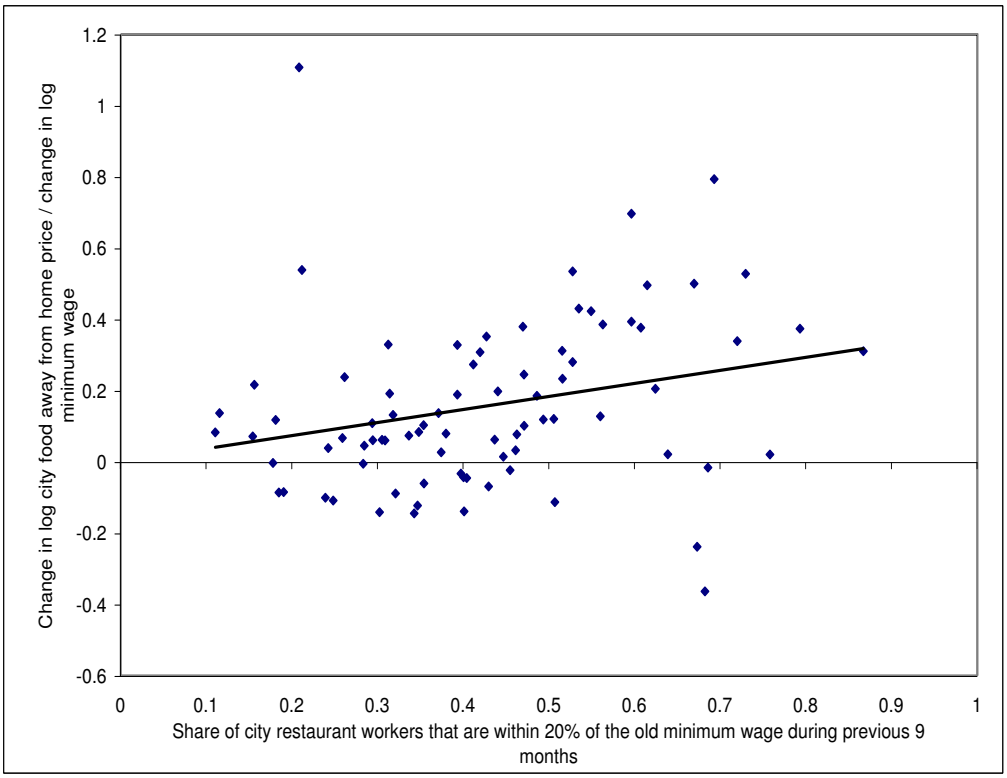

Figure 1: City level price increases

the old minimum wage during the nine months prior to the minimum wage enactment date. However, the results are not sensitive to picking reasonable thresholds other than 120 percent or time frames other than nine months. The vertical axis displays $\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}}$, the ratio of the log change in city food away from home prices to the log change in the city's minimum wage. The price data is the CPI for food away from home. The price changes are measured from two months before to two months after the minimum wage is enacted. $\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}}$ is adjusted for year fixed effects to account for inflation and other secular changes in national labor market conditions.

The most noteworthy aspect of figure 1 is the positive correlation between the two series. The regression coefficient γ is 0.36 with a robust, city clustered-corrected standard error of 0.24.[32] Not only is the sign of this coefficient consistent with competition but the magnitude is as well. Assuming perfect competition in the labor market, the regression coefficient should equal labor's share. Recall from section 3, labor's share is approximately 30 to 35 percent.

Note also the abundance of observations on $\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}}$ that are positive. Of the 82 city-year observations, 19 are negative, including only 2 of the largest 30 price responses, defined as when $|\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}}| > 0.20$. These two are interesting in that they come from the same city, Denver, over consecutive years, 1980 and 1981. Unfortunately, we have little information as to what was happening in Denver during this time but we can highlight it for being the main example where city-level prices fall quickly in response to a minimum wage change.[33]

The most plausible alternative explanation for these price responses is that they are driven by shocks to demand that happen to be correlated with changes to the minimum wage. We tried two ways to test this possibility. First, we estimated equation 2 without year fixed effects but included changes in the city CPI in the x_{it} vector. The intercept from this specification is not statistically different from zero, suggesting that prices do not rise after minimum wage hikes in areas where the minimum wage does not bind. This finding is consistent with the

[32] Using the share of minimum wage workers within 110 percent of the old minimum, rather than 120 percent, the point estimate (and adjusted standard error) is 0.42 (0.24). A Huber biweight regression procedure implies a point estimate of 0.28 (0.16) and 0.42 (0.15) using the 110 and 120 percent minimum wage share thresholds. Finally, out of concern that inflation, even over this short period, are driving our results, we tried two things. First, when we include city price deflators as controls on the right hand side, the point estimates are roughly 0.50 with t-statistics of roughly 2 to 2.5. Alternatively, we look at price changes only over the two months after the minimum wage change. In this case, the point estimate is between 0.20 and 0.30, again with t-statistics of roughly 2 to 2.5.

[33] Since Denver is one of the 12 cities surveyed semiannually starting in 1986, we do not include the 1990 and 1991 Denver data points in the figure. However, they are both positive, albeit small: 0.19 for 1990 and 0.09 for 1991.

view that demand shocks are not confounding our estimates because if they were, we would expect that prices would rise in areas where the minimum wage does not bind.

As a second check, we searched for alternative measures of p_{it} that vary by local demand conditions, are available for our city panel, and, most importantly, are relatively unaffected by low wage labor costs. By far, the two best candidates are housing and medical care. Therefore, we reran the regression described above, but substituted food away from home prices with these two indices. As expected, we find no evidence that minimum wage hikes are associated with price hikes for housing and medical care.[34]

Finally, we can compare the estimates in table 2 to predicted pass-through under competition and constant returns to scale technology using metro variation in restaurant wage distributions from the outgoing rotation files of the CPS. Under competition, the relationship between these predicted price responses and the share of restaurant workers impacted by minimum wage laws should correspond to labor share.

To conduct this test, we define two groups of cities. High wage cities are those with an average hourly wage among the top fifth of all metropolitan areas in 1997-98. Low wage cities are those with average wages among the bottom fifth of all metropolitan areas in 1997-98. Among high (low) wage cities, 34 (59) percent of all restaurant workers are paid within 120 percent of the minimum wage, our rough measure of the share of workers impacted by such laws. Based on the estimates in column (4) of table 2, the average predicted price response for low and high wage cities is 0.097 and 0.090, respectively.[35] The relationship between these variables (slope of the line connecting high and low wage cities) is $\frac{.97-.90}{.59-.34} = 0.28$, slightly lower than, and statistically indistinguishable from, observed labor share in the restaurant industry. Furthermore, we get a similar labor share prediction when we estimate equation (2) using the store-level data.[36]

[34] For housing, γ (and its adjusted standard error) is -0.31 (0.33). For medical care, it is -0.06 (0.33).

