New Findings Regarding the Effect of the Minimum Wage on Prices

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Abstract: We find restaurant prices grow by 0.35% for every 10% increase in the minimum wage, which is about half the previously accepted elasticity. Price increases occur almost entirely in the month the minimum wage is implemented; however, restaurants *lower* prices about 4 months ahead of the hike, which implies an overly narrow focus on the immediate months of a minimum wage increase might miss important effects. Small minimum wage increases are associated with price reductions, a finding not previously reported and consistent with the existence of monopsonistic competition in restaurant labor markets.

Keywords: minimum wage, restaurant prices, monopsony, monopsonistic competition

JEL Codes: J38, J42, J31, E31

1. Introduction

Partly due to the failure of Congress to raise the federal minimum wage, many states and cities have passed their own minimum wage laws. In turn, these state and city laws promoted a renaissance in the study of the employment effect of minimum wage hikes for two reasons. First, they created a greater number of minimum wage changes to be studied. Second, by increasing geographical variation in minimum wage policy, these state and city minimum wage laws produced something akin to "natural experiments" whereby the employment statistics in a state or city that increased its minimum wage could be compared to those in neighboring areas that did not increase their minimum wage. One strand of the new minimum wage literature found—contrary to the previously accepted belief—that some minimum wage hikes led to either no decline in employment or to a slight increase in employment (e.g., Card and Krueger 1994, 1995; Dube, Lester, and Reich 2010). In contrast, a second strand of this literature continued to find evidence supporting the claim that minimum wage hikes reduced employment, at least in some circumstances and for some types of workers (e.g., Neumark 2001; Neumark and Wascher 2002, 2007, 2008).

Although the minimum wage literature most often focuses on employment effects, a full consideration of the impact of minimum wage hikes must also consider how such hikes affect *prices*. Such price changes are important both in themselves, as one direct effect of minimum wage hikes, and because price changes affect the welfare of consumers. However, the impact of minimum wages on prices has been little studied. Wessels (1980) and Card and Krueger (1995) were among the first to study price increases following minimum wage hikes, but the most influential of these studies has been a series of papers by Daniel Aaronson and coauthors. Aaronson (2001), Aaronson and French (2006), Aaronson and French (2007), and Aaronson, French, and MacDonald (2008) studied the effect of minimum wage hikes on prices, and then used their empirical results to shed light on the competitive nature of restaurant labor markets and the employment impact of minimum wage hikes.

Academic research, policy proposals, and the news media have cited the above studies by Aaronson and co-authors as authoritative.[1] However, these studies should be updated for four reasons. First, these studies rely on data from no later than 1997, but the United States has seen a profusion of state and city minimum wage laws since then. In particular, whereas Aaronson (2001) studied the impact of 34 minimum wage increases, we are able to study the impact of 160 minimum wage hikes, which permits us to measure the effect of minimum wages on prices more accurately than before. Second, the nature of minimum wage hikes has changed since 1997. Small and regularly scheduled annual hikes have increasingly replaced large hikes coming at the whim of Congress. We cannot simply assume the two types of minimum wage hikes have the same effects. Third, we use the data in ways that differ from the previous literature to extract greater insight into the process through which minimum wages lead to price increases. For instance, while Aaronson (2001) used a combination of monthly and bimonthly data to determine monthly price dynamics, we use only monthly data to study monthly dynamics, which is a more appropriate procedure. Fourth, by including data after 1997, we can use CPI data, not available to Aaronson (2001), that are less affected by various biases (such as substitution bias) and, so, that should generate more accurate estimates of the price effects of minimum wage hikes.

The payoffs for this reconsideration of the effect of minimum wage hikes on prices are several. First, we find that minimum wage hikes cause prices to rise much less than previously reported. Second, we show that the dynamics of the price effect—that is, when price changes occur—also differ from what was previously accepted. Third, we present evidence that although

[1] Among recent economic research, policy proposals, and news media citing Aaronson (2001) as authoritative are, respectively, MaCurdy (2015), Dube (2014), and Kirkham and Hsu (2014).

large minimum wage hikes boost prices, small minimum wage hikes are associated with price declines. The latter finding is consistent with the claim that labor markets in the restaurant industry have some degree of monopsony, which supports the claim that minimum wage hikes need not always reduce employment.

2. Literature Review

Minimum wage hikes increase the costs of businesses employing minimum wage workers and some of this increased cost might be passed through to customers of these businesses in the form of higher prices. The size of the so-called "pass-through" is measured by a particular elasticity: the percent change in output prices associated with a 1 percent increase in the minimum wage.

The oft-cited study by Aaronson (2001) estimated that the size of the pass-through—in particular, the cumulative minimum wage-price elasticity from 3 months before up to 3 months after a minimum wage hike—was 0.072.[2] In other words, a 10% increase in the minimum wage—one that is somewhat larger than the size of the typical minimum wage increase—caused the average $10.00 restaurant item to become a $10.07 item. Aaronson also found that this price increase was spread out over the three months centered on the month the hike was implemented. The coefficients he reported for T-1, T, and T+1, where T was the month the hike was implemented, were: 0.022, 0.028, and 0.014, each of which was statistically significant. According to these point estimates, less than half of the price increase occurred in the actual

[2] Aaronson (2001), table 4, regression 2. Throughout this paper we will use the results from Aaronson's regression 2 as representing his findings.

month the minimum wage hike was implemented. Aaronson also found that prices dropped four months in advance of the minimum wage hike, but he did not discuss the possible causes nor implications of such a price drop. These results were generated using metropolitan-area food away from home (FAFH) CPI data between 1978 and 1995.

Aaronson, French, and MacDonald (2008) used micro-level restaurant price data for the period between 1995 and 1997, during which two changes to the federal minimum wage were implemented, to generate a wage-price elasticity of, once again, about 0.07. Although the empirical literature is somewhat limited outside of these two formative works, other studies have found similar results in other countries and other cases.[3]

In turn, Aaronson and French (2007) argued that the estimated size of the pass-through was consistent with perfect competition in low-wage labor markets. Using a calibrated model of labor demand, they estimated that a 10% increase in the minimum wage in a perfectly competitive industry would lead to approximately a 0.7% increase in output prices, which was exactly what the existing empirical work had found. Therefore, Aaronson and French concluded that low-wage labor markets in the restaurant industry are best characterized as perfectly competitive. At stake here is the employment impact of a minimum wage increase. If low-wage labor markets are

[3] Other studies include Fougère, Gautier, Bihan (2010) who studied France, Lemos (2006) who studied Brazil, and Wadsworth (2010) and Draca, Machin, and Van Reenen (2011) who both studied the U.K. Another national-level study that focuses on the prices of a few restaurant items (burgers, chicken, pizza) is Basker and Khan (2013). Lemos (2008) reviews many of the relevant studies.

indeed perfectly competitive, an increase in the minimum wage will boost prices, lower output as the industry moves up the demand curve, and reduce industry employment. This work on the effect of minimum wage on prices, then, speaks to the ongoing controversy about the employment impact of minimum wage increases.

