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How Do Retail Prices React to Minimum Wage Increases?

A textbook consequence of an industry-wide cost shock is that it will be passed on to consumers through an increase in prices. The minimum wage offers a compelling natural experiment of such a cost shock, particularly among industries that employ low-wage labor.

We assess the effect of recent minimum wage increases on restaurant prices, using specific item prices collected by the Bureau of Labor Statistics (BLS). We find that price responses follow textbook expectations in several dimensions. First, restaurant prices rise, by amounts that are broadly consistent with the modest costs imposed by minimum wage increases. Second, prices respond rather quickly, within a six-month window around the wage increase. Third, price increases are greater among fastfood outlets and in low-wage locations, where minimum wage increases would be expected to have greater effects on costs.

But other elements of the price response are more complicated. A restaurant does not raise all of its prices by amounts reflecting the costs of minimum wage increases (from 0.3 to 1.8 percent, depending on outlet type and location). Rather, it raises fewer prices (up to 25 percent of its items), but by 3 to 6 percent, on average. That response suggests that there may be some item-specific costs to changing price, or that demand elasticities vary across items. Furthermore, we find that items at certain prices (such as fastfood items with prices ending in 99 cents) are less likely to be raised in the face of a minimum wage increase, and that outlets with recent price reviews are less likely to respond to minimum wage changes.

This pattern of restaurant price responses relates closely to notions of price stickiness. As has been known since Keynes, the dynamics of a price change can be complicated; in particular, prices may not react instantaneously to cost changes. Figure 1 exemplifies the issue; it plots monthly changes in the Food Away from Home component of the Consumer Price Index (CPI) from 1995 to 1997, along with Producer Price Indexes for two key restaurant inputs: pork and ground beef. While input prices are quite volatile, the Away from Home CPI moves slowly and methodically. Figure 1 suggests that prices may respond very slowly to cost changes, and the literature suggests conditions under which prices may not react at all. We present further evidence of sluggish price changes in our analysis below and relate those findings to other firm level results on nominal price rigidity. We then contrast the usual co-movements of prices and costs with the rapid and full response to a large, national, well-identified cost shock like the minimum wage.

Recent Analyses of Minimum Wage Effects on Prices

Recently, Card and Krueger (1995), and Aaronson (2001) analyzed the impact of the 1980s and 1990s minimum wage legislation on restaurant prices.[1] Card and Krueger (1995) used their own surveys of fast food outlets in New Jersey and Pennsylvania, and reported that prices increased in New Jersey outlets, but not in Pennsylvania's, after an increase in New Jersey's minimum wage. But within New Jersey, they did not find larger increases in areas that were more likely to be affected by the changes. They did, however, find the expected incidence at the national level when they looked at 1990 and 1991 Federal minimum wage changes. Using the Food Away from Home CPI indexes for 27 large metropolitan areas, Card and Krueger found that prices rose more, between 1989 and 1992, in those cities with higher proportions of low-wage workers in 1989. But the before-and-after nature of their comparisons meant that they could not evaluate issues of timing, and their limited data generated improbable estimates of magnitude.[2]

Figure 2 updates Aaronson's (2001) findings, based on the CPI Food Away from Home index for 1978-97.[3] The graph reports regression-adjusted price changes in the four months surrounding a minimum wage increase. The solid line captures the mean predicted change, while the dashed lines include the 95% confidence interval. Prices rise by statistically significant amounts in periods around minimum wage increases, and the increases approximate the expected cost effects of the minimum wage changes. The striking feature of the graph concerns timing: most of the price response occurs within a month or two of the minimum wage change. Despite the fact that minimum wage laws are often passed well before the enactment date, there is only a small price increase in anticipation of the "event," and no additional price reaction at long lags.

Price Data

Our dataset is the distinctive feature of our study. We use prices sampled over a three year period (January, 1995 through December, 1997) to construct the Food Away from Home component of the CPI. BLS field personnel collected prices for nearly 7,500 food items at over 1,000 different outlets. Outlets were drawn from 88 Primary Sampling Units (PSUs), which in turn included 76 Metropolitan Statistical Areas and 12 other areas representing the urban non-metro United States. PSUs were assigned to one of three reporting cycles; outlets in the five largest were surveyed each month, while others were surveyed in two bimonthly cycles of odd and even numbered months. Because most prices were collected bimonthly, we compare price changes over two month periods.[4]

We make use of a BLS "type of business" code assigned to each outlet. In "limited service" (LS) outlets meals are served for on or off premises consumption, patrons typically pay at the register before they eat, and patrons do not typically place orders while seated at a table, booth, or counter. In "full service" (FS) outlets, food is sold primarily for on premises consumption, waiter/waitress service is provided, orders are taken while patrons are seated at a table, booth or counter, and patrons typically pay after they eat. We group all other outlet types, such as schools, department stores, convenience stores, gas stations, or vending machines, into an "other" category. Full service outlets account for about half of all price quotes, while LS outlets account for about 29%.[5] LS outlets employ higher proportions of teenage and unskilled workers; thus their prices should be more sensitive to minimum wage increases.

Enumerators price multiple items at most outlets—about half of the sample's outlets have seven price quotes, and about a third have eight. Once an outlet is selected, specific items are selected for pricing with probability proportional to sales. During our 1995-97 period, an "item" most commonly was a meal, as BLS aimed to price complete meals as typically purchased at an outlet (for example, a meal item at an LS outlet might consist of a hamburger, french fries, and a soft drink). Our dataset codes items broadly, as breakfast, lunch, dinner, or snacks (corresponding to BLS "entry level item" codes). Because we did not obtain exact meal descriptions for our dataset, we can't compute meaningful measures of average prices across outlets. But BLS strives to price identical items over time, and codes in our database describe temporal item substitutions due to discontinuances and alterations. Our analysis focuses on price changes for identical items; we do not use price comparisons where one quote reflects BLS item substitution.[6]

Aaronson (2001) and Card and Krueger (1995) used published BLS price indexes for 27 large metro areas, whereas we use actual prices in 88 urban and metro areas. Consequently, we can assess more observations over a wider range of labor market conditions. We can also identify prices in LS outlets, whereas published price indexes cover all outlets. But the database has some clear limitations. Because BLS introduced a complete outlet and item resampling in January 1998, we only use data through December 1997, and cannot look for long lags in response to the September, 1997 Federal minimum wage increase. And because our dataset contains no specific item descriptions, we cannot tie price changes to item-specific measures of input price changes (such as ground beef or chicken price indexes).

