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The Effects of Increasing the Minimum Wage on Prices: Analyzing the Incidence of Policy Design and Context

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Abstract

We analyze the price pass-through effect of the minimum wage and use the results to provide insight into the competitive structure of low-wage labor markets. Using monthly price series, we find that the pass-through effect is entirely concentrated on the month that the minimum wage change goes into effect, and is much smaller than what the canonical literature has found. We then discuss why our results differ from that literature, noting the impact of series interpolation in generating most of the previous results. We then use the variation in the size of the minimum wage change to evaluate the competitive nature of low-wage labor markets. Finally, we exploit the rich variation in minimum wage policy of the last 10 to 15 years—including the rise of state- and city-level minimum wage changes and the increased use of indexation—to investigate how the extent of price pass-through varies by policy context. This paper contributes to the literature by clarifying our understanding of the dynamics and magnitude of the pass-through effect and enriching the discussion of how different policies may impact that effect.

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1. Introduction

In recent years, in the face of federal inaction to raise the minimum wage, states and cities have increasingly passed their own minimum wage laws. These state and city laws promoted a renaissance in the study of the employment effect of minimum wage hikes as they not only created greater numbers of minimum wage changes to be studied using then-standard techniques but also created "natural experiments" that permitted alternative techniques to be used to identify the employment impact of minimum wages. Two branches of minimum wage research developed starting in the 1990s. One branch found that, contrary to the previously accepted belief, some minimum wage hikes led to either no decline in employment but potentially increased employment (e.g., Card and Krueger 1994, 1995; Dube, Lester, and Reich 2010). A second branch found evidence supporting the claim that minimum wage hikes reduced employment (e.g., Neumark 2001; Neumark and Wascher 2002, 2007, 2008).[1]

An additional important, although less-studied, question about the impact of minimum wage hikes is the impact such hikes have on output prices, the so-called "pass-through" effect. Early studies include Wessels (1980) and Card and Krueger (1995). The most influential of these studies, however, have been a series of papers by Aaronson and coauthors. Aaronson (2001), Aaronson and French (2006), Aaronson and French (2007), and Aaronson, French, and MacDonald (2008) find evidence that minimum wage hikes increase output prices and that the size of this pass-through suggests that the increased cost associated with a minimum wage hike is completely passed along to consumers.[2] Aaronson and coauthors also developed an argument that their findings support the claim that low-wage labor markets are highly competitive and, by implication, that minimum wage hikes necessarily lower employment. This literature on pass-though, then, is important both in itself and because it sheds indirect light on the on-going debate over the employment effect of minimum wage hikes.

This paper contributes to the literature on price pass-through by presenting more accurate estimates of the pass-through effect than found in the previous literature, and by using these results to give insight into the competitive structure of low-wage labor markets. We find that the extent of pass-through is much smaller than previously reported and that the behavior of this pass-through is more consistent with monopsonistic competition than it is with perfect competition. Additionally, we exploit the rich variation in minimum wage policy—the rise of state- and city-level minimum wages, as well as the increased use of indexation of the minimum

[1] Explanations for small negative or positive employment effects included the existence of various market frictions arising from imperfect competition or search (e.g. Bhaskar and To 1999; Lang and Khan 1998).
[2] The studies citied above are for the US. Lemos (2008) provides a survey of the literature.

wage to the CPI in areas such as Florida, Washington, Ohio, and San Francisco—to investigate how the extent of pass-through varies by policy context.

2. Literature Review and Contribution to the Literature

Previous empirical studies have documented that minimum wage hikes produce substantial price pass-through effects. The oft-cited study by Aaronson (2001) estimated the magnitude of the pass-through using metropolitan-area food away from home (FAFH) CPI data between 1978 and 1995. In the base specification (pg. 162 of his article) which included only monthly and yearly controls, the cumulative wage-price elasticity from 3 months before up to 3 months after a minimum wage hike was estimated at about 0.07, meaning that a 10% increase in the minimum wage is associated with a 0.7% increase in FAFH prices.

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, again, about 0.07. Behind this average price increase was substantial variation: prices for some restaurant items grew faster than this average, while prices for other items grew slower than the average, and some prices even fell after a minimum wage hike. The price increase was also higher in limited-service restaurants than it was in full-service restaurants. Though the empirical literature is somewhat limited outside of these two formative works (see Lemos 2008 for a review), other studies have found similar results in other countries and other cases.[3] The magnitude of the pass-through has been presented as being consistent with what models of a perfectly competitive labor market would predict about the size of the pass-through, assuming standard values for demand elasticities of fast-food and capital-labor elasticities (Aaronson and French 2007; Aaronson, French, and MacDonald 2008).[4] An increase in the minimum wage increases the marginal cost of labor, reduces employment, lowers output, and raises prices. This work on the pass-through therefore speaks to the on-going controversy about the competitive structure of low-wage labor markets.

[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.

[4] Although the overall thrust of the existing empirical literature on minimum wage hike pass-though is to support the claim that labor markets for restaurants are best characterized by competition, the evidence is not unambiguous. For instance, Aaronson and French (2007, p.696) write after their analysis of BLS micro price data for restaurants, "Given that some restaurants do not increase their prices after minimum wage hikes, but restaurants that do raise their prices usually do by more than 0.7 percent, it is difficult to compare the observed price response to the competitive prediction."

Policy and academic work has frequently cited the above studies by Aaronson and co-authors as the authoritative studies on minimum wages and pass-through.[5] However, these studies deserve to be updated for a couple of reasons.

First, these studies rely on data from no later than 1997 but since then we have seen an increase in the variation of minimum wage policy across several dimensions.[6] For instance, since 1997 we have seen a profusion of state and city minimum wage laws whose effect we cannot assume are identical to federal minimum wage hikes. Further, some states and cities have implemented laws that provide for scheduled increases in their minimum wage often indexed to some measure of price inflation. In this way, these new policies differ from the majority of minimum wages investigated by Aaronson and co-authors which were often large, one-shot increases implemented with relatively little warning to businesses. Again, we cannot presume these new types of minimum wage hikes affect prices, or more generally the economy, in the same way minimum wage changes implemented before 1997 did. Indeed, one contribution of our study is to present a comparative analysis of different types of minimum wage policies within a common data and econometric setting.

Table 1 details the differences between the minimum wages considered by Aaronson and co-authors with those we consider in this study. The table shows that state-level minimum wage increases are much more common—and federal-level increases much less common—after 1998. Other variations in policy such as indexed, city minimum wages, or perpetually scheduled minimum wage increases were absent or nearly absent from the period considered by the previous studies.

[Table 1]

Second, we use the data differently than how Aaronson (2001) used it to give additional insight into the process of pass-through. For instance, we treat monthly and bimonthly price series separately (instead of combining them as did Aaronson (2001)) to better reveal the dynamics of pass-through pricing. Furthermore, by embracing the complicating factor of multiple-state metropolitan areas (instead of avoiding it as did Aaronson (2001)) we more accurately measure the impact of different types of minimum wage increases, and are thereby able to shed additional light on the competitive nature of low-wage labor markets. Finally, by using data after 1997 we are able to use CPI data that are less affected by various biases (such as

[5] Most of the later pass-through literature cites this paper as the canonical example, as well as much of the rest of the literature on the effects of the minimum wage such as Dube et al. (2010) and MaCurdy (2015).

[6] The use of data from this period continues up to present studies, as seen in MaCurdy (2015) , who uses data from 1996, and from a single federal minimum wage increase, to draw conclusions about all minimum wages.

substitution bias) that was not available to Aaronson (2001); this will again permit us to generate more accurate estimates of the extent of pass-through.

Looking ahead to the results, our first main finding is that pass-through is primarily concentrated on the month that the minimum wage hike goes into effect, with no appreciable impact on the month before or after. This finding contradicts the previous work. Second, we estimate wage-price elasticities are notably lower than reported in previous work: we find prices grow by 0.36% for every 10% increase in the minimum wage, which is almost half of the previously accepted 0.7%.[7] Third, we find the behavior of pass-though is consistent with market power on the demand-side of low-skilled labor markets (e.g. monopsony or monopsonistic competition), which sheds light on one of the more contentious issues in the debates over the employment impact of minimum wage hikes. We also find that not all minimum wage hikes are the same. For instance, small, scheduled minimum wage hikes have smaller impacts on prices than large, one-time minimum wage hikes. Yet, we find no significant differences between state- and federal-level minimum wage increases, even though we might expect business flight to have a larger impact in the case of state-level minimum wage changes.

