PRICE PASS-THROUGH AND THE MINIMUM WAGE

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Abstract—This paper tests a textbook consequence of competitive markets: that an industry-wide increase in the price of labor is passed on to consumers through an increase in prices. Using several data sources on restaurant prices, I explore the price impact of minimum-wage hikes in Canada and the United States. Particular attention is paid to the timing of these price responses to gauge the "stickiness" of minimum-wage cost shocks. I find that restaurant prices generally rise with changes in the wage bill and that this response is concentrated in the quarter surrounding the month during which the legislation is enacted.

I. Introduction

A textbook consequence of competitive markets is that an industry-wide increase in the price of inputs is passed on to consumers through an increase in prices. Researchers interested in who bears the burden of taxation and exchange-rate fluctuations have explored this fundamental implication. However, little attention has focused on the price implications of minimum-wage hikes. From a policy perspective, this is an oversight: welfare analysis of minimum-wage laws should not ignore consumers. Furthermore, estimates of price shifting can have important implications for wage-push inflation stories and is a key factor in understanding employment responses to minimum-wage hikes.[1]

The Bureau of Labor Statistics (BLS) conducted early studies of minimum-wage price shifting in the 1960s.[2] More recently, Katz and Krueger (1992) and Card and Krueger (1995) collected price information from respondent restaurants in their research on fast-food restaurants. Due to the imprecision of many of their estimates, their results are mixed and difficult to interpret. Generally, they find little evidence of price inflation in their sample of Texas fast-food restaurants but more (yet still mixed) evidence from their New Jersey-Pennsylvania sample. However, this work is limited to restaurants in three states and two minimum-wage episodes. Other work by Card and Krueger involves a broader cross-sectional sample of U.S. states using some of the same data that are used in this paper, but their work is limited to the period surrounding the early-1990s federal increases. Using input-output techniques, Lee and O'Rourke (1999) compute the impact of a minimum-wage increase on food and restaurant prices, but their model assumes full pass-through.

This paper uses several data sources on restaurant prices to more thoroughly examine the impact of minimum-wage hikes in Canada and the United States. Particular attention is paid to the timing of these price changes in order to gauge the stickiness of prices. Although the results are not robust to every specification check, the majority of the evidence in both the United States and Canada suggest that restaurant prices rise with increases in the wage bill that result from minimum-wage legislation. Some evidence suggests that the price response may have been higher in the high-inflation period of the late 1970s and early 1980s, consistent with the work of Cecchetti (1986). Furthermore, the price responses are concentrated in the quarter surrounding the month during which the legislation is enacted. Although minimum-wage legislation is typically enacted many months in advance, there is no price response leading up to the hike and little adjustment in the months subsequent to the hike, excepting the few months around the enactment date.

II. The Price Effect of a Minimum-Wage Increase

Although work on the price impact of minimum-wage increases is limited, a large literature exists on the price response to changes in other industry-wide costs, such as sales taxes and exchange rates.[3] In the standard perfect-competition analysis (assuming constant marginal cost and demand elasticity), an increase in an input price is fully shifted to consumers. Firms set output at a level at which price is equal to marginal cost, and changes in the minimum wage are fully passed on to the consumer. This implication arises from many models, including those in Poterba (1996) and Besley and Rosen (1999).

However, more generally, the degree of shifting depends on a variety of factors, including the magnitude of the demand elasticity, the convexity of demand, the elasticity of marginal cost with respect to output, and the degree of competition. These findings have been generalized by Stern (1987), Besley (1989), and Delipalla and Keen (1992) to allow for free entry and various cost structures. They show that, under specific conditions, even overshifting can occur in imperfectly competitive markets.[4] Undershifting is more commonly observed, reflecting the elastic demand of many

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[1] Most market structures, such as the standard competitive model, predict that minimum-wage increases leads to a simultaneous rise in prices and a fall in employment. One notable exception is Card and Krueger's (1995) monopsony model, which predicts no (or increasing) employment and no (or falling) price inflation.

[2] See Wessels (1980) for a description.

[3] On taxes, see Sumner (1981), Sullivan (1985), Katz and Rosen (1985), Karp and Perloff (1989), Besley and Rosen (1999), and Poterba (1996). The latter two papers describe the literature. Recent examples of exchange-rate pass-through include Gron and Swenson (1996), Lee (1997), and Yang (1997).

[4] Empirically, overshifting of ad valorem taxes has been found by, among others, Besley and Rosen (1999) in the retail apparel industry and Karp and Perloff (1989) in the Japanese television market.

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products. For example, Yang (1997) shows that product differentiation plays an important role in partial pass-through of exchange-rate fluctuations, suggesting that pass-through is generally smaller with products that have a high degree of substitutability. Furthermore, industry demand and supply conditions may lead to dynamic issues about the difference between long- and short-run responses. For example, on the supply side, prices may not react instantaneously to cost changes (Ball & Mankiw, 1994). Furthermore, firms may alter capital-labor input ratios in response to a cost shock such as the minimum wage, resulting in undershifting of prices. On the demand side, undershifting of minimum-wage hikes can occur if consumers cross state borders to purchase goods. Therefore, empirically, it is not surprising that pass-through predictions vary across industries and, even within industries, over time and across studies. Consequently, the impact of an industry-wide cost change, such as a minimum-wage hike, on price behavior is very much an empirical issue.

TABLE 1.—U.S. MINIMUM-WAGE INCREASES, BY STATE AND YEAR, 1978–1995[1]

	All		Not Including Changes Within 6 Months of a Federal Increase	
	Number of Increases	Average Percent Increase	Number of Increases	Average Percent Increase
	(1)	(2)	(3)	(4)
U.S. increases	6	4.4		
By state				
Alaska[2]	7	3.9	0	
California	1	10.3	1	10.3
Connecticut	2	5.0	2	5.0
Hawaii	3	5.1	3	5.1
Iowa	3	4.7	1	3.9
Maine	5	1.2	4	1.2
Massachusetts	3	1.6	3	1.6
Minnesota	4	2.6	3	2.4
New Hampshire	5	1.1	3	1.2
New Jersey	1	7.5	1	7.5
Oregon	3	5.1	3	5.1
Pennsylvania	1	4.3	1	4.3
Rhode Island	5	2.5	5	2.5
Vermont	6	1.4	5	1.6
Washington	3	5.5	3	5.5
Wisconsin	1	3.7	1	3.7
Total	53	3.3	39	3.3
State increases, by year[3]				
1978–1984[4]	0		0	
1985	1	1.3	1	1.3
1986	4	1.9	4	1.9
1987	6	1.8	6	1.8
1988	8	4.0	8	4.0
1989	9	3.3	9	3.3
1990[4]	7	2.7	3	3.2
1991[4]	5	3.0	2	3.4
1992	3	5.4	3	5.4
1993	1	4.3	1	4.3
1994	1	6.2	1	6.2
1995	1	2.5	1	2.5

Notes:
[1] Does not include Washington, D.C. The state minimum wage is taken as the maximum of the federal and state level.
[2] Alaska law requires the state minimum wage to be 0.50 above the federal level.
[3] Not including Alaska or federal increases.
[4] Year with federal minimum-wage hike. Annual federal increases occurred between 1978 and 1981.