Characteristics of the data

We report descriptive statistics to introduce this distinctive data, to provide support for some of our modeling decisions, and to compare restaurant pricing behavior to evidence of nominal price rigidity reported in other firm-level data sets.

<u>Frequency of price changes</u>

Prior studies of nominal pricing rigidities often find that prices remain fixed for long periods (Carlton 1986, Cecchetti 1986, Kashyap 1995). That pattern is likely to hold in restaurants, where firms often review prices at regular time intervals, and change item prices only at those intervals (Hershey, 2000). If restaurants change prices only at designated review times, then we will need to account for the possibility of lead and lagged responses to minimum wage increases.

Consider figure 3, which shows the proportion of prices in our dataset that remained unchanged in each bimonthly period. Full service meals show remarkable stability—on average, 87.4 percent of prices remained unchanged in any two month period, and that share shows little variation over time. Aside from the sharp changes immediately following minimum wage increases in 1996 and 1997, LS meal prices also appear to be quite stable; across all outlet types and bimonthly comparison periods, 86.6 percent of prices remained unchanged.

Another way to look at the incidence of price changes is to consider a price's typical life span. Duration descriptions are limited by the short sample period (with resultant left and right censoring of duration measures), but we can still offer some useful information. We looked at LS items whose prices increased just after the minimum wage increase in October 1996. Ten months later, 56% of those new prices remained unchanged. The pattern was quite similar among items that didn't change price just after the minimum wage increase: 49% of those prices remained unchanged 10 months later. The duration measures are consistent with figure 3, and the findings together suggest that the half-life of an LS restaurant price in this low-inflation period was about 10 months.[7]

<u>The distribution of price changes</u>

During months without a minimum wage increase, the mean bimonthly sample-wide price change was 0.37 percent, and varied little (from 0.36 to 0.38) across outlet types. During months with a minimum wage hike, the mean price change more than doubled in limited service outlets, rising to 0.88 percent, but rose only slightly (to 0.45) in all other outlets. We return to quantifying this minimum wage effect later in the paper.

Since most prices are unchanged in bimonthly comparisons (figure 3), sample-wide price increases must combine zero increases for most items with substantial mean price increases among those items whose prices change.[8] Table 1 reports on the size distribution of price changes. The top panel summarizes the results for limited service restaurants, while the bottom panel covers full service. Columns report data separately for price increases and decreases, and for months with and without minimum wage changes.

Several features are noteworthy. First, there are both "large" and "small" changes in price but increases are clearly skewed to the right. Among limited service outlets, roughly one-quarter of all price increases are less than 2 percent and an additional 30 percent are between 2 and 4 percent. A tail of larger increases raises the mean price increase, conditional on an increase, to 5.3 percent when minimum wages aren't increasing and 4.8 percent when they are. Similar magnitudes obtain among full service outlets: just over half of all increases are less than 4 percent, while mean increases are just under 5 percent.[9]

One-ninth of price increases exceed 10 percent, with a somewhat smaller incidence during minimum wage increases. At first glance, these wide distributions seem in contrast to some (S,s) models that predict firms will move when a threshold between optimal and actual prices is crossed. However, it is possible that the cost of changing prices varies cross-sectionally or should be modeled as a random variable as in Caballero and Engel (1999), where the cost of changing prices is sometimes low, leading to small price adjustments, and sometimes high, leading to large price adjustments.

Consistent with cross-sectional variation in (S,s) pricing, the mean size of price increases is remarkably stable over our short time frame. Figure 4 reports mean bimonthly price increases between March 1995 and December 1997, among quotes increasing price. Aside from a sharp spike among limited service outlets in the summer of 1997 related to large increases in bacon usage and pork belly prices, the mean bimonthly price increase stays within a range of 4 to 6 percent. The standard deviation of bimonthly mean price increases is 0.5 percent for full service outlets and 1.3 percent for limited service outlets, with the bacon episode accounting for half the difference in standard deviations. Furthermore, there's little change in the distribution of LS price increases when minimum wages are increasing, and later analyses find no minimum wage effect on the mean size of price increases.[10] Lastly, although price cuts are rare, they tend to be larger in magnitude than price increases (table 1). Over one quarter of price cuts exceed 10 percent, and cuts frequently exceed 20 percent. Although small price cuts occur, less than half are below 4 percent

Because mean price increases vary little from month to month, the evidence suggests that firms may respond to industry-wide cost shocks by adjusting the likelihood of price increases rather than their size. Moreover, because many price cuts reflect sales, and hence are of larger size but limited duration, we may need to explicitly account for prior price changes in our later modeling of minimum wage effects.

Across- and within-store synchronization

An important feature of modeling the inflation process is the mechanism by which prices move together. The sticky price literature generally assumes that price changes are staggered across stores but synchronized within-stores. Lach and Tsiddon (1992,1996) confirm this pattern using micro data from the food component of the Israeli CPI.

Kashyap (1995) notes that catalog price changes are not synchronized with common price shocks, such as changes in the money supply, suggesting prices are staggered across-stores. Although we do not provide any formal tests, some simple descriptive statistics suggest that staggering probably occurs both across- and within-stores in the U.S. restaurant industry.

Figure 3, which shows the monthly proportion of quotes that do not change price for the previous two months, provides evidence of across-store staggering. If stores perfectly synchronized their price changes and price durations over 10 months, we would expect to see a series of eight 0s followed by a 1. This would imply a standard deviation of 33 percent, approximately 7 to 16 times the size of the standard deviations that we see in the data. But instead, figure 3 shows remarkable stability in the share of no-change quotes. The standard deviation of this proportion is 4.8 percent for limited service outlets (2.8 percent without the minimum wage months included) and 2.0 percent for full service outlets. For this pattern to occur, and given the price durations noted above, price changes must be staggered, not synchronized, across outlets.