[35] The 20th percentile restaurant wage in the high wage cities is \$5.25, compared to \$4.82 in the low wage cities.

[36] To derive γ from the micro data, we use the regression results from column 4 of table 2, which gives the relationship between $\frac{\Delta \ln p_{it}}{\Delta \ln w_{min,it}}$ and the 20th percentile of city market wage. Next, we regress the share of workers in a city's restaurant industry that are paid the minimum wage, $prob(w_{it} = w_{min,it})$, on the 20th percentile of that city's market wage. Using the 27 major CPI cities during 1995 to 1997, the point estimate from this latter regression is $prob(w_{it} = w_{min,it}) = 0.839 - 0.068 \times \text{wage20}_{it} + \nu_{it}$, where wage20_{it} is the 20th percentile of city i's wage distribution at time t. From these two regression equations, we can solve $\gamma = 0.24$. Ideally we could precisely estimate equation (2) using the micro data. Unfortunately, as we note in footnote 25, the number of observations in the CPS for individual cities can be small. Consequently, $prob(w_{it} = w_{min,it})$ cannot be precisely estimated. But $wage20_{it}$ can.

4 Theory

In this section, we show how our results contribute to the debate on the employment effects of the minimum wage. Assorted models offer differing explanations for why the estimated employment responses to minimum wage hikes are small. As we point out below, however, most of these models imply that prices either do not change or fall in response to a minimum wage hike. Therefore, our results provide evidence against models that have been used to explain the small employment responses found in the minimum wage literature.

Throughout the discussion we assume that all firms are profit maximizers and thus set the level of employment, L, at the point where the marginal cost of the last worker hired $MC(L)$ is equal to the extra revenue she produces (her marginal revenue product of labor, or $MRP(L)$). Appendices A and B contain the formal details of the model.

4.1 The Competitive Model

We begin by briefly considering the textbook competitive model. If a minimum wage is introduced (or increased) beyond the market-clearing wage in a competitive labor market, the marginal cost of hiring a worker increases. In response, holding all else equal, firms will move along their downward sloping marginal revenue product of labor curve until they reach the point where $MRP(L)$ is again equated to marginal cost. Higher $MRP(L)$ can only be obtained by reducing the workforce. Why? One important reason is that fewer workers imply less output. Even if an additional worker produced the same amount as the previous worker, reduced output increases output prices, marginal revenue, and thus the marginal revenue product of labor. Therefore, minimum wage hikes cause prices to rise and employment to fall in a competitive labor market environment.

4.2 The Textbook Monopsony Model

However, under monopsony, increasing the minimum wage can cause employment to rise. The fundamental reason is tied to the link between the wage, the marginal cost of labor, and the product price. In this section we describe the textbook model of monopsony, where firms are monopolists in the product market and monopsonists in the labor market. In Section 4.3 we show that the key results hold under the more realistic scenario of monopolistic competition in the product market and monopsonistic competition in the labor market.