III. Data

A. *Minimum Wages*

The minimum-wage histories of the United States and Canada are obtained from several sources. The U.S. legislation is described in the January issues of Monthly Labor Review. This source is corroborated with state minimum-wage histories reported by Neumark and Wascher (1992). Table 1 reports some descriptive statistics on the size and frequency of these changes by state and year. A state's minimum wage is taken as the maximum of the federal and state minimum wage. Notice that only sixteen states had minimum-wage levels above the federal level at any time between 1978 to 1995. Furthermore, most of the state increases occur between 1986 and 1992.

On the other hand, Canada has a very active minimum-wage history. As shown in table 2, there were 97 province-specific increases over the period 1978 to 1995. The most active provinces, Quebec and Ontario, had fifteen minimum wage hikes each over these eighteen years. The minimum-wage time series are obtained from Labor Canada (1996).

B. *Prices*

The restaurant price data used in this study come from three sources: the Bureau of Labor Statistics (BLS), the American Chamber of Commerce Researchers Association (ACCRA), and Statistics Canada (StatCan).

The BLS collects price information on a monthly and bimonthly basis for 27 cities.[5] I use the CPI for food eaten away from home as the restaurant index. In the analyses of U.S. law changes, the overall CPI, the food eaten at home CPI, and specific food CPIs—such as beef, chicken, potatoes, tomatoes, bread, and cheese—are used to control for city-level and national price trends. The city panel runs from 1978 to 1995, encompassing six federal and 39 state minimum-wage hikes. The primary advantage of the BLS data is its frequency; monthly data allow detailed analysis on the timing of price changes relative to minimum-wage increases. However, degrees of freedom are lost because many of the states that passed minimum-wage laws during the 1980s and 1990s are not represented by the BLS cities. Unfortunately, much of the identification is limited to the

[5] Currently, only fifteen cities are collected at this frequency. New York City, Philadelphia, Chicago, Los Angeles, and San Francisco are collected monthly; and Boston, Pittsburgh, Detroit, St. Louis, Cleveland, Washington, D.C., Dallas, Baltimore, Houston, and Miami are collected bimonthly. Prior to 1986, twelve additional cities were collected bimonthly: Buffalo, Minneapolis, Milwaukee, Cincinnati, Kansas City, Atlanta, Seattle, San Diego, Portland, Honolulu, Anchorage, and Denver. After 1986, the BLS reduced the frequency of data collection to a semiannual basis in these twelve cities. Therefore, they are included in the sample through only 1986.

TABLE 2.—CANADIAN MINIMUM-WAGE INCREASES BY PROVINCE AND YEAR, 1978–1995

	Number of Increases	Average Percent Increase
	(1)	(2)
By province		
Alberta	5	4.4
British Columbia	10	3.7
Manitoba	10	2.5
New Brunswick	8	3.1
Newfoundland	7	4.0
Nova Scotia	8	3.4
Ontario	15	2.7
PE Island	8	3.1
Quebec	15	1.9
Saskatchewan	11	2.3
Total	97	2.9
By year		
1978	6	1.7
1979	5	2.6
1980	11	3.7
1981	12	3.0
1982	5	4.9
1983	1	3.6
1984	2	2.9
1985	5	2.8
1986	5	4.2
1987	4	1.9
1988	6	2.4
1989	7	2.8
1990	6	2.5
1991	7	3.0
1992	6	2.9
1993	3	2.1
1994	2	1.7
1995	4	2.4

federal minimum-wage hikes, as only seven state hikes occur in BLS cities. Furthermore, Besley and Rosen (1999) argue that results based on the BLS's broad categorization of commodities may be difficult to interpret if individual components within categorizations are weighted differently across time or area.

The Canadian version of the BLS's CPI data is the StatCan database. The main difference between the BLS's CPI and StatCan's CPI is that the unit of observation in Canada is the province. The price index is food at restaurants; overall, food that is eaten at home and specific food CPI indices are again employed to gauge province-specific and national price trends. The province panel runs from 1978 to 1995, an active period for Canadian minimum-wage legislation. Like the BLS data, a primary advantage of the StatCan data is its frequency. Furthermore, unlike the American data, the Canadian data encompass the entire country, and therefore all minimum-wage hikes can be included in the analysis. Given the frequency of Canadian minimum-wage adjustments, this data set is particularly attractive. However, like the BLS data, broad item categorizations are a concern.

Finally, the ACCRA data alleviate concern about sample size and broad categorization of items by gathering detailed price data on hundreds of U.S. cities. These data are collected from quarterly publications of ACCRA's Cost of Living Index for 1986 to 1993. Each quarterly publication contains a sample of cities that varies from issue to issue.[6] The primary sample includes cities that report price information for 90% of the quarters between 1986 and 1993.[7] However, I also report results using all data available. Of the 542 cities that appear in at least one quarter during the eight years, 107 cities, representing 35 states, appear in the 90% sample.[8] Of the 3,317 possible observations (107 cities times 31 quarters), more than 93% or 3,097 city-quarter observations are available. Quarters with missing data are dropped. Unfortunately, some states in which minimum-wage activity is abundant, particularly in New England, are not represented.

Besides the breadth of cities (and states) represented in this publication, a further advantage of the ACCRA data is that prices for three specific products of the fast-food industry are assembled:

1. Hamburger sandwich: quarter-pound patty with cheese (McDonald's Quarter-Pounder where available)
2. Pizza: 12 in.–13 in. thin-crust cheese pizza (Pizza Hut or Pizza Inn, where available)
3. Fried chicken: thigh and drumstick (Kentucky Fried Chicken or Church's, where available)

These products have remained homogenous through time and across jurisdictions.

However, the ACCRA data has three primary problems. First, the Chamber of Commerce warns that the index does not measure inflation, because the number and mix of the

[6] The set of cities reported is based on whether local Chamber of Commerce personnel participate in a given quarter. This sample selection process is unlikely to bias the estimates.

[7] Others users of this data, such as Parsley and Wei (1996), apply this criterion as well.

[8] The 107 cities are Birmingham, AL; Dothan, AL; Huntsville, AL; Mobile, AL; Fairbanks, AK; Juneau, AK; Phoenix, AZ; Fayetteville, AR; Fort Smith, AR; Jonesboro, AR; Blythe, CA; Indio, CA; Palm Springs, CA; Riverside, CA; Visalia, CA; Boulder, CO; Colorado Springs, CO; Denver, CO; Fort Collins, CO; Grand Junction, CO; Dover, DE; Wilmington, DE; Americus, GA; Atlanta, GA; Augusta, GA; Macon, GA; Decatur, IL; Quad Cities, IL; Rockford, IL; Springfield, IL; Anderson, IN; Bloomington, IN; Indianapolis, IN; South Bend, IN; Cedar Rapids, IA; Mason City, IA; Garden City, KS; Lexington, KY; Louisville, KY; Lake Charles, LA; Monroe, LA; New Orleans, LA; Benton Harbor, MI; St. Cloud, MN; St. Paul, MN; Columbia, MO; Kirksville, MO; St. Louis, MO; Hastings, NE; Lincoln, NE; Omaha, NE; Reno, NV; Albuquerque, NM; Binghamton, NY; Glens Falls, NY; Syracuse, NY; Charlotte, NC; Greenville, NC; Raleigh, NC; Winston-Salem, NC; Akron, OH; Canton, OH; Youngstown, OH; Oklahoma City, OK; Salem, OR; Harrisburg, PA; Lancaster, PA; Philadelphia, PA; Wilkes-Barre, PA; Columbia, SC; Greenville, SC; Myrtle Beach, SC; Spartanburg, SC; Rapid Cities, SD; Vermillion, SD; Chattanooga, TN; Knoxville, TN; Memphis, TN; Morristown, TN; Nashville, TN; Abilene, TX; Amarillo, TX; Dallas, TX; El Paso, TX; Houston, TX; Kerrville, TX; Killeen, TX; Lubbock, TX; Odessa, TX; San Antonio, TX; Waco, TX; Salt Lake City, UT; Roanoke, VA; Richland, WA; Seattle, WA; Spokane, WA; Tacoma, WA; Yakima, WA; Appleton, WI; Fond Du Lac, WI; Green Bay, WI; Janesville, WI; Lacrosse, WI; Manitowoc, WI; Marinette, WI; Wausau, WI; and Casper, WY.