We looked at within–store staggering by looking at price changes among those outlets with many quotes--7 or 8 price comparisons in a bimonthly period. Since most sample outlets have 7 or 8 quotes, we still had a lot of data, 8,809 outlet-months and 64,750 bimonthly price comparisons. Table 2 sorts the data by outlet type, number of quotes, and time (whether or not a minimum wage is increasing), and for each category reports the distribution of within-outlet price increases. If perfect synchronization occurs within stores, observations should bunch at zero and the maximum number of price increases (7 or 8). Price changes do bunch at zero—most outlets change no sample prices in most months, even when minimum wages increased. But when outlets do increase prices, they clearly stagger price increases across items—there's no clear bunching at the maximum number of price increases.

Table 2 also provides some suggestive evidence of how prices change in response to minimum wage increases. First, note there is little obvious change among FS outlets, but a sharp change among LS outlets: far fewer LS outlets keep all prices stable during minimum wage changes, and far more LS outlets increase many prices.

Price points

The final noteworthy element is the existence of psychological price points (Kashyap 1995). Almost 12 percent of LS item prices end in 99 cents. Over 30 percent end in 9 (about three times random chance). Only 1 percent of full service restaurant prices end in 99 cents, but 8 percent end in 95 cents and the three most common price endings--95, 00, and 25 cents-- encompass 20 percent of all observations.

Price changes also cluster: changes of 5, 10, 20, and 30 cents account for half of all observed LS price changes (10 cents alone accounts for 26 percent), while those four, plus 25 cents, 50 cents, and a dollar, account for half of all FS price changes.

Nominal price points suggest a discontinuous threshold whereby firms will lose marginal revenue if prices are raised.[11] If price points are important, we would expect to see quote durations that are longer and price changes that are higher when starting from these thresholds. Furthermore, critical masses for price changes suggest that outlets may put off price changes until costs rise to a price changing point; if true, then price increases will lag cost increases.

Restaurant price data are generally consistent with other studies that report evidence of nominal price rigidity. First, price durations are long and price changes are rare. Second, the size of price changes varies cross-sectionally, with both small and large price increases common, but the cross-section distributions are stable over our three year period. Third, price staggering appears to occur both within- and across-stores. Finally, the use of certain price points is common and, as we show later, results in long quote durations and larger price increases. We next look more closely at how restaurants might respond to increases in minimum wages.

Measuring the Effects of Minimum Wage Changes

President Clinton signed a bill raising the minimum wage on August 20, 1996, after a debate that lasted through the previous six months. The increase took place in two stages: an October 1, 1996 increase from $4.25 an hour to $4.75 an hour (11.8%), and a second increase 11 months later, on September 1, 1997, to $5.15 (8.4%).

Price responses to the change should display temporal and geographic variations. Consider the timing of the 1996-97 Federal increase. When the law was passed (August 20, 1996) businesses knew minimum wages would be increased in 6 weeks (October 1, 1996) and again in a year (September 1, 1997). An outlet that reviewed prices quarterly might raise prices just before a minimum wage rise, or it might wait until the next review several months later.[12] Our analysis will consider lead and lag temporal effects.

Changes in the Federal minimum affect specific geographic areas in different ways, because some states maintain higher minimums, and because prevailing market wages vary widely. States with no minimum wage standards and those with minimums pegged to the Federal level experience the full change in the Federal minimum, while states with minimums above the new Federal level remain unaffected by the policy changes. Some State minimums fall between old and new Federal minimums; these states experience increases that are less in percentage terms than the first group of states. Finally, some State minimums changed during 1995-97, and businesses in those states faced minimum wage increases at times that differed from other states.

We capture varying geographic effects by defining the measure MW_{it}: the percentage increase in the effective minimum wage in state i during month t.[13] Because of State policies, percentage increases in the Federal minimum wage exceed sample mean values of MW_{it}, which is 10.7% in October, 1996, compared to the 11.8% Federal increase, and 7.0% in September, 1997, compared to the 8.4% Federal increase.

Market wages can vary widely across and within states. In some areas, prevailing low-skill wages may exceed minimum wages, and increases in the minimum wage should have little effect on market wages for low-skilled labor, and consequently little effect on costs and prices. In other regions, minimum wages may exceed market wages, and changes in the minimum wage will have stronger effects on observed wages, costs and prices. We expect increases in effective minimum wages to have greater impacts on costs and hence on prices in low-wage areas.

Our database includes unique outlet codes and precise outlet locations (addresses, zip codes, and telephone numbers) so we can link items to outlets and we can link outlet records to related geographic information. We use data from the Current Population Survey (CPS) to summarize hourly wage distributions in the outlet's Metropolitan Statistical Area (MSA).[14] We use observations for all of 1996 for each MSA, and calculate the distribution of MSA hourly wages. We use the 20th percentile hourly wage (WAGE20) as our measure of low-skill wages in a metro area.

Model and Basic Empirical Results

Our basic statistical model can be summarized as follows:

$$1) \qquad \ln (P_{kj,t}/P_{kj,t-2}) = f(\text{PPI, IPRICE, MEALTYPE, MW}).$$

$P_{kj,t}$ is the price of the item k at outlet j in month t. The dependent variable is therefore the percentage change in price over a bimonthly period. Table 3 details the explanatory variables. PPI, a vector of two-month percentage changes in the Producer Price Index for Processed Foods, measures input price shocks faced by sample outlets, and specific measures include contemporaneous as well as one- and two-period lags. The vector IPRICE captures some dynamics of changes in the item's price over time, reflecting pricing strategies as well as seasonal price patterns in restaurants. Some sale prices fall by substantial amounts in our sample (table 1), and are likely to increase substantially some time after the promotion. Similarly, items with price increases frequently decline again at some later period, because the original increase may have reflected seasonal cost increases or because rivals didn't match an increase. In order to capture these dynamics, we enter a variable, IPUP, equal to the percentage increase in the item's price in a previous period, or zero if the price did not increase. Similarly, IPDOWN is the percentage decrease in price in a previous period, or zero if there was no decrease. We enter one-, two- and three-period lags on the item price variables.[15]

The BLS uses ELI (entry level item) codes in item selection, and Food Away from Home codes identify dinner, lunch, snack, and breakfast/brunch items. We dropped snack observations (a small part of the database), since they were disproportionately offered through unusual outlet types, like vending machines or gas stations. The model includes dummy variables for lunch and breakfast meals, with dinner as the base.