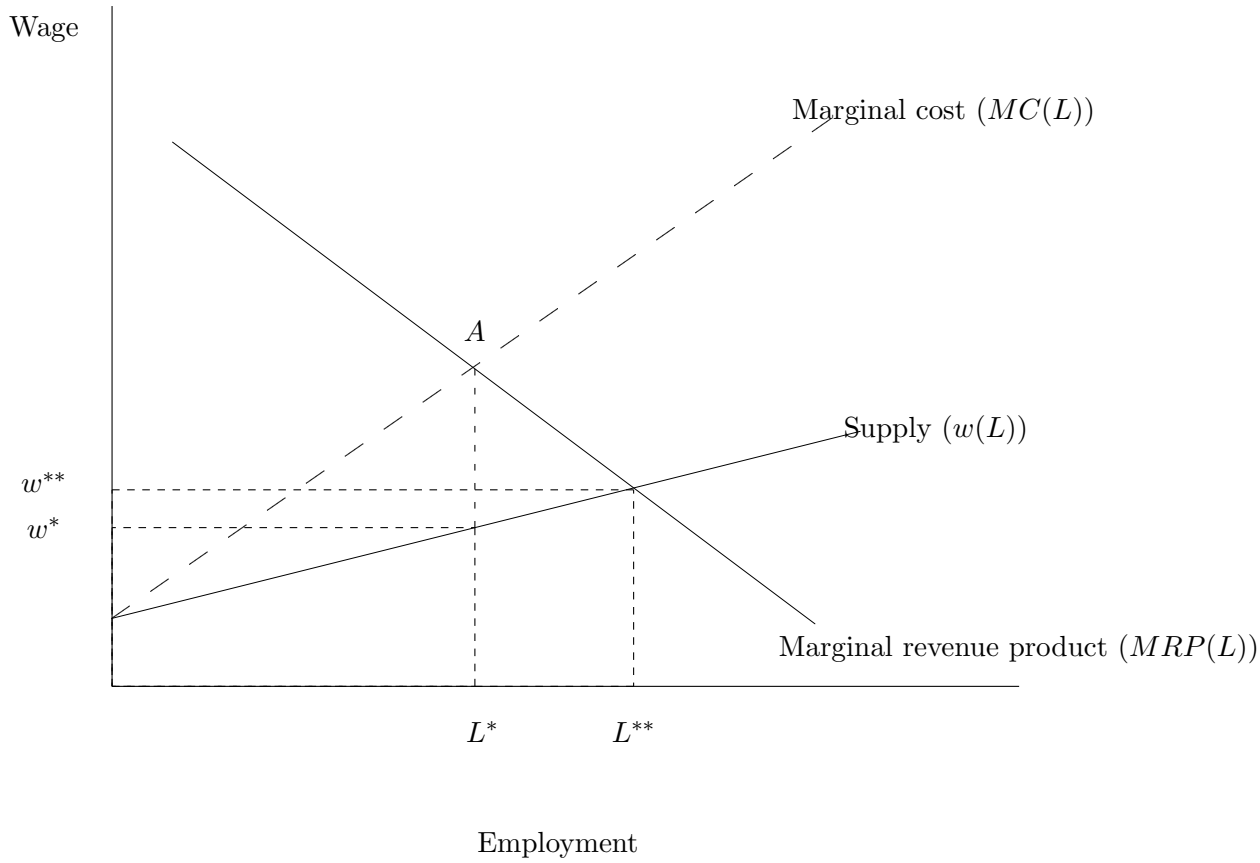

Figure 2: Illustration of monopsony equilibrium

Unlike the competitive firm, which pays the prevailing market wage regardless of how much labor it demands, if the monopsonist wants to expand its labor force, it has to raise the wage of its current workers as well.[37] Therefore, the marginal cost of hiring a worker is greater than the new worker's wage. Figure 2 shows the wage the firm would have to pay in order to attract an additional worker, $w(L)$, and the marginal cost of hiring that worker $MC(L)$. When the monopsonist maximizes profits by setting marginal costs equal to marginal product, shown at point A, total market employment, L^*, is lower than the "competitive case" (L^{**}) where employment is set based on the prevailing market wage w^{**}.

A properly placed minimum wage, set somewhere between the wage

[37] This is true if the monopsonist cannot perfectly discriminate between workers with high reservation wages and low reservation wages.

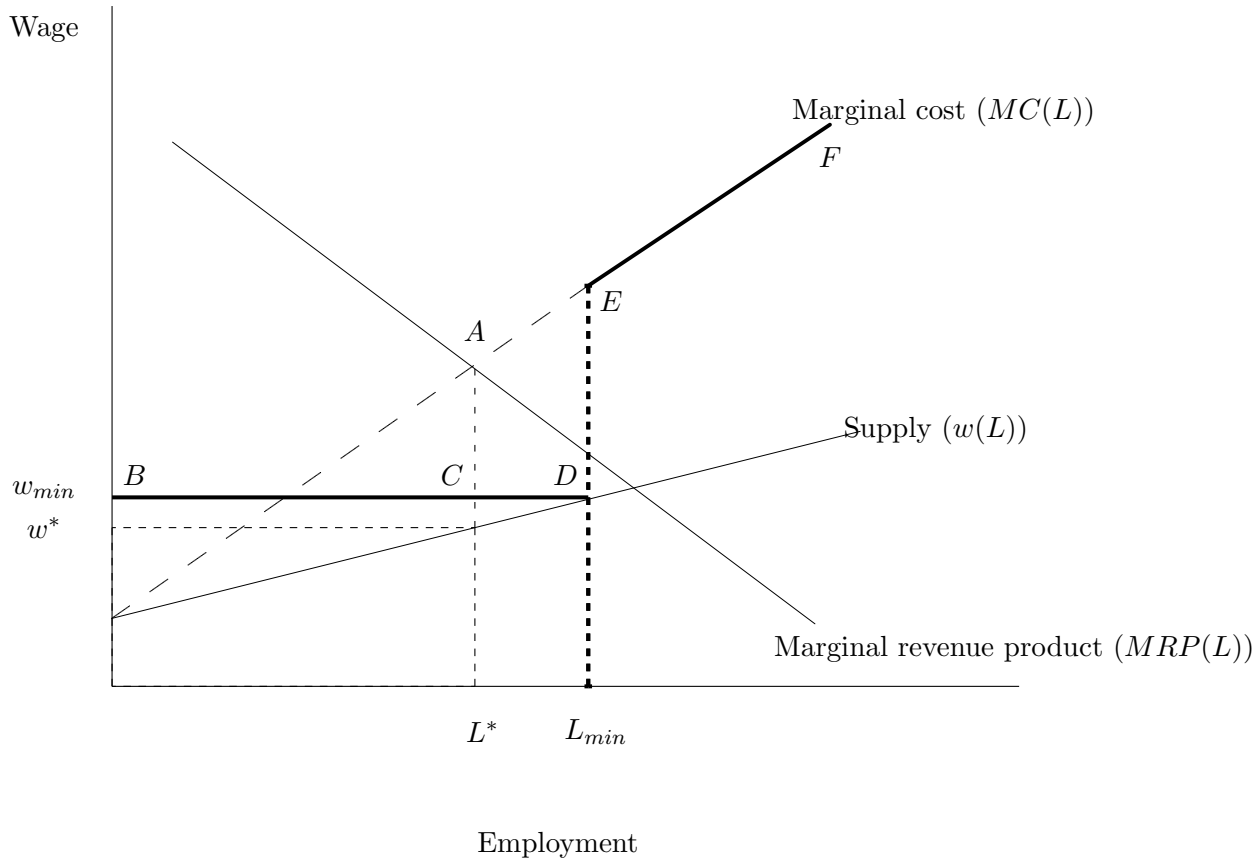

Figure 3: Illustration of monopsony equilibrium with minimum wage (Bold line denotes $MC(L)$ curve)

paid by a monopsonist (w^*) and the wage paid by a perfect competitor (w^{**}), will increase employment. The intuition for this result, displayed in figure 3, is that although the minimum wage *increases* the firm's average cost of labor, it *reduces* the marginal cost of labor. Recall that, absent a minimum wage, the marginal cost of hiring that last worker (at point A) lies above the wage paid by the monopsonist (because everyone's wage has to be raised in order to induce a marginal worker to join the firm). If the minimum wage is set above the monopsony equilibrium wage but below the marginal cost of hiring a worker, the new marginal cost of hiring a worker falls from point A to point C (the new marginal cost of labor curve is $BCDEF$); the marginal cost of hiring additional workers is now just the minimum wage. Because the firm must pay all workers at least the minimum wage, regardless of employment

level, the firm does not have to increase the pay of its existing workforce to attract more employees (so long as employment is below L_{min}). This reduction in the marginal cost of hiring additional labor causes firms to expand output and employment in response to the minimum wage increase.

Moreover, employment and prices are negatively related because the fall in the marginal cost of labor causes the marginal cost of producing an extra unit to fall. Consequently, the product price falls as well.

It is important to note that if firms are monopsonists in the labor market but the minimum wage is set sufficiently high (above w^{**} in Figure 2), employment is determined by the intersection of the minimum wage and the marginal product of revenue curve. In this case, an increase in the minimum wage increases prices and reduces employment, just like in a competitive labor market. Thus our empirical results cannot necessarily disprove the existence of monopsony labor markets in cases where the minimum wage is set high (above competitive market-clearing levels). However, we believe our empirical results should temper enthusiasm for monopsony power being the explanation for the negligible employment responses found in the literature.