TABLE 3.—DESCRIPTIVE STATISTICS ON INFLATION MEASURES[1]

Dataset	Series	Mean (1)	Standard Deviation (2)	Sample Size (3)
U.S. BLS	*1978–1995*			
	Food away from home	0.418	0.606	4,486
	CPI all	0.455	0.523	4,486
	1982–1995			
	Food away from home	0.290	0.482	3,192
	CPI all	0.285	0.414	3,192
	CPI core[2]	0.332	0.441	3,155
Canada's StatCan	*1978–1995*			
	Food at restaurants	0.438	0.744	2,070
	CPI all	0.406	0.473	2,070
	CPI core	0.406	0.431	2,070
	1982–1995			
	Food at restaurants	0.339	0.717	1,680
	CPI all	0.301	0.426	1,680
	CPI core	0.319	0.390	1,680
U.S. Chamber of Commerce	*1986–1993, Smoothed*[3]			
	Hamburger	0.833	2.35	3,085
	Pizza	0.456	3.03	3,082
	Chicken	0.539	3.63	3,065
	1986–1993, Raw data			
	Hamburger	0.848	5.24	3,097
	Pizza	0.519	5.18	3,097
	Chicken	0.658	8.24	3,097

Notes:
[1] BLS data is monthly at the city level. There are 27 cities up through 1986 and fifteen after 1986. StatCan data is monthly at the province level. There are ten provinces. Chamber of Commerce data is quarterly at the city level. There are 107 cities that are in the 1986 to 1993 sample period for at least 90% of the quarters.
[2] City-level CPI core index begins in 1982.
[3] See text for more detail about the hamburger, pizza, and chicken products. Smoothed data eliminates temporary (less than two quarters) and large (>5% quarterly change) spikes in the Chamber of Commerce price data through linear interpolation. Sample sizes vary between the smoothed and raw data, because spikes that occur in the first two and last two quarters of the sample are discarded.

participants vary from quarter to quarter. Second, because the data are collected on a quarterly basis, it is more difficult to determine the exact timing of those price changes that result from specific events. Third, local Chamber of Commerce staff undertakes the data collection. Therefore, data quality may vary across cities. According to Parsley and Wei (1996), between five and ten prices are collected for each product in each city and averaged to obtain the raw data reported in the publication. Because of the small samples and uneven data quality, the signal-to-noise ratio may be low. To improve the data quality, I smoothed the time series to eliminate large, inexplicable spikes where prices change by more than 5% in a quarter before returning to their original level within two quarters. However, as much as measurement error is limited to the left-side variables and is uncorrelated with the right-side variables, these spikes should not bias the results. Nevertheless, it can cause a loss of efficiency. To assess the importance of measurement issues, the smoothed results are compared with findings using the raw data. Other smoothing techniques—such as averaging across states and robustness techniques that weigh outlier residuals—are also reported.

Table 3 reports descriptive statistics on the key price variables for each data set. Not surprisingly, restaurant inflation is more variable than are broader CPI measures. The smoothed ACCRA data have roughly the same variance as the BLS and StatCan restaurant inflation variables, after accounting for the difference in frequency in the data. However, the standard deviations of the raw ACCRA data are approximately twice as high as the smoothed data. The chicken data are especially noisy relative to the other food products, but the standard deviation is reduced from 8.24 to 3.63 by the smoothing techniques.

IV. Empirical Strategy and Results

The empirical strategy is to relate price changes in the restaurant industry at time t in location i to changes in the minimum wage. Attempts to estimate structural models of tax incidence are presented in Sumner (1981), Sullivan (1985), and Karp and Perloff (1989). To estimate the relationship between taxes and price, these models make heavy data demands as well as require functional-form assumptions about cost and demand in the industry. Instead, like much of the recent tax-incidence literature, this study exploits the time and spatial variation in minimum-wage laws to estimate reduced-form equations of the general form:

$$\pi_{it}^{r} = \alpha + \sum_{t=-T_1}^{T_2} \beta_t w_{it} + \phi \pi_{it} + \gamma E_{it} + \epsilon_i + \epsilon_t + \epsilon_{it} \tag{1}$$

where $\pi_{it}^{r} = \Delta_t \ln (p_{it}^{r})$,
p_{it}^{r} is the restaurant price level at time t for location i,
$w_{it} = \Delta_t \ln (m_{it})$, and
m_{it} is the minimum wage level for location i at time t.

Many theories suggest that firm prices will not respond instantaneously to changes in costs. Therefore, the impact of wage changes is allowed to encompass a finite time period ($-T_1$ to T_2) around the enactment date. This period is set to four months before and after the hike for much of the analysis (or, equivalently, one quarter before and after), but other results are described that allow longer time frames. City- (or province-) and year-fixed effects control for intertemporal and spatial differences that might otherwise bias β.[9] The estimating equations also include monthly or quarterly dummies to control for seasonal behavior in the inflation rate. Alternatively, the national inflation rate (π_t) plays a similar role as year dummies; therefore, specifications are employed with and without these price trends. Controls for the price inflation of specific food products that are common inputs to the restaurant industry (beef, chicken, potatoes, tomatoes, bread, and cheese) are also included in some specifications. BLS and StatCan national food prices are used, because these products are typically sold in national markets. Finally, I also include overall city or province-specific inflation (π_{it}) and state employment conditions (E_{it}) to control for local inflation trends. However, local price

[9] Besley and Rosen include specific measures of time-varying costs that might influence price levels. They find that these measures—including proxies for rental, wage, and energy costs—do not affect their results.

TABLE 4.—THE IMPACT OF MINIMUM-WAGE INCREASES ON INFLATION BLS CITY PRICE DATA, 1978–1995 (HUBER STANDARD ERRORS IN PARENTHESES)