Finally, MW is the change (increase) in the effective minimum wage in period t at outlet j. We also use lead and lag measures of MW to assess price effects one period before and one period after periods including minimum wage increases. We found no evidence of longer leads or lags, either with these two month periods or with analyses of those prices at outlets that are surveyed monthly.

We present the basic analysis of price effects in Table 3, using all outlets. BLS generally collected several prices at an outlet in 1995-97 (the mode is 7); because quotes from the same outlet are unlikely to be statistically independent, standard error calculations in all analyses account for quote clustering, using Huber-White robust estimation techniques. Adjustment matters here, as robust standard errors are generally two to three times larger than OLS standard errors.

Item price effects are important and statistically significant in table 3; prior price cuts lead to current period price increases, and past price increases indicate current price cuts (responses are dampened, however; full reversion to a previous price would imply coefficient values, in absolute terms, of 1, while these fall well below 1 and usually below 0.1). PPI effects are small--food is 20 percent of outlet costs, and outlets purchase many different combinations of food products--while unreported meal type effects were small and not statistically significant.

We multiply all reported minimum wage coefficients and standard errors by 10, to save space (results should therefore be read as the effects of 10 percent, rather than 1 percent, minimum wage increases). In equation 1 of table 3, the minimum wage effect is positive and highly significant (t=5.09), suggesting that outlet prices rise by 0.33 percent, on average, for contemporaneous 10 percent increases in the minimum wage. Equation (2) adds lead and lag effects for the minimum wage change ($MW_{i,t-1}$ measures price responses to prior period minimum wage increases while $MW_{i,t+1}$ tests for price responses to next period minimum wage increases). Each effect is positive; the one period lag effect is highly significant (t=3.18), while the one period lead is significantly greater than zero at the 90% confidence level (t=1.72). Lead and lag effects raise the estimate on the contemporaneous effect, and the full effect is more than double that in equation 1.[16]

The Incidence of Minimum Wage Price Effects

Table 4 examines minimum wage effects in more detail, with models that account for differences in outlet types and in prevailing area wages. The table reports minimum wage effects, but the models retain all other explanatory variables in table 3.[17]

Equations 1-3 explore outlet type effects, running regressions for full (equation 1) and limited service outlets (equations 2 and 3). Full service minimum wage effects are statistically significant but quite small in a specification that mirrors table 3's equation 2. In contrast, the effects in limited service outlets are much larger (equations 2 and 3). The coefficient on MW_{it} in equation (2), positive and highly significant (t=6.44), suggests that LS prices rise by 0.8 percent in response to contemporaneous 10 percent increases in minimum wages (four times the FS effect). Equation (3) lead and lag effects are positive, large, and statistically significant (t=2.25 for the lead effect and t=2.56 for the lag effect).[18] Again, inclusion of lead and lag effects raises the contemporaneous estimate.

Equations 4-6 add an interaction term between MW_{it} and WAGE20, the hourly wage at the 20th percentile of an area's wage distribution. High values of WAGE20 should indicate high wage areas, which may be less strongly affected by changes in the minimum wage. In the all-outlet sample (equation 4) the main MW coefficient remains positive and highly significant, while the interaction term is negative, fairly large, and statistically significant (t=2.07). Minimum wage increases have larger effects on prices in low wage areas. Among full service outlets (equation 5), the coefficients on MW and the interaction term are of the expected sign but are only marginally significant.

Equation (6) adds the interaction term to the LS sample. As in the all-outlet sample, LS wage effects are negative, substantive, and statistically significant (t=2.09). Price effects are larger in low wage areas. Finally, equation (7) adds interaction terms between WAGE20 and the lead and lag MW effects. Inclusion of those measures has no effect on the contemporaneous effects. The lagged effects have the expected sign (positive on $MW_{i,t-1}$ and negative on the interaction) but are not quite significant. The lead effects lose all power, suggesting that there is no interaction and that inclusion of it simply created multicollinearity with the main effect.

How Large are Minimum Wage Price Effects?

So far, we've found that restaurant prices increased when minimum wages rose, that increases were larger among limited service outlets and in low-wage areas, and that the response occurred in a six month window around the wage increase. The patterns suggest that minimum wages affect outlet costs in different ways, and that price changes vary in response to those different cost effects. We have no information on factor shares for the outlets in our sample, and no data that would allow for estimation of factor-substitution possibilities. As a result, we can't directly estimate the effects of minimum wages on outlet costs, and therefore cannot directly identify the extent to which costs are passed through in the form of higher prices.

Instead, we rely on Lee and O'Roark's (1999) calculations of likely pass-through under various assumptions about factor shares, substitution possibilities, and firm behavior. They used 1992 Current Population Survey wage and employment data, and U.S. Input-Output tables, to calculate the direct and indirect (through purchases) shares of low wage labor in "eating and drinking places" (SIC 581). Assuming full pass-through of costs to prices, no substitution, and no spillovers of minimum wage increases to other wages, a 10 percent minimum wage increase would lead to a 0.74 percent price increase among eating and drinking places in Lee and O'Roark's model (table 5).

Eating and drinking places include bars that serve no food, and our transaction-based sample includes supermarkets that serve food eaten on premises, schools, gas stations, and the like. But there's still a wide overlap between the Food Away from Home all-outlet sample and SIC 581, and there's a strikingly close correspondence between the Lee-O'Roark estimate of the full pass-through price increase and our estimated price increase for the all-outlet sample—0.74 vs. 0.73 in table 5. We base our calculation on the sum of the lead, lag and contemporaneous coefficients, the most comparable approach.[19] Further comparison can be found in Aaronson (2000) who used aggregated

1978-95 BLS indexes to estimate a long-run price response of 0.72, quite close to our 0.73 for disaggregated 1995-97 BLS data.