3. Data and Data Transformations

The dependent variable in this study is the change in the log of food away from home CPI (FAFH CPI), a price index generated by the Bureau of Labor Statistics for select U.S. metropolitan areas and available on their website. Aaronson (2001) used the same dependent variable. FAFH includes food purchased and consumed outside of the home, and, for the most part, includes items sold at full- and limited-service restaurants.[4] The restaurant industry is more dependent on minimum wage workers than any other U.S. industry,[5] and so it makes sense to focus on that industry when studying the impact of the minimum wage on prices. We include in our analysis all 28 metropolitan areas that have either monthly or bimonthly FAFH CPI data for

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[4] FAFH also includes ready-to-eat food purchased at motels and restaurants, food provided at employer and school sites, along with food purchased at vending machines and from mobile vendors (Bureau of Labor Statistics, 2015). For conciseness, we will refer in the text to "restaurants" when we talk about the group of sites selling food away from home as restaurants make up the majority of sites whose prices enter into FAFH CPI.

[5] Almost half of all workers in the U.S. who earn the federal minimum wage, or less, are found in food preparation and food serving related occupations (Bureau of Labor Statistics, 2016, table 10).

at least part of the period of our study, 1978-2015.[6] We begin our analysis in 1978, as did Aaronson (2001), as the 1978 federal minimum wage hike was the first such hike after changes in the Fair Labor Standards Act that directly affected the restaurant industry (for instance, a restructured tip credit process and a repeal of the partial exemption of restaurant employees from overtime rules).

One characteristic of the CPI data requires comment. In January 1999, the Bureau of Labor Statistics switched to a geometric mean formula when they calculated CPI price indexes. This switch was prompted by arguments that the BLS's previous method for calculating the CPI

[6] Using the major city within the area to identify them, the metropolitan areas included in our study are: Anchorage (bimonthly, until 1986), Atlanta (bimonthly, full time period), Baltimore (bimonthly, until 1995), Boston (bimonthly, full period), Buffalo (bimonthly, until 1986), Chicago (monthly, full period), Cincinnati (bimonthly, until 1986), Cleveland (bimonthly, full period), Baltimore/Washington D.C. (bimonthly, since 1995), Washington D.C. (bimonthly, until 1995), Dallas (bimonthly, full period), Denver (bimonthly, until 1986), Detroit (monthly until 1986, then bimonthly for rest of period), Honolulu (bimonthly, until 1986), Houston (bimonthly, full period), Kansas City (bimonthly, until 1986), Los Angeles (monthly, full time period), Miami (bimonthly, full period), Milwaukee (bimonthly, until 1986), Minneapolis (bimonthly, until 1986), New York City (monthly, full period), Philadelphia (monthly until 1997, then bimonthly for rest of period), Pittsburgh (bimonthly, until 1997), Portland (bimonthly, until 1986), San Diego (bimonthly, until 1986), San Francisco (monthly between 1987 and 1997, bimonthly for the rest of the series), Seattle (bimonthly until 1986 and then from 1997 for the rest of the period), and St. Louis (bimonthly until 1997).

produced an upward bias to the CPI and its subcomponents. The new geometric mean formula mimics consumers' substitution between the products they buy as they respond to changes in relative prices, something the previously-used Laspeyres formula did not take into account (Dalton, Greenlees, and Stewart, 1998). If the CPI was biased upwards before 1999, then any study of the size of the pass-through that uses only pre-1999 CPI data, such as Aaronson (2001), would generate estimates of the pass-through that are potentially biased upwards. Our study, which uses data for 1978-2015, can use the more accurate geometric mean-based CPI for the second half of the period and, so, generates more accurate estimates of the pass-through.

The independent variable of interest in our regressions is the change in (binding) minimum wage rates.[7] Our data on minimum wages comes from various issues of the *Monthly Labor Review*, state Department of Labor reports, and city and county ordinances for San Francisco, San Jose, Oakland, Berkeley, Washington, D.C., and Prince George's and Montgomery counties. The years 1978-2015 saw 11 federal minimum wage increases, 126 binding state minimum wage increases, and 23 city minimum wage increases. Table 1 reports the month and year of passage for all of these increases.

[Table 1]

We also include month, year, and metropolitan area fixed-effects. One additional control is "CPI-All" (Urban Consumers) for each metropolitan area, included to take into account

[7] Not all minimum wage hikes were binding in all metropolitan areas as some federal minimum wage hikes failed to increase the minimum wage above an already existing state minimum wage.

unknown determinants of FAFH CPI inflation.[8] The latter control variable might rob some of the influence from minimum wage changes as this control variable is affected by inflation in the FAFH sector.

Some metropolitan areas in this study encompass single states, such as that for San Francisco. Others include territory from multiple states. For multistate metropolitan areas, the BLS generates FAFH CPI by using prices from a sample of restaurants drawn from all parts of the metropolitan area. For example, the FAFH CPI for the New York-Northern New Jersey-Long Island metropolitan area is constructed from prices taken from a sample of restaurants drawn from four states: New York, Pennsylvania, New Jersey, and Connecticut. Therefore, the FAFH CPI for this single multistate metropolitan area is potentially affected by minimum wage hikes implemented by four different states. Table 2 (below) provides information about the metropolitan areas in our sample that include territory from more than a single state.

[Table 2]

The existence of multistate metropolitan areas benefits this study. We can include in our dataset many more state minimum wage changes than would have been the case if, say, the New York metropolitan area only included territory from New York State alone. However, we need to transform a single-state minimum wage increase affecting only restaurants in one portion of a multistate metropolitan area into a variable measuring its impact on average FAFH prices for the full multistate area. We will assume that a 10% state minimum wage hike that affects only 20%

[8] Published by the BLS and available for download online for most years and areas. Some of this data is only available in printed BLS reports.

of the restaurants in a metropolitan area (that is, those restaurants in that state) will have an impact on metropolitan restaurant prices equal to a 2% (that is, 10% x 20%) minimum wage hike for the whole metropolitan area. We define, then, the "restaurant-weighted minimum wage change" (RMW) as,

$$\Delta\log(mw^*)_{it} = \sum_s[\lambda_{ist} * \Delta\log(mw)_{st}] \quad (1)$$

where i is the metropolitan area, s is the state, t is the month, λ_{ist} is the proportion of restaurants from state s in month t in metropolitan area i, and $\Delta\log(mw)_{st}$is the minimum wage change in state s in time t.[9] When a metropolitan area includes only a single state, λ_{ist} will equal 1 and RMW for any minimum wage will simply be the change in the associated state minimum wage.[10]

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[9] For example, consider the metropolitan area of District of Columbia which, in 2009, included selected parts of surrounding counties from Maryland, Virginia, and West Virginia. Factoring in the number of restaurant establishments in each of these subsamples of counties as a percent of the total establishments in those counties gives the following weight to apply to each state's minimum wage in order to construct the District of Columbia minimum wage variable: D.C. (0.164), Maryland (0.344), Virginia (0.471), and West Virginia (0.020). Thus, if Maryland increased its minimum wage in January 2009 by 10%, this would be a full metropolitan area equivalent minimum wage change of 3.44% (=10% x 0.344).