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. FAFH includes food purchased and consumed outside of the home, and for the most part includes items sold at full- and limited-service restaurants.[8] This data is available at the BLS website. We include in our analysis all metropolitan areas that have either monthly or bimonthly FAFH data for at least part of the period of our study, 1977-2015, which gives us 28 series.[9]

[7] This 0.036 elasticity is similar to what was found by Card and Krueger (1995, p. 54) in their study of a single minimum wage increase in New Jersey.

[8] Additionally, FAFH 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. See BLS, *Handbook on Methods*, Chapter 17. For conciseness, we will refer in the text to "restaurants" when we talk about the group of sites considered as selling food away from home.

[9] 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

We begin our analysis in 1978 because that is the year that Aaronson (2001) started his analysis. The minimum wage increase in 1978 was also the first one after the implementation of 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) along with the expansion of the minimum wage to all covered, non-exempt employees. Thus, 1978 was the first year in which minimum wage changes would affect all minimum wage workers regardless of occupational status or industry, giving our estimates more consistency than if we relied on earlier data where different minimum wages affected different subsets of workers.[10]

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 method for calculating the CPI before 1999 produced an upward bias to the CPI and its subcomponents. The new geometric mean formula could mimic consumers' substitution between the products they buy in response to changes in relative prices, something the previously used Laspeyres formula did not do.[11] If the CPI was biased upwards before 1999, then any study of the size of the pass-through that uses pre-1999 CPI data, such as Aaronson (2001), presents estimates of the pass-through that are potentially biased upwards. Our study, which uses data for 1978-2015, is able to use the more accurate geometric mean-based CPI for the second half of the period and, so, is able to generate more accurate estimates of pass-through.

The main independent variable of interest in our regression is the change in (binding) minimum wage rates. Our data on minimum wages comes from various issues of the *Monthly Labor Review*, state Department of Labor reports, and, for San Francisco, San Jose, Oakland, Berkeley, Washington, D.C., and Prince George's and Montgomery counties, city and county ordinances. As indicated in Table 2 (below), the years 1978 to 2015 saw 11 federal minimum wage increases, 126 binding state minimum wage increases, and 23 city minimum wage increases. Table 2 reports the month and year of passage for all of these increases.

[Table 2]

(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), St. Louis (bimonthly until 1997).

[10] See, for instance, http://www.dol.gov/whd/minwage/coverage.htm.

[11] Dalton, Greenlees, and Stewart (1998) provides an overview of this change.

We also include, in most of our regressions, control variables such as month, year, and a metropolitan area fixed-effects. One additional control is "CPI-All" (Urban Consumers),[12] included to take into account various unknown determinants of FAFH CPI inflation. The inclusion of 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. As will be seen, however, this does not seem to be a problem, as when CPI-All is included in our regressions it has virtually no effect on our main coefficients of interest.

The BLS generates FAFH CPI for multistate metropolitan areas by using prices from restaurants located in more than one state. For example, in the case of the New York-Northern New Jersey-Long Island metropolitan area, the FAFH CPI is constructed from prices taken from a sample of restaurants located in 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 3 (below) provides information about the metropolitan areas in our sample that include territory from more than a single state.

[Table 3]

The existence of multistate metropolitan areas provides a benefit to this study. We are able to 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. We need a way to transform a single-state minimum wage increase affecting only restaurants in one portion of in a multistate metropolitan area into a variable measuring its impact on average FAFH prices in the full metropolitan area. We will tentatively presume that a 10% state minimum wage hike that affects only 20% of the restaurants in a metropolitan area (that is, those restaurants in that state) will be equal to a 2% (10% x 20%) minimum wage hike for the whole metropolitan area. We will define the "restaurant-weighted state minimum wage change" (RSMW) as,

$$\Delta_t \log(mw_{it}^*) = \sum_s \lambda_{ist} * \Delta_t \log(mw_{ist}) \quad (1)$$

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 mw_{st} is the minimum wage change in state s in time t.[13]

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[12] Published by the BLS and available on their website.
[13] For example, consider the District of Columbia in 2009. That series is composed partly of counties in 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

When a metropolitan area includes only a single state, λ_{ist} will equal 1 and the RSMW for any minimum wage will simply be the change in the associated state minimum wage. The number of restaurant establishments in the various state subsections of multistate metropolitan areas comes from County Business Patterns while information about the particular towns and cities included in each state subsection of a metropolitan area comes from the definitions of these metropolitan areas provided by the Office of Management and Budget.[14]

An additional noteworthy characteristic of our data is that some of the price series are available monthly while other price series are only available bimonthly. (The same holds true for the data used on Aaronson (2001) and related studies.) Table 4 (below) breaks down the total number of binding minimum wage hikes in our sample by whether the affected price series reports monthly or bimonthly observations.

[Table 4]

As can be seen, the monthly price series has connected with them a range of federal and state minimum wage increases, but the number of monthly observations is much less than the number of observations we have for the bimonthly data. Good reason exists, then, to use the information included in the bimonthly data in this study as it permits us to take into account a far wider range of minimum wage increases. Yet, the bimonthly data is not granular enough to permit a consideration of details about the dynamic (here, monthly) impact of the pricing process set in motion by a minimum wage hike.

In summary, we estimate price pass-through due to the minimum wage by using the food away from home price index for 28 cities between 1978 and 2015. We also use each city's CPI-All as a control variable. Since some cities are in fact composed of multiple states, we are able to incorporate additional minimum wage changes into our analysis. We apply a weighting scheme to our minimum wage change variable that draws on County Business Pattern data on the number of restaurant establishments in each city's sample area. We will use both monthly data

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minimum wage in order to construct the District of Columbia minimum wage variable: D.C. (0.164), Maryland (0.344), Virginia (0.471), 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). We tentatively propose, in this case, that a 10% increase in the minimum wage in Maryland would have the same impact on prices in the wider District of Columbia metropolitan area as would a 3.44% increase in the federal minimum wage. The equality of these two impacts is, of course, debatable and we address it below.

[14] The BLS's *Handbook on Methods*, Chapter 17, describes in general terms the way that they select outlets to use as their source of prices. The BLS attempts to select these 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 regional differences in average consumer restaurant bills might lead the distribution of restaurant purchases to vary from the regional distribution of restaurant establishments. We also used population weights in place of restaurant establishment weights, but the results we got from using population weights did not different much from what is reported in the text.

and bimonthly data in our study. In the following section, we discuss our empirical model and present preliminary results using monthly data on food away from home prices.

4. Estimates of Pass-Through with Monthly Data

Our two initial tasks are to: (1) estimate the extent of pass-through and (2) discover *when* this pass-through occurs (i.e., either only contemporaneously with the imposition of the minimum wage hike or also in the months before and/or after the hike is imposed). We can accomplish both these tasks simultaneously if we limit ourselves to monthly price series only. The bimonthly price series are not granular enough to reveal the detailed monthly dynamics of the pass-through process and so we temporarily set them aside. The downside of this approach is that we are only able to consider the impact of 82 of the 354 minimum wage hikes appearing in our full sample (see Table 4) and limit ourselves to using less than half the total data we have available.

The subsample used in this section comes from the three metropolitan areas (New York, Chicago, and Los Angeles) that have monthly data for the entire period and from three additional metropolitan areas (San Francisco, Philadelphia, and Detroit) that have monthly data for some subset of the period 1978-2015. Monthly observations were reported for San Francisco between 1986 and 1998, for Philadelphia before 1998, and for Detroit before 1987. We do not use the bimonthly data from these metropolitan areas from outside these years. Together, these metropolitan areas account for only about 20% of all federal-level minimum wage increases and about 30% of all state-level minimum wage increases in our sample.

We estimate the equation below, which has Food Away from Home (FAFH) inflation as the dependent variable and, as independent variables, the weighted log difference in the minimum wage $mw*$ (defined in Equation 1), overall metropolitan area CPI inflation, along with metropolitan area, month, and year fixed effects as independent variables:

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

This regression includes leads and lags of four months as we want to capture the impact of a minimum wage hike on prices in the months both preceding and following the month on which a minimum wage hike is implemented.

Table 5 reports our findings. As we go from regression 1 to regression 3, we add month and year dummies along with the metropolitan area's overall CPI as controls. Regression 3, which adds City CPI-All and shows that it is significant, is used as the basis for the discussion below.

[Table 5]

In regression 3 the contemporary elasticity is 0.039, a value that is statistically significant at the 99% confidence level. We also get a statistically significant negative coefficient four months before the minimum wage is imposed, but no other coefficients achieve statistical significance in either regression 2 or 3.[15] According to the monthly data, then, a minimum wage hike leads to a price increase *only* in the month it is imposed. In that month, a 10% increase in the minimum wage is associated with a 0.39% increase in the FAFH CPI. We also find that prices also grow slower four months ahead of a minimum wage hike, as indicated by the statistically significant (p-value of 0.015) coefficient of -0.014 for T-4. When we take into consideration the net effect on prices over the 9-month period centered on the minimum wage hike, we find a 10% increase in the minimum wage leads to a net increase in FAFH CPI of 0.25%.