Table 5 also reports price effects by outlet type and by prevailing wage level. In each case, we report short tun (current period) estimates as well as longer run estimates that add lead and lag effects. LS price effects are more than twice as large as the all-outlet estimate—a 1.56 percent price increase in response to a 10 percent minimum wage increase, while FS price effects are much lower, 0.40 percent if we exclude the lead effect (negative and not significant). In turn, LS effects are larger still in low wage areas—1.83 percent when wages of 5.50 an hour set the area 20th percentile.

How Do Outlets Raise Prices?

We've seen that prices at LS outlets rise by nearly 1.6 percent in response to a 10 percent increase in minimum wages. Restaurants can arrange that increase in a variety of ways. They could raise all prices by 1.6 percent, or they could raise fewer prices by greater amounts. If they choose the latter, they've got to decide which item prices to increase. We pursue that question in this section.

We begin by assessing the incidence of price increases, rather than the average price response to minimum wage change. Table 6 reports minimum wage effects for the LS sample, based on logit estimates of the model used throughout, but with the dependent variable now a dummy variable set equal to one for price increases.

The contemporaneous minimum wage effect is positive and highly significant (t=3.67) in equation 1, while the lagged effect is positive and marginally significant (t=1.76) and the lead effect is small and not significant. The wage interaction is statistically significant (t=2.29) and works in the expected direction, with a smaller incidence of price increases, following a minimum wage increase, in high wage areas.

In equation (2), we investigate the linkage between discrete restaurant price reviews and the formation of leads and lags in price responses. Most outlets change no sampled prices in most months: that decision may reflect the underlying economics, but it may also reflect periodic rather than continuous price reviews. If price reviews are periodic, then outlets that have recently changed prices may not respond to a minimum wage increase until another price review.

We define a variable P_REV, and set it equal to one if the outlet changed price for any sampled item in the previous period (that is, prices were reviewed in the previous period). We then interacted P_REV with each of our MW terms. The results are striking. Minimum wage changes are less likely to lead to a price increase if the outlet had reviewed prices in the period prior to the increase. In the lagged period, firms are considerably more likely to raise price if the outlet raised price just after the minimum wage changes.[20]

Equations 3-6 explore the role of pricing points. We define a variable PP99 and set it equal to one if the LS item price at the beginning of the period ended in 99 cents. The coefficient on PP99 in equation (3) is negative and quite significant—firms are considerably less likely to raise a price ending in 99 cents. Equation (4) interacts PP99 with our minimum wage variable, and the results indicates that the logit coefficient on PP99 changes very little when minimum wages increase. We add an interaction with the actual value of the beginning period price in equation (5); the marginally significant price suggests that restaurants are particularly unlikely to raise prices on low priced 99 cent items (priced at 0.99 or 1.99). The last equation combines the PP99 and P_REV measures and finds that results are robust to the inclusion of all.[21]

Because the coefficients in logit models can't be directly interpreted, table 7 reports the predicted probabilities of a price increase under different conditions. We set the base probability of a price increase to that predicted on the basis of the intercept (no changes in PPI or IPRICE measures over the previous several periods). The base probability of a price increase is 10.3 percent for items that aren't price points, and 10 percent minimum wage increases have large contemporaneous effects, raising probabilities to 23 and 33 percent in high and low wage areas, respectively. Probabilities decline noticeably if the outlet had just had a prior period price review, to 16 and 24 percent, respectively. Lagged effects are of interest; if the outlet changed no prices in the immediate wake of a minimum wage increase, the probability of a lagged increase is quite small, 12.9 percent. But for outlets that responded immediately with price increases on some items then there's a substantial lagged possibility (21 percent) of a price increase. Price point effects are quite large. The probability that an item's price will be increased if price ends in 99 cents falls by 35 percent, to 6.7 percent, in the base case. While probabilities rise after a minimum wage increase, price point items remain substantially more stable than other items.

Table 8 considers the second aspect of price construction—how much are item prices increased, when outlets decide to raise price? The sample is restricted to LS items with price increases and the dependent variable is the log difference in prices. Equations (1) and (2) support the cruder inferences drawn on figure 4 and table 2: minimum wage effects have virtually no effects on the size of price increases, as the coefficients are small and not nearly significant. As table 2 suggests, firms respond to minimum wage increases by raising more prices, not by raising prices more.

In equation (3), the coefficient on PP99 is positive and significant; when firms raise prices on price point items, they raise them by more, although the coefficient is not that large (suggesting that PP99 item prices rise by 7.2 percent, on average). When we interact PP99 with its actual beginning price, we find that typical percentage price increases are larger among low-priced items (about 8.9% for 99 cent items, and 8.3% for $1.99 items, versus 5% for $6.99 items), when they are increased. Finally, we can see that P_REV plays a role, too; price increases are somewhat smaller for those items with prior month price reviews.[22]

Discussion

Food prices respond quickly to minimum wage increases; most of the observed response (about 60%) occurs in the two-month period immediately after a minimum wage increase, while the rest occurs in the periods two months after or two months immediately preceding the wage increase. The timing is close to that observed by Aaronson using aggregated BLS data over the 1978-95 period; he found that about 56% of the price increase occurred in the two months immediately after the increase. Actual price changes could occur even more quickly than our estimations suggest, if they occur early in the lead and lag periods (footnote 18). Since other evidence indicates that restaurant prices respond only slowly to cost changes, the minimum wage response suggests that prices will respond quickly to common and permanent industry cost shocks.

Prices rose by amounts that were predicted on the basis of full pass-through of costs. While the finding does not prove that costs are fully passed through (wage spillovers, along with less than full pass-through, would give the same result), it does provide greater confidence for the results. Moreover, the incidence of increases accords with what we know of the low-wage labor market. Prices rose far more in LS outlets than in other restaurants, reflecting greater reliance on low wage labor in the former, and prices rose more in low wage areas than in high wage areas.

Our paper's title implies an interest in the size, speed, and incidence of price increases following minimum wage increases, but it also connotes an interest in the mechanics of how restaurant increase prices. Our results imply that prices in limited service restaurants rise by 1.56 percent, on average, in response to a 10 percent minimum wage increase (and by 0.73 percent in the all-outlet sample). But outlets don't construct that response by raising all prices 1.56 percent; instead they increase price by about 5 percent, on average, on a much smaller set of items. That pattern suggests one source of stickiness, item-specific costs of changing prices. We found two other specific sources, a marked reluctance to change price on items with prices at a psychological pricing point (ending in 99 cents), and a temporally discrete approach to price review.