4.3 Price and Employment Responses when there is Monopsonistic Competition in the Labor Market and Monopolistic Competition in the Product Market

The results in the previous section were based on a very stylized model. In this section, we show that those results are quite general under weak assumptions about technology, the product market, and the labor market. Specifically, the results hold when firms have a production function with substitutability between labor and other inputs, monopolistic competition in the product market, and monopsonistic competition in the labor market.

Several researchers have argued that monopsonistic competition in the labor market is the relevant case (Bhashkar and To (1999), Dickens et al. (1999)). That is, workers are not indifferent between employers, even if all employers pay the same wage. One plausible explanation is geography. Employers are located in different places and transportation costs are large relative to earnings of minimum wage workers. Thus a worker is willing to take a low paying job in order to be relatively near home. Alternatively, teenagers may want to

work at the same restaurant as their friends. More generally, certain aspects of one employer may be disagreeable to some workers but not others.

In order to simplify the problem, we make six assumptions beyond the usual axioms of firm behavior:

Assumption 1 *There is a fixed number N of firms.*

Assumption 2 *All firms have an identical production function, $Q_n = Q(K_n, L_n)$ where Q_n, K_n, L_n are output, an aggregator of non-labor input (that includes capital), and labor at the nth firm.*

Assumption 3 *The production function is increasing in all inputs, concave, continuous and twice differentiable.*

Assumption 4 *K and L are complementary inputs in the production function ($Q_{12} > 0$).*

Assumption 5 *The utility function of the representative consumer is $U = \left((1-\alpha)Q_0^{1-\eta} + \alpha\tilde{Q}^{1-\eta}\right)^{\frac{1}{1-\eta}}$, where Q_0 is the numeraire good, α is close to zero, $\tilde{Q} \equiv \left(\sum_{n=1}^{N} Q_n^{1-\eta_Z}\right)^{\frac{1}{1-\eta_Z}}$, and Q_n denotes output at the nth restaurant. Concavity implies $\eta > 0$ and $\eta_Z \in [0, 1)$.*

Assumption 6 *The firm is a price taker in the capital market and purchases K_n at price r. However, the firm is potentially a monopsonist in the labor market. The quantity of labor supplied to the firm is $L_n^S = L(w_n, w_{-n})$, where w_{-n} is the average wage paid by all other firms, $\left.\frac{dL(w_n,w_{-n})}{dw_n}\right|_{w_{-n}=w_{min}} > 0$, and $\left.\frac{dL(w_n,w_{-n})}{dw_n}\right|_{w_{-n}=w_n} > 0$.*

Under these assumptions, firms sell their products at a price $p(Q)$ and choose inputs to maximize profits π :

$$\pi_n(K_n, L_n) = p(Q_n)Q_n - w_n L_n - rK_n. \tag{3}$$

These assumptions are standard, although a few require some elaboration. Assumption 5 gives rise to monopolistic competition in the product market. Markets are perfectly competitive if $\eta_Z = 0$ and firms operate as monopolists if $\eta_Z = \eta$. Assumption 6 states that the quantity of labor supplied to the firm need not be perfectly elastic and, therefore, firms face a

monopsonistically competitive labor market. Consequently, firms face a weakly upward sloping inverse labor supply curve, $w(L_n)$, where $\frac{dw(L_n,w_{-n})}{dL_n} \geq 0$. However, because the minimum wage potentially binds, the offered wage is $w_n = \max\{w(L_n, w_{-n}), w_{min}\}$.[38]

Firms may be price takers in the labor market because the labor supply curve that the firm faces is perfectly elastic or the minimum wage is sufficiently high that it destroys the firm's monopsony power. Either way, if firms are price takers in the labor market, Theorem 1 holds.

Theorem 1 *Given the assumptions above, and if firms are price takers in the labor market, the industry level demand curve for labor slopes down.*

Proof: see Appendix A. Theorem 1 is more general than the discussion in Section 4.1. There, it is presumed that firms make employment decisions given a fixed $MRP(L)$ curve, an assumption that is appropriate for monopolists. But, under monopolistic competition, minimum wage changes potentially shift the $MRP(L)$ curve by altering the decision of other firms, and thus influencing aggregate prices. Theorem 1 also differs from the Weak Axiom of Profit Maximization, which assumes perfect competition in both the product and factor markets.[39]

Section 4.2 discusses why employment can rise under monopsony. Given assumption 6, this is true under monopolistic competition as well. Together with Theorem 1, this shows that minimum wage hikes cause employment to fall under competition and rise under monopsony.

The next theorem shows that we can use prices to infer the importance of monopsony power in the labor market.

Theorem 2 *Given an increase in a binding minimum wage, prices rise under perfect competition and, so long as $w_{min} < w^{**}$, prices fall under monopsony.*

[38]The assumption of capital and labor being complementary inputs (i.e. $Q_{12} > 0$) rules out situations where the profit maximizing choice would be to switch from a capital intensive, high output technology to a labor intensive, low output technology. An example of this is a firm that is capital efficient only up to a certain size. After this size, capital cannot be efficiently used. For example, suppose $p = 1$, $r = 1$, $Q = K^{.5}L^{.5}$ if $L < 10$ and $Q = L^{.5}$ if $L \geq 10$. Increasing L from 9 to 10 would reduce output but depending on the cost of labor, could increase profits. However, this rather extreme counter-example appears to go against the empirical evidence. For example, it is difficult to reject the hypothesis that production functions are constant returns to scale, and constant returns production functions implicitly assume $Q_{12} > 0$.

[39]See Varian (1984) and Kennan (1998) for proofs of Theorem 1 under competition and monopoly, respectively. We have also proved the theorems in this section for the case where firms are Cournot competitors in the output market. Proofs are available from the authors.

Proof: see Appendix A. The intuition for Theorem 2 was discussed in Section 4.2.

Finally, there is the quantitative importance of price pass through when there is monopolistic competition in the product market. Theorem 3 shows that if the production function is constant elasticity of substitution, then firms still push 100% of the higher labor costs onto consumers in the form of higher prices.

Theorem 3 *If $Q(.,.)$ is a constant elasticity of substitution aggregator, and if firms are price takers in the labor market, then $\frac{d \ln p}{d \ln w_{min}}$ = minimum wage labor's share.*

Proof: see Aaronson and French (2007). The intuition for this result is straightforward. Given the assumptions above, firms have a constant marginal cost and thus have a horizontal supply curve. Thus, in a perfectly competitive market, all higher labor costs will be pushed onto consumers in the form of higher prices. In the case of monopolistic competition in the product market, there is a constant mark-up over marginal cost. Thus the supply curve is still horizontal and all labor costs are pushed onto consumers in the form of higher prices.