These findings are different from what Aaronson (2001) reported. For instance, he reports statistically significant price increases in the month before and the month after a minimum wage hike is imposed whereas we find no such effect in those months. Aaronson also reports a much larger pass-through than we do: he finds that in the 9 months surrounding a minimum wage hike a 10% increase boosts prices by 0.67%.[16] Our finding of 0.25% is less than half of what Aaronson found. We will defer further comment on these differences until we discover what our full sample (including both monthly and bimonthly data) says about these differences.

We have one interesting finding in common with Aaronson (2001): we both find a statistically significant negative coefficient four months in advance of a minimum wage hike. The elasticities we find are nearly identical, -0.014 for us and -0.013 for Aaronson.[17] That prices grow slower in advance of a minimum wage is hard to square with a perfectly competitive setting, in which businesses only respond to actual changes in costs. Further, that an anticipated increase in future costs might lead to a moderating of price increases ahead of this increase is quite interesting and we can only speculate about the mechanism behind this behavior. If this finding—of slower growth in prices in advance of a minimum wage increase—is confirmed by regressions using our full sample, one implication might be that studies of the impact of the minimum wage (either on prices or even on employment) that limit their focus to a couple of

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[15] The finding that only a single lead or lag in regressions 2 or 3 achieves statistical significance is evidence against the potential claim of endogeneity—i.e., that minimum wage policy is partly a response to inflation. Because the dependent variable is the percentage change in FAFH prices, a potential endogeneity problem reflects the idea that minimum wage hikes occur during periods of *escalating* inflation. The fact that the majority of coefficients for the leads and lags are not statistically significant from zero indicates that this sort of endogeneity is not an issue in our regressions.

[16] Table 4, regression 2 in Aaronson (2001)

[17] Aaronson (2001), Table 4, Regression 2. Aaronson has little to say about this statistically significant coefficient.

months before and after the minimum wage hike might be missing part of the response they are trying to measure.

5. Using Interpolated Data

Following Aaronson (2001), we will join our monthly and bimonthly series to create a larger single dataset. By joining these two types of data, we expand the number of minimum wages we account for from 82 to 354. The first step in joining these two types of data is transforming, through a process of interpolation, the underlying bimonthly data into monthly series before that data is logged and then joined with logged values of the (actual) monthly series. Performing this joining increases the number of observations we have from 1852 to 8124.[18]

In much of the econometric literature, interpolation involves transforming quarterly data into monthly data or transforming annual data into quarterly data. Further, the interpolation often involves using related higher frequency data to inform the 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. Therefore, we interpolate by simply averaging the neighboring bimonthly data and, where appropriate, splicing information about the minimum wage hikes that occurred (contemporaneously, with leads, or with lags) onto the interpolated monthly series.

Any interpolation process creates something akin to measurement error in the resulting interpolated data points. In our case, by interpolated values for some metropolitan areas for FAFH CPI and City CPI-All, we must treat the dependent variable and one independent variable as if they were measured with error. This raises the possibility that both the coefficients and standard errors produced by regressions using this data are biased. The precise nature of these

[18] The 8124 observations include 1852 monthly observations, 3136 bimonthly observations, and 3136 interpolated observations. However, the degrees of freedom used to calculate standard error in regressions using all these observations will be less than the number of observations. In general, the degrees of freedom is equal to the number of independent pieces of information that goes into the estimation of a parameter. Some of our interpolated data is not independent as it has been generated from a linear combination of the bimonthly data on either side of it and, so, such interpolated data does not add independent information. However, *some* of our interpolated data might be seen as adding new information. For instance, when we generate a monthly observation for January by interpolating bimonthly FAFH data for December and February in some cases we add to this observation new information: that a minimum wage hike occurred in January. Arguably, the latter type of interpolated data does add some new information, and so it might be seen to add an additional degree of freedom to our regression procedures. Yet, this new information is embedded in some not-new information (the interpolated part). We take the conservative approach by assuming that none of the interpolated data contributes degrees of freedom to our estimates of standard errors. So, for instance, if a regression uses the largest dataset (8124 observations) we will use 4988 (=1852+3136) as the starting point for our determination of the degrees of freedom for the standard errors for the coefficients for these regressions.

biases will depend, of course, on the nature of the measurement error and the particular estimation technique used. We will consider each in turn.

Interpolation will likely generate "pseudo-measurement" errors for FAFH CPI that are positive both for the month preceding a minimum wage hike (T-1) and for the month following such hikes (T+1). Interpolation will also likely generate pseudo-measurement errors that are negative for the month of a minimum wage hike. The argument that the pseudo-measurement errors have these signs (on average) is simple. First, we assume that the impact of minimum wages on prices in a metropolitan area is unrelated to whether the BLS collects monthly or bimonthly FAFH CPI data for that metropolitan area. If that is the case, we can use the results of our monthly regressions above to say that in metropolitan areas that collect bimonthly data minimum wage hikes lead to increases in prices on the month of the hike but not in the month before or after.

The upper-half of Figure 1 portrays a stylized pattern of FAFH prices when a minimum wage hike is imposed. In this figure, we presume prices growth smoothly except for in the month of the minimum wage hike (on month 0). Now assume that only bimonthly data was collected for this metropolitan area (on month -2, month 0, and month +2, etc). In the lower-half of Figure 1 the points a, c, and d are the actual data we have. If we linearly interpolate between a and c (indicated by the plus sign) we can see our interpolated value for -1 to exceeds the actual data point b. As a result of this, the growth rate in FAFH prices from -2 to -1 will be larger than it really is while that from -1 to 0 will be smaller than it really is. If, on the other hand, we have bimonthly data for months -1 and +1 then the interpolated data point for month 0 will be lower than it really is, and as a result the growth rate of FAFH prices from -1 to 0 will be lower than it really is and from 0 to +1 the growth rate of prices will be higher than it really is. If we have a mix of the two types of bimonthly data, and generate a monthly series for the growth of FAFH prices then this will tend to create, in regressions that use this interpolated data, upward biases for the coefficients for T-1 and T+1 and a downward bias for T=0. The conclusion is that interpolation shifts the apparent price increases away from the month in which it was imposed onto both the month before and the month after. The same shifting, for the same reason, will occur from T-4 to T-3 because of the positive coefficient for T-4 in the monthly regressions above.[19]

[19] Pseudo-measurement errors might also be correlated with our monthly dummies because of predictable seasonal movements of prices. If prices typically grow rapidly in, say, April and we interpolate between February and April CPI data points then the interpolated value for March will tend to be greater than it really is as will the resulting value for the grow rate of prices in March. Similarly, the growth rate of prices between March and April, using the interpolated data, will be downward biased. If this seasonal issues does occur our monthly coefficients might be

[Figure 1]

We now turn to the second issue: the impact of the interaction between the particular data we use in this study and the particular estimation technique we use. We gain insight into the consequences of interpolating the bimonthly data by, again, making use of our monthly data. We note again that how restaurants respond to minimum wage hikes should not depend on whether the BLS generates monthly or bimonthly FAFH CPI series for their metropolitan area. This suggests the following experiment: for the metropolitan areas that do have monthly data, we can simulate what the data would have been if it was actually have been collected bimonthly and then use this data to run our regressions. We can then compare the regression results generated from this simulated bimonthly data with the results produced by the true monthly data. The differences we discover in this experiment using fabricated bimonthly data should be transferable to metropolitan areas for which we have only bimonthly data.

We return, then, to the six series for which we have full monthly data, deleting half of each city's FAFH and CPI-All observations, and then linearly interpolating each series to create observations to replace those we deleted. 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. We then logged and first-differenced each of the fabricated bimonthly (with interpolation) series to obtain our measure of inflation, and estimated a regression model based on Equation 2.

[Table 6]

Regression 4 in Table 6 reports the result of using the fabricated bimonthly (with interpolation) data. As predicted above, interpolation spreads out the contemporaneous impact of the minimum wage hike to the month preceding and the month following the hike. As we move from regression 3 (from Table 5) to regression 4, the contemporaneous impact falls from 0.039 to 0.021 while the coefficients for T-1 and T+1 rise (and achieve significance or near-significance). The sum of the coefficients for T-1 to T+1 is identical in regressions 3 and 4. Once we get to the sum of T-4 to T+4, that for regression 4 does exceed that for regression 3 but this increase is due mostly to what happened for T+4. In most, but not all, cases the standard errors fell, as expected, but the magnitude of these changes were not large enough to (alone) cause estimated coefficients to achieve significance.[20]

systematically biased. But this additional factor does not affect the estimated coefficients for the variables of interest to us in this study and, so, we ignore it here.