Despite evidence of stickiness, the rather small cost increases associated with minimum wage increases did generate a quick aggregate price response. As a widely publicized and permanent industry-wide cost shock, we suspect that minimum wage changes do not generate the sort of coordination failures that other cost or demand changes produce.

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Table 1: Bimonthly Price Changes, by Outlet Type and Minimum Wage Change

	(1)	(2)	(3)	(4)
Month w/ Min Wage ↑?	No	Yes	No	Yes
Type of Price Change	Increase	Increase	Decrease	Decrease
Limited Service Outlets				
% of Monthly Quotes	11.4%	22.4%	2.7%	2.6%
Mean Price Change	5.31%	4.83%	8.36%	8.19%
Incidence of Price Changes		--Share of Quotes (%)--		
0-2%	25.5	22.8	29.3	24.5
2-4%	30.6	34.1	18.3	15.3
4-6%	16.8	18.9	11.8	15.3
6-8%	9.4	8.0	7.0	11.2
8-10%	6.3	6.7	4.6	6.1
10-20%	7.7	7.9	16.2	18.4
>20%	3.8	1.5	12.8	9.2
Full Service Outlets				
% of Monthly Quotes	10.8%	11.2%	1.7%	1.7%
Mean Price Change	4.76%	4.88%	7.45%	9.29%
Incidence of Price Changes		--Share of Quotes (%)--		
0-2%	32.7	26.2	25.3	21.4
2-4%	25.1	27.6	19.8	14.3
4-6%	15.3	18.8	12.2	8.9
6-8%	10.1	10.7	11.3	9.8
8-10%	6.2	7.1	7.4	12.5
10-20%	9.2	8.4	17.1	21.4
>20%	1.6	1.2	6.8	11.6

Notes: A month with a minimum wage increase is one immediately (within two months) following any state or federal change that effectively raised the minimum wage in a state.

Table 2: Incidence of Price Increases Among Outlets with Many Quotes

	(1)	(2)	(3)	(4)
Price Quotes at Outlet	7	7	8	8
Minimum Wage ↑ ?	No	Yes	No	Yes
Limited Service Outlets				
Outlet-Months	1852	287	1070	162
Bimonthly Price Pairs	12,964	2,009	8,568	1,296
Number of Price Increases	-Share of Outlet-Months (%)-			
0	77.1	60.3	67.5	59.3
1	7.6	6.6	12.1	14.2
2	3.3	5.6	5.7	4.9
3-5	6.7	12.5	9.7	11.1
6	1.2	4.5	2.1	1.2
7	4.2	10.5	1.7	5.6
8			1.1	3.7
Full Service Outlets				
Outlet-Months	3086	505	1,578	269
Bimonthly Price Pairs	21,602	3,535	12,624	2,152
Number of Price Increases	-Share of Outlet-Months (%)-			
0	80.3	81.2	81.0	78.1
1	5.7	5.3	5.6	5.9
2	2.4	3.0	2.7	2.2
3-5	6.0	5.8	4.6	6.6
6	1.8	0.8	1.1	1.9
7	3.8	4.0	2.0	1.9
8			3.0	3.3

Notes: Months with increasing minimum wages are October and November, 1996, and September and October, 1997.

Table 3: Magnitude and Timing of Price Responses to Minimum Wage Increase

Variable	Description	Coefficients (s.e) (1)	(2)
Intercept		0.379 (0.026)	0.315 (0.032)
PPI_t	% change in the Producer Price Index for Processed Foods, period t	0.050 (0.029)	0.070 (0.030)
PPI_{t-1}	% PPI change, period t-1	-0.070 (0.037)	-0.040 (0.038)
PPI_{t-2}	% PPI change, t-2	0.061 (0.028)	0.046 (0.030)
$IPUP_{t-1}$	% increase in item price, one period lag; zero if no increase	-0.089 (0.013)	-0.090 (0.013)
$IPDOWN_{t-1}$	% decrease in item price, one period lag; zero if no decrease	0.388 (0.045)	0.389 (0.045)
$IPUP_{t-2}$	% increase in item price, two period lag; zero if no increase	-0.070 (0.012)	-0.070 (0.012)
$IPDOWN_{t-2}$	% decrease in item price, two period lag; zero if no decrease	0.096 (0.023)	0.096 (0.023)
$IPUP_{t-3}$	% increase in item price, three period lag; zero if no increase	-0.036 (0.010)	-0.036 (0.009)
$IPDOWN_{t-3}$	% decrease in item price, three period lag; zero if no decrease	0.060 (0.030)	0.060 (0.030)
MW_{it}	% change in minimum wage in state i, period t.	0.331 (0.065)	0.406 (0.063)
$MW_{i,t-1}$	% minimum wage change, state i, period t-1.		0.207 (0.006)
$MW_{i,t+1}$	% minimum wage change, state i, period t+1		0.115 (0.007)
R^2		0.065	0.066
N		68,887	68,887

Note: equations also included categorical meal type variables

Table 4: Incidence of Price Response to 10% Minimum Wage Increase

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Outlet Type	FS	LS	LS	All	FS	LS	LS
Variables				Coefficients and standard errors			
MW_{it}	0.191	0.805	0.937	1.449	1.048	2.698	2.696
	(0.086)	(0.125)	(0.134)	(0.531)	(0.692)	(0.883)	(0.883)
$MW_{i,t-1}$	0.212		0.305	0.200	0.227	0.263	1.222
	(0.083)		(0.119)	(0.068)	(0.088)	(0.121)	(0.797)
$MW_{i,t+1}$	-0.062		0.319	0.124	-0.041	0.296	-0.262
	(0.075)		(0.142)	(0.070)	(0.080)	(0.155)	(0.810)
MW_{it}*WAGE20				-0.162	-0.129	-0.278	-0.278
				(0.078)	(0.099)	(0.133)	(0.133)
$MW_{i,t-1}$*WAGE20							-0.147
							(0.120)
$MW_{i,t+1}$*WAGE20							0.086
							(0.128)
N	35,759	21,064	21,064	61,716	32,822	18,024	18,024
R^2	.017	.151	.152	.068	.018	.164	.165

Notes: Each regression also includes other variables listed in table 1. FS refers to full service outlets, while LS refers to limited service outlets. $MW_{i,t-1}$ and $MW_{i,t+1}$ refer to 1 period before and 1 period after minimum wage changes. All standard error estimates are adjusted for nonindependence within outlets.