Furthermore, Aaronson and French give predicted price and employment responses under monopsonistic competition. They show that if the employment response is large and positive, then the price response will be large and negative. For example, if the employment elasticity is +0.2, which is possible under monopsony, then the price response will be -0.05. These price responses vary notably from what is reported in Table 2.

The only assumption that we view as not innocuous is the first. Although the size of a business is allowed to change in response to a higher minimum wage, firm exit or entry is precluded. We think this is a reasonable assumption given the existing, albeit rather meager, empirical evidence.[40] Moreover, in this paper, we are interested in a short-term response that likely severely limits entry and exit decisions.

The main reason for assuming no exit is that under monopsony, minimum wage hikes increase employment per restaurant, but likely reduce the total number of restaurants. There-

[40] Card and Krueger (1995) and Machin and Wilson (2004) find no effect in the U.S. and U.K., respectively. We have done some analysis of restaurant entry and exit using the Census' Longitudinal Business Database. Consistent with the literature, our preliminary findings suggest negligible entry and exit effects in the year following a minimum wage change. These results stand in contrast to those of Campbell and Lapham (2004), who find a significant amount of retail net entry along the U.S.-Canada border within a year of exchange rate movements. We suspect these different results reflect the importance of exchange rates relative to minimum wage levels in terms of firm costs.

fore, the industry level employment response is ambiguous. In this sense, we view the assumption of no exit as supporting the monopsony argument.[41]

4.4 Efficiency Wage Models

Efficiency wage models (where increased wages increase effort or reduce turnover costs), often give monopsony like predictions. Manning (1995), Rebitzer and Taylor (1995), and Deltas (1999) all present models where increases in the minimum wage can increase employment. None of these models allows for capital-labor substitutability, and only Deltas (1999) allows for endogenous prices. Below we present an efficiency wage model with endogenous prices and capital-labor substitutability.

We follow Solow (1979), who argues that the wage affects morale and effort amongst other things, and let the wage enter the production function directly. If an increase in the wage causes increased effort, more meals can be produced with the same amount of labor and capital. Therefore, it is not necessarily more costly to produce meals when the minimum wage increases. This can attenuate the disemployment effects of the minimum wage. However, if employment does not fall (and other factors do not fall either) and productivity rises, total output will rise and product prices will fall.

Let the production function be:

$$Q = Q(K, L, w) = \left((1-\alpha)K^{\rho} + \alpha(Lw^{\theta})^{\rho}\right)^{\frac{1}{\rho}}, \tag{4}$$

where $\sigma \equiv \frac{1}{1-\rho}$ is the partial elasticity of substitution between K and Lw^{θ} in the production of Q, and Lw^{θ} is "effective labor". The parameter $0 \leq \theta < 1$ may be greater than 0 because increases in the wage increase effort, which could happen for a variety of reasons. Furthermore, assume that Assumptions 1 to 5 in Section 4.3 hold. Then the price and

[41] Nevertheless, we also understand that, given our assumed market structure, entry and exit can change the market price for a given industry output. This is potentially important because Bashkar and To (1999) argue that an increase in the minimum wage could reduce the number of firms in the market but increase employment per restaurant, causing total employment to potentially increase. Because the number of firms decrease, market power of survivors increase. Consequently, both prices and output may increase. However, given the small observed exit rates in response to minimum wage hikes, we doubt that these effects would be large enough to overturn the basic presumption that market output and price move in opposite directions.

employment responses are:

$$\frac{d\ln p}{d\ln w} = s(1-\theta) \tag{5}$$

$$\frac{d\ln L}{d\ln w} = -(1-\theta)\big((1-s)\sigma - s\eta\big) - \theta \tag{6}$$

where s is the share of total costs going to labor. When $\theta = 0$, equations (5) and (6) give the textbook response to the minimum wage. However, when $\theta > 0$ (and all else is equal), a one percent increase in the wage increases effective labor θ percent, causing the marginal cost of effective labor to only rise $1-\theta$ percent. This has implications for both prices and employment responses. The price response is attenuated, rising by $s(1-\theta)$ percent, a fraction $(1-\theta)$ less than without endogenous work effort. The employment response is also muted by $(1-\theta)$ due to an increase in the marginal cost of labor. However, the same amount of effective labor and output can be produced with fewer bodies, lessening the need for labor (the $-\theta$ term at the end of equation (6)).

Regardless, while a smaller employment response relative to the case without endogenous work effort is possible, the model clearly predicts a smaller price response as well. However, our estimates indicate large price responses to the minimum wage. Thus, our price results provide evidence against the hypothesis that endogenous work effort is responsible for the small observed employment responses to minimum wage hikes.

4.5 Other Models of the Employment Effects of the Minimum Wage

We have argued that the price responses to minimum wage hikes are useful for distinguishing between competition and monopsony in the labor market. Likewise, our estimated price responses help shed light on other explanations of the small employment response found in the minimum wage literature.

Some researchers (e.g. Kennan 1995, MaCurdy and OBrien-Strain 2000) suggest higher income resulting from a minimum wage increase causes low wage workers to buy more minimum wage products, attenuating the disemployment effect of the minimum wage. For the restaurant industry, Kennan refers to this possibility as the "hungry teenager hypothesis." The price response reported in this paper is a key parameter for such a calculation. Aaronson

and French (2006) write down a model that allows for such demand-induced feedbacks and show that increases in the minimum wage reduce real income for non-minimum wage workers (because prices rise) but increase real income of minimum wage workers (because their wage rises). They find that if minimum wage workers spend a large fraction of their income on fast food, then the rise in incomes for fast food workers can at least partly offset the disemployment effect of the minimum wage. Using data from the Consumer Expenditure Survey and US Department of Agriculture's Continuing Survey of Food Intake by Individuals, they find that minimum wage workers spend between 20 and 100% more of their income on fast food as those who are not minimum wage workers.[42] Given these estimates (and other calibrated parameters) the increased income going to minimum wage workers can offset between 25 and 40% of the output loss, and 10 to 30% of the employment loss of a model that does not account for the increased income and increased expenditures of minimum wage workers.

