I. Introduction

A textbook consequence of competitive markets is that an industry-wide increase in the price of inputs will be passed on to consumers through an increase in prices. This fundamental implication has been explored by researchers interested in who bears the burden of taxation and exchange rate fluctuations. 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, as well as potentially provide an explanation for the small short-run employment effects that have been found in some of the minimum wage literature.

The only work that explores minimum wage price shifting is Card and Krueger (1995), who collected price information from respondent restaurants in their research on fast food restaurants. Their results are mixed and difficult to interpret due to the imprecision of many of their estimates. Generally, they find little evidence of price inflation in their Texas fast food restaurant sample 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 that they present looks at a broader cross-sectional sample of U.S. states but is also limited to the period surrounding the early 1990s federal increases.

This paper uses several data sources on restaurant prices to more thoroughly examine the impact of minimum wage hikes in Canada and the U.S. Particular attention is paid to the timing of these price changes and their overall impact relative to full or competitive price pass-through predictions. Section II briefly describes the price implications of minimum wage hikes, including the predicted elasticity that would be consistent with full price shifting. Three data sources on prices are described in section III. Section IV presents the empirical strategy and results. The results suggest that restaurant prices rise roughly one-for-one with increases in the wage bill that result from minimum wage legislation. Furthermore, the price responses are concentrated in the quarter surrounding the month that 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. If anything, there is some evidence that minimum wage price effects dissipate over time. The magnitude of these findings is roughly the same in the U.S. and Canada, and is fairly robust to changes in data, specification, and estimation techniques. However, because of small predicted elasticities, it is difficult to draw inferences about the price impact on broader indices or industries that have a small share of low wage labor.

II. The Price Effect of a Minimum Wage Increase

While there has been little work on the price impact of minimum wage increases, there is a literature on the price response to changes in other industry-wide costs, such as sales taxes and exchange rates.[1] In the standard perfect competition analysis, assuming constant marginal cost and demand elasticity, an increase in the price of an input is fully shifted to consumers. This implication arises from many models, but can be easily seen by looking at the comparative statics from a conjectural variations model, where the firm chooses an output x_i to maximize

(1) $$q(X)x_i - C(x_i) - \delta x_i .$$

q(X) is the inverse demand function for the industry, X is total industry output, $C(x_i)$ is the industry cost of production, and δ is the additional cost of labor due to minimum wage laws. The interpretation of δ is discussed below. The first order condition for a profit maximizer is

(2) $$q(X) + q'(X)x_i - C_{x_i} - \delta + q'(X)x_i\alpha_i = 0$$

where $\alpha_i = dX / dx_i$ is the firm i's conjectural variation or the amount by which a one unit change in firm i's output will affect total industry output. Differentiating (2), it can be shown that

[1] On taxes, see Sumner (1981), Sullivan (1985), Katz and Rosen (1985), Karp and Perloff (1989), Besley and Rosen (1994), and Poterba (1996). The latter two papers describe the literature. Recent examples of exchange-rate passthrough include Gron and Swenson (1996), Lee (1997) and Yang (1997).

the impact of an industry-wide cost change on price depends critically on industry market structure (α_i) and demand elasticities (ε):

(3) $$\frac{dq}{d\delta} = q'(X)\Sigma_{i=1}^{N} \frac{\partial x_i}{\partial \delta}$$

$$\text{where } \frac{\partial x_i}{\partial \delta} = \frac{1}{2-(1+\alpha_i)\varepsilon + \alpha_i - C_{x_i x_i} / q'} \text{ and } \varepsilon = -\frac{x_i q''}{q'}$$

and N is the number of firms in the industry. In the simple case of linear costs ($C_{x_i x_i} / q' = 0$) and a perfectly competitive market ($\alpha_i = -1$), firms set output at a level at which price is equal to marginal cost. Changes in the minimum wage are fully passed on to the consumer.

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.[2] Undershifting is more commonly observed, reflecting the elastic demand of many products. For example, Yang (1997) shows that product differentiation plays a key 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.[3] Gron and Swenson (1996) find that automobile producers transfer their production across national borders to deal with exchange rate fluctuations, leading to undershifting of prices relative to exchange rate movements. In a similar vein, undershifting of

[2] Empirically, overshifting of ad valorem taxes has been found by, among others, Besley and Rosen (1994) in the retail apparel industry and Karp and Perloff (1989) in the Japanese television market.
[3] See Ball and Mankiw (1994) on sticky prices.

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, across studies. As a result, the impact of an industry-wide cost change, such as a minimum wage hike, on price behavior is very much an empirical issue.

<u>What is the Full Price Pass-Through Prediction?</u>

Further confounding the measurement of the extent of price shifting from minimum wage legislation is that the full shifting amount must be estimated. In the simplest case, a competitive market with inverse demand function q=q(Y) 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. Assume affected labor is computed as the product of the share of labor cost (L) in total operating cost (C) and the fraction of workers that are affected by the minimum wage level (E_{mw} / E). The increase in wages will lead to an increase in prices that is proportional to the share of affected labor:

$$(4) \qquad \delta = \frac{L}{C}\frac{E_{mw}}{E}\frac{w_t}{w_{t-1}}$$

However, Equation (4) ignores two factors. First, demand may fall when restaurants raise prices, thus biasing upward the full pass-through prediction of (4).[4] But in the case of industry-wide price shifts, such as when the minimum wage increases, the relevant output elasticity is the industry's, which is likely to be much smaller than any individual restaurant's elasticity. Brown (1990) estimates that the price elasticity of demand for all restaurants is -0.2 but is -1.0 for fast food restaurants. Therefore, the relatively inelastic demand for restaurant food suggests that the output bias to (4) is likely to be small for restaurants but may be problematic for fast food products.

[4] Partly counterbalancing this effect is the possibility that workers whose wages are increased by a minimum wage hike increase their demand for restaurant products.