[10] The Bureau of Labor Statistics (2015) describes how the BLS selects outlets to use as a source of prices. The BLS selects outlets so they reflect where people are buying their food. We use the regional distribution of restaurant establishments as a proxy for the regional distribution of restaurant purchases. This is an imperfect proxy as regional differences in restaurant sizes and

To check the appropriateness of using RMW, we ran our regressions with a subsample of series that contained only data from a single state (such as San Francisco, Los Angeles, Atlanta, and Detroit). The coefficients in these regressions did not differ substantially from the ones based on the full sample (and using RMW).

Following the lead of Aaronson (2001), we use a dataset that combines monthly and bimonthly data. As can be seen in table 3 (below), by combining these two types of data we greatly increase both the number of observations and the number of binding minimum wage increases that appear in our dataset. Later, we will discuss the benefits and costs of combining these two types of data.

[Table 3]

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regional differences in average consumer restaurant bills might lead the distribution of restaurant purchases to vary from the regional distribution of restaurant establishments. When population weights are used in place of restaurant establishment weights, the results for our regressions did not differ much from what is reported in the text. For the calculation of RMW, the number of restaurant establishments in the various state subsections of multistate metropolitan areas came from County Business Patterns while information about the particular towns and cities included in each state subsection of a metropolitan area came from the definitions of these metropolitan areas provided by the Office of Management and Budget.

4. When Do Prices Change?

If a restaurant increases its prices in response to a minimum wage hike, it will likely increase them in the month the minimum wage hike is imposed. Additionally, a restaurant might also adjust its prices in the months after the hike due to menu costs and/or other frictions in price adjustments. Conceivably, prices might even rise ahead of an announced minimum wage hike as restaurants, say, seek to spread price increases over multiple months in order to insulate customers from price shocks. As noted above, Aaronson (2001) found prices changes grew in each of the three months centered on the month the minimum wage was implemented, a result that might be explained by the factors mentioned above.

Because bimonthly series are not granular enough to reveal monthly pricing behavior, we temporarily set aside our bimonthly price series. We will use our monthly series to discover how price changes, associated with minimum wage hikes, are spread out over time. This will not only deepen our understanding of the dynamics of minimum wage-induced price changes but is also, as will be seen, a necessarily preliminary step in our analysis of price changes that uses our full dataset, comprising both monthly and bimonthly data.

In this section, we use data from three metropolitan areas (New York, Chicago, and Los Angeles) that have monthly data for the entire period along with data from three additional metropolitan areas (San Francisco, Philadelphia, and Detroit) that have monthly data for various subsets of the period 1978-2015.[11] Together, these metropolitan areas account for only about

[11] Monthly observations were reported for San Francisco between 1986 and 1998, for Philadelphia before 1998, and for Detroit before 1987.

20% of all federal-level minimum wage increases, about 30% of all state-level minimum wage increases, none of the local-level minimum wage increases, and only 1,850 of 4,988 total observations found in our full sample.

To discover the impact of minimum wage hikes on restaurant prices in the months surrounding the implementation of the hike, we estimate the following equation:

$$\Delta \log(FAFH)_{it} = \alpha + \sum_{t=-4}^{4} \beta_t * \Delta\log(mw^*)_{it} + \theta * \log(cityCPI)_{it} + c_i + \epsilon_{it}. \quad (2)$$

This equation has Food Away from Home (FAFH) inflation as the dependent variable and, as independent variables, the restaurant-weighted minimum wage change, mw* (defined in Equation 1), overall metropolitan area CPI inflation, along with fixed effects for metropolitan areas, months, and years. This regression includes leads and lags of four months to capture the impact of minimum wage hikes on prices in the months both preceding and following the month in which a minimum wage hike is implemented. We also include city-level fixed effects (c_i) to account for time-invariant unobserved heterogeneity in FAFH inflation between different cities, and include city-level inflation to account for other factors shaping FAFH inflation. In addition, we include month and year fixed effects.

Throughout this paper, we use cluster-robust standard errors for statistical inference—where the clusters are metropolitan areas—because of autocorrelation within the residuals of our regressions. Further, because of the small number of clusters in our data, we base our statistical inference on critical values taken from T(G-1), where G equals the number of clusters (Cameron and Miller 2015). The latter gives substantially higher p-values for a given standard error than approaches based on degrees of freedom.

Table 4 reports the results of our estimation of Equation 2. As we go from regression 1 to regression 3, we add month and year dummies along with metropolitan area overall CPI as controls. We use regression 3 as the basis for our discussion below.

[Table 4]

In regression 3, the contemporary elasticity is 0.039, a value that is statistically significant. In the three months surrounding the minimum wage hike, the total elasticity rises to 0.047, although this value is not statistically significant (p-value=0.12). We will defer, however, full consideration of the coefficients and sums of coefficients in our regressions until we discover what our full sample (including both monthly and bimonthly data) says about their values.

Our interest in regression 3 is in the monthly dynamics it reveals. In this regression, only the coefficient for the month in which the minimum wage hike was imposed, T, achieved statistical significance.[12] The coefficients for the two months surrounding the month the hike was imposed, T-1 and T+1, fell far short of statistical significance. This latter result differs from the canonical Aaronson (2001), which found statistically significant price increases for not only the month the minimum wage hike was implemented but also for the month before and the month after. Our finding that restaurants increase prices only in the month the minimum wage hike occurs suggests that menu costs, the desire to insulate customers from price shocks, and other real world

[12] The finding that none of the leads or lags in the regressions appearing in table 4 achieves statistical significance is evidence against the potential claim of endogeneity—i.e., that minimum wage policy is partly a response to inflation.

factors are possibly not important enough to spread price increases away from the month the minimum wage hike is imposed.[13]

Regression 3 also failed to reproduce one intriguing finding of Aaronson (2001): a statistically significant *decline* in prices four months ahead of US minimum wage hikes. Our coefficient for T-4, -0.014, was almost identical to that reported by Aaronson (-0.013), but ours fell short of statistical significance (p-value = 0.15). However, as will be seen below, regressions that use our full dataset reveal that prices do indeed fall several months in advance of a minimum wage hike. However, before we use our full dataset we need to consider how it is constructed.