Our price responses are less consistent with the idea that that minimum wage laws merely cause firms to reshuffle compensation packages from non-wage benefits to wages. For example, fast food restaurants could stop giving workers free meals after minimum wage hikes. Hashimoto (1982) and Neumark and Wascher (2001) argue that firms may reduce training after minimum wage hikes. As a result, there is no increase in the cost of labor faced by firms. However, if the minimum wage does not increase the cost of labor, it is unclear why there are price increases after minimum wage changes. Although shifting compensation packages from non-wage to wage benefits may occur, our results indicate that firms still bear a sizeable fraction of the cost of minimum wage hikes. In this sense, our findings are consistent with Card and Krueger (1995), who also find very little substitution between wage and non-wage benefits after minimum wage hikes and also find no evidence of minimum wage hikes on training.

Finally, our analysis has not considered search models (e.g. Burdett and Mortensen 1998), which also give monopsony implications for employment. A full analysis of the variety of search models is well beyond the scope of this paper, especially since most of them would need to be modified to account for endogenous product prices and capital labor substituability. However, it seems likely that employment and prices would move in opposite directions in most standard applications of search. For example, given that the Burdett and Mortensen

[42] MaCurdy and OBrien-Strain (2000) also find that low income individuals spend a greater fraction of their budget on products produced by minimum wage workers.

production technology is linear in labor (i.e. no substitution among inputs), increases in employment increase output and should presumably reduce the product price were it endogenized. It is also worth noting that most estimated search models, including Van den Berg and Ridder (1998), Flinn (2006), Ahn and Arcidiacono (2003), find some disemployment in response to a minimum wage increase.

Arguably, our empirical results themselves - that marginal cost shocks are passed onto consumers through higher output prices - may be consistent with small disemployment effects. If restaurants face factors that limit their ability to raise prices, say because it is costly to switch prices, or because the price elasticity of demand for food away from home is infinitely elastic, the predicted disemployment effects of a minimum wage increase would be larger than if these factors did not hinder price behavior. If firms cannot pass cost increases onto consumers, then profits will be squeezed and firms may sharply cut their workforce. Given that we find rather large price increases in response to minimum wage hikes, firms seem to be able to push costs onto consumers, and are not having their profits greatly reduced.

Instead, we interpret our results to be consistent with the moderate disemployment effects reported in Aaronson and French (2007). They calibrate a structural model of labor demand that incorporates the price responses found here to show that a 10 percent increase in the minimum wage reduces restaurant employment by 2 to 3 percent, a short run response that is within the range of estimates found in the literature. Moreover, the total (low plus high skill) restaurant employment response may be as small as 1 percent.

5 Conclusion

Much work has looked at the employment implications of raising the minimum wage, with a range of estimates reported in the literature. We offer new empirical evidence using output prices both at the store-level and aggregated to the city-level. In both cases, prices unambiguously increase in response to a minimum wage change. Furthermore, the results are similar across three sources of variation in the data: cross-state differences in the size of the minimum wage change, cross-restaurant type differences in the tendency to pay at or near the minimum wage, and cross-metro differences in the fraction of workers paid at or near the minimum wage. There is no evidence that prices fall in response to a minimum wage increase.

We interpret these findings within a simple yet quite general model of employment determination that shows that monopsony and perfect competition have opposite implications for not only employment but output prices as well, so long as the minimum wage is not set too high. In particular, under monopsony, an increase in a binding minimum wage causes employment to rise and output prices to fall. Under competition, employment falls and output prices rise. Therefore, our price results appear to provide evidence against the hypothesis that monopsony power is important for understanding the small observed employment response to minimum wage changes. Indeed, our estimated price responses provide evidence against other explanations of the small employment response, including the potential substitution of nonwage for wage compensation and the importance of endogenous work effort. Rather, we interpret our price results to be fairly consistent with the textbook model of labor demand.

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Appendix A: Proofs

Proof of Theorm 1

Dixit and Stiglitz (1977) point out that two-stage budgeting techniques can be used to analyze the consumer demand problem described in the text. In the second stage the consumer solves:

$$\max_{\{Q_n\}_{n=1}^N} \left(\sum_{n=1}^N Q_n^{1-\eta_Z} \right)^{\frac{1}{1-\eta_Z}} \tag{7}$$

subject to

$$\sum_{n=1}^N p_n Q_n = X, \tag{8}$$

where X is total expenditure on $\tilde{Q}$. The consumer's first order condition for utility maximization yields

$$p_n = \lambda^{-1} \left(\frac{Q_n}{\tilde{Q}} \right)^{-\eta_Z} \tag{9}$$

where λ is the Lagrange multiplier on the budget constraint for expenditure on $\tilde{Q}$, i.e., it is the multiplier on the first stage constraint. If N is sufficiently large, then Q_n is small relative to $\tilde{Q}$ and thus the firm does not take into account the effect of Q_n on $\tilde{Q}$ when assessing the effect of Q_n on p_n.

Inserting equation (9) into equation (3) yields

$$\pi(K_n, L_n) = \Omega Q_n^{1-\eta_Z} - w_n L_n - r K_n \tag{10}$$

where $\Omega = \lambda^{-1} \tilde{Q}^{\eta_Z}$. The first order condition for maximization of the profit function (10) with respect to K_n yields

$$\Omega(1-\eta_Z) Q_n^{-\eta_Z} Q_{n,1} = r \tag{11}$$

where $\frac{\partial Q_n}{\partial K_n} \equiv Q_{n,1}$. If the firm is a price taker in the labor market, then employment is

determined by maximization of equation (10) with respect to L_n, which yields

$$\Omega(1-\eta_Z)Q_n^{-\eta_Z}Q_{n,2} = w. \tag{12}$$

where $\frac{\partial Q_n}{\partial L_n} \equiv Q_{n,2}$. First, we discuss the case where the firm is a price taker in the labor market. Given the assumptions above, all firms are of equal size, so $\tilde{Q} = N^{\frac{1}{1-\eta_Z}}Q_n$. Because of this, equation (9) shows that

$$\frac{d\ln\lambda}{d\ln w} = -\frac{d\ln\tilde{p}}{d\ln w}, \tag{13}$$

where the price index associated with $\tilde{Q}$ is $\tilde{p} \equiv \left(\sum_{n=1}^{N} p_n^{\frac{\eta_Z-1}{\eta_Z}}\right)^{\frac{\eta_Z}{\eta_Z-1}}$. Using these definitions and Assumption 5, one can show that $\frac{d\ln\tilde{Q}}{d\ln\tilde{p}} = -\eta$.