Second, equation (4) does not take into account the possibility that firms substitute capital or high wage labor for low wage labor when minimum wages increase. However, consistent with Shephard's lemma, this bias is likely to be negligible. Evidence on this proposition is exemplified by the contentious literature on employment effects of minimum wage increases. This debate seems to revolve around whether there is a small effect (Neumark and Wascher 1992) or no effect (Card and Krueger 1995) of minimum wage increases on low wage employment. Even using Neumark and Wascher's employment elasticity does not substantially alter the estimate of δ. In addition, to quantify the importance of capital-labor substitution bias, I computed labor shares (L/C) from McDonald's annual reports for 1988 to 1995. For this specific company, the share of labor does not change, despite two federal minimum wage increases and 28 state increases. In 1988, labor share is 31.1 percent of company-operated expenses and in 1995 it is 31.2 percent. Between 1989 and 1992 (encompassing two federal increases that raised the nominal minimum wage by 27%), labor share remained constant at 31.1 percent. Of course, this particular evidence does not allow us to preclude the possibility that McDonald's shifts labor costs away from low wage workers and toward higher wage workers in response to minimum wage increases. Therefore, while I am confident that using equation (4) will closely approximate the full shifting minimum wage price elasticity, the actual pass-through prediction, which is detailed next, should be considered an upper bound estimate.

Labor Share in the Restaurant Industry

To compute labor share in the restaurant industry, two data sources are employed. First, I searched the SEC's EDGAR database of annual and quarterly company reports using five keywords: restaurant, steak, seafood, hamburger, and chicken. Of the 17 restaurant companies that appeared, the mean and median 1995 payroll expense to total cost ratio is approximately 30 percent, roughly consistent with McDonald's payroll to expense ratio of 31 percent.[5]

[5] The standard deviation is 5.0 percent, the minimum is 20.7 percent (Food Quest), and the maximum is 41.7 percent (Cooker Restaurants).

Second, the Internal Revenue Service summarizes operating costs by industry in the Statistics on Income Bulletin. This data come from a sampling of corporate income tax forms. Because operating costs are broken down by category, it is possible to calculate labor share.[6] Unfortunately, I must look at partnerships rather than corporations because, according to communication with the IRS, labor cost is notoriously difficult to decompose for corporations. The IRS analysts do not believe this is a serious problem for smaller firms. Nevertheless, the figure reported in the IRS sample is roughly the same as in the SEC sample; the labor share of operating costs among eating place partnerships is 33 percent. The labor share for all industries is 19 percent.

The Fraction of Minimum Wage Workers in the Restaurant Industry

Estimating E_{mw}/E is more precarious. The primary complication arises because the minimum wage may have a spillover effect on the wage distribution (Grossman 1983, Card and Krueger 1995, and Green and Paarsch 1997).[7] In particular, Card and Krueger estimate that approximately 33 to 40 percent of Texas fast food restaurants increased the wage of workers who were already above the old minimum (but below the new minimum) to a level beyond the new minimum after the new law took effect. As for workers already well above the new minimum wage, only nine percent of restaurants with a starting wage equal to the minimum increased the pay of their high-wage workers. However, 60 percent that start their employees at a much higher level bumped up their employees to maintain a wage differential from other restaurants.

To calculate the fraction of the restaurant population affected by a minimum wage hike, I use the March rotation of the 1986 to 1993 Current Population Surveys. Table 1 reports some of the findings. The top row reports the fraction of workers in four industry classifications - all

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6 Labor cost is the sum of wages and salaries, cost of labor in the cost of sales and operations, employee benefit plans, half of repairs, and pension, profit sharing, and other annuities. Total operating cost is the sum of cost of sales and operations, rent, interest, taxes, bad debts, repairs, depreciation, depletion, guaranteed payments to partners, wages and salaries, employee benefit plans, and pension, profit sharing, and other annuities.

7 Workers above the new minimum may get raises because of shifts in demand away from unskilled and toward skilled workers after a minimum wage increase. Grossman (1983) presents an alternative model that relies on a correlation between skilled worker effort and the difference between skilled and unskilled wages.

private, retail trade, restaurant, and food service workers -- who are at or below their state's minimum wage (columns 1 to 4) or above but within one dollar of the minimum wage level (columns 5 to 8). Hourly earnings, which includes tips, is defined as weekly earnings divided by weekly hours. Approximately 24 percent of all restaurant employees in the U.S. are at or below the minimum wage and an additional 37 percent are within one dollar. When the sample includes only those who are in food service, about 29 percent are at or below the minimum and an additional 40 percent are within one dollar. However, these numbers vary dramatically across states. Very few workers earn the minimum wage in Alaska, even though state law requires the minimum to be 50 cents above the federal threshold. However, in West Virginia, over one-half of all food service workers are at the minimum wage and close to 90 percent are within one dollar. The figures are somewhat similar when using fraction of total hours worked by low wage workers; the slightly smaller numbers account for the fewer hours worked by low wage workers.

Consequently, using equation (4), a one percent increase in the minimum wage level will increase restaurant prices by 0.11 percent when additional labor costs are fully passed through to consumers. This figure assumes that all workers below and one-third of workers within one dollar of the new minimum wage have their pay increased by the same amount as a result of a minimum wage change, and there are no additional effects from a change in the price of intermediate goods. If only minimum wage workers are affected by an increase, the full price shifting elasticity is 0.075. The corresponding elasticity for broader CPI measures is substantially smaller; the predicted pass-through is in the order of 0.015 to 0.020, reflecting that five percent of private industry workers are below the minimum wage, an additional 10 percent are within one dollar, and labor share is roughly 20 percent for all industries.[8] However, these elasticities are best considered upper bound estimates because of the possibility of output and substitution biases.

[8] 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 (1994). However, since labor share in retail apparel is 13 percent and only 10 percent of workers earn the minimum wage (although 30 percent earn within one dollar of the minimum), it is difficult to differentiate full from zero pass-through given the predicted elasticity of 0.013 to 0.026 and the size of the standard errors.

For example, if one assumes that low wage employment declines by two percent as a result of a 10 percent minimum wage hike (Neumark and Wascher's estimate), the implied minimum wage price elasticities should be reduced by approximately 0.01.

III. Data

Minimum Wages

The minimum wage histories of the U.S. 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 in Neumark and Wascher (1992). Table 2 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 16 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 3, there were 97 province-specific increases over the period 1978 to 1995. The most active provinces, Quebec and Ontario, had 15 minimum wage hikes each over these 18 years. The minimum wage time series are obtained from Labor Canada (1996).