5. Interpolation: Friend and Foe

Following the lead of Aaronson (2001), we join our monthly and bimonthly series to create a single, larger dataset. The first step in joining these two types of data is transforming, through a process of interpolation, our bimonthly series into monthly series. We interpolate data for both FAFH CPI and City CPI-All for those metropolitan areas and periods that have only bimonthly data. The major benefit of joining monthly data with the bimonthly (interpolated) data is that we

[13] On the one hand, our conservative approach of using critical values from T(G-1) works against our finding support for Aaronson's (2001) finding of statistically significant price increases in the months before and after a minimum wage hike. On the other hand, our point estimates for the price effects on these months (-0.001 and 0.008) fall short of those reported by Aaronson (0.022 and 0.014) so even if we used a less conservative approach for determining critical values our findings would have been at odds with Aaronson's finding that a substantial part of the price increases following a minimum wage hike occurred on the months before and after the hike.

increase the number of observations in our sample from 1852 to 8124[14] and expand the number of individual minimum wage changes we consider from 82 to 354.

In much of the econometric literature, interpolation involves creating monthly data from quarterly data or creating quarterly data from yearly data. Additionally, interpolation often involves using related higher frequency data to inform this process (e.g., Chow and Lin, 1971). In our study, the frequency change is much smaller (from bimonthly to monthly), and we transform the data in a setting in which no related higher frequency data exists. Thus, we interpolate by averaging the neighboring bimonthly price data.

Interpolation causes all sorts of mischief. It produces measurement error in the interpolated data points, in our cases for both the dependent variable (FAFH CPI) and an independent variable (City CPI-All). Interpolation also increases first-order correlation between residuals. Both of these facts lead to coefficients and standard errors that are biased. The nature of the bias depends on the nature of the measurement error and the particular estimation technique used.

[14] The 8124 observations include 1852 monthly observations, 3136 bimonthly observations, and 3136 interpolated "observations." For statistical inference that depends on the degrees of freedom, it is inappropriate for use to use as the starting point in the determination of the number of the degrees of freedom the total number of observations (here, 8124) entering into the regression. This is because our interpolated data is not independent: it has been generated from a linear combination of the bimonthly data on either side of it. However, the determination of the correct degrees of freedom is a moot issue in this study as we use critical values from T(G-1), where G equals the number of clusters, for statistical inference.

Consider, first, how interpolation affects the estimation of the growth rates of FAFH CPI. Assume we have a series of CPI levels, $[x_1,x_2,x_3] = [x_1, x_1(1+g_2), x_2(1+g_3)]$. If we estimate x_2 by linearly interpolating between x_1 and x_3, we get $x_2^i = [x_1 + x_1(1+g_2)(1+g_3)]/2$. In this case, the rates of growth between x_1 and x_2^i and between x_2^i and x_3—respectively, g_2^i and g_3^i—will equal the true rates of growth if and only if $g_2 = 1/(1-g_3)-1$.[15] On the other hand, if $g_2 < 1/(1-g_3)-1$, then $g_2^i > g_2$ and $g_3^i < g_3$. Finally, if $g_2 > 1/(1-g_3)-1$, then $g_2^i < g_2$ and $g_3^i > g_3$. In short, using interpolation to generate FAFH inflation rates can generate systematic errors, and the sign of the errors—in this case, in the estimation of the impact of a minimum wage increase on x_2—depends on the relationship between the growth rates for the underlying monthly data, g_2 and g_3.

The problem is that bimonthly data only reveals the product of $(1 + g_2)(1 + g_3)$, and not g_2 and g_3 separately. Bimonthly data itself, then, does not give us the information we need to determine the errors caused in calculating growth rates by using interpolated data. Luckily, we have a way to discern the unseen monthly dynamics behind the bimonthly data because of the patterns we have observed in the monthly data. Regression 3, generated by monthly data, provides estimates of the growth rates of FAFH CPI in the months surrounding a minimum wage hike. For the months [T-1, T, T+1] these estimates were [-0.001, 0.039, 0.008]. However, suppose we did not have the actual monthly FAFH CPI data but, instead, only had bimonthly data. With the bimonthly data falling on even months, […,T-2, T, T+2, …], we can let $g_2 = -0.001$ and $g_3 = 0.039$. In this case, $1/(1-g_3)-1 = 0.041 > -0.001$. Then, according to the previous

[15] If $g_2 = g_2^i$ then $g_2 = \frac{[x_1+x_1(1+g_2)(1+g_3)]/2}{x_1}$. By rearranging the latter equality you get $g_2 = [1/(1-g_3)-1]$.

paragraph, $g_2^i > g_2$ and $g_3^i < g_3$. That is, the growth rates calculated using the interpolated data will exceed the actual growth rate in T-1 and will fall short of the actual growth rate for T. If, however, we assume the bimonthly data was collected in odd months we can let $g_2 = 0.039$ and $g_3 = 0.008$. In this case, $1/(1-g_3)-1 = 0.008 < 0.039$. In this second case, $g_2^i < g_2$ and $g_3^i > g_3$. That is, the growth rates calculated using the interpolated data will fall short of the actual growth rate in T and will exceed the actual growth rate for T+1.

In summary, the above analysis shows that if we had only bimonthly data but some of it was for even months while others was for odd months, any regressions using the bimonthly data (with the monthly gaps filled by interpolation) would give too high coefficients for T-1 and T+1 and a too low coefficient for T.

We confirmed this conclusion by comparing the results seen in regression 3 with a regression that used bimonthly data in place of monthly data. We started with the monthly data used to estimate regression 3, and then fabricated bimonthly data from it by discarding every other observation. For half of the series we delete the December/February/April/… FAFH price index observations, and for the other half we delete the January/March/May/... observations. Next, we interpolated between these fabricated bimonthly data points to fill in the gaps we had just created, giving us a monthly series once again. Next, we estimated equation 2 using this artificial monthly series. The estimated coefficients for [T-1, T, T+1] in this new regression were [0.010, 0.022, 0.015]. This compares to the results from regression 3 of [-0.001, 0.039, 0.008]. Although none of the coefficients for this new regression achieved statistical significance, interpolation reduced the contemporaneous impact of a minimum wage hike and spread the price effect to the month preceding and the month following the hike. This is as predicted. Notably, the estimated total effect over [T-1,T+1] was almost identical for the two regressions, 0.047 versus 0.046.

By transforming our bimonthly data into monthly data through the process of interpolation, and then joining this with our real monthly data, we can investigate the effect of many more minimum wage hikes than otherwise would have been the case. Yet, we must use caution when interpreting the individual coefficients generated by regressions using the combined monthly/bimonthly data. But although interpolated bimonthly price data might lead us astray when we try to discover *when* the price impact occurs, they retain accurate information about *the total effect* of minimum wages on prices when we sum up the coefficients over, say, the three months surrounding the minimum wage hike.