	Dependent Variable: Log Monthly Change in Food Away From Home				
	(1)	(2)	(3)	(4)	(5)
min.-wage hike ($t-4$)	−0.013*	−0.013*	−0.008	−0.002	0.000
	(0.006)	(0.006)	(0.005)	(0.003)	(0.004)
min.-wage hike ($t-3$)	−0.006	−0.006	−0.003	−0.001	0.001
	(0.005)	(0.005)	(0.006)	(0.004)	(0.004)
min.-wage hike ($t-2$)	0.008	0.008	0.005	0.004	0.003
	(0.007)	(0.007)	(0.007)	(0.008)	(0.008)
min.-wage hike ($t-1$)	0.022*	0.022*	0.021*	0.006	0.003
	(0.007)	(0.007)	(0.007)	(0.005)	(0.005)
min.-wage hike (t)	0.028*	0.028*	0.028*	0.021*	0.019
	(0.007)	(0.007)	(0.007)	(0.009)	(0.011)
min.-wage hike ($t+1$)	0.013*	0.014*	0.009	0.014*	0.008
	(0.006)	(0.006)	(0.006)	(0.007)	(0.007)
min.-wage hike ($t+2$)	−0.001	−0.001	−0.007	0.003	0.001
	(0.005)	(0.005)	(0.005)	(0.005)	(0.006)
min.-wage hike ($t+3$)	0.006	0.006	−0.001	0.003	0.001
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
min.-wage hike ($t+4$)	0.007	0.007	−0.002	−0.006	−0.007
	(0.005)	(0.005)	(0.005)	(0.004)	(0.004)
Sum of coefficients					
[$t-4$, $t+4$]	0.066*	0.067*	0.042	0.042	0.029
	(0.024)	(0.024)	(0.023)	(0.022)	(0.023)
[$t-3$, $t+3$]	0.071*	0.072*	0.052*	0.051*	0.037
	(0.020)	(0.021)	(0.021)	(0.020)	(0.021)
[$t-2$, $t+2$]	0.070*	0.071*	0.056*	0.049*	0.035
	(0.017)	(0.017)	(0.017)	(0.017)	(0.019)
[$t-3$, t]	0.052*	0.053*	0.051*	0.030*	0.027
	(0.015)	(0.015)	(0.015)	(0.015)	(0.016)
[t, $t+3$]	0.047*	0.047*	0.029*	0.042*	0.030
	(0.014)	(0.014)	(0.014)	(0.015)	(0.016)
Adjusted R-squared	0.150	0.151	0.177	0.035	0.052
Sample size	4,486	4,486	4,486	2,868	2,868
Time period	1978–95	1978–95	1978–95	1983–95	1983–95
Controls[1]					
Month	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes
City	no	yes	yes	yes	yes
City overall inflation	no	no	yes	no	yes
City food at home inflation	no	no	yes	no	yes
U.S. overall inflation	no	no	yes	no	yes
U.S. food at home inflation	no	no	yes	no	yes
U.S. specific food inflation[2]	no	no	yes	no	yes
Change in log unemployment	no	no	yes	no	yes
Change in log labor force	no	no	yes	no	yes

Notes:
* = significantly different from zero at the 5% level.
[1] Price and unemployment controls include current and lagged monthly variables.
[2] Specific U.S. food inflation include BLS city averages for beef, chicken, bread, potatoes, cheese, and tomatoes.

trends may be affected by the minimum-wage increase and therefore could lead to an understatement of β.

The top of table 4 reports the minimum-wage parameters from regressions using the BLS CPI food-away-from-home inflation rate as the dependent variable, and city and U.S. inflation and employment rates as control variables. All standard error calculations use Huber's formula to account for arbitrary forms of heteroskedasticity.[10] The bottom of the table gives the sum of coefficients for various periods around the minimum-wage hike.

A striking result from the BLS data is the price spike that occurs at the month of the minimum-wage hike.[11] This result is robust to different price controls. In the months prior to the hike, prices drop slightly before jumping significantly during the month before, the month of, and the month after the hike. There is little price adjustment in subsequent months, although F-tests of the null that the lags and leads outside of one month sum to zero are rejected at

[10] To correct for possible autocorrelation in area-specific inflation rates, Newey-West standard errors are also computed. However, because the Huber and Newey-West standard errors are similar, the latter are not reported.

[11] In a specification without lags or leads, the coefficient on the contemporaneous minimum-wage variable is 0.025 (0.008). With a single lag, the coefficients on the contemporaneous and lag variables are 0.027 (0.005) and 0.022 (0.006), respectively. Adding additional lags and leads has little impact on the contemporaneous coefficient. The same result holds for the Canadian and the ACCRA data.

the 1% level for every specification reported in table 4. In the three months ($t - 3$ to $t + 3$) surrounding the wage change, a 1% increase in the minimum wage increases restaurant prices by approximately 0.072 (with a standard error of 0.021), as shown in columns (1) and (2).[12] The elasticity does drop to roughly 0.052 (0.021) when price and employment controls are included in column (3). However, because inflation trends might be affected by the minimum-wage increase, including these controls can bias downward the minimum-wage coefficients. Finally, excluding the high-inflation period of 1978–1982 reduces the pass-through estimate to 0.051 (0.020) when city- and time-fixed effects are included and 0.037 (0.021) with a full set of price and employment controls.[13]

Table 5 reports analogous findings for the Canadian restaurant measure. Similar to the U.S. findings, there is significant price pass-through in the quarter of the minimum-wage increase, although the monthly pattern is a bit different between the two data sources. The StatCan price response is very small leading up to and including the month of the enactment date. The price changes begin occurring the month after the minimum-wage change ($t + 1$) and continue through the third month ($t + 3$). The $t + 3$ coefficient is roughly the same magnitude as the U.S. month t coefficient. Nevertheless, the overall impact is roughly the same size in Canada as the United States, an approximately 0.074% (0.032) increase for every 1% increase in the minimum wage when city- and time-fixed effects are controlled.[14] The pass-through estimate increases slightly when province and Canadian price and employment trends are controlled. But, as with the BLS data, excluding the late 1970s and early 1980s reduces the sum of coefficients to the point of not being statistically significant. Therefore, the high-inflation late 1970s and early 1980s, in part, drives the significant pass-through results in the United States and Canada. The ability of restaurant firms to pass-through minimum wage-increases may have declined in the intervening years.

Table 6 shows the results using the ACCRA price data. Because the data are reported quarterly, only a single lag and lead is included, but these three quarters encompass the same amount of time as the four-month lags and leads of the previous tables. Three sets of results are reported for each of the three food products. Columns (1), (5), and (9) report the results when using the raw data published in ACCRA's Cost of Living Index. Columns (2), (6), and (10) adjust for the temporary and large time-series spikes by smoothing out any quarterly price change that exceeds 5% and does not persist for at least three quarters. The final two columns for each food item use the smoothed data but control for U.S. price trends in food at home and overall inflation or the specific food products already noted. All regressions include month-, year-, and city-fixed effects.

For hamburgers, the raw data show roughly the same-size sum of coefficients as the more aggregated CPI restaurant measures. Furthermore, like the BLS data, nearly all of the inflation response occurs within the quarter of the law's enactment. The pizza and chicken responses are zero and, in some cases, negative. However, smoothing the data to eliminate the large spikes results in much larger hamburger and fried chicken estimates of the price pass-through. These regressions suggest a 0.12% to 0.16% increase in hamburger and chicken prices for every 1% increase in the minimum wage.[15]

The different price responses in the ACCRA and the BLS/StatCan data have a number of possible explanations. First, these findings are consistent with the different price responses found in the tax-incidence studies of Poterba (1996), who finds full shifting using the BLS apparel indices, and Besley and Rosen (1999), who find overshifting

[12] Because year dummies incorporate a potentially misspecified step function, I also ran the regressions with an additional quadratic time trend. This made very little difference to the results. I also added a full set of year-month dummies. However, the year-month dummies capture the impact of the federal increases; identification is limited to the state increases. In the city-fixed effect specification, the three-month lag and lead coefficients sum to 0.032 (0.025).

[13] How do these estimates relate to "full" price shifting? It is difficult to measure the extent of full pass-through from minimum-wage legislation. In the simplest case, a competitive market with inverse demand function and all firms facing the same constant returns to scale production technology, an increase in the wage increases the industry's selling price by the affected labor's share of operating cost. However, the affected labor share estimate is sensitive to a number of assumptions, including the fraction of workers who are affected by the legislation, the share of operating costs these workers represent, and whether firms and consumers respond to potential price and costs shifts. See Aaronson (1998) for a discussion of these considerations.