Prices

There are three sources of restaurant price data used in this study: 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.[9] I use the CPI for food eaten away from home as the restaurant index. In the analyses of U.S. law

[9] Currently only 15 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 DC, Dallas, Baltimore, Houston, and Miami are collected bimonthly. Prior to 1986, an additional 12 cities were collected bimonthly. They are Buffalo, Minneapolis, Milwaukee, Cincinnati, Kansas City, Atlanta, Seattle, San

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 l990s are not represented by the BLS cities. Unfortunately, much of the identification is limited to the federal minimum wage hikes, as only seven state hikes occur in BLS cities. Furthermore, Besley and Rosen (1994) suggest that the BLS's broad categorization of commodities can hide underlying information that varies over time and location.

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 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 minimum wage legislation in Canada. 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 dataset is particularly attractive. However, like the BLS data, the price indices may be prone to the aggregation bias noted in Besley and Rosen.

Finally, the ACCRA data alleviate concern about the BLS sample size and aggregation bias by gathering detailed price data on hundreds of U.S. cities. It is collected from quarterly publications of ACCRA's *Cost of Living Index* for 1986 to 1993. Each quarterly publication

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Diego, Portland, Honolulu, Anchorage, and Denver. After 1986, the BLS reduced the frequency of data collection to a semiannual basis in these 12 cities. Therefore, they are included in the sample only through 1986.

contains a sample of cities that varies from issue to issue.[10] In an attempt to construct a relatively balanced panel of cities, only those that report price information in 90 percent of the quarters between 1986 and 1993 are included. Of the 542 cities that appear in at least one quarter during the eight years, 107 cities, representing 35 states, appear in the requisite number of periods.[11] Unfortunately, some key states where 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. They are:

1. Hamburger sandwich- ¼ pound patty with cheese. McDonald's Quarter-Pounder where available,
2. Pizza - 12-13" 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, there are three primary problems with the ACCRA data. First, the Chamber of Commerce warns that the index does not measure inflation since the number and mix of the participants vary from quarter to quarter. Second, since the data are collected on a quarterly basis, it is more difficult to determine the exact timing of price changes resulting from specific

[10] 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.

[11] 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.

events. Third, the data collection is undertaken by local Chamber of Commerce staff. 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 then averaged to obtain the raw price data reported in the publication. As a result 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 over five percent in a quarter before returning to their original level within two quarters. However, as much as measurement error is limited to the left hand side price variables and is uncorrelated with the right hand side variables, these spikes should not bias the results. Nevertheless, it can cause a loss of efficiency.[12] To assess the importance of this measurement issue, I compare the smoothed results with regressions using the raw data. Other smoothing techniques, such as averaging the city data across states and using robustness techniques that weigh outlier residuals, are also reported.

Appendix 1 reports descriptive statistics on the key price variables for each data set. Not surprisingly, restaurant inflation is more variable than 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 deviation 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

[12] Suppose the measured dependent variable (π_{it}) is the sum of the true value (π^*_{it}) and measurement error (μ_{it}). The variance matrix of the OLS estimator is then

incidence are presented in Sumner (1981), Sullivan (1985), and Karp and Perloff (1989). These models make heavy data demands as well as require functional form assumptions about cost and demand in the industry in order to estimate the relationship between taxes and price. 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:

$$\pi_{it}^{r} = \alpha + \sum_{t=-T_1}^{T_2} \beta_t w_{it} + \varphi\pi_{it} + \gamma E_{it} + \varepsilon_i + \varepsilon_t + \varepsilon_{it} \tag{5}$$

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 allow longer time frames and a geometric lag structure that introduces an infinite, but geometrically weighted, lag length. City (or province) and year fixed effects control for intertemporal and spatial differences that might otherwise bias β.[13] 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 and 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

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$\mathrm{Var}(\beta) = E[(X'X)^{-1} X'(\varepsilon+\mu)(\varepsilon+\mu)' X(X'X)^{-1}]$. Assuming $\sigma_{\varepsilon\mu} = 0$, the efficiency loss from measurement on the left hand side is equal to $\sigma_{\mu}^{2}(X'X)^{-1}$.

[13] 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.

Since 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 any of the results.

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 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.[14] The bottom of the table gives the sum of coefficients and tests the significance of the sums from 0 and 0.075 for various time periods around the minimum wage hike. The latter test is a lower bound estimate of full price shifting.

A striking result from the BLS data is the price spike that occurs at the month of the minimum wage hike. 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. In the three months (t-3 to t+3) surrounding the wage change, a one percent increase in the minimum wage increases restaurant prices by approximately 0.070 in columns (1) to (4) and (6). The elasticity does drop to roughly 0.06 when the specific food inputs are controlled in column (5). Nevertheless, these estimates are significantly different from zero and statistically within the full price shifting effect of 0.075 to 0.110.

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. Furthermore, the impact is roughly the same size in Canada as the U.S., approximately 0.07 to 0.08 percent for every one percent increase in the minimum wage. However, an unique feature of the Canadian price response is the monthly pattern. The impact is very small leading up to and including the month of the wage legislation's starting date. The price changes begin occurring the month after the minimum wage change (t+l) and continue through the third month (t+3). The t+3 coefficient is roughly the same magnitude as the U.S. month t coefficient. Therefore, assuming the fraction of minimum wage workers and the share of labor cost is the same in Canada as the U.S., there appears to be evidence of full cost shifting in Canada.

Table 6 shows the results using the ACCRA price data. The data are reported quarterly and therefore 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 five percent 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 noted already. 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 estimates of the impact of minimum wage changes on hamburger and fried chicken prices. These regressions suggest a 0.12 to 0.16 percent increase in hamburger and chicken prices for every one percent increase in the minimum wage.

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

There are a number of explanations for the different price responses found in the ACCRA and BLS/StatCan data. 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 (1994), who find overshifting using the ACCRA clothing indices.[15] Second, fast food restaurants, like McDonald's and Kentucky Fried Chicken, may have more workers affected by a change in the minimum wage than restaurants in general. These establishments tend to comply with minimum wage laws and do not allow tipping, so there is likely to be a larger impact on prices. If 50 percent of workers at these fast food chains were impacted by these wage laws, the 0.15 finding would be consistent with full pass-through.