Table 5: Comparing estimates of the price effect of a 10% minimum wage increase

Authors	Scope of estimate	% Price Rise
Lee & O'Roark	All outlets	0.74
Aaronson	All outlets	0.72
MacDonald & Aaronson	All outlets (lr, sr)	0.73, 0.41
	Full Service (lr, sr)	0.34, 0.19
"	Limited Service (lr, sr)	1.56, 0.94
"	Limited Service, high wage (lr, sr)	1.31, 0.65
"	Limited Service, low wage (lr, sr)	1.83, 1.17

Notes: 'lr' refers to long run response, while 'sr' refers to short run (one period) response. Hi wage cities have WAGE20 equal to 7.37 an hour, while low wage cities have WAGE20 equal to 5.50 an hour (median of cities in top and bottom quartiles, respectively.

Table 6: Logit Model Coefficients: Probability of Price Increase, Limited Service

Variables	Coefficients and standard errors					
	(1)	(2)	(3)	(4)	(5)	(6)
MW_{it}	2.939	2.983	2.854	2.852	2.841	2.896
	(0.800)	(0.802)	(0.799)	(0.799)	(0.798)	(0.801)
$MW_{i,t-1}$	0.313	0.220	0.308	0.308	0.308	0.212
	(0.178)	(0.193)	(0.179)	(0.179)	(0.179)	(.194)
$MW_{i,t+1}$	0.055	0.224	0.048	0.048	0.049	0.214
	(0.176)	(0.206)	(0.176)	(0.176)	(0.176)	(0.205)
MW_{it}*WAGE20	-0.284	-0.273	-0.272	-.270	-0.269	-0.260
	(0.124)	(0.125)	(0.125)	(0.125)	(0.124)	(0.124)
MW_{it}*P_REV		-0.496				-0.482
		(0.278)				(0.279)
$MW_{i,t-1}$*P_REV		0.616				0.632
		(0.313)				(0.301)
$MW_{i,t+1}$*P_REV		-0.461				-0.454
		(0.304)				(0.305)
PP99			-0.485	-0.464	-0.726	-0.482
			(0.129)	(0.137)	(0.204)	(0.130)
PP99*MW_{it}				-0.098		
				(0.287)		
PP99*P1					0.062	
					(0.036)	

Notes: Standard errors are in parentheses. Estimates drawn from logit estimation; full results reported in Appendix. Dependent variable is dummy variable equal to one if price increased in period. MW^{+1} and MW^{-1} refer to 1 period before and 1 period after minimum wage change. P_REV is a dummy variable equal to 1 if any outlet prices were changed in the previous period. PP99 is a dummy variable equal to 1 for items with beginning prices ending in .99. P1 is the beginning period price. WAGE20 is the 20th percentile wage in the sampling area.

Table 7: Predicted probabilities that an item's price will increase

Events	(1)	(2)	(3)	(4)
Price Review Last Period?	No	Yes	No	Yes
P1 ends in 99 cents?	No	No	Yes	Yes
	Predicted Probability of Price Increase (%)			
Base Probability	10.3	10.3	6.7	6.7
10% Minimum Wage Increase				
In high wage area	23.4	15.9	15.9	10.5
In low wage area	33.1	23.5	23.5	15.9
1 Period After Increase	12.4	21.0	8.0	14.1
1 Period Before Increase	12.4	8.2	8.0	5.2

Notes: Predicted probabilities from logit model of price increase. Full model reported in appendix table 1. Low wages are set to \$5.50 an hour; high wages are set to \$7.37.

Table 8: Minimum Wage Price Effects, Among Items that Raised Price, Limited Service

Variables	Coefficients and standard errors				
	(1)	(2)	(3)	(4)	(5)
MW_{it}	0.020	0.750	0.909	0.837	0.646
	(0.338)	(1.966)	(1.977)	(1.965)	(1.958)
$MW_{i,t-1}$	-0.007	-0.052	-0.076	-0.114	-0.148
	(0.431	(0.494)	(0.495)	(0.493)	(0.495)
$MW_{i,t+1}$	0.844	0.926	0.918	0.876	0.940
	(0.589)	(0.696)	(0.700)	(0.702)	(0.696)
MW_{it}*WAGE20		-0.095	-0.123	-0.107	-0.084
		(0.310)	(0.312)	(0.309)	(0.306)
P_REV					-0.691
					(0.303)
PP99			1.712	4.083	4.195
			(0.448)	(0.924)	(0.930)
PP99*P1				-0.592	-0.620
				(0.165)	(0.170)
R^2	0.460	0.476	0.483	0.487	0.489
N	2841	2,316	2,316	2,316	2,316

Notes: All equations also include other variables listed in table 1. MW^{+1} and MW^{-1} refer to 1 period before and 1 period after minimum wage change. P_REV is a dummy variable equal to 1 if any outlet prices were changed in the previous period. PP99 is a dummy variable equal to 1 for items with beginning prices ending in .99. P1 is the beginning period price. WAGE20 is the 20^{th} percentile wage in the sampling area.