A first-order estimate of affected labor share can be obtained by looking at the fraction of workers who are affected by a minimum-wage increase and the share of their labor costs on total operating costs. In the 1979-to-1995 CPS outgoing rotation files, approximately 23% of all restaurant employees in the United States are at or below the minimum wage. However, minimum-wage workers comprise only 16% of restaurant annual income, because their hourly wage is roughly 62% of those workers above the minimum. As for labor share, according to the Internal Revenue Services' Statistics on Income Bulletin, labor share in the restaurant industry is 33%. This suggests that full pass-through should be roughly 0.05. However, there are reasons to believe this is an underestimate. Grossman (1983), Card and Krueger (1995), and Green and Paarsch (1997) show that the minimum wage has a spillover effect on the wage distribution above the new minimum. Adding these workers could increase the pass-through estimate to 0.075. Furthermore, the restaurant labor share does not include these labor costs that are associated with intermediate goods that might also be affected by the minimum wage. Nevertheless, these rough calculations suggest that our estimates are plausibly within the range of full pass-through.

[14] With a full set of year-month dummies and province indicators, the sum of the three-month lag and lead coefficients is 0.069 (0.023). These results are similar to the year- and month-fixed effect estimator, because most Canadian minimum-wage changes are made at the province level.

[15] With a full set of year-quarter dummies and city indicators, the sums of the quarter lag and lead coefficients are very similar: hamburger 0.164 (0.052), chicken 0.166 (0.061), and pizza 0.012 (0.062).

Without the 90% time-period restriction, the $[t - 1, t + 1]$ sum of coefficients with price controls and quarter-, year-, and city-fixed effects are hamburger 0.160 (0.033), chicken 0.079 (0.047), and pizza 0.008 (0.042).

TABLE 5.—THE IMPACT OF MINIMUM-WAGE INCREASES ON INFLATION STATISTICS CANADA PROVINCE PRICE DATA, 1978–1995 (HUBER STANDARD ERRORS IN PARENTHESES)

	Dependent Variable: Log Monthly Change in Food Away at Restaurants				
	(1)	(2)	(3)	(4)	(5)
min.-wage hike ($t-4$)	−0.003	−0.003	−0.001	−0.014	−0.004
	(0.009)	(0.009)	(0.008)	(0.008)	(0.008)
min.-wage hike ($t-3$)	0.007	0.007	0.003	0.026	0.015
	(0.017)	(0.017)	(0.012)	(0.029)	(0.018)
min.-wage hike ($t-2$)	0.004	0.004	−0.002	0.007	0.003
	(0.011)	(0.011)	(0.010)	(0.015)	(0.009)
min.-wage hike ($t-1$)	0.011	0.011	0.013	−0.004	−0.001
	(0.009)	(0.009)	(0.007)	(0.008)	(0.010)
min.-wage hike (t)	0.000	0.000	0.004	−0.010	−0.009
	(0.010)	(0.010)	(0.011)	(0.011)	(0.009)
min.-wage hike ($t+1$)	0.013*	0.012*	0.013	0.002	0.004
	(0.006)	(0.006)	(0.007)	(0.007)	(0.008)
min.-wage hike ($t+2$)	0.011	0.011	0.016*	0.002	0.014*
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
min.-wage hike ($t+3$)	0.030*	0.029*	0.033*	0.026	0.027*
	(0.014)	(0.014)	(0.011)	(0.019)	(0.013)
min.-wage hike ($t+4$)	−0.004	−0.004	−0.010	−0.003	−0.007
	(0.006)	(0.005)	(0.006)	(0.006)	(0.006)
Sum of coefficients					
[$t-4$, $t+4$]	0.069*	0.067	0.069*	0.032	0.043
	(0.034)	(0.035)	(0.034)	(0.047)	(0.036)
[$t-3$, $t+3$]	0.076*	0.074*	0.080*	0.048	0.053
	(0.031)	(0.032)	(0.032)	(0.043)	(0.033)
[$t-2$, $t+2$]	0.039	0.038	0.044	−0.004	0.011
	(0.023)	(0.023)	(0.023)	(0.028)	(0.023)
[$t-3$, t]	0.022	0.022	0.018	0.019	0.008
	(0.026)	(0.026)	(0.026)	(0.038)	(0.027)
[t, $t+3$]	0.054*	0.052*	0.066*	0.019	0.036
	(0.021)	(0.021)	(0.021)	(0.026)	(0.021)
Adjusted R-squared	0.188	0.185	0.354	0.171	0.424
Sample size	2,070	2,070	2,070	1,560	1,560
Time period	1978–95	1978–95	1978–95	1983–95	1983–95
Controls[1]					
Month	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes
Province	no	yes	yes	yes	yes
Province core inflation	no	no	yes	no	yes
Province food at home inflation	no	no	yes	no	yes
Canada core inflation	no	no	yes	no	yes
Canada food at home inflation	no	no	yes	no	yes
Canada specific food inflation[2]	no	no	yes	no	yes
Change in log unemployment	no	no	yes	no	yes
Change in log labor force	no	no	yes	no	yes

Notes:
* = significantly different from zero at the 5% level.
[1] Price and unemployment controls include current and lagged monthly variables.
[2] Specific Canadian food inflation includes beef, chicken, potatoes, and tomatoes.

using the ACCRA clothing indices.[16] Second, fast-food restaurants such as McDonald's and Kentucky Fried Chicken tend to comply with minimum-wage laws and have more workers affected by the minimum wage than do restaurants in general, so a larger price effect is likely.

Given the noisiness of the ACCRA data, another possibility is that outliers drive the larger coefficients. To gauge the importance of outliers, two different estimation methods are used. First, the regressions are rerun using a robustness technique that weights observations based on an initial regression and which therefore is not affected by outliers. Observations with large residuals are assigned lower weights. Those with small residuals receive weights approaching 1.[17] The results are reported in panel A of table 7. Regressions with city- and year-fixed effects show total elasticities ranging from 0.035 (0.013) for pizza to 0.073 (0.024) for hamburgers. However, because of the noisiness of this data, up to 500 observations receive weights of less

[16] It is not clear why aggregation matters. Besley and Rosen argue that the BLS indices comprise a variety of products that vary over time and across areas, making the results more difficult to interpret.

[17] The estimation technique calculates Huber and biweights (Berk, 1990). Huber weights are used as starting values for the biweight iteration. Both weights are used because Huber has trouble dealing with extreme outliers and biweights sometimes do not converge. Iterations stop when the maximum change in weights drops below a tolerance level.