Given the noisiness of the ACCRA data, another possibility is that the larger coefficients are driven by outliers. Therefore, the regressions were rerun using a robustness technique that weights observations based on an initial regression. Observations with large residuals are assigned lower weights. Those with small residuals receive weights approaching one.[16] The results are reported in table 7. Regressions with city and year fixed effects show total elasticities ranging from 0.035 for pizza to 0.073 for hamburgers. However, note that because of the noisiness of this data, up to 500 observations receive weights of less 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 robustness regression results are in line with the findings using the more aggregated BLS and StatCan data.

Curiously, the impact on pizza prices reported in table 6 is zero or negative even after smoothing out the inexplicable spikes in the data. Part of this surprising finding is due to outliers, as shown by the results in table 7. Further experimentation suggests that much of the

[15] 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.

[16] The estimation technique calculates Huber weights and biweights (see Berk 1990 for a description). Huber weights are used as a starting value 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.

inconsequential pizza price response is driven by the April 1991 federal minimum wage increase.[17] 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 to 1990 and 1992 to 1993. In 1991, pizza and chicken inflation were *lower* in the quarters without 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 0.084, still below the hamburger and chicken price effects, but roughly in line with the full pass-through prediction. 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 for the federal increases and 0.148 for the state increases. The state pizza elasticity is roughly the same magnitude as the chicken and hamburger findings. 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. 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. The bottom of table 8 confirms Katz and Krueger's finding of small, and even negative, price responses in 1991 among hamburger and chicken (but not pizza) restaurants in nine Texas cities using the ACCRA data.[18] Furthermore, smaller April 1991 price effects also occur in the BLS

[17] Another explanation for the smaller Pizza Hut findings is that their production plan is different: labor share or the fraction of workers affected by minimum wage legislation is lower and, correspondingly, price responses are lessened. Labor share appears to be the same; American Restaurant Partners, one of the restaurants in my SEC sample and owners of 60 Pizza Huts throughout the U.S., has labor share of 29 percent in 1995 (compared to 30 percent for the entire restaurant sample). However, it is plausible that the fraction of low wage workers is different. Many Pizza Huts are sit down establishments where some employees are tipped.

[18] Two of the 11 Texas cities in the sample are missing data for the second quarter of 1991.

data. The price elasticity using the 1982 to 1995 time period is approximately 0.048, but the elasticity rises to 0.064 if 1991 data is removed.

Several other robustness checks are made of the ACCRA results. First, since there are multiple cities in each state, 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. The results are very similar to those reported in table 6. When the sample is restricted to those states in the sample 90 percent of the quarters, the three quarter sum of coefficients are 0.152, -0.062, and 0.134 for hamburgers, pizza, and chicken. Using only those states that appear in all 32 quarters does not change any inferences; the sums are 0.169, 0.001, and 0.112.

Second, I deleted cities that are on the borders of other states. Border cities could be a problem since they are under the influence of legislation from multiple states. As a result, demand elasticities may be different for border and nonborder cities if consumers can cross borders to purchase products. Furthermore, some restaurants may be influenced by the new legislation to raise prices, while others are not affected by the law and are geographically sufficiently separated enough from those that are that they do not to have to raise prices. This situation could mechanically lower price estimates even when full shifting is occurring. In the 107 city sample, there are 20 border cities but only 12 that have differences in minimum wage levels at any time between 1986 and 1993. When the equations are rerun without these 12 cities, the impact on the results is minimal. There is a slight increase in the hamburger results but the pizza and chicken findings are essentially the same.

Longer Run Estimates

Baker *et al* (1995) show that the difference between short and long run responses can reconcile different findings on minimum wage employment effects. Likewise, four months might not be enough time to capture the entire price response to the new law. Therefore, table 9 displays two alternative specifications: an unconstrained lag structure that extends the lag and lead time around the enactment month to nine months (i.e. T_1 and T_2 in equation (5) are set to 9) and a geometric lag structure that allows an infinite, weighted lag structure. The geometric lag structure is intended to estimate an equation like

$$(6) \qquad \pi^{r}_{it} = \alpha + \beta(\eta w_{it} + \eta^{2} w_{it-1} + \eta^{3} w_{it-2} ..) + \phi\pi_{it} + \varepsilon_{t} + \varepsilon_{i} + \varepsilon_{it}$$

where $\eta < 1$ is the weight assigned to the minimum wage covariates. It is easy to show that equation (6) can be rewritten as

$$(7) \qquad \begin{aligned} \pi^{r}_{it} &= \alpha(1-\eta) + \beta\eta w_{it} + \eta\pi^{r}_{it-1} + \phi\pi_{it} - \phi\eta\pi_{it-1} + u_{t} \\ u_{t} &= \varepsilon_{it} - \eta\varepsilon_{it-1} + \varepsilon_{t} - \eta\varepsilon_{t-1} + (1-\eta)\varepsilon_{i} \end{aligned}.$$

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 (7) is estimated using instrumental variables, where the instruments are w_{it-1} and the other right hand side variables.

The top panel shows the parameters from the nine month unconstrained lag structure. The BLS and ACCRA results suggest little additional impact outside the three month time frame discussed in earlier tables. If anything, the price response drops over time. Surprisingly, the Canadian results show that the long run price response is zero. While a substantial price response occurs within three months of the enactment period, prices drop before and after this period, resulting in no overall impact on Canadian restaurant prices. Furthermore, if the BLS time period is extended beyond nine months, the estimated impact converges toward zero as well. Part of the explanation for this finding may be due to the diminished precision of the estimated price elasticity as more lags are added. However, the ACCRA elasticities do not diminish when the lag structure is extended beyond nine months.

The geometric lag structure confirms that the price response in the U.S. 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; the long run price response is within the range that one would expect to see in full pass-through situations. 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 approximately 0.03 and are statistically different from 0 and 0.075, suggesting partial pass-through in the long run. The ACCRA hamburger, chicken, and pizza coefficients are 0.060, 0.074, and -0.006, respectively. The former two are different from zero and the latter is different from 0.075 at the five percent level. Therefore, the evidence in table 9 is mixed. The unconstrained and some of the geometric U.S. and Canadian CPI results suggest that price shifting dissipates over time; other findings, especially using the ACCRA data, confirm full or close to full price pass-through predictions.