[20] The reason why not all standard errors fall is because we use Huber-White robust standard errors.

Generalizing, interpolation in the context of this study tends to reduce the estimated contemporaneous price increase, shifts some of the contemporaneous impact to the month before and after the minimum wage hike, and should be assumed to reduce standard errors. Still, a regression using some interpolated data does provide useful information about the total effect of minimum wage hikes on the FAFH CPI.

Although we cannot say for sure what caused Aaronson (2001) to find statistically significant increases in prices in month before and after minimum wage hikes, the above discussion about the impact of interpolation suggests that Aaronson's results were at least partly (and maybe fully) due to his use of interpolation.

For comparison, regression 5 in Table 6 presents the results using data coming only from those metropolitan areas for which the BLS generates bimonthly price data. No monthly data was used. The regressions were generating from series using bimonthly (with interpolation) data. For some cities, the BLS releases their FAFH price index on a January/March/May/… cycle, while others follow the alternate cycle of December/February/April/…. In order to estimate elasticities using these series, we linearly interpolated the original FAFH price index as well as the city CPI-All. This new series, now made up of a combination of the actual bimonthly data and data interpolated between the bimonthly data, was logged and first-differenced to construct the measure of FAFH inflation that serves as our dependent variable.

The results seen in regression 5 are very similar to those seen in regression 4, but with greater significance on certain coefficients possibly due to the higher number of observations used to estimate regression 5. One difference seen is that the slowdown in the price increase (ahead of the minimum wage hike) shifted forward one month to T-3. The various sums of coefficients are very similar to those found in regressions 3 and 4.

The results of regression 5 is exactly what one would expect if the true underlying monthly data (if it existed) was just like that which generated the results in regression 3. When properly interpreted, the results of regressions using interpolated data give insight into the impact of minimum wage hikes on prices. We turn next to combining monthly and bimonthly (with interpolation) data to consider the impact of minimum wage hikes along with other issues relevant to policy design.

6. Main Results: How Do Prices Respond? Are the Results Consistent with Perfectly Competitive Low-Skilled Labor Markets?

We now pool together monthly and bimonthly (interpolated) data for the 1978-2015 period. Table 7 presents the results. We focus on the results of regression 7 which includes City CPI-All as a control.

[Table 7]

According to regression 7, a 10% increase in the minimum wage boosts prices by 0.45% in the three months centered on the month the hike is imposed. However, based on the discussion in the previous section, we can say that regression 7 likely overstates the size of the price increases on the month before and after the minimum wage hike is imposed and understates the size of the price increase on the month the hike is actually imposed, though the sum of these coefficients likely does indicate the fully impact of these three months. The sum of the coefficients [T-1, T+1] in this regression, 0.045, is almost identical to that found in regression 3 (which used only monthly data).[21]

As before, we also find minimum wage hikes lead restaurants to moderate their price increases 3 to 4 months ahead of the hike. In regression 7, the coefficients for T-3 and T-4 are both negative and statistically significant. A portion of the price decline assigned to T-3 in this regression is likely due to a shifting of price increases occurring in T-4 by the process of interpolation. The sum of the coefficients for these two months is 0.015, which is identical the sum of coefficients of the same two months in regressions 3 and 5.

The *total* effect of minimum wage hikes in the 9 months centered on the month the hike is imposed is 0.036, a number close to that seen in regression 5 but somewhat larger than seen in regression 3. So, considering the full period over which a minimum wage affects prices, we find that a 10% increase in the minimum wage leads to a 0.36% net increase in prices. That is, if a $10.00 item experienced this average price increase, it would become a $10.04 item.

The size of the price increase (and so the implied welfare loss) we find is lower than previously reported: Aaronson (2001) reports a 10% increase in the minimum wage causes a net 0.67% increase in the nine months centered on the month the minimum wage hike is imposed.[22] We find a price increase for the same period close to half of that reported by Aaronson (0.36% vs

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[21] Although the interpolation process generates standard errors that are biased downwards (as discussed above), the p-values for most of these coefficients in regression 7 are so small that it is hard to believe that the reported statistical significance was due simply to interpolation.
[22] Table 4, regression 2.

0.67%), and so our findings suggest a lower welfare loss to consumers following a minimum wage hike.

The importance of our findings go beyond finding a reduce welfare impact on consumers when a minimum wage hike is imposed. Building on a set of reasonable assumptions about the operation of restaurants in a hypothetical perfectly competitive market, Aaronson and French (2007) 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. As they find, in various regressions, elasticities near 0.07 they conclude that low-wage restaurant labor markets are best characterized as perfectly competitive. The implication of this, in turn, is that any minimum wage increase will necessarily reduce employment.

However, we get results *inconsistent* with highly competitive low-wage labor markets in the restaurant industry: our elasticity of 0.036 for the nine months centered on the month of a minimum wage hike and of 0.043 for the much narrower period of [T-1,T+1] fall short of the 0.07 hypothesized as being consistent with perfect competition. However, our findings do not provide positive support for any particular alternative structure of low-wage labor markets. In the next section we consider whether the data we have provides positive support for one alternative labor market structure, monopsonistic competition.

7. Monopsonistic Competition in Low-Wage Labor Markets: Theory and Evidence

Monopsonistic competition has been offered in recent years as an alternative model for *some* labor markets.[23] Most notably, Card and Krueger (1995) proposed that monopsony-like conditions in low-skilled labor markets might explain their finding that minimum wages increased employment. Since then, Burdett and Mortensen (1998), Bhaskar and To (1999), Bhaskar, Manning, and To (2002), and others have proposed different causes for imperfect competition on the buyer-side of labor markets, and developed formal models that drew out the potential consequences of monopsonistic competition. All of these formal models of monopsonistic competition, however, are consistent with Stigler's (1946) observation of the impact of a minimum wage when businesses have market power in labor markets: the impact of a minimum wage on employment (and so on output prices) is context dependent. More narrowly, Stigler pointed out that when employers had power over wages, a small rise in a minimum wage generates increased employment (and, implied by this, increased output and reduced prices)

[23] Few argue that pure monopsony in labor markets has been found outside of a few unusual labor markets (for instance, in the market for professional baseball players in the United States before the ending of the reserve clause). Many economists, however, persist in using the term "monopsony" as shorthand for monopsonistic competition.

while a large increase in the minimum wage reduces employment (and, implied, reduces output and higher prices).

This can be seen in the standard model of monopsony in the labor market. The monopsonist has market power and, therefore, faces an upward-sloping labor supply curve. To attract more workers, the monopsonist needs to increase the wage, which necessitates increasing the wages of those already hired. This implies the marginal cost of labor for the monopsonist is greater than the wage and so, the marginal cost of labor curve is upward-sloping and rises faster than the labor supply curve.

In Figure 2(a) (below) the equilibrium wage for the monopsonist, in the absence of a minimum wage, is at W_m while employment stands at L_m. This equilibrium wage is below what it would have been in a perfectly competitive setting,W_{pc}.

[Figure 2]

Figure 2(b) shows the impact of a "small" minimum wage increase. Suppose, just for the sake of convenience, that initially the minimum wage stood at W_m. Suppose, next, that a new minimum wage is implemented and the size of the increase is "small." The new minimum wage is established at W_{smw}, which stands above W_m but below W_x, where labor supply equals labor demand. The marginal cost of labor now includes the horizontal solid line starting at W_{smw}. The new marginal cost curve will induce the monopsonist to expand employment up to L_{smw} as each worker below that level of employment will now have a marginal cost below his/her value of marginal product (given by the labor demand curve). As drawn, the small increase in the minimum wage will increase employment as $L_{smw} > L_m$. In turn, this increased employment will (given plausible assumptions) lead to higher output (at least in the short-run) and, so, lower prices.

Figure 2(c) shows the impact of a "large" increase in the minimum wage. With a "large" increase, the minimum wage pushes the wage from W_m to above W_x , and employment falls as $L_{lmw} < L_m$. Under reasonable assumptions, this decline in employment is associated with a decline in output and prices.

This context-dependent nature of the impact of minimum wage hikes on employment, output, and prices within monopsonistic competition contrasts starkly with the prediction of a model of perfect competition. In perfect competition, an increase in the minimum wage—no matter what its size—will lead to a price hike that fully passes along the higher labor costs onto consumers and will cause lower employment. Further, the perfectly competitive labor market model gives no reason to suppose that the wage-price elasticity would vary systematically with the size of a minimum wage change: the wage-price elasticity associated with a small minimum wage increase should not systematically vary from the wage-price elasticity associated with a large minimum wage increase.