Table 6.—The Impact of Minimum-Wage Increases on Inflation[1] American Chamber of Commerce Price Data, 1986–1993 (Huber standard errors in parentheses)

	Log Quarterly Change in McDonald's Hamburger Price				Log Quarterly Change in Kentucky Fried Chicken Chicken Price				Log Quarterly Change in Pizza Hut Pizza Price			
	Not Smoothed	Smoothed			Not Smoothed	Smoothed			Not Smoothed	Smoothed		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
min.-wage hike	0.007	0.021	0.041*	0.013	−0.097	0.031	0.002	0.014	−0.006	0.020	−0.010	0.008
($t-1$)	(0.028)	(0.017)	(0.020)	(0.024)	(0.053)	(0.022)	(0.026)	(0.031)	(0.030)	(0.021)	(0.024)	(0.031)
	0.076*	0.092*	0.087*	0.063*	0.074	0.094*	0.080*	0.125*	−0.015	0.005	−0.010	−0.041
min.-wage hike (t)	(0.023)	(0.018)	(0.018)	(0.023)	(0.061)	(0.028)	(0.029)	(0.038)	(0.032)	(0.021)	(0.022)	(0.033)
min.-wage hike	−0.009	0.045*	0.039	0.079*	−0.002	0.029	0.040	0.023	−0.060	−0.011	0.012	0.042
($t+1$)	(0.038)	(0.022)	(0.024)	(0.040)	(0.067)	(0.022)	(0.023)	(0.027)	(0.038)	(0.025)	(0.027)	(0.036)
Sum of coefficients												
	0.074	0.158*	0.167*	0.155*	−0.025	0.154*	0.122*	0.162*	−0.081	0.014	−0.008	0.009
[$t-1$, $t+1$]	(0.063)	(0.042)	(0.040)	(0.053)	(0.124)	(0.050)	(0.053)	(0.062)	(0.063)	(0.048)	(0.050)	(0.064)
	0.083*	0.113*	0.128*	0.076*	−0.023	0.125*	0.082	0.139*	−0.021	0.025	−0.020	−0.033
[$t-1$, t]	(0.037)	(0.028)	(0.030)	(0.036)	(0.090)	(0.039)	(0.043)	(0.051)	(0.047)	(0.034)	(0.037)	(0.049)
	0.067	0.137*	0.126*	0.142*	0.072	0.123*	0.120*	0.148*	−0.075	−0.006	0.002	0.001
[t, $t+1$]	(0.053)	(0.033)	(0.034)	(0.045)	(0.100)	(0.039)	(0.040)	(0.050)	(0.056)	(0.038)	(0.038)	(0.052)
Adjusted R squared	0.011	0.049	0.051	0.077	0.005	0.021	0.023	0.040	0.001	0.009	0.015	0.033
Sample size	3,097	3,085	3,085	3,085	3,097	3,065	3,065	3,065	3,097	3,082	3,082	3,082
Controls[2]												
Quarter	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
City	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes	yes
U.S. core inflation	no	no	yes	yes	no	no	yes	yes	no	no	yes	yes
U.S. food at home infl	no	no	yes	yes	no	no	yes	yes	no	no	yes	yes
U.S. specific food infl[3]	no	no	no	yes	no	no	no	yes	no	no	no	yes

Notes:
* = significantly different from zero at the 5% level.
[1] See text for more detail on the hamburger, pizza, and chicken products. Not smoothed columns (1, 5, 9) use raw data from ACCRA publications. The data used in the smoothed columns eliminates temporary (less than two quarters) and large (>5% quarterly change) spikes in the price data through linear interpolation. Sample sizes vary because spikes that occur in the first two and last two quarters of the sample are discarded.
[2] Price controls include current and lagged quarterly variables.
[3] Specific U.S. food inflation include ACCRA city averages for beef, chicken, bread, potatoes, cheese, and tomatoes.

than 0.1 in some regressions. This is so even after the data have been smoothed to eliminate the extreme, temporary price spikes. Nevertheless, the robust regression results are in line with the BLS and StatCan data.

Second, the equations were rerun using median regression, which minimizes the sum of the absolute residuals rather than the sum of the square of the residuals. Because median regression is concerned with estimating the median of the dependent variable, it is less sensitive to outliers. Panel B of table 7 displays the results. The median regression coefficients are similar to the robust regression results, with a notable difference between the hamburger elasticities, which are roughly 0.10 to 0.11 (0.015), and the chicken-pizza elasticities, which are closer to 0.02 to 0.03.

Are there particular minimum-wage episodes that drive these outlier results? In the case of pizza prices, some evidence suggests that the result in table 6 is due to the April, 1991, federal minimum-wage increase. Table 8 decomposes the quarterly price changes by whether there is a minimum-wage hike. Columns (1) and (2) show the mean price change for 1991 and columns (3) and (4) for 1986–1990 and 1992–1993. In 1991, the pizza and chicken inflation were lower in the quarters with a minimum-wage hike, whereas the remaining years show the expected pattern of higher price growth in quarters with such labor cost changes. If this 1991 federal increase is excluded from the sample, the three-quarter sum of coefficients for pizza is 0.075 (0.054). The chicken price coefficients rise slightly as well when 1991 data are excluded. Alternatively, if I rerun the pizza regressions with separate federal and state minimum-wage change variables, the total elasticity is −0.134 (0.066) for the federal increases and 0.148 (0.078) for the state increases. The state pizza elasticity is roughly the same magnitude as the chicken and hamburger findings, although again less precisely estimated. The state-federal classification has no effect on the hamburger or chicken elasticities or on the BLS food-away-from-home elasticity.

It is difficult to know why the 1991 price response was different, especially for pizza and chicken restaurants.[18]

[18] One possibility for the absence of pass-through is the weak state of the economy in 1991. To test this hypothesis, the BLS and StatCan regressions were rerun on subsamples of "high" and "low" state or province unemployment rate periods, where the threshold for a "low" unemployment rate period is 6% in the United States (the average unweighted state unemployment rate in our sample is 7.2%) and 9% in Canada (the average unweighted province unemployment rate is 10.8%). The U.S. results were similar across the two states of the economy for the 1978–1995 period when price controls are included; the three-month lag-lead pass-through is 0.053 (0.029) in low-unemployment states and 0.054 (0.030) in high-

TABLE 7.—ALTERNATIVE ESTIMATES OF THE IMPACT OF MINIMUM-WAGE INCREASES ON INFLATION, AMERICAN CHAMBER OF COMMERCE PRICE DATA, 1986–1993 (HUBER STANDARD ERRORS IN PARENTHESES)

	McDonald's Hamburger		KFC Chicken		Pizza Hut Pizza	
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Robust regression[1]						
min.-wage hike ($t-1$)	−0.010	−0.003	0.011	0.017	0.022*	0.015*
	(0.011)	(0.012)	(0.013)	(0.015)	(0.006)	(0.007)
min.-wage hike (t)	0.067*	0.061*	0.032*	0.029*	0.015*	0.012
	(0.012)	(0.012)	(0.013)	(0.014)	(0.006)	(0.007)
min.-wage hike ($t+1$)	0.016	0.003	0.008	0.007	−0.002	0.005
	(0.011)	(0.012)	(0.013)	(0.014)	(0.006)	(0.007)
Fraction of observations with a weight <0.1[2]	7.1%	7.2%	11.3%	11.0%	18.6%	19.1%
Sum of coefficients						
[$t-1$, $t+1$]	0.073*	0.061*	0.051	0.053	0.035*	0.032*
	(0.024)	(0.025)	(0.028)	(0.029)	(0.013)	(0.014)
[$t-1$, t]	0.057*	0.058*	0.043*	0.046*	0.037*	0.027*
	(0.018)	(0.019)	(0.021)	(0.022)	(0.010)	(0.011)
[t, $t+1$]	0.083*	0.064*	0.040	0.036	0.013	0.017
	(0.018)	(0.018)	(0.021)	(0.022)	(0.010)	(0.010)
Panel B: Median regression						
min.-wage hike ($t-1$)	0.002	0.016*	0.006	0.006	0.015*	0.008
	(0.006)	(0.008)	(0.008)	(0.008)	(0.005)	(0.007)
min.-wage hike (t)	0.087*	0.086*	0.024*	0.024*	0.013*	0.010
	(0.006)	(0.008)	(0.009)	(0.009)	(0.005)	(0.007)
min.-wage hike ($t+1$)	0.015*	0.012	0.003	0.002	−0.004	0.003
	(0.006)	(0.008)	(0.009)	(0.009)	(0.005)	(0.006)
Sum of coefficients						
[$t-1$, $t+1$]	0.104*	0.114*	0.033	0.032	0.024*	0.021
	(0.013)	(0.017)	(0.018)	(0.018)	(0.011)	(0.013)
[$t-1$, t]	0.089*	0.102*	0.030*	0.030*	0.028*	0.018
	(0.010)	(0.013)	(0.013)	(0.014)	(0.008)	(0.010)
[t, $t+1$]	0.102*	0.098*	0.027	0.026*	0.009	0.013
	(0.010)	(0.013)	(0.014)	(0.013)	(0.008)	(0.010)
Controls[3]						
quarter	yes	yes	yes	yes	yes	yes
year	yes	yes	yes	yes	yes	yes
city	yes	yes	yes	yes	yes	yes
U.S. core inflation	no	yes	no	yes	no	yes
U.S. food at home inflation	no	yes	no	yes	no	yes