<u>Effects on Broader Price Measures</u>

Table 10 reports the minimum wage coefficients from a regression using the BLS's and StatCan's CPI core and CPI for all products. As explained in section II, the predicted elasticity for broader CPI measures is approximately 0.015 to 0.020. Column (1) displays the overall BLS CPI regressions using data from 1978 to 1995. Surprisingly, the impact is about 0.04, suggesting substantial overshifting of prices. This finding may be driven by the January 1979 federal minimum wage increase, which coincided with the beginning of the OPEC oil shock. However, controlling for energy inflation makes little difference. Alternatively, I reran the regressions but started the data in 1982, well after reverberations from the oil shock and the resulting recession. This time period is also useful because city-specific core CPI prices began to be computed in 1982. Columns (2) and (3) show that this shorter period suggests smaller minimum wage elasticities. When controlling for energy inflation, the minimum wage price pass-through is approximately 0.012, not statistically different from the predicted full shifting effect. Using the core CPI rate, negative price responses are found. These results are similar when looking at longer run responses using unconstrained or geometric lag structures. The difference in the 1978-1995 and 1982-1995 findings is consistent with the assertion in Cecchetti (1986) that firms are more able to adjust prices during periods of high inflation.

The Canadian results are similar to the BLS results in some important respects. When exploring a time frame that begins in the late 1970s (columns 4 and 6), the impact on the core and overall CPI is roughly 0.010, in line with the expected impact, although much smaller than the U.S. elasticity. When the volatile early years are excluded (columns 5 and 7), the effect drops and is even negative for the overall CPI. The magnitude of these latter findings is consistent with the U.S. CPI results. However, because of the much smaller predicted effect of 0.015 to 0.02 and the size of the standard errors, it is difficult to statistically differentiate a zero and full shifting effect in the U.S. or Canadian parameters. The most that can be said about these results is that some cost shifting may take place in the overall economy, but these effects are not nearly as large as they once were in the late 1970s and early 1980s (this is also true of the restaurant sample) and statistically cannot be distinguished from a zero pass-through scenario.[19]

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. As a result, 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 the evidence on firm behavior presented in Cecchetti (1986) 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 looking at inflation patterns before the enactment of minimum wage legislation (i.e. when legislation is debated and passed). Fortunately, in the BLS and StatCan

[19] Because of the imprecision of the parameters, it is dangerous to make too much of the small pass-through found for the broader CPI measures. However, there are reasons to think that pass-through would be partial in these cases. First, restaurants tend to be local; therefore, every firm faces a similar cost structure. In national markets, firms must compete with others who do not experience an increase in minimum wages, and therefore it may be more difficult to pass-through minimum wage hikes. Second, there is little shifting of production in the restaurant industry due to minimum wage changes. For example, Card and Krueger (1995) find no minimum wage impact on McDonald's openings or closings. However, Gron and Swenson (1996) find that this is a strategy employed by international firms to avoid price increases resulting from exchange rate fluctuations. National firms that are hit hard by minimum wage changes may employ this production shifting strategy as well.

data, there is virtually no evidence that inflation is higher in the two years prior to the legislation's enactment date. The possible exception to this finding is when U.S. state increases are analyzed separately. The endogeneity problem may be more severe in the case of state legislation. However, reestimating the BLS and 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: an industry-wide increase in the price of inputs will be passed on to consumers through an increase in prices. Estimates of price shifting can have important implications for wage-push inflation stories, as well as potentially provide an explanation for the small short-run employment effects that have been found in some of the minimum wage literature. The results suggest that restaurant prices rise roughly one-for-one with increases in the wage bill that result from minimum wage legislation. Furthermore, the price responses are concentrated in the quarter surrounding the month that 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. If anything, there is some evidence that minimum wage price effects dissipate over time. The magnitude of these findings is roughly the same in the U.S. and Canada, and is fairly robust to changes in data, specification, and estimation techniques. However, because of small predicted elasticities, it is difficult to draw inferences about the price impact on broader indices or industries that have a small share of low wage labor.

Table 1
Coverage of Minimum Wage and Near-Minimum Wage Workers, by Industry and State
Mean (standard deviation) of 1986-1993 coverage[1]

	Fraction at or below minimum wage				Additional fraction within $1 of minimum wage			
	All private workers	Retail trade	Restaurant	Food service	All private workers	Retail trade	Restaurant	Food service
	(1)	(2)	(3)	(4)	(5)	(7)	(6)	(8)
A. Fraction of all employees that are low wage workers								
U.S.	5.2	12.8	**24.4**	**28.7**	10.2	27.0	**36.6**	**40.0**
	(1.0)	(3.1)	**(7.4)**	**(9.4)**	(0.6)	(1.4)	**(5.5)**	**(7.5)**
State with lowest fraction of restaurant minimum wage workers								
Alaska	2.5	3.9	6.1	7.2	4.7	14.7	26.8	29.5
	(0.5)	(1.2)	(2.1)	(2.2)	(0.9)	(2.6)	(3.7)	(4.4)
State with highest fraction of restaurant minimum wage workers								
West Virginia	10.8	26.4	45.9	52.8	13.1	32.1	37.3	38.1
	(1.7)	(5.3)	(8.9)	(9.0)	(1.0)	(3.7)	(7.6)	(7.3)
f								
B. Fraction of total hours worked by low wage workers								
U.S.	4.1	10.1	**20.7**	**25.9**	8.3	22.6	**33.2**	**38.1**
	(0.7)	(2.5)	**(6.5)**	**(8.7)**	(0.6)	(1.1)	**(4.5)**	**(6.6)**
State with lowest fraction of restaurant minimum wage hours								
Alaska	2.1	3.1	5.1	6.6	3.6	11.7	22.0	25.1
	(0.4)	(0.9)	(1.5)	(1.8)	(0.7)	(2.5)	(3.8)	(5.0)
State with highest fraction of restaurant minimum wage hours								
West Virginia	8.8	22.6	42.3	51.3	11.3	29.5	37.4	38.8
	(1.4)	(4.7)	(9.0)	(9.0)	(0.8)	(2.9)	(7.2)	(6.8)

Notes:
[1] Source: Current Population Survey, NBER Labor Extracts, 1986-1993. All descriptive statistics are weighted by the CPS final weights. Wages are defined as weekly earnings divided by weekly hours. Earnings includes tips.