We will implement a rough test of the claim that low-wage labor markets in the restaurant industry are best characterized as monopsonistically competitive by seeing whether “small” increases in minimum wages have a different effect on FAFH prices than do “large” minimum wage increases. We separate the minimum wage changes in our sample into two groups, “small” and “large” increases depending whether the minimum wage change is below or above the average minimum wage increase in our sample, 6.8%. We cannot be sure, of course, that this average is close to W_x in our diagram.

Table 8 (regression 8) presents a regression based on these two types of minimum wage changes. The standard controls appear in this regression.

[Table 8]

As can be seen, for the small minimum wage hikes a single coefficient achieves statistical significance, that for [T-4], and this coefficient is negative. The sum of coefficients for the months immediately surrounding the small minimum wage increase, [T-1,T+1], is also negative although statistically insignificant. The sum of coefficients for the full nine-month period surrounding small minimum wage hike, [T-4,T+4], is negative and significant.

These findings are inconsistent with the perfectly competitive model, which would deny a higher minimum wage could be associated with no price increase and certainly not with a price decline. These findings are consistent with a model of monopsonistic competition as the only statistically significant coefficient for small minimum wage hike is negative. Also in support of a monopsonistic competition model is the fact that the sum of coefficients for [T-4, T+4] in the “small” cases is statistically significant and negative. Note however that this finding should be viewed with some degree of caution because of the effect that interpolation has on standard errors, thus possibly causing us to reject null hypotheses more often than is warranted.

In contrast, the coefficients for “large” minimum wage hikes are statistically significant and positive for all of T, [T-1,T+1], and [T-4,T+4], with elasticities that closely match the results reported for the full dataset in Table 7. The combined finding that large minimum wage hikes boost prices while small minimum wage hikes have no (and possibly a negative) effect on prices is consistent with the model of monopsonistic competition discussed earlier. Taken together, regression 7 provides evidence against perfect competition in low-wage labor markets while regression 8 provides evidence that such labor markets are monopsonistically competitive.[24]

8. Policy Contexts Matter, Sometimes

Minimum wage policies can differ along many dimensions. For instance, a law could provide for a single, large increase in the minimum wage or a series of small, annual increases with no ending date. The impact on prices and employment of these two laws *might* be different as the second type of law permits more long-term planning by businesses and that, in turn, might lead to different consequences for prices, employment, and output. Policy details might matter for the effect that a minimum wage hike has on output prices.

Further, minimum wages laws might also differ by the competitive context of businesses facing a minimum wage increase. For instance, we might treat a *federal* minimum wage hike, as far as the restaurant industry goes, as if it was implemented in a closed economy: cross-national trade and capital mobility relevant to the restaurant industry is relatively unimportant. On the other hand, a *city* minimum wage increase might be treated as if it involved an open economy: the movement of customers and restaurants over the city boundary to or from a neighboring area could be potentially be large enough to affect how businesses respond to the city minimum wage hike.

In this section, we first consider whether the competitive context matters for the level of pass-through. We then turn to consider whether policy details have systematic effects on the level of pass-through.

8.1 Competitive Context: Federal vs State vs City Minimum Wage Hikes

A minimum wage hike might induce cross-border movement of restaurants. For instance, a restaurant facing a minimum wage hike might believe relocating outside the area implementing the minimum wage hike might bring it higher profits than if it had stayed. Or, a restaurant owner might relocate outside the area implementing the minimum wage hike because of opposition to what she/he might see as inappropriate government intervention into the local restaurant industry.

In the case of a federal minimum wage hike, such relocation would likely be rare due to the high cost of relocating outside the United States. State or city minimum wage hikes are much

[24] Bhaskar, Manning, and To (2002) reviews the empirical work associated with monopsonistic competition while Staiger, Spetz, and Phibbs (2010) is a recent study showing strong evidence of monopsonistic competition in the nursing labor market.

more likely to lead to such relocation of restaurants. Further, it seems plausible that the shorter the distance a restaurant must travel to find an area that has not increased its minimum wage, the more likely a restaurant would relocate. If so, then relocation would seem more likely for restaurants facing state minimum wage hikes that are located in metropolitan areas along a state border or for restaurants facing a city minimum wage hike.

If sufficient relocation of restaurants does occur, this might lead to less competition within the area that had increased its minimum wage, which, in turn, could potentially lead to a higher increase in output prices than otherwise would have been the case. This effect would be largest with city minimum wage hikes and in metropolitan areas along a state border facing state minimum wage hikes, less in metropolitan areas away from state borders facing state minimum wage hikes, and the least for a federal minimum wage hike.

A minimum wage hike might also induce cross-border movement of customers.[25] If the relative prices of restaurant meals in two different areas changed, customers would be inclined to cross a border to eat at restaurants whose relative prices had fallen. The cost to consumers to cross an international border (US-Mexico, US-Canada) seeking relatively lower-cost restaurant meals would likely be much higher than the cost to customers crossing state or city borders. It seems plausible to, once again, presume that this effect would be largest with city minimum wage hikes and in metropolitan areas along a state border facing state minimum wage hikes, less in metropolitan areas away from state borders facing state minimum wage hikes, and the least for a federal minimum wage hike.

The impact of cross-border movement on restaurant prices depends on the structure of output and labor input markets. In the case of highly competitive markets, a minimum wage hike raises prices and some customers (who used to eat in the local area) would be inclined to seek restaurant meals outside the area implementing the higher minimum wage. This would shift the demand curve to the left for restaurant meals in the area that has raised the minimum and, everything else remaining equal, this will cause restaurant prices to grow by less in the area that had implemented the minimum wage hike than otherwise would have been the case. This lower (than otherwise) price would be more likely in the case of a city minimum wage hike than in the case of a federal minimum wage hike.

But, if imperfect competition existed (say, monopsonistic competition in low-skilled labor markets) the impact depends on the size of the minimum wage hike. If the minimum wage hike was "small," restaurant prices would fall after the hike, and this might draw customers living

[25] Cross-border movement of labor is also possible but we believe such movement would have only a small effect, if any, on output prices following a minimum wage hike.

outside the area implementing this hike. This increased demand for restaurant meals would keep restaurant meal prices from falling as much as they otherwise would have. Alternatively, if the minimum wage hike was "large," restaurant prices would rise and any cross-border movement of customers would be out of the area and this would lead prices to rise less than they otherwise would have, just as in the case of the highly competitive situation.[26] This cross-border movement of customers would make prices higher or fall not as much as they otherwise would fall.

The net effect of the consequences of cross-border movement of businesses and of customers cannot be known *a priori*. It is an empirical matter.

Regression 9 in Table 9 shows the results of separating federal, state, and city minimum wage hikes. What stand out immediately are the major elasticities seen in the "city" cases relative to the federal and state coefficients. This is consistent with the idea of exit of restaurants from cities implementing minimum wage hikes, yet we believe it is premature to see these results as providing support for this effect. The main reason to be cautious is that the dominant city-level minimum wage appearing in our dataset is San Francisco, which is a special case, for reasons we discuss below. However, one possible implication of the very different results we get for cities in regression 9 is that it might be wrong to simply presume that a study of minimum wage hikes in cities reveals what occurs with federal or state minimum wage hikes. We will temporarily set aside a concern with city minimum wage hikes, but will return to them below using a different methodology that might better reveal the true impact of city minimum wage hikes on output prices.

[Table 9]

The federal and state results presented in regression 9 differ, suggesting that each might have slightly different impacts on FAFH CPI. The sums of [T-1,T+1] for federal and state hikes are, respectively, 0.048 and 0.036, both of which are significant. We also see that federal minimum wage hikes lead to slower growth in prices several months in advance of the hike, which reduces the total difference between federal and state minimum wage hikes, as seen by the sums of [T-4,T+4] of 0.033 and 0.025 respectively. However, F-tests of the equality of coefficients across federal and state cases reveals that none of these differences are statistically significant. We can say, however, that this regression fails to provide evidence that the net effect

[26] Restaurant meals are much closer to a pure service then they are to a tradable good. While home delivery of meals can cross borders (city, state, or even international) much like a good, the delivery area is typically quite small. Similarly, customers can, and do, travel many miles for restaurant meals (perhaps, again, crossing borders) but typically the distance travelled is far shorter than a good would be if shipped across a border. Our discussion of the impact of a minimum wage hike on prices is therefore only relevant to an industry like the near-pure-service restaurant industry, and not necessarily relevant for minimum wage hikes that affect goods-producing industries.

of cross-border movement in the case of state minimum wage hikes leads to any predictable change in prices.[27]

8.2 City Minimum Wage Hikes

The results of regression 9 (above) indicated that city minimum wage hikes are either different cases than federal or state hikes or we need to study city minimum wage hikes using a different approach.