Because price changes are responses to supply shocks, price changes identify the demand curve for the aggregate product, so $\frac{d\ln\tilde{Q}}{d\ln\tilde{p}} = -\eta^*$ for some $\eta^* \in (0,\eta]$, where $\eta^* = \eta$ if the production technology is constant returns. Therefore, we may write

$$\frac{d\ln\tilde{p}}{d\ln w} = \frac{\frac{d\ln\tilde{Q}}{d\ln w}}{\frac{d\ln\tilde{Q}}{d\ln\tilde{p}}} = -\frac{1}{\eta^*}\left(\frac{d\ln Q_n}{d\ln w}\right) \tag{14}$$

where we are assuming that $\frac{d\ln N}{d\ln w} = 0$. Taking logs and differentiating equations (11) and (12) with respect to $\ln w$ and inserting equations (13) and (14) yields

$$A_n\frac{dK_n}{dw} + C_n\frac{dL_n}{dw} = 0 \tag{15}$$

$$B_n\frac{dK_n}{dw} + D_n\frac{dL_n}{dw} = 1 \tag{16}$$

and differentiating equation (11) and (12) with respect to $\ln r$ and inserting equations analogous to (13) and (14) yields

$$A_n\frac{dK_n}{dr} + C_n\frac{dL_n}{dr} = 1 \tag{17}$$

$$B_n\frac{dK_n}{dr} + D_n\frac{dL_n}{dr} = 0 \tag{18}$$

where

$$A_n \equiv (-\frac{wQ_{n,1}}{\eta^* Q_n} + \frac{wQ_{n,11}}{Q_{n,1}}) \tag{19}$$

$$B_n \equiv (-\frac{wQ_{n,1}}{\eta^* Q_n} + \frac{wQ_{n,12}}{Q_{n,2}}) \tag{20}$$

$$C_n \equiv (-\frac{wQ_{n,2}}{\eta^* Q_n} + \frac{wQ_{n,12}}{Q_{n,1}}) \tag{21}$$

$$D_n \equiv (-\frac{wQ_{n,1}}{\eta^* Q_n} + \frac{wQ_{n,22}}{Q_{n,2}}). \tag{22}$$

In matrix form this can be rewritten as

$$\begin{pmatrix} A_n & C_n \\ B_n & D_n \end{pmatrix} \begin{pmatrix} \frac{dK_n}{dr} & \frac{dK_n}{dw} \\ \frac{dL_n}{dr} & \frac{dL_n}{dw} \end{pmatrix} = \begin{pmatrix} 1 & 0 \\ 0 & 1 \end{pmatrix}$$

which can be rewritten as

$$\begin{pmatrix} \frac{dK_n}{dr} & \frac{dK_n}{dw} \\ \frac{dL_n}{dr} & \frac{dL_n}{dw} \end{pmatrix} = \frac{1}{A_n D_n - B_n C_n} \begin{pmatrix} D_n & -C_n \\ -B_n & A_n \end{pmatrix}$$

where solving for $A_n D_n - B_n C_n$ shows that the term is positive if the production function is concave.

Solving for $\frac{dK_n}{dw}$ and $\frac{dL_n}{dw}$ yields

$$\frac{dL_n}{dw} = \frac{A_n}{A_n D_n - B_n C_n} \tag{23}$$

$$\frac{dK_n}{dw} = \frac{-C_n}{A_n D_n - B_n C_n}. \tag{24}$$

The term A_n is negative. Therefore, the quantity of labor demanded falls when the price rises. The market employment change is N times the firm level employment change.

We consider the monopsony case next. Because both the production and utility functions are smooth, small changes in the minimum wage result in small changes in the $MRP(L_n)$, denoted in equation (12). Therefore, employment is determined by labor supply, not labor demand. Figure 3 clarifies this point. Given the assumed upward sloping labor supply curve, employment rises in response to an increase in the minimum wage.

QED.

Proof of Theorem 2

Consider the case of perfectly competitive labor markets first. Define p, Q and L as the market price, output and employment. Because all firms produce the same amount, $p = p_n$, $Q = NQ_n$ and $L = NL_n$ for all n. The price response to the wage change is

$$\frac{dp}{dw} = \frac{dp}{dQ}\frac{dQ}{dw} \tag{25}$$

where a Taylor's series expansion shows

$$\frac{dQ}{dw} = N\left(Q_{n,1}\frac{dK_n}{dw} + Q_{n,2}\frac{dL_n}{dw}\right). \tag{26}$$

Inserting equations (23) and (24) into equation (26) yields

$$\frac{dQ}{dw} = N\left(\frac{-Q_{n,1}C_n + Q_{n,2}A_n}{A_nD_n - B_nC_n}\right). \tag{27}$$

The numerator of this expression is negative and the denominator is positive if $Q_{n,12} > 0$. Therefore, $\frac{dQ}{dw} < 0$. Since $\frac{dp}{dQ} < 0$ by assumption, $\frac{dp}{dw} > 0$ under perfectly competitive labor markets.

Now consider the monopsony case. Recall that under the monopsony case, labor is determined by labor supply, not labor demand. Therefore, we can take labor supply as exogenous. Therefore, we may rewrite equation (25) as:

$$\frac{dp}{dw} = \frac{dp}{dQ}\frac{dQ}{dw} = \frac{dp}{dQ}\left(\frac{dQ}{dL}\frac{dL}{dw}\right). \tag{28}$$

Since by assumption, $\frac{dp}{dQ} < 0$ and $\frac{dL}{dw} > 0$, all that remains is to prove that $\frac{dQ}{dL} > 0$ when dL is taken as exogenous in order to prove that $\frac{dp}{dw} < 0$ under monopsony.