Notes:
* = significantly different from zero at the 5% level.
[1] Regressions use robust techniques that weight observations based on the size of the residuals in a first-stage regression. Price data smoothes out one- and two-quarter spikes. See text for more explanation.
[2] The fraction of observations that receive a weight of 0.1 or less using Huber and biweighting functions.
[3] Price and unemployment controls include current and lagged monthly variables.

However, it appears to be a recurring finding. Katz and Krueger's (1992) independent survey of Kentucky Fried Chicken, Burger King, and Wendy's restaurants in Texas also found little price pass-through due to the April, 1991, federal minimum-wage increase. However, when this analysis was redone on a sample of nine Texas cities in 1991, the results were mixed: there appears to be small, and even negative, price responses among chicken restaurants but not hamburger or pizza restaurants.[19] Smaller April, 1991, price effects also occur in the BLS data. The price elasticity using the 1983–1995 time period is approximately 0.051 (0.020), but the elasticity rises to 0.064 (0.023) if 1991 data are removed.

Several other robustness checks are made of the ACCRA results. First, because each state has multiple cities, I averaged data across states and reran the equations using state-level prices. This can be thought of as another smoothing filter on the data. Second, the sample is restricted to those states in the sample 90% and 100% of the 32 quarters. Neither of these changes affected the elasticities reported in table 6 by more than 0.02.

Third, I deleted cities that are on the borders of other states. Such border cities could be a problem because they are under the influence of legislation from multiple states. Therefore, demand elasticities may be different for border and nonborder cities if consumers can cross borders to purchase products. Furthermore, some restaurants may be

unemployment states. If 1991 is excluded, the results are 0.056 (0.029) for the low-unemployment states and 0.065 (0.041) for the high-unemployment states. The comparable Canadian results are 0.111 (0.044) for low-unemployment provinces and 0.079 (0.035) for high-unemployment provinces. Therefore, there appears to be no evidence of a business-cycle component to the pass-through results.

[19] Two of the eleven Texas cities in the sample are missing data for the second quarter of 1991.

TABLE 8.—CHAMBER OF COMMERCE MEAN PRICE CHANGES, BY YEAR (STANDARD DEVIATION IN PARENTHESES)

	1991		1986–1990, 1992–1993	
	Hike	No Hike	Hike	No Hike
	(1)	(2)	(3)	(4)
Pizza	0.27	0.59	0.73	0.43
	(2.54)	(3.32)	(2.67)	(3.03)
Hamburger	1.01	0.37	2.24	0.81
	(2.21)	(1.73)	(2.16)	(2.40)
Chicken	0.04	0.31	1.66	0.53
	(3.43)	(2.86)	(4.25)	(3.68)
Sample size	89	304	122	2569
Texas only				
Pizza	0.69	−0.13	0.80	0.46
	(0.82)	(5.57)	(1.58)	(3.71)
Hamburger	0.73	0.49	2.78	0.78
	(1.03)	(2.53)	(1.87)	(1.87)
Chicken	−0.50	−0.13	2.44	0.34
	(4.37)	(3.19)	(4.76)	(3.88)
Sample size	9	31	9	263

influenced by the new legislation to raise prices, while others are not affected by the law and are geographically sufficiently separated from those that are that they do not have to raise prices. This situation could mechanically lower price estimates even when full shifting is occurring. The 107-city sample includes twenty border cities, but only twelve have different minimum-wage levels between 1986 and 1993. When the equations are rerun without the twelve cities, the impact is minimal. There is a slight increase in the hamburger results, but the pizza and chicken findings are essentially the same.

A. *Specification Checks*

This section describes several more specification checks on the results. First, I experimented with the length of the window around the minimum-wage hike. Four months might not be enough time to capture the entire price response to the new law. Therefore, table 9 displays an alternative specification that allows an infinite, weighted lag structure.[20] The geometric lag structure is intended to estimate an equation such as

$$\pi^r_{it} = \alpha + \beta(\eta w_{it} + \eta^2 w_{it-1} + \eta^3 w_{it-2} \,.\,.) + \phi\pi_{it} + \epsilon_t + \epsilon_i + \epsilon_{it}, \tag{2}$$

where $\eta < 1$ is the weight assigned to the minimum wage covariates. It is easy to show that equation (2) can be rewritten as

$$\pi^r_{it} = \alpha(1-\eta) + \beta\eta w_{it} + \eta\pi^r_{it-1} + \phi\pi_{it} - \phi\eta\pi_{it-1} + u_t$$
$$u_t = \epsilon_{it} - \eta\epsilon_{it-1} + \epsilon_t - \eta\epsilon_{t-1} + (1-\eta)\epsilon_i. \tag{3}$$

The estimated long-run response is $\beta/(1 - \eta)$. However, the presence of the lagged dependent variable causes OLS parameter estimates to be biased and inconsistent in the presence of serial correlation in the errors. Therefore, equation (3) is estimated using instrumental variables, where the instruments are w_{it-1} and the other right-side variables.

The geometric lag structure confirms that the price response in the United States is concentrated in a short period around the hike and no additional price increase occurs before or after this short window. The Canadian results suggest a long-run impact that is in line with the findings from earlier tables and the U.S. results. However, these parameters are not well estimated. Only the BLS coefficients are statistically different from zero. If OLS is used instead of IV, the BLS and StatCan long-run coefficients are 0.037 (0.007) and 0.030 (0.009). The ACCRA hamburger, chicken, and pizza coefficients are 0.060 (0.014), 0.074 (0.028), and −0.006 (0.021), respectively.