Two series in our sample have seen their own minimum wage laws: Washington, D.C. and San Francisco (the latter has actually seen multiple cities, including Berkeley, Oakland, and San Jose, pass their own minimum wage laws). We have good reason to believe that, at the least, the city of San Francisco represents a unique case that requires special treatment. San Francisco is the only case of indexing the minimum wage to the *local*—i.e. city—CPI, making wage-price elasticities especially difficult to estimate because of the potential two-way influence between minimum wage hikes and city price increases. Furthermore, a strong housing market, tourism industry, and the rise of Silicon Valley have all led to unusually high rates of increase in the cost of living and in restaurant prices in particular. On top of that, San Francisco also implemented (starting in 2008) a health care ordinance that directly increased the costs of the restaurant industry. Washington, D.C. is also unique for its tourism industry, presence of a large group of young professional workers and public officials, and overall strong demand in the restaurant industry.

For these reasons, we adopt an "event study" approach where we compare FAFH inflation in the two cities to the average FAFH inflation in all other cities that did not see a minimum wage change in that month. The "events" include all the minimum wage increases as well as the change in costs due to the health care ordnance in San Francisco that went into effect in April 2008 and which are subsequently increased in January of each year. Previous studies of citywide mandates have used a similar approach when focusing on a single city, where it is convenient to compare a single case to a reference group (Dube, Naidu, and Reich 2007; Colla, Dow, Dube, and Lovell 2014). The pass-through effects are modeled in Equation 3, with dummy variables mw_change_{it} indicating the month t that a minimum wage affects that city i and with dummy

[27] A possible criticism of this analysis is that for some minimum wage increases we will not be able to capture evidence of business flight because the affected series samples from several states. Thus, for the New York City price series, a firm affected by a minimum wage increase in New York may move to a part of New Jersey that is still sampled in the New York City price series. Thus, no effect would be registered in the New York City series. To address this criticism, as a robustness check we ran a second regression that restricted our sample to series that only contain samples from a single state. The results of this second regression did not differ much from that reported above.

variables $mw\_reference_{it}$ indicating cities that, in that same month, did not experience a minimum wage increase.

$$\Delta \log(FAFH)_{it} = \alpha + \sum_{t=-5}^{5} \delta_t * mw\ change_{S.F.or\ D.C.,t} + \sum_{t=-5}^{5} \gamma_t * mw\ reference_{it} + \theta * \log(cityCPI)_{it} + c_i + \epsilon_{it} \quad (3)$$

The overall effect of the minimum wage change in a particular city can then be calculated by subtracting each γ_t from δ_t. The results are reported in regression 10.

[Table 10]

The cumulative T-1 through T+1 coefficient is equal to 0.0028 for San Francisco (p-value of 0.0624) and 0.0014 for Washington, D.C., though the latter is not significant (p-value of 0.2085). Since the cumulative coefficient for the reference group is -0.0006, this implies an overall effect for San Francisco of about 0.0031, or a 0.31% increase in FAFH prices relative to cities that did not see a minimum wage increase. While these appear to be in line with the main wage-price elasticities reported in Table 6, note that because of indexing, changes in San Francisco's minimum wage have been significantly less than 10%, aside from the large initial increase in January 2004. In other words, this coefficient suggests a slightly larger pass-through effect than what was found in the main results. The lack of significance for the case of Washington, D.C. suggests that our findings reflect the unique aspects of San Francisco's economy mentioned above.[28]

In sum, the event study approach illustrates that after accounting for the unique nature of the low-wage labor market in each of these cities, the pass-through effect seems more consistent with what was found in the "federal" and "state" regions. After accounting for local context and using a different (though commonly-applied) methodology, we find that indexation of the minimum wage to the local CPI does not lead to any kind of "wage-price" spiral.

8.3 Indexed vs Scheduled vs One-Shot

The characteristics making federal, state, and city minimum wage hikes potentially different from each other (discussed above) do not exhaust potentially relevant differences between

[28] Some of the coefficients outside of the T-1 to T+1 range are also significant, but in all of these cases, we argue that they are not economically significant. For example, the coefficient for T-3 in the Washington, D.C. case is significant but after accounting for the inflation in the reference cases for that month, the measured impact is minor.

minimum wage hikes. The actual details of the minimum wage law can potentially also be important.

Some minimum wage laws have provided for one-shot increases, where at some date in the future the minimum wage is increased and then the (current) law provides for no further increases (until, perhaps, a new law then provides for another one-shot increase). Other minimum wage laws have provided for a series of increases, perhaps occurring for a small number of years, after which again they provide for no further increases beyond that. Other laws have provided for a perpetual increases that do not end, and are (after perhaps an initial set increase) tied to some cost-of-living index.

We now take advantage of this variation in minimum wage policy caused by recent examples of indexation to compare that approach to the traditional minimum wage "hike"—or the other popular approach of scheduling that "hike" across several years (the strategy adopted in most of the federal minimum wage changes, for example). Since minimum wage increases are usually not voted on or announced more than a few months before the proposed increase is planned to go into effect, more predictable changes (due to scheduling or indexation) may allow business owners to better prepare for and take account of increases in labor costs. Also, more moderate changes (due to indexation which—after the initial large increase—generally results in smaller changes in the minimum wage) could also allow firms to more easily absorb the increase in costs. Reflecting on the findings in the previous section that two "low" minimum wage changes are not the same as one "high" one, moderation along this dimension could temper the contemporary pass-through effect. At any rate, since the competitive model would clearly not predict any difference in wage-price elasticities across different kinds of policies, any evidence of difference may suggest the presence of non-competitive elements.

The results are reported in regression 11 in Table 11, where we compare the cases of indexed minimum wages, excluding the indexed minimum wage changes San Francisco (for the reasons discussed above), with "one shot" cases in which the minimum wage increases a single time, as well as "scheduled" cases in which the minimum wage increase is spread out over a number of years. For both "scheduled" and "one shot" cases, the sum of the T-1 through T+1 coefficients is significant and much higher than the "indexed" case. For the "indexed" case, the sum of the coefficients is not significant. An F-test of a comparison of the equality of coefficients across the "indexed" and "scheduled" cases provides evidence to support the rejection of the null hypothesis that the coefficients are the same. These results are consistent with our finding earlier that moderate minimum wage changes do not lead to significant increases in FAFH prices, and they provide additional evidence that indexation may temper the pass-through effect.

[Table 11]

9. Summary

Among the findings of this paper are the following. First, the impact of minimum wage hikes on output prices (more precisely, on the FAFH CPI) is substantially smaller than previously reported. Whereas the commonly accept elasticity of prices to minimum wage changes is 0.07, we find a value almost half of that, 0.036. The value we found, 0.036, falls far short of what would be expected if low-wage labor markets were perfectly competitive. Second, increases in prices following minimum wage hikes generally occur on the month the minimum wage hike implemented (and not in the month before or the month after). Previous research had reported notable increases in prices the month before and the month after, but we present evidence that such a finding was likely an artifact of interpolation.

Third, small minimum wage hikes have much lower (verging on zero) output price elasticities than do large minimum wage hikes. Such a finding is consistent with the claim that low-wage labor markets are monopsonistically competitive. If such labor markets are monopsonistically competitive, then small increases in minimum wages might lead to increased employment. Our study of restaurant pricing, then, indirectly addresses one of the more contentious issues associated with the employment impact of minimum wage hikes. Fourth, we find no evidence suggesting that exit of restaurants fleeing state minimum wage hikes is large enough to affect output prices. [More on this based on extra regression.]

Fifth, we find evidence that the particulars of a minimum wage policy (indexed, one-shot, scheduled) might affect how price changes occur within the relevant area. These results can be used to design future minimum wage policies that best temper the pass-through effect.

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Tables and Figures

Table 1
Characteristics of Minimum Wages Considered in this Study

Characteristic	1978-1997 (Aaronson et al)	1998-2015	1978-2015
Federal	8	3	11
State	25	101	126
City	1	22	23
Indexed	0	43	43
One or two in series[1] of increases	20	25	45
Perpetually scheduled[2]	0	21	21

[1] Four or less consecutive yearly minimum wage increases.
[2] More than four consecutive yearly minimum wage increases.