Taking the log of equation (11) and differentiating with respect to $\ln L$ yields

$$\frac{d\ln\Omega}{d\ln L} + \frac{d\ln Q_{n,1}}{d\ln L} = 0 \tag{29}$$

Because $\frac{d\ln\lambda}{d\ln L} = -\frac{d\ln\tilde{p}}{d\ln L}$ and for reasons discussed in the proof of Theorem 1, equation (29) may be rewritten as

$$\frac{d\ln\tilde{p}}{d\ln L} + \frac{d\ln Q_{n,1}}{d\ln L} = 0 \tag{30}$$

Using $\frac{d\ln\tilde{p}}{d\ln L} = \frac{\frac{d\ln\tilde{Q}}{d\ln L}}{\frac{d\ln\tilde{Q}}{d\ln\tilde{p}}}$ we may rewrite equation (30) as

$$\frac{-1}{\eta^*}\left(\frac{d\ln Q_n}{d\ln L} + \frac{d\ln N}{d\ln L}\right) + \frac{d\ln Q_{n,1}}{d\ln L} = 0 \tag{31}$$

for some $\eta^* \in (0, \eta]$, where $\frac{d\ln N}{d\ln L} = 0$. Using the chain rule,

$$\frac{dQ_n}{dL} = Q_{n,1}\frac{dK_n}{dL} + Q_{n,2}, \tag{32}$$

$$\frac{dQ_{n,1}}{dL} = Q_{n,11}\frac{dK_n}{dL} + Q_{n,12}. \tag{33}$$

Inserting equations (32) and (33) into (31) yields

$$\left(-\frac{L}{\eta^* Q_n}\right)\left(Q_{n,1}\frac{dK_n}{dL} + Q_{n,2}\right) + \left(\frac{L}{Q_{n,1}}\right)\left(Q_{n,11}\frac{dK_n}{dL} + Q_{n,12}\right) = 0 \tag{34}$$

Solving for $\frac{dK_n}{dL}$ and inserting this into equation (32) yields

$$\frac{dQ_n}{dL} = Q_{n,1}\left(\frac{\frac{1}{\eta^* Q_n}Q_{n,2} - \frac{1}{Q_{n,1}}Q_{n,12}}{-\frac{1}{\eta^* Q_n}Q_{n,1} + \frac{1}{Q_{n,1}}Q_{n,11}}\right) + Q_{n,2} \tag{35}$$

where the term in parentheses is $\frac{dK_n}{dL}$. The denominator of this term is negative. The first term in the numerator is positive but the second term is non-positive, making the sign of the numerator term in parentheses ambiguous. However, $\frac{dQ_n}{dL}$ is positive. To see this, note that

$|\big(-\frac{Q_{n,1}}{\eta^* Q_n}\big)| < |\big(-\frac{Q_{n,1}}{\eta^* Q_n}+\frac{Q_{n,11}}{Q_{n,1}}\big)|$. Therefore, $|\left(\frac{\frac{Q_{n,2}}{\eta^* Q_n}}{-\frac{Q_{n,1}}{\eta^* Q_n}}\right)| > |\left(\frac{\frac{Q_{n,2}}{\eta^* Q_n}}{-\frac{Q_{n,1}}{\eta^* Q_n}+\frac{Q_{n,11}}{Q_{n,1}}}\right)|$ and thus

$$\frac{dQ_n}{dL} > Q_{n,1}\left(-\frac{\frac{Q_{n,2}}{\eta^* Q_n}}{\frac{Q_{n,1}}{\eta^* Q_n}}\right) + Q_{n,1}\left(\frac{-\frac{Q_{n,12}}{Q_{n,1}}}{-\frac{Q_{n,1}}{\eta^* Q_n}+\frac{Q_{n,11}}{Q_{n,1}}}\right) + Q_{n,2} = Q_{n,1}\left(\frac{-\frac{Q_{n,12}}{Q_{n,1}}}{-\frac{Q_{n,1}}{\eta^* Q_n}+\frac{Q_{n,11}}{Q_{n,1}}}\right) \tag{36}$$

which is greater than 0 if $Q_{n,12}$ is greater than 0. Therefore, $\frac{dQ_n}{dL} > 0$ and thus $\frac{dQ}{dL} > 0$ and $\frac{dp}{dw} < 0$ under monopsony. QED.

Appendix B: Price and Employment Responses in Models with Endogenous Work Effort

In this appendix we assume that firms are price takers in all markets (so $\eta_Z = 0$), although using the results in Aaronson and French (2004) it is straightforward to account for the case where $\eta_Z > 0$. That paper provides a similar and more detailed derivation than the one below, although they assume $\theta = 0$.

Maximization of equation (3) with respect to K and L, and taking logs, yields:

$$\sigma(\ln r - \ln p) = \sigma \ln(1-\alpha) + \big(\ln Q - \ln K\big), \tag{37}$$

$$\sigma(\ln w - \ln p) = \sigma \ln \alpha + \big(\ln Q - \ln L\big) + \sigma\rho\theta \ln w. \tag{38}$$

Assume that product prices potentially change in response to a change in the minimum wage, but the price of capital does not. Differentiating equations (37) and (38) with respect to $\ln w$, and using results in Aaronson and French (2007), yields:

$$-\sigma\big(\frac{1}{\eta}\frac{d\ln Q}{d\ln w}\big) = -\frac{d\ln Q}{d\ln w} + \frac{d\ln K}{d\ln w}, \tag{39}$$

$$-\frac{d\ln L}{d\ln w} = -\frac{d\ln Q}{d\ln w}\big(1-\frac{\sigma}{\eta}\big) + (1-\sigma)\theta + \sigma, \tag{40}$$

where η is the elasticity of demand for the product (in absolute value), as in the text. Furthermore, differentiating the production function with respect to the wage yields:

$$\frac{dQ}{dw} = \frac{\partial Q}{\partial K}\frac{dK}{dw} + \frac{\partial Q}{\partial L}\frac{dL}{dw} + \frac{\partial Q}{\partial w}. \tag{41}$$

Solving for $\frac{\partial Q}{\partial w}$ as a function of $\frac{\partial Q}{\partial L}$, noting that $\frac{\partial Q}{\partial L} = \frac{w}{p}$, $\frac{\partial Q}{\partial K} = \frac{r}{p}$, and denoting $s = \frac{wL}{pQ}$ as

labor's share, some algebra shows that equation (41) can be rewritten as:

$$\frac{d\ln Q}{d\ln w} = (1-s)\frac{d\ln K}{d\ln w} + s\frac{d\ln L}{d\ln w} + s\theta \tag{42}$$

Using equations (39), (40), and (42), we can solve for the unknowns:

$$\frac{d\ln Q}{d\ln w} = -\eta s(1-\theta), \tag{43}$$

$$\frac{d\ln K}{d\ln w} = s(1-\theta)(\sigma-\eta), \tag{44}$$

$$\frac{d\ln L}{d\ln w} = -(1-\theta)\big((1-s)\sigma - s\eta\big) - \theta. \tag{45}$$

The price response will be

$$\frac{d\ln p}{d\ln w} = s(1-\theta). \tag{46}$$

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