Alternatively, the bottom of table 9 reports results with lower frequency data. Baker, Benjamin, and Stanger (1999) report that the frequency of the data can reconcile different findings on minimum-wage employment effects. Furthermore, data that ignores the noisy monthly changes may improve the signal-to-noise ratio. With this in mind, the monthly and quarterly data are recomputed as annual data, and two sets of results are reported. The first row under the "annual average" panel uses differenced data, as has been employed throughout the paper. The magnitude of these results are similar to those reported in earlier tables, although, as in the geometric lag structure findings, the standard errors of the estimates increase enough so that most of the results are no longer significant at the 5% level. The second row reports results using log price levels rather than differences, because there may be concern that the differencing is eliminating important long-run relationships between minimum wages and prices. The levels data show a strong correlation between restaurant prices and minimum wages in the Canadian and the ACCRA hamburger and chicken data, but little relationship in the BLS data due to the decreased precision of these estimates. Therefore, the evidence in table 9 seems to suggest that there are no further price increases outside of the initial increase that occurs within a quarter of the minimum-wage enactment date. Given the size of the standard errors, it is difficult to identify any long-run price dynamics with this data.

Finally, it might be expected that cities with more minimum- and low-wage workers experience larger price effects from minimum-wage increases. To explore this possibility, the 27 BLS cities are ranked by average restaurant industry wages using the 1979–1995 CPS outgoing rotation files. I identify the six cities with the lowest and highest average

[20] I also experimented with extending the lag and lead around the enactment month to nine and twelve months. Similar to table 9, those results have much larger standard errors than do tables 4 to 7 and therefore tend not to be significant at the 5% level. These results are available upon request.

TABLE 9.—SPECIFICATION CHECKS (HUBER STANDARD ERRORS IN PARENTHESES)

	BLS Food Away From Home	Statistics Canada Food at Restaurants	Chamber of Commerce		
			Hamburger	Chicken	Pizza
	(1)	(2)	(3)	(4)	(5)
Geometric lag structure[1]					
Lagged dependent variable	0.615*	0.529	0.594	0.352	2.200
	(0.172)	(0.394)	(0.419)	(0.323)	(10.570)
min.-wage hike (t)	0.023*	0.035*	0.063*	0.066*	−0.039
	(0.006)	(0.015)	(0.023)	(0.031)	(0.160)
long-run coefficient	0.060*	0.075	0.155	0.102	0.033
	(0.026)	(0.071)	(0.158)	(0.061)	(0.167)
Annual averages[2]					
price inflation [t, t + 1]	0.062	0.040	0.174*	0.116	0.068
	(0.040)	(0.041)	(0.046)	(0.072)	(0.084)
price levels [t, t + 1]	0.042	0.097*	0.195*	0.207*	0.066
	(0.060)	(0.028)	(0.041)	(0.056)	(0.069)
Controls[3]					
Month	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes
City	yes	yes	yes	yes	yes
City overall inflation	yes	yes	no	no	no
City food at home inflation	yes	yes	no	no	no
Time period	1978–95	1978–95	1986–93	1986–93	1986–93

Notes:
* = significantly different from zero at the 5% level.
[1] Geometric lag structure allows for an infinite, weighted lag structure. The lagged dependent variable is instrumented by the lag of the minimum-wage change. The geometric lag structure estimates for columns (3) and (4) use the minimum wage in month t + 3 instead of month t.
[2] Annual data, rather than monthly or quarterly data, is used.
[3] Price controls include current and lagged monthly or quarterly variables. Annual data does not include control lags.

wages.[21] The rest of the cities are deemed "middle-wage cities," but the results are not especially sensitive to adjusting the number of cities in each group. The basic specifications are rerun in two ways. First, I include interactions between the high- and low-wage city indicators and the minimum-wage measures. These results, reported in the top panel of table 10, show that pass-through is significantly less in high-wage cities, as expected. Second, the regressions are run on stratified samples of the three wage groupings. These results, reported at the bottom of table 10, show a monotonic decline in the pass-through estimates but no statistical difference between the three groups of cities. Therefore, the results of the wage test are mixed.

B. Is Minimum Wage Legislation Endogenous?

It is plausible that state and federal legislators may become more concerned with the deteriorating real value of minimum wages during periods of high inflation. Therefore, the estimated minimum-wage elasticity may be biased upward; persistently high inflation rates may cause an increase in the minimum wage rather than the other way around. However, this pattern is also consistent with evidence presented by Cecchetti (1986) that firms are more able to adjust prices during periods of high inflation and therefore is not prima facie evidence of an endogeneity problem. Furthermore, time- and city-fixed effects should account for unusually high inflation periods. Nevertheless, I tested for the possibility of an endogeneity problem by examining inflation patterns before the enactment of minimum-wage legislation (that is, when legislation is debated and passed) between 1982 and 1995. Fortunately, the BLS and StatCan data reveal virtually no evidence that inflation is higher in the two years prior to the legislation's enactment date. A possible exception is if state increases are analyzed separately; the endogeneity problem may be more severe in the case of state legislation. However, reestimating the BLS and

[21] The low-wage cities are St. Louis, Pittsburgh, Milwaukee, Cleveland, Baltimore, and Cincinnati. These are the only six cities with an average restaurant wage below the U.S. mean. The high-wage cities are Anchorage, Honolulu, Boston, New York City, Washington, D.C., and San Francisco. Ranking by fraction of minimum-wage workers changes the list of cities a little, but the results remain unchanged.

TABLE 10.—HIGH WAGE VERSUS LOW WAGE CITIES BLS CITY PRICE DATA, 1978–1995 (HUBER STANDARD ERRORS IN PARENTHESES)

	Sum of Coefficients		Sample Size
	−4, +4 Lags	−3, +3 Lags	
Interaction specification			
Low-wage cities	0.070*	0.070*	4,486
	(0.036)	(0.031)	
Middle-wage cities	0.098*	0.099*	4,486
	(0.031)	(0.028)	
High-wage cities	0.005	0.024	4,486
	(0.033)	(0.030)	
Stratified sample			
Low-wage cities	0.085	0.087*	1,072
	(0.048)	(0.041)	
Middle-wage cities	0.064	0.071*	2,342
	(0.037)	(0.033)	
High-wage cities	0.058	0.064*	1,072
	(0.036)	(0.032)	

Notes:
* = significantly different from zero at the 5% level.

ACCRA regressions with separate federal and state minimum-wage covariates shows no difference in the state or federal coefficients, except among the ACCRA pizza parameters. Therefore, I conclude that there is little reason to be concerned about endogeneity in this analysis.

V. Conclusions

Using a variety of data sources on restaurant prices, this paper tests a textbook consequence of competitive markets (that an industry-wide increase in the price of inputs is passed on to consumers through an increase in prices).[22] Estimates of price shifting can have important implications for wage-push inflation stories and estimates of the response of employment to minimum-wage increases. Although the results are not robust to every specification check, the majority of the evidence suggests that restaurant prices rise with increases in the wage bill that result from minimum-wage legislation. Some evidence suggests that the price response may have been higher in the high-inflation period of the late 1970s and early 1980s, consistent with the work of Cecchetti (1986). Furthermore, the price responses are concentrated in the quarter surrounding the month during which the legislation is enacted. Although minimum-wage legislation is typically enacted many months in advance, there is no price response leading up to the hike and little adjustment in the months subsequent to the hike, excepting the few months surrounding the enactment date.

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[22] The predicted elasticity for overall CPI measures is substantially smaller (on the order of 0.01) and therefore difficult to identify given the size of the standard errors. It would be interesting to analyze other industries, but few have sizable low-wage labor costs. For example, this analysis might shed light on the different tax-incidence findings in Poterba (1996) and Besley and Rosen (1999). However, because labor share in retail apparel is 13% and only 10% of workers earn the minimum wage (although 30% earn within one dollar of the minimum), it is difficult to differentiate full from zero pass-through given the predicted elasticity of 0.01 to 0.02 and the size of the standard errors.

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