Table 2
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[1]
Federal (11 total, leading to 193 binding minimum wage increases)	1/1978, 1/1979, 1/1980, 1/1981, 4/1990, 4/1991, 10/1996, 9/1997, 8/2007, 8/2008, 8/2009
State (131 total binding minimum wage increases)	Alaska (1978, 1979, 1980, 1981) [2]
	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[3] (1/2007, 1/2008, 1/2009, 1/2011 1/2012, 1/2013, 1/2014, 1/2015)

[1] In some cases, the effective month of the minimum wage change is shifted to the following month because the law did not go into effect until later in that month. 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.
[2] During these years, Alaska set its minimum wage at $0.50 greater than the federal minimum wage.
[3] Starting in 2007, Ohio indexed its minimum wage to the national CPI.

	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[4] (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)
	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[5] (1/2004, 1/2005, 1/2006, 1/2007, 1/2008, 1/2009, 1/2010[6], 1/2011, 1/2012, 1/2013, 1/2014, 1/2015)

[4] Starting in 2005, Florida indexed its minimum wage to the South's regional CPI.
[5] San Francisco indexes its minimum wage to the city's CPI.
[6] 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.

	San Jose (3/2013, 1/2014, 1/2015)
	Oakland (3/2015)
	Berkeley (10/2014)

Table 3
Series with Sample Areas in Multiple States

Series for the FAFH Price Index	Sample Areas Used for Restaurant Weights
Boston	Massachusetts, New Hampshire, Maine (post-1998), Connecticut (post-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 (post-1998)
Philadelphia-Wilmington-Atlantic City	Pennsylvania, New Jersey, Delaware (post-1998), Maryland (post-1998)

Note: for the individual counties and towns used for 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 4
Minimum Wage Hikes by Series Periodicity

Periodicity	Observations	Minimum Wage Hikes			
		Federal	State	Local	Total
Monthly	1852	40	42	0	82
Bimonthly	3136	150	101	21	272
Both	4988	190	143	21	354

Table 5

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 *
	(0.005)	(0.006)	(0.006)
T-3	0.006	0.000	0.000
	(0.007)	(0.007)	(0.007)
T-2	0.012	0.003	0.001
	(0.010)	(0.009)	(0.009)
T-1	0.008	-0.002	-0.001
	(0.005)	(0.005)	(0.005)
T	0.052 **	0.039 **	0.039 **
	(0.010)	(0.010)	(0.010)
T+1	0.022 **	0.008	0.008
	(0.008)	(0.008)	(0.008)
T+2	0.012	-0.002	-0.002
	(0.007)	(0.006)	(0.006)
T+3	0.012	-0.002	-0.004
	(0.007)	(0.006)	(0.006)
T+4	0.010	-0.002	-0.002
	(0.006)	(0.005)	(0.005)
[T-1,T+1]	0.081 **	0.044 **	0.046 **
[T-3,T+3]	0.121 **	0.043	0.041
[T-4,T+4]	0.127 **	0.027	0.025
City CPI-All	---	---	0.113 **
			(0.031)
City fixed effects	Yes	Yes	Yes
Month, Year Controls	No	Yes	Yes
Observations	1852	1852	1852
Cities	6	6	6
R^2	0.043	0.162	0.170
Adj. R^2	0.036	0.133	0.141

* $p<0.05$, ** $p<0.01$. Huber-White standard errors reported.

Figure 1
Interpolation and a Stylized Minimum Wage Hike

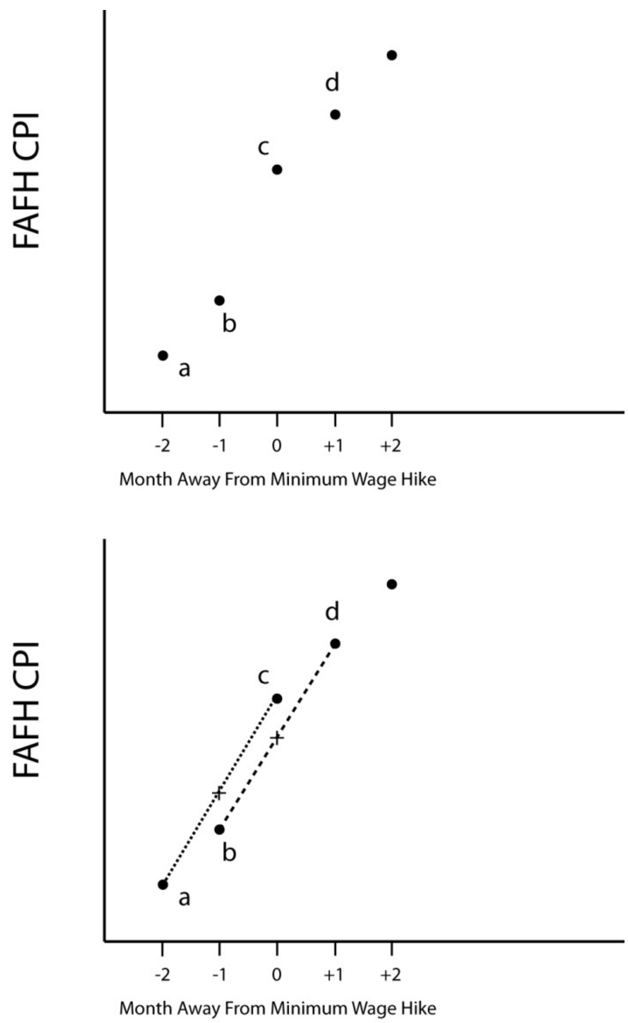

Table 6
Illustrating the Effect of Interpolation
Dependent variable: FAFH inflation

	(4)	(5)
Minimum Wage Change:		
T-4	-0.012 **	-0.007
	(0.004)	(0.003)
T-3	0.001	-0.008 **
	(0.004)	(0.003)
T-2	0.005	-0.003
	(0.005)	(0.003)
T-1	0.010	0.013 **
	(0.006)	(0.004)
T	0.021 **	0.017 **
	(0.007)	(0.005)
T+1	0.015 *	0.015 **
	(0.007)	(0.005)
T+2	-0.003	0.002
	(0.005)	(0.004)
T+3	-0.006	0.006
	(0.004)	(0.004)
T+4	0.005	0.004
	(0.006)	(0.003)
[T-1, T+1]	0.046 **	0.045 **
[T-3, T+3]	0.043 *	0.042 **
[T-4, T+4]	0.036	0.039 **
City CPI-All	0.084 **	0.132 **
	(0.031)	(0.020)
City fixed effects	Yes	Yes
Month, Year Controls	Yes	Yes
Observations	1851	6272
Metropolitan Areas	6	25
R^2	0.285	0.189
Adj. R^2	0.260	0.178

* p<0.05, ** p<0.01. Huber-White standard errors adjusted for smaller degrees of freedom reported (see footnote 18).

Table 7
Estimate of Pass-Through, Full Dataset
Dependent variable: FAFH inflation

	(6)	(7)
Minimum Wage Change:		
T-4	-0.010 **	-0.009 **
	(0.003)	(0.003)
T-3	-0.005 *	-0.006 *
	(0.003)	(0.003)
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 **
	(0.005)	(0.005)
T+1	0.013 **	0.013 **
	(0.004)	(0.004)
T+2	0.001	0.001
	(0.003)	(0.003)
T+3	0.004	0.003
	(0.003)	(0.003)
T+4	0.002	0.002
	(0.003)	(0.003)
[T-1, T+1]	0.044 **	0.045 **
[T-3, T+3]	0.043 **	0.043 **
[T-4, T+4]	0.035 **	0.036 **
City CPI-All	---	0.130 **
		(0.017)
City fixed effects	Yes	Yes
Month, Year Controls	Yes	Yes
Observations	8124	8124
Metropolitan Areas	28	28
R^2	0.170	0.180
Adj. R^2	0.161	0.171

* $p<0.05$, ** $p<0.01$. Huber-White standard errors adjusted for smaller degrees of freedom reported (see footnote 18).