Finally, we should point out that the finding of Aaronson (2001)—that minimum wage hikes generated price increases both in the month before and the month after the hike—might be at least partly explained by his use of interpolated bimonthly data to draw conclusions about monthly price dynamics. Our finding in the previous section—that prices did not increase in those months—is likely more compelling as our finding was generated by using monthly data alone.

6. How Sensitive Are Prices to Minimum Wage Increases? Are the Results Consistent with Perfect Competition in Labor Markets?

We now pool monthly and bimonthly (with interpolation) data for the 1978-2015 period and estimate equation 2. We will use the results of regression 5 in table 5, which includes City CPI-All as a control, as the basis for our discussion.

[Table 5]

According to regression 5, the contemporaneous price elasticity is 0.023 while that for the month before and the month after are, respectively, 0.010 and 0.013. All these coefficients are statistically significant. However, because some of the data used to estimate regression 5 was interpolated, we must assume that the coefficient for T is downward biased in this regression while those for T-1 and T+1 are upward biased.[16] As a result, we do not have evidence that minimum wage hikes actually increase prices the month before or the month after the hike is implemented despite the statistically significant coefficients for T-1 and T+1. Some of the price increase that regression 5 gives to T-1 and T+1 likely came from T.

The sum of elasticities for [T-1,T+1] is 0.046, which is statistically significant and almost identical to the 0.047 we found for the same period for regression 3. Minimum wage hikes do, on average, lead to price increases around the month the hike is implemented; much, if not most, of this increase might occur in the month the hike is implemented.

[16] The coefficients for the three months surrounding the minimum wage hike in regression 5 are almost identical to those seen, [0.010, 0.021, and 0.015], when we used our monthly data to generate bimonthly data (then filled out by interpolation) to estimate equation 2. This is additional evidence that, despite the statistically significant increase in prices for T-1 and T+1 provided by regression 5, prices might not have actually increased in those months. The statistically significant coefficients might have been an artifact of the interpolation process that spreads price increases from T to the month before and the month after.

Regression 5 indicates that prices fall in advance of a minimum wage hike: we get statistically significant negative coefficients of -0.009 and -0.006 for T-4 and T-3. Regression 3 produced coefficients for the same months of -0.014 and 0.000 but the former fell short of statistical significance (p=0.15).

Interestingly, Aaronson (2001) also found a decline in prices in the months ahead of a minimum wage hike although he said little about it. The mechanism leading to such a price decline is unknown. Importantly, the total decline in prices seen for T-4 and T-3 (-0.015) is almost one-third the size of the price increase occurring during the three months surrounding the minimum wage increase (0.046). An overly narrow focus on price changes in the immediate months surrounding a minimum wage hike might, therefore, lead researchers and policymakers to overlook other price changes—in this case, *lower* prices—that occur in the months ahead of the minimum wage hike and so over-estimate the effect of minimum wage hikes on prices.[17]

The total effect of a minimum wage on prices over the full person considered, [T-4,T+4], is 0.035, a value that achieved statistical significance. Therefore, a 10% increase in the minimum wage causes a $10.00 item to become, on average, a $10.04 item over the nine months centered on the minimum wage hike. This response of prices to a minimum wage hike is about one-half that reported in Aaronson (2001).[18]

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[17] If prices change in the months ahead of a minimum wage hike then the possibility exists that employment might also change ahead of such a hike.

[18] Aaronson (2001), table 4, regression 2.

Our finding of a much smaller pass-through than previously reported is noteworthy in itself, but it also calls into question one bit of evidence offered in support of the claim that restaurant labor markets are perfectly competitive. Aaronson and French (2007), building on a set of assumptions about the operation of restaurants in a hypothetical perfectly competitive market, argue that restaurants in perfectly competitive markets will fully pass through any increase in the minimum wage and that the full pass-through elasticity will be equal to approximately 0.07. Since they find, in various regressions, elasticities near 0.07, they conclude that restaurant labor markets are best characterized as perfectly competitive. Further, as a wage increase should cause employment decline in a perfectly competitive market, this finding that restaurant labor markets are perfectly competitive provides indirect evidence that a minimum wage hike should reduce employment.

In contrast, our estimated elasticity of 0.035 falls short of the 0.07 that Aaronson and French (2007) argue is consistent with perfect competition. Further, that prices fall several months in advance of a minimum wage hike is hard to square with a perfectly competitive setting, in which higher costs for businesses should only cause them to increase their prices, and do so only in response to increases in *current* costs. We fail to find support here, then, for the claim that labor markets in the restaurant industry are perfectly competitive.

7. The Response of Prices to Minimum Wages in Monopsonistic Competitive Labor Markets: Theory and Evidence

Stigler (1946) argued, based on a numerical example, that "If any employer has a significant degree of control over the wage rate he pays for a given quality of labor, a skillfully-set minimum wage may increase his employment and wage rate and, because the wage is brought closer to the value of marginal product, at the same time increase aggregate output" (p. 360). If we assume that an increase in industry output pushes down prices, a "skillfully-set" minimum wage should lower prices.

The standard model of monopsony in the labor market can illustrate Stigler's point. The monopsonist has market power and faces an upward-sloping labor supply curve. To attract more workers, the monopsonist must increase the wage, which implies the marginal cost of labor for the monopsonist is greater than the wage. In Figure 1(a) (below), W_x identifies where the marginal cost of labor equals the value of the marginal product of workers. In the absence of a binding minimum wage, a monopsonist employer will hire N_1 employees, paying them W_1. The wage paid by the monopsonist will fall short of the market-clearing wage, W_C.

[Figure 1]

Figure 1(b) shows one Stiglerian "skillfully-set" minimum wage. Suppose the wage is initially at W_1, but then the government imposes a minimum wage equal to W_{mw1}. For the monopsonist, the marginal cost of labor now includes a horizontal line starting at W_S.[19] The new marginal cost curve will induce the monopsonist to increase employment up to N_S as each worker below that level of employment has a marginal cost below the worker's value of marginal product. As can be seen, if one starts from W_1, any minimum wage below W_x—in Stigler's terms, a skillfully-set one—will lead the employer to hire more workers. If indeed

[19] The new marginal cost of labor curve will be a horizontal line extending from W_S to the labor supply curve, at which point it will jump up to the previous rising marginal cost curve. For the sake of simplicity, we do not include the full marginal cost of labor curve in Figure 1(b) or 1(c).

greater employment leads to greater output and greater output pushes down prices, a skillfully-set minimum wage leads to lower prices. Alternatively, if the minimum wage hike is too large—that is, pushes the wage above W_x—employment will fall, output will likely shrink and, as a result, prices will rise.