Table 2
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 (1)	Average percent increase (2)	Number of increases (3)	Average percent increase (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	#REF!	39	0.0
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 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.

Table 3
Canadian Minimum Wage Increases[1]
by Province and Year, 1978-1995

	Number of increases (1)	Average percent increase (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

Notes:
[1] The province minimum wage does not include information on the federal level.

Table 4
The Imact of Minimum Wage Increases on Inflation
BLS City Price Data, 1978-1995
dependent variable: log monthly change in food away from home
minimum wage variable: log change in state or federal minimum wage 1
(Huber standard errors in parentheses)

	food away from home	food away from home	food away from home	food away from home	food away from home	food away from home
	1	2	3	4	5	6
min. wage hike (t-4)	-0.013 *	-0.013 *	-0.011	-0.011	-0.006	-0.012 *
	0.006	0.006	0.005	0.006	0.005	0.006
min. wage hike (t-3)	-0.006	-0.006	-0.007	-0.005	-0.005	-0.005
	0.005	0.005	0.005	0.005	0.005	0.005
min. wage hike (t-2)	0.008	0.008	0.005	0.006	0.007	0.008
	0.007	0.007	0.007	0.007	0.007	0.007
min. wage hike (t-1)	0.022 *	0.022 *	0.022 *	0.022 *	0.022 *	0.022 *
	0.007	0.007	0.007	0.007	0.007	0.007
min. wage hike (t)	**0.028 ***	**0.028 ***	**0.031 ***	**0.031 ***	**0.029 ***	**0.029 ***
	0.007	**0.007**	**0.007**	**0.007**	**0.007**	**0.007**
min. wage hike (t+1)	0.013 *	0.014 *	0.013 *	0.013 *	0.009	0.014 *
	0.006	0.006	0.006	0.006	0.006	0.006
min. wage hike (t+2)	-0.001	-0.001	-0.003	-0.005	-0.007	-0.002
	0.005	0.005	0.005	0.005	0.005	0.005
min. wage hike (t+3)	0.006	0.006	0.003	0.004	0.002	0.005
	0.005	0.005	0.005	0.005	0.005	0.005
min. wage hike (t+4)	0.007	0.007	0.005	0.006	0.000	0.007
	0.005	0.005	0.005	0.005	0.005	0.005
Controls 2						
Month	yes	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes	yes
City	no	yes	no	yes	yes	yes
City overall inflation	no	no	yes	no	no	no
City food at home infl	no	no	yes	no	no	no
U.S. overall inflation	no	no	no	yes	no	no
U.S. food at home infl	no	no	no	yes	no	no
U.S. specific food infl 3	no	no	no	no	yes	no
Ch. log unemployment	no	no	no	no	no	yes
Ch. in log labor force	no	no	no	no	no	yes
Adjusted R-squared	0.150	0.151	0.166	0.156	0.161	0.153
Sample size	4,486	4,486	4,483	4,486	4,486	4,486
Time period	1978-95	1978-95	1978-95	1978-95	1978-95	1978-95
Sum of coefficients						
[t-4,t+4]	0.066 *	0.067 *	0.057 *	0.062 *	0.052 *	0.065 *
[t-3,t+3]	**0.071** *	**0.072** *	**0.063** *	**0.067** *	**0.058** *	**0.071** *
[t-2,t+2]	0.070 *	0.071 *	0.067 *	0.068 *	0.060 *	0.071 *

[t-3,t]	0.052 *	0.053 *	0.050 *	0.054 *	0.054 *	0.053 *
[t,t+3]	0.047 *#	0.047 *#	0.043 *#	0.044 *#	0.034 *#	0.047 *#

Notes:
1 *=significantly different from 0 at the 5 percent level.
#=sum of coefficients are significantly different from 0.075 at the 5 percent level.
2 Price and unemployment controls include current and lagged monthly variables.
3 Specific U.S. food inflation include BLS city averages for beef, chicken, bread, potatoes, cheese, and tomat

Table 5
The Imact of Minimum Wage Increases on Inflation
Statistics Canada Province Price Data, 1978-1995
dependent variable: log monthly change in food away at restaurants
minimum wage variable: log change in provincal minimum wage[1]
(Huber standard errors in parentheses)

	food at restaurants (1)	food at restaurants (2)	food at restaurants (3)	food at restaurants (4)	food at restaurants (5)	food at restaurants (6)
min. wage hike (t-4)	-0.003	-0.003	-0.004	-0.003	-0.002	-0.003
	(0.009)	(0.009)	(0.008)	(0.009)	(0.009)	(0.009)
min. wage hike (t-3)	0.007	0.007	0.006	0.005	0.005	0.006
	(0.017)	(0.017)	(0.014)	(0.013)	(0.017)	(0.017)
min. wage hike (t-2)	0.004	0.004	0.002	-0.002	0.004	0.003
	(0.011)	(0.011)	(0.011)	(0.010)	(0.011)	(0.011)
min. wage hike (t-1)	0.011	0.011	0.010	0.013	0.012	0.011
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
min. wage hike (t)	**0.000**	**0.000**	**-0.002**	**0.005**	**-0.001**	**0.001**
	(0.010)	**(0.010)**	**(0.011)**	**(0.010)**	**(0.011)**	**(0.010)**
min. wage hike (t+1)	0.013	0.012	0.011	0.010	0.016	0.012
	(0.006)	(0.006)	(0.006)	(0.007)	(0.007)	(0.006)
min. wage hike (t+2)	0.011	0.011	0.015 *	0.018 *	0.009	0.010
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
min. wage hike (t+3)	**0.030 ***	**0.029 ***	**0.034 ***	**0.032 ***	**0.029 ***	**0.029 ***
	(0.014)	**(0.014)**	**(0.012)**	**(0.012)**	**(0.014)**	**(0.014)**
min. wage hike (t+4)	-0.004	-0.004	-0.008	-0.008	-0.005	-0.004
	(0.006)	(0.005)	(0.005)	(0.006)	(0.006)	(0.005)
Controls [2]						
Month	yes	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes	yes
Province	no	yes	no	yes	yes	yes
Province core inflation	no	no	yes	no	no	no
Prov. food at home inf	no	no	yes	no	no	no
Canada core inflation	no	no	no	yes	no	no
Canada food at home	no	no	no	yes	no	no
Can.specific food infl[3]	no	no	no	no	yes	no
Ch. log unemploymen	no	no	no	no	no	yes
Ch. in log labor force	no	no	no	no	no	yes
Adjusted R squared	0.188	0.185	0.291	0.320	0.193	0.185
Sample size	2,070	2,070	2,070	2,070	2,070	2,070
Time period	1978-95	1978-95	1978-95	1978-95	1978-95	1978-95
Sum of coefficients [1]						
[t-4,t+4]	0.069 *	0.067 *	0.064 *	0.070 *	0.066 *	0.065
[t-3,t+3]	**0.076 ***	**0.074 ***	**0.076 ***	**0.081 ***	**0.073 ***	**0.072 ***
[t-2,t+2]	0.039	0.038	0.036	0.044 *	0.039	0.037
[t-3,t]	0.022 #	0.022 #	0.016 #	0.021 #	0.020 #	0.021 #
[t,t+3]	0.054 *	0.052 *	0.058 *	0.065 *	0.053 *	0.052 *