Figure 1: Comparing Changes in the Away From Home CPI and Meat PPIs

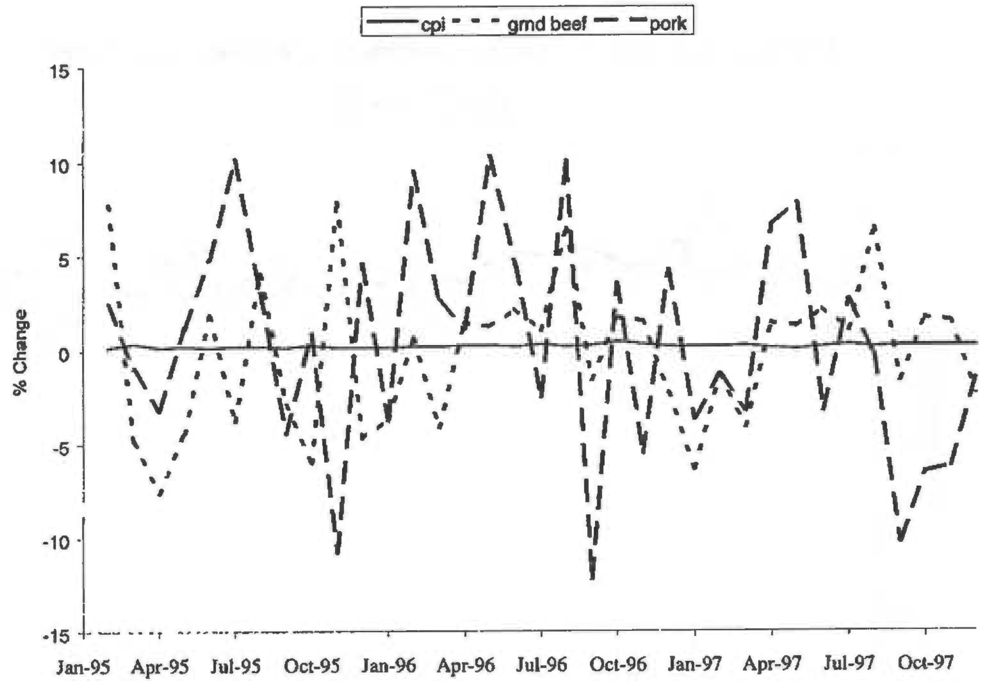


Figure 2: Response of Food Away from Home CPI to Minimum Wage Increase, 1978-97

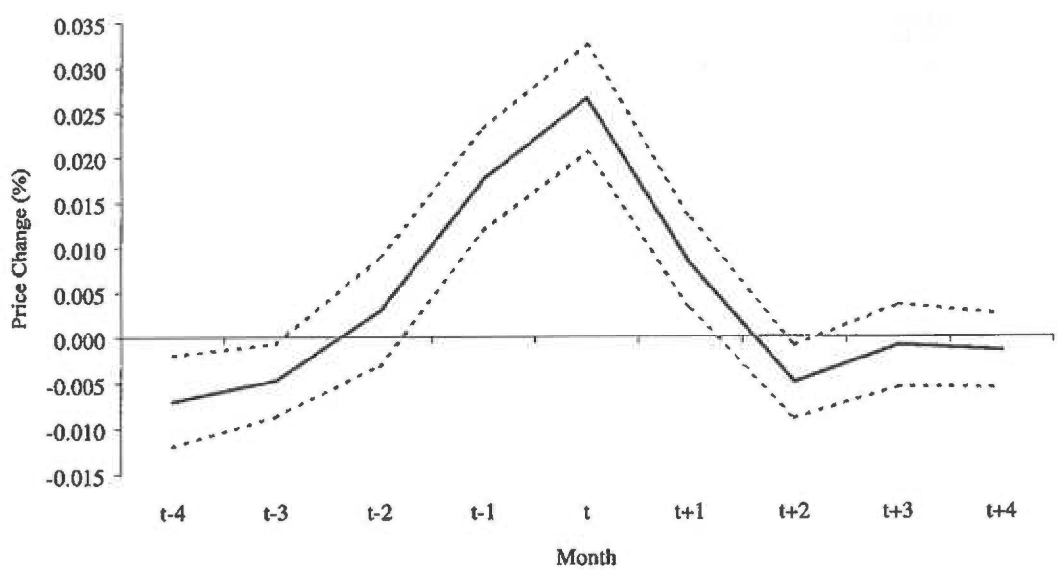


Figure 3: Percent of Items with Unchanged Price, Full and Limited Service Outlets

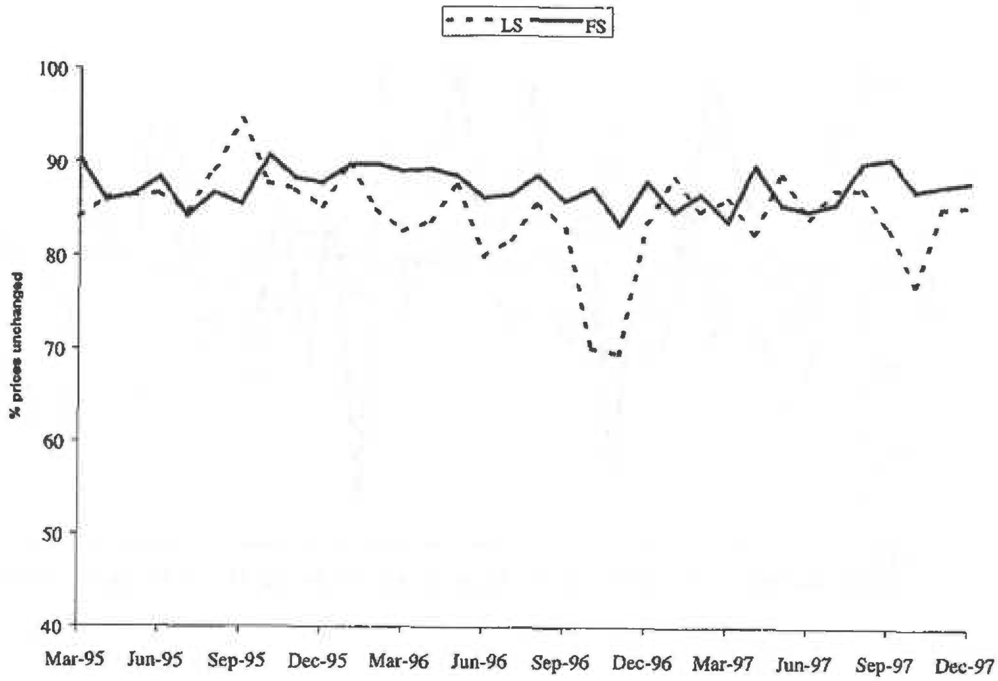