Even if a prior binding minimum wage has been imposed, but keeps the wage below W_c, a further increase in the minimum wage can still lead to expanded employment and, so, to lower prices if the minimum wage increase is small enough. Here, "small enough" is defined as a minimum wage increase smaller than the difference between the pre-increase wage and the value of the marginal product of workers at the current level of employment. But, once the minimum wage exceeds W_c, any further increase in the minimum wage, no matter how small, should lead to lower employment and higher prices. With a prior minimum wage equal to or exceeding W_c, it would be impossible to find a Stiglerian skillfully-set minimum wage.

However, a model of monopsony is unlikely to portray the restaurant industry, in which numerous firms compete and the threats of entry and exist are ever-present. Bhaskar and To (1999) developed a model of monopsonistic competition that introduces competition and entry/exit, and then considered the impact of a minimum wage in such a setting. In their model, starting from the unconstrained optimal wage for the monopsonistic competitor, a minimum wage that is small enough—in their model, below the mean of the value of the net revenue product of labor and a measure of the reservation wage of laborers—causes either a price increase or price reduction depending on whether factor substitutability is high or low. On the other hand, a large minimum wage increase will cause prices to grow.[20]

The above prediction that some "small enough" minimum wage increases can cause prices to fall differentiates models that introduce elements of monopsony in the labor market from the model of perfect competition.[21] In perfect competition, *any* binding increase in the minimum wage, no matter how small, lowers employment, lowers output, and increases output prices.

If prices in our dataset fall when a minimum wage hike is "small enough" this would be positive support for the claim that restaurant labor markets have elements of monopsony. Unfortunately, we can only determine which minimum wage increases fulfill this condition if we

[20] According to Bhaskar and To (1999), "…if the minimum wage is less than the mean of the net revenue product of labor (φ) and the high reservation wage (v), aggregate employment is increasing in the minimum wage, and is decreasing otherwise" (p. 195). And then, "If we allow for more general production functions, minimum wages will always raise prices whenever employment falls. When employment rises, prices may rise or fall, depending upon whether the production function allows for more or less factor substitutability" (p. 199).

[21] For the sake of completeness, we note that the model presented by Sweezy (1939) leads to a similar conclusion: small increases in wages might lead firms to not raise their prices while large wage increases will lead to price increases. His conclusion is based on the claim that the optimal response to small wage increases for interdependent businesses that are unable to collude is to keep prices the same if the price elasticity of demand is high enough.

have information about the net revenue product of labor, the reservation wage of workers, and the level of factor substitutability. We do not have such information.

But a possible way to move forward exists. We split the minimum wage changes in our sample into two groups, "small" and "large" depending on whether the minimum wage change is below or above the average minimum wage increase in our sample, 6.9%. Although we cannot know whether any particular minimum wage increase satisfies the condition necessary for it to increase or reduce prices according to, say, Bhaskar and To (1999), the small group in our sample should capture more minimum wage changes that satisfy these conditions than will the large group. If that is indeed the case, we can then estimate an equation based on regression 6, but separate the minimum wage increases into these two groups, small and large. If the individual and sum of coefficients for the T_i's for the small group are smaller than those for the large group, this would be consistent with the predictions of the monopsony model of the labor market.

Splitting our sample this way reduces the chance we find evidence consistent with monopsonistic competition. First, some of the minimum wage changes we include in our small group will fail to meet the conditions that lead to price declines. For instance, they might be small in percentage terms but they might be greater than the conditions specified by Bhaskar and To. Alternatively, the prior minimum wage might have already exceeded the market clearing wage. Second, some of the minimum wage changes we include in our large group might lead to price declines because, for instance, the starting minimum wage is far below what it would have been in a hypothetical perfectly competitive market, although perhaps above what it would have been in unconstrained monopsonistic competition.

Regression 6, found in table 6, presents the results of splitting the minimum wage hikes in our sample into two groups, small and large. As can be seen, for small minimum wage increases the coefficient for T-4 is negative and statistically significant. The effect of small minimum wage increases on prices over three, seven, and nine months centered on the minimum hike are likewise negative, although none of these effects are statistically significant. In contrast, large minimum wage hikes generate positive and statistically significant increases for T-1, T, and T+1. Likewise, the sum of coefficients for [T-1,T+1], and [T-4,T+4] for large increases are also positive and statistically significant. Small and large minimum wage increases appear to have different effects on prices.

[Table 6]

The results from regression 6 are consistent with the existence of monopsonistic competition in restaurant labor markets. Previous research has proposed that the *employment* response to minimum wage hikes supports the existence of elements of monopsony in low-skilled labor markets (e.g., Card and Krueger, 1995). We have shown that the response of *prices* to minimum wage hikes also supports the existence of elements of monopsony in such labor markets.

8. Conclusion

We used BLS price data to measure how restaurant prices respond to minimum wage hikes over 1978-2015. We found that prices grew by 0.35% for every 10% increase in the minimum wage, which is about half the previously accepted elasticity. In other words, a 10% increase in the minimum wage causes the average $10.00 restaurant item to become a $10.04 item.

Price *increases* occur almost entirely in the month the minimum wage is implemented. However, price *reductions* occur about 4 months in advance of a minimum wage hike. Therefore, any study of the effect of minimum wage hikes that focuses only on the months immediately surrounding the hike might miss important effects of the hike.

Large and small minimum wage hikes affect prices differently. Whereas large minimum wage hikes push prices up, small hikes are associated with lower prices. These findings are consistent with the existence of monopsonistic competition in restaurant labor markets and, so, provide indirect support for the claim that minimum wage hikes need not always reduce employment.

Among other things, this study shows the benefit of not treating all minimum wage hikes—for instance, large and small—as if they had the same effect. We might speculate that recent state and city laws that provide for regular, small hikes in the minimum wage have different effects—on prices, on employment, on capital-labor substitution, and the like—than federal laws that provide for infrequent, large hikes in the minimum wage. When it comes to the design of minimum wage laws, the details might matter.

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Tables and Figures

Table 1

City-, State-, and Federal-level Minimum Wage Changes Affecting Cities in Our Sample, 1977-2015

Political Unit Passing Minimum Wage Increase	Month/Year of Increase[22]
Federal (11 total, leading to 193 binding minimum wage increases in our dataset)	1/1978, 1/1979, 1/1980, 1/1981, 4/1990, 4/1991, 10/1996, 9/1997, 8/2007, 8/2008, 8/2009

[22] For a small number of minimum wage hikes, we considered as the month the hike was imposed the month following the actual month of the hike. Before 2004, the BLS generated CPI values from price data collected no later than the 18th of the month. But if a minimum wage hike was imposed on the 31st of the month, we would expect to see higher prices only in the data for the following month. After 2004, with the shift to computer-assisted data collection, the BLS stared collecting price data as late as the last day of the month, but the vast majority of price data for that month would have been collected before the imposition of a minimum wage hike on, say, the 31st. In this case, it is once again appropriate to shift, in our data, the imposition date of the minimum wage to the following month. See Bureau of Labor Statistics (2015), p. 16. We used a cutoff date of the 24th day of the month: any minimum wage change that occurred on or after that day was assumed to affect prices beginning the following month.