Notes:
[1] *=significantly different from 0 at the 5 percent level.
#=sum of coefficients are significantly different from 0.075 at the 5 percent level.
[2] Price and unemployment controls include current and lagged monthly variables.
[3] Specific Canadian food inflation includes beef, chicken, potatoes, and tomatoes.

Table 6

The Imact of Minimum Wage Increases on Inflation

American Chamber of Commerce Price Data, 1986-1993

dependent variable: log quarterly change in fast food hamburger, pizza, and chicken prices 1

minimum wage variable: log change in state or federal minimum wage 2

(Huber standard errors in parentheses)

	McDonald's hamburger				Kentucky Fried Chicken chicken				Pizza Hut pizza			
	not smoothed	smoothed			not smoothed	smoothed			not smoothed	smoothed		
	1	2	3	4	5	6	7	8	9	10	11	12
min. wage hike (t-1)	0.007	0.021	0.041 *	0.012	-0.097	0.031	0.002	0.012	-0.006	0.020	-0.010	0.008
	0.028	0.017	0.020	0.020	0.053	0.022	0.026	0.026	0.030	0.021	0.024	0.025
min. wage hike (t)	**0.076 ***	**0.092 ***	**0.087 ***	**0.073 ***	**0.074**	**0.094 ***	**0.080 ***	**0.093 ***	**-0.015**	**0.005**	**-0.010**	**0.004**
	0.023	**0.018**	**0.018**	**0.022**	**0.061**	**0.028**	**0.029**	**0.034**	**0.032**	**0.021**	**0.022**	**0.027**
min. wage hike (t+1)	-0.009	0.045 *	0.039	0.062 *	-0.002	0.029	0.040	0.032	-0.060	-0.011	0.012	0.001
	0.038	0.022	0.024	0.026	0.067	0.022	0.023	0.025	0.038	0.025	0.027	0.029
<u>Controls 3</u>												
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
US core inflation	no	no	yes	no	no	no	yes	no	no	no	yes	no
US food at home infl	no	no	yes	no	no	no	yes	no	no	no	yes	no
US specific food infl	no	no	no	yes	no	no	no	yes	no	no	no	yes
Adjusted R squared	0.011	0.049	0.051	0.060	0.005	0.021	0.023	0.023	0.001	0.009	0.015	0.019
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
<u>Sum of coefficients 2</u>												
[t-1,t+1]	**0.074**	**0.158 ***	**0.167 ***	**0.147 ***	**-0.025**	**0.154 ***	**0.122 ***	**0.138 ***	**-0.081 #**	**0.014 #**	**-0.008 #**	**0.013**
[t-1,t]	0.083 *	0.113 *	0.128 *	0.085 *	-0.023	0.125 *	0.082 *	0.105 *	-0.021 #	0.025 #	-0.020 #	0.011
[t,t+1]	0.067	0.137 *	0.126 *	0.135 *	0.072	0.123 *	0.120 *	0.125 *	-0.075 #	-0.006 #	0.002 #	0.005

Notes:

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 2 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 thrown out.

2 *=significantly different from 0 at the 5 percent level. #=sum of coefficients are significantly different from 0.075 at the 5 percent level.

3 Price controls include current and lagged quarterly variables.

4 Specific U.S. food inflation include BLS city averages for beef, chicken, bread, potatoes, cheese, and tomatoes.

Table 7

Robust Regression : The Imact of Minimum Wage Increases on Inflation[1]

American Chamber of Commerce Price Data, 1986-1993

dependent variable: log quarterly change in fast food hamburger, pizza, and chicken prices

minimum wage variable: log change in state or federal minimum wage[2]

	McDonald's hamburger		KFC chicken		Pizza Hut pizza	
	(1)	(2)	(3)	(4)	(5)	(6)
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)
Controls [4]						
quarter	yes	yes	yes	yes	yes	yes
year	yes	yes	yes	yes	yes	yes
city	yes	yes	yes	yes	yes	yes
US core inflation	no	yes	no	yes	no	yes
US food at home infl	no	yes	no	yes	no	yes
Fraction of observations with a weight < 0.1[5]	7.1%	7.2%	11.3%	11.0%	18.6%	19.1%
Sum of coefficients [2,3]						
[t-1,t+1]	0.073 *	0.061 *	0.051	0.053	0.035 *#	0.032 *#
[t-1,t]	0.057 *	0.058 *	0.043 *	0.046 *	0.037 *#	0.027 *#
[t,t+1]	0.083 *	0.064 *	0.040	0.036	0.013 #	0.017 #

Notes:

[1] Regressions use robust techniques that weight observations based on the size of the residuals in a first stage regression. Price data smoothes out 1 and 2 quarter spikes. See text for more explanation.

[2] *=significantly different from 0 at the 5 percent level.

[3] #=significantly different from 0.075 at the 5 percent level.

[4] Price and unemployment controls include current and lagged monthly variables.

[5] The fraction of observations that receive a weight of 0.1 or less using Huber and biweighting functions.