Figure 2

Impact of Minimum Wage Increase in Monopsonistic Competition

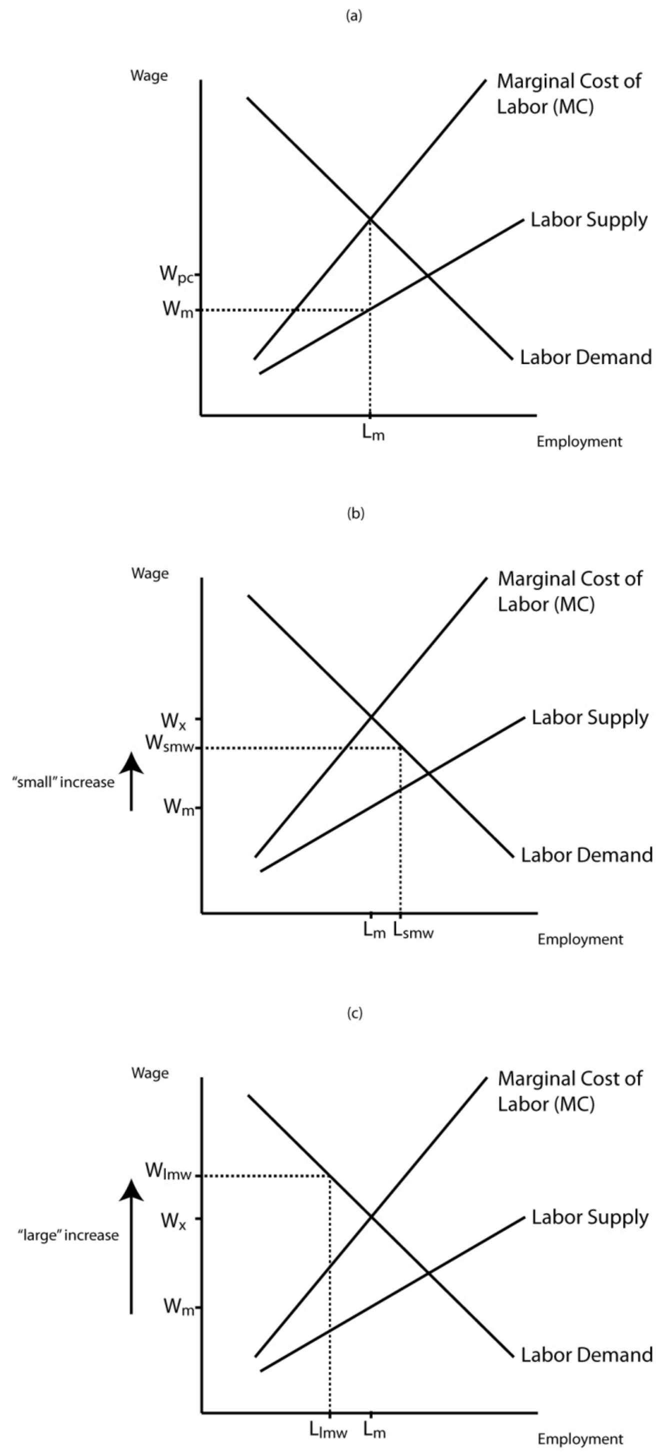

Table 8
Estimate of Pass-Through, Full Dataset
Dependent variable: FAFH inflation

	(8)	
Minimum Wage Change	Small	Large
T-4	-(0.035) *	-(0.007) **
	(0.011)	(0.003)
T-3	-0.011	-0.006 *
	(0.011)	(0.003)
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 **
	(0.013)	(0.005)
T+1	-0.005	0.014 **
	(0.011)	(0.004)
T+2	-0.002	0.001
	(0.011)	(0.003)
T+3	-0.015	0.005
	(0.011)	(0.003)
T+4	-0.001	0.002
	(0.010)	(0.003)
[T-1, T+1]	-0.003	0.048 **
[T-3, T+3]	-0.033	0.045 **
[T-4, T+4]	-0.069 *	0.040 **
City CPI-All	0.132**	
	(0.017)	
City fixed effects	Yes	
Month, Year Controls	Yes	
Observations	8124	
Metropolitan Areas	28	
R^2	0.178	
Adj. R^2	0.172	

* p<0.05, ** p<0.01. Huber-White standard errors adjusted for smaller degrees of freedom reported (see footnote 18).

Table 9
Pass-Through Effects by Policy Context
Dependent variable: FAFH inflation

	(9)		
Minimum Wage Change	Federal	State	City
T-4	-0.014 **	-0.001	0.008 *
	(0.004)	(0.003)	(0.003)
T-3	-0.008 *	-0.003	0.008 *
	(0.004)	(0.003)	(0.003)
T-2	-0.001	-0.006	0.009
	(0.005)	(0.004)	(0.005)
T-1	0.011 *	0.005	0.004
	(0.005)	(0.003)	(0.006)
T	0.023 **	0.022 **	0.012
	(0.006)	(0.006)	(0.008)
T+1	0.014 **	0.010	0.014 *
	(0.005)	(0.005)	(0.007)
T+2	0.000	0.000	0.019 *
	(0.004)	(0.004)	(0.009)
T+3	0.005	-0.002	0.016
	(0.004)	(0.004)	(0.009)
T+4	0.002	0.004	-0.004
	(0.004)	(0.003)	(0.007)
[T-1, T+1]	0.048 **	0.036 *	0.030 *
[T-3, T+3]	0.044 **	0.026 **	0.082 **
[T-4, T+4]	0.033	0.025 **	0.086 **
City CPI-All	0.128 **		
	(0.017)		
City fixed effects	Yes		
Month, Year Controls	Yes		
Observations	8124		
Cities	28		
R^2	0.181		
Adj. R^2	0.170		

* p<0.05, ** p<0.01. Huber-White standard errors adjusted for smaller degrees of freedom reported (See footnote 18).

Table 10
Pass-Through Effects by Policy Context
Dependent variable: FAFH inflation

	(10)[1]		
Minimum Wage Change	Washington, D.C.	San Francisco	Reference Group
T-4	-0.0017 **	0.0009	0.0002
	(0.0007)	(0.0006)	(0.0002)
T-3	0.0007	-0.0001	0.0001
	(0.0005)	(0.0004)	(0.0002)
T-2	0.0006	-0.0005	0.0001
	(0.0007)	(0.0005)	(0.0002)
T-1	0.0004	-0.0001	-0.0003
	(0.0006)	(0.0004)	(0.0002)
T	0.0000	0.0015	-0.0004 *
	(0.0007)	(0.0009)	(0.0002)
T+1	0.009	0.0014	0.0001
	(0.0009)	(0.0009)	(0.0002)
T+2	0.0023 **	0.0004	0.0003
	(0.0008)	(0.0009)	(0.0002)
T+3	-0.0001	0.0001	0.0001
	(0.0005)	(0.0009)	(0.0002)
T+4	0.0000	0.0001	0.0001
	(0.0009)	(0.0008)	(0.0002)
[T-1, T+1]	0.0014	0.0028	-0.0006
[T-4, T+4]	0.0032	0.0036	0.0004
City CPI-All	0.124 **		
	(0.0172)		
City fixed effects	Yes		
Month, Year Controls	Yes		
Observations	8124		
Cities	28		
R^2	0.172		
Adj. R^2	0.162		

* p<0.05, ** p<0.01. Huber-White standard errors adjusted for smaller degrees of freedom reported (see footnote 18).

[1] Coefficients are based on dummy variables, and therefore do not measure wage-price elasticities directly.

Table 11
Pass-Through Effects by Policy Context
Dependent variable: FAFH inflation

	(11)		
Minimum Wage Change	Indexed	Scheduled	One-Shot
T-4	0.006 *	-0.014 **	-0.002
	(0.003)	(0.004)	(0.003)
T-3	0.001	-0.008	-0.004
	(0.004)	(0.004)	(0.003)
T-2	-0.001	0.000	-0.006
	(0.005)	(0.005)	(0.005)
T-1	0.008	0.012 *	0.003
	(0.005)	(0.005)	(0.003)
T	0.011 *	0.024 **	0.025 **
	(0.005)	(0.007)	(0.007)
T+1	0.001	0.015 **	0.014 **
	(0.007)	(0.006)	(0.005)
T+2	0.003	0.001	0.001
	(0.007)	(0.004)	(0.004)
T+3	0.010 *	0.005	-0.004
	(0.004)	(0.004)	(0.004)
T+4	0.005	0.002	0.001
	(0.004)	(0.004)	(0.004)
[T-1, T+1]	0.020 *	0.051 **	0.040 **
[T-4, T+4]	0.044 **	0.037	0.023 *
City CPI-All	0.128 **		
	(0.017)		
City fixed effects	Yes		
Month, Year Controls	Yes		
Observations	8124		
Cities	28		
R^2	0.181		
* p<0.05, ** p<0.01. Huber-White standard errors adjusted for smaller degrees of freedom reported (see footnote 18)			

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Table 9: Tests of the Equality of Coefficients Across Policy Contexts					
	Low vs. High	Indexed vs. Scheduled	Indexed vs. One-shot	Federal vs. State	Federal vs. State (robustness check)
p-value (equality of contemporaneous coefficients)	0.4863	0.1478	0.1071	0.9350	0.9203
p-value (equality of T-1 through T+1 coefficients)	0.0149	0.0432	0.0925	0.3838	0.1894

Table 9, Continued		
	S.F. vs. Reference Group	D.C. vs. Reference Group
p-value (equality of contemporaneous coefficients)	0.0414	0.7256
p-value (equality of T-1 through T+1 coefficients)	0.0227	0.0942