State (131 total binding minimum wage increases)	Alaska (1978, 1979, 1980, 1981) [23]
	Massachusetts (7/1986, 7/1987, 7/1988, 1/1996, 1/1997, 1/2000, 1/2001, 1/2007, 1/2008, 1/2015)
	New Hampshire (1/1987, 1/1988, 1/1989, 1/1990, 1/1991, 9/2007, 9/2008)
	Connecticut (10/1987, 10/1988, 1/1999, 1/2000, 1/2001, 1/2002, 1/2003, 1/2004, 1/2006, 1/2007, 1/2009, 1/2010, 1/2014, 1/2015)
	Maine (1/2002, 1/2003, 10/2004, 10/2005, 10/2006, 10/2007, 10/2008, 10/2009)
	Wisconsin (7/1989, 6/2005, 6/2006)
	Illinois (1/2004, 1/2005, 7/2007, 7/2008, 7/2009, 7/2010)
	Ohio[24] (1/2007, 1/2008, 1/2009, 1/2011 1/2012, 1/2013, 1/2014, 1/2015)
	West Virginia (7/2006, 7/2007, 7/2008, 1/2015)
	Maryland (1/2007, 1/2015)
	Michigan (10/2006, 7/2007, 7/2008, 9/2014)
	California (7/1988, 3/1997, 3/1998, 1/2001, 1/2002, 1/2007, 1/2008, 7/2014)
	Florida[25] (2/2005, 1/2006, 1/2007, 1/2008, 1/2009, 6/2011, 1/2012, 1/2013, 1/2014, 1/2015)
	New Jersey (4/1992, 10/2005, 10/2006, 1/2014, 1/2015)
	New York (1/2005, 1/2006, 1/2007, 1/2014, 1/2015)

[23] During these years, Alaska set its minimum wage at $0.50 higher than the federal minimum wage.

[24] Starting in 2007, Ohio indexed its minimum wage to the national CPI.

[25] Starting in 2005, Florida indexed its minimum wage to the South’s regional CPI.

	Pennsylvania (2/1989, 1/2007, 7/2007)
	Delaware (4/1996, 1/1997, 5/1999, 10/2000, 1/2007, 1/2008, 6/2014)
	Washington (1/1989, 1/1990, 1/1999, 1/2000, 1/2001, 1/2002, 1/2003, 1/2004, 1/2005, 1/2006, 1/2007, 1/2008, 1/2009, 1/2011, 1/2012, 1/2013, 1/2014, 1/2015)
City/County	Washington, D. C. (10/1993, 1/2005, 1/2006, 8/2008, 8/2009, 7/2014)
	San Francisco[26] (1/2004, 1/2005, 1/2006, 1/2007, 1/2008, 1/2009, 1/2010[27], 1/2011, 1/2012, 1/2013, 1/2014, 1/2015)
	San Jose (3/2013, 1/2014, 1/2015)
	Oakland (3/2015)
	Berkeley (10/2014)

[26] San Francisco indexes its minimum wage to the city's CPI.

[27] While the minimum wage did not increase in San Francisco this year, there was a change to labor costs due to the Health Care Security Ordinance (an employer spending mandate) that went into effect starting April 2008 (July 2008 for businesses with 20-49 employees), requiring employers to pay at an hourly rate per employee. For more information on the ordinance, see https://www.wageworks.com/media/179290/2903-SFHCSO-Compliance-Alert.pdf. The change in labor costs resulting from this act has been factored into all relevant years.

Table 2

Series with Sample Areas in Multiple States

Series for the FAFH Price Index	Sample Areas Used for Restaurant Weights
Boston	Massachusetts, New Hampshire, Maine (starting in 1998), Connecticut (starting in 1998)
Chicago-Gary-Kenosha	Illinois, Indiana, Wisconsin
Baltimore-Washington, D.C.	Washington, D.C., Maryland, Virginia, West Virginia
New York City-Northern New Jersey-Long Island	New York, New Jersey, Connecticut, Pennsylvania (starting in 1998)
Philadelphia-Wilmington-Atlantic City	Pennsylvania, New Jersey, Delaware (starting in 1998), Maryland (starting in 1998)

Notes: for the individual counties and towns covered each area, see the sources below. Restaurant establishment data (according to the individual county and town information) found using the County Business Patterns Census Database (http://censtats.census.gov/cgi-bin/cbpnaic/cbpsect.pl).

Sources: *Metropolitan Areas and Components, 1998* (published through the U.S. Census), http://www.census.gov/population/metro/files/lists/historical/93mfips.txt. 1993 edition: http://www.census.gov/population/metro/files/lists/historical/83mfips.txt.

Table 3

Minimum Wage Hikes by Series Periodicity

Periodicity	Observations	Minimum Wage Hikes			
		Federal	State	Local	Total
Monthly	1,852	40	42	0	82
Bimonthly	3,136	150	101	21	272
Both	4,988	190	143	21	354

Sources: *Monthly Labor Review*, state reports, and relevant city and county ordinances.

Table 4

Estimates of Pass-Through Using Monthly Data

Dependent Variable: FAFH Inflation

	(1)	(2)	(3)
Minimum Wage Change			
T-4	-0.004	-0.014	-0.014
	(0.006)	(0.008)	(0.008)
T-3	0.006	0.000	0.000
	(0.005)	(0.004)	(0.003)
T-2	0.012	0.003	0.001
	(0.012)	(0.009)	(0.008)
T-1	0.008	-0.002	-0.001
	(0.004)	(0.004)	(0.004)
T	0.052*	0.039*	0.039*
	(0.014)	(0.013)	(0.013)
T+1	0.022	0.008	0.008
	(0.012)	(0.014)	(0.014)
T+2	0.012	-0.002	-0.002
	(0.009)	(0.008)	(0.008)
T+3	0.012	-0.002	-0.004
	(0.007)	(0.005)	(0.005)
T+4	0.010	-0.002	-0.002
	(0.005)	(0.005)	(0.005)
[T-1,T+1]	0.081*	0.045	0.047
[T-3,T+3]	0.122**	0.044	0.042
[T-4,T+4]	0.127**	0.028	0.025
City CPI-All	---	---	0.113
			(0.166)
City fixed effects	Yes	Yes	Yes
Month, Year Controls	No	Yes	Yes
Observations	1,852	1,852	1,852
Cities	6	6	6
R^2	0.043	0.162	0.170
Adj. R^2	0.036	0.133	0.141

Notes: * $p<0.05$, ** $p<0.01$. Regressions use monthly data from Los Angeles, Chicago, and New York City between 1978 and 2015, as well as San Francisco (1987-1997), Detroit (through 1986), and Philadelphia through 1997. The T-4 coefficient indicates the partial effect of the minimum wage change on FAFH inflation 4 months prior to the date of the minimum wage change. Cluster-robust standard errors reported in parentheses. Significant levels based on T(G-1).