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

Table 9

The Longer Run Imact of Minimum Wage Increases on Inflation

dependent variable: log monthly change in price inde

minimum wage variable: log change in state, federal, or provincial minimum wage[1]

		BLS		Statistics Canada		Chamber of Commerce		
		food away from home	food away from home	food at restaurants	food at restaurants	Hamburger	Chicken	Pizza
		(1)	(2)	(3)	(4)	(5)	(6)	(7)
Unconstrained lag structure (sum of coefs):[2]								
Monthly Data	Quarterly Data							
[t-9,t+9]		0.058	0.044	-0.024 #	-0.024 #			
[t-6,t+6]	[t-2,t+2]	0.052	0.041	-0.001	0.009	0.116	0.122	0.002
[t-3,t+3]	[t-1,t+1]	0.066 *	0.056 *	0.054	0.061 *	0.145 *	0.141 *	0.014
[t-9,t]	[t-2,t]	0.018 #	0.011 #	-0.036 #	-0.045 #	0.101 *	0.046	-0.072 #
[t,t+9]	[t,t+2]	0.067 *	0.062 *	0.009 #	0.023	0.094	0.160 *	0.072
Geometric lag structure[3]								
Lagged dep. var.		0.639 *	0.656 *	0.341	0.597	0.594	0.352	2.200
		(0.233)	(0.176)	(0.365)	(0.377)	(0.419)	(0.323)	(10.570)
min wage hike (t)		0.021 *	0.021 *	0.030 *	0.034 *	0.063 *	0.066 *	-0.039
		(0.007)	(0.006)	(0.010)	(0.011)	(0.023)	(0.031)	(0.160)
long run coef.		0.058	0.061 *	0.046	0.084	0.155	0.102	0.033
Controls[4]								
Month		yes	yes	yes	yes	yes	yes	yes
Year		yes	yes	yes	yes	yes	yes	yes
City		yes	yes	yes	yes	yes	yes	yes
National overall infl.		no	yes	no	yes	yes	yes	yes
National food inflation		no	yes	no	yes	yes	yes	yes
Time period		1978-95	1978-95	1978-95	1978-95	1986-93	1986-93	1986-93

Notes:

[1] *(#) =Significantly different from 0 (0.075) at the 5 percent level.

[2] Unconstrained lag structure allows 9 leads and lags for monthly data and 2 leads and lags for quarterly data.

[3] Geometric lag stucture 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.

[4] Price controls include current and lagged monthly or quarterly variables.

Table 10
The Imact of Minimum Wage Increases on Inflation
BLS and Statistics Canada CPI and CPI Core Price Data
dependent variable: log monthly change in CPI Core and CPI for all products
minimum wage variable: log change in state-federal or provincal minimum wage 1
(Huber standard errors in parentheses)

	US BLS			Statistics Canada			
	CPI all	CPI all	CPI core 2	CPI all	CPI all	CPI core	CPI core
	1	2	3	4	5	6	7
min. wage hike (t-4)	-0.004	0.003	0.003	0.005	-0.002	0.000	-0.009
	0.005	0.006	0.005	0.005	0.006	0.004	0.005
min. wage hike (t-3)	0.004	0.003	0.001	0.004	0.008	0.003	0.002
	0.005	0.006	0.007	0.007	0.010	0.005	0.007
min. wage hike (t-2)	0.026 *	0.021 *	0.015 *	0.000	-0.006	0.004	-0.000
	0.004	0.004	0.005	0.005	0.006	0.004	0.005
min. wage hike (t-1)	0.008	-0.000	0.003	0.003	-0.002	0.002	0.004
	0.005	0.005	0.006	0.005	0.006	0.006	0.009
min. wage hike (t)	**-0.012 ***	**-0.010 ***	**-0.011**	**0.001**	**0.000**	**0.004**	**0.008**
	0.003	**0.005**	**0.006**	**0.004**	**0.005**	**0.004**	**0.005**
min. wage hike (t+1	-0.003	-0.005	-0.004	0.001	-0.006	0.001	-0.003
	0.003	0.005	0.006	0.005	0.005	0.004	0.004
min. wage hike (t+2	0.001	-0.003	-0.013 *	-0.002	-0.007	-0.003	-0.007
	0.003	0.005	0.006	0.005	0.005	0.003	0.004
min. wage hike (t+3	0.016 *	0.007	-0.001	-0.003	-0.004	-0.003	-0.002
	0.004	0.004	0.006	0.006	0.008	0.006	0.008
min. wage hike (t+4	0.013 *	-0.002	0.007	0.003	0.002	0.006	0.007
	0.004	0.003	0.004	0.004	0.005	0.004	0.005
Controls 3							
Month	yes	yes	yes	yes	yes	yes	yes
Year	yes	yes	yes	yes	yes	yes	yes
City	yes	yes	yes	yes	yes	yes	yes
Energy inflation	yes	yes	no	yes	yes	no	no
Adjusted R-squared	0.388	0.191	0.126	0.425	0.302	0.406	0.266
Sample size	4,488	3,192	3,155	2,070	1,680	2,070	1,680
Time period	1978-1995	1982-1995	1982-1995	1978-1995	1982-1995	1978-1995	1982-1995
Sum of coefficients							
[t-4,t+4]	0.049 *	0.013	0.001	0.013	-0.017	0.014	-0.000
[t-3,t+3]	**0.040 ***	**0.012**	**-0.009**	**0.005**	**-0.016 #**	**0.008**	**0.002**
[t-2,t+2]	0.020	0.002	-0.009 #	0.004	-0.021 #	0.008	0.001
[t-3,t]	0.027 *	0.014	0.008	0.009	0.000	0.013	0.014
[t,t+3]	0.002 #	-0.012 #	-0.028 *#	-0.002 #	-0.016 #	-0.001 #	-0.004 #

Notes:
1 *=significantly different from 0 at the 5 percent level.
#=sum of coefficients are significantly different from 0.02.
2 The U.S core CPI is not available by city before 1982.
3 Price controls include current and lagged monthly variables.

Appendix 1
Descriptive Statistics on Inflation Measures 1

Dataset	Series	Mean (1)	Std. Dev (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 1986. StatCan data is monthly at the province level. There are 10 provinces. Chamber of Commerce data is quarterly at the city level. There are 107 cities 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. Sr
data eliminates temporary (less than 2 quarters) and large (> 5% quarterly cha
spikes in the Chamber of Commerce price data through linear interpolation. Sa
sizes vary between the smoothed and raw data because spikes that occur in the
two and last two quarters of the sample are thrown out.

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