Table 5

Estimate of Pass-Through, Full Dataset

Dependent variable: FAFH inflation

	(4)		(5)	
Minimum Wage Change				
T-4	-0.010	**	-0.009	**
	(0.003)		(0.003)	
T-3	-0.005	*	-0.006	*
	(0.002)		(0.002)	
T-2	0.000		-0.002	
	(0.003)		(0.003)	
T-1	0.010	**	0.010	**
	(0.003)		(0.003)	
T	0.022	**	0.023	**
	(0.004)		(0.004)	
T+1	0.013	*	0.013	*
	(0.005)		(0.005)	
T+2	0.001		0.001	
	(0.003)		(0.002)	
T+3	0.004		0.003	
	(0.003)		(0.003)	
T+4	0.002		0.002	
	(0.002)		(0.002)	
[T-1, T+1]	0.045	**	0.046	**
[T-3, T+3]	0.043	**	0.042	**
[T-4, T+4]	0.035	**	0.035	**
City CPI-All	---		0.130	**
			(0.029)	
City fixed effects	Yes		Yes	
Month, Year Controls	Yes		Yes	
Observations	8,124		8,124	
Metropolitan Areas	28		28	
R^2	0.170		0.180	
Adj. R^2	0.161		0.171	

Notes: * p<0.05, ** p<0.01. Regressions 4 and 5 use the full dataset (i.e., pooled monthly data with the bimonthly, interpolated, data).

The T-4 coefficient indicates the partial effect of the minimum wage change on FAFH inflation 4 months prior to the date of the minimum wage change. Standard errors corrected for arbitrary forms of heteroscedasticity are reported in parentheses. Significance levels based on T(G-1).

Figure 1

Impact of Minimum Wage Increase in Monopsonistic Competition

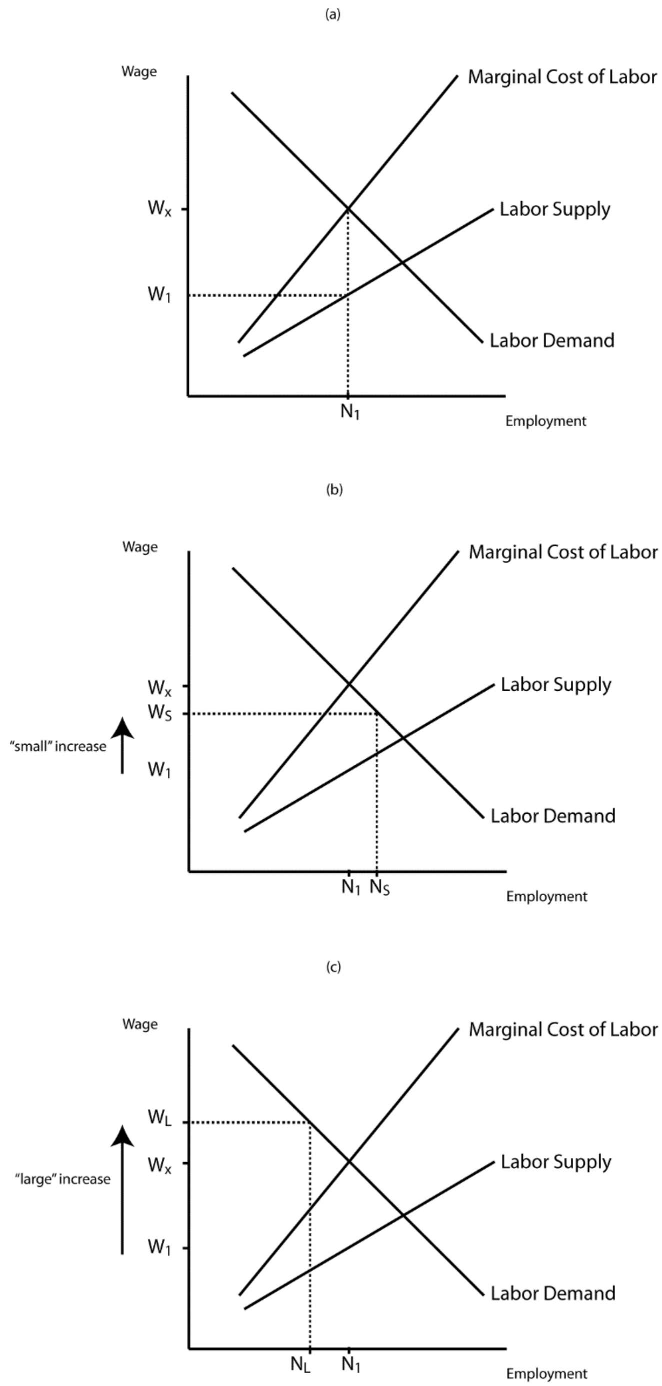

Table 6

Estimate of Pass-Through, Full Dataset

Dependent Variable: FAFH Inflation

	(6)	
Minimum Wage Change	Small	Large
T-4	-(0.036)*	-(0.007)**
	(0.015)	(0.003)
T-3	-0.011	-0.006*
	(0.013)	(0.002)
T-2	-0.002	-0.003
	(0.011)	(0.003)
T-1	-0.011	0.011**
	(0.010)	(0.003)
T	0.013	0.023**
	(0.015)	(0.005)
T+1	-0.005	0.014**
	(0.011)	(0.005)
T+2	-0.002	0.001
	(0.012)	(0.003)
T+3	-0.015	0.005
	(0.011)	(0.003)
T+4	-0.001	0.002
	(0.012)	(0.002)
[T-1, T+1]	-0.005	0.046**
[T-3, T+3]	-0.038	0.045**
[T-4, T+4]	-0.073	0.039**
City CPI-All	0.132**	
	(0.030)	
City fixed effects	Yes	
Month, Year Controls	Yes	
Observations	8,124	
Metropolitan Areas	28	
R^2	0.182	
Adj. R^2	0.172	

Notes: * p<0.05, ** p<0.01. Regression 6 uses the full dataset (i.e. the monthly data pooled with the bimonthly interpolated data).

The T-4 coefficient indicates the partial effect of the minimum wage change on FAFH inflation 4 months prior to the date of the minimum wage change. Standard errors corrected for arbitrary forms of heteroscedasticity are reported in parentheses. Significance levels based on T(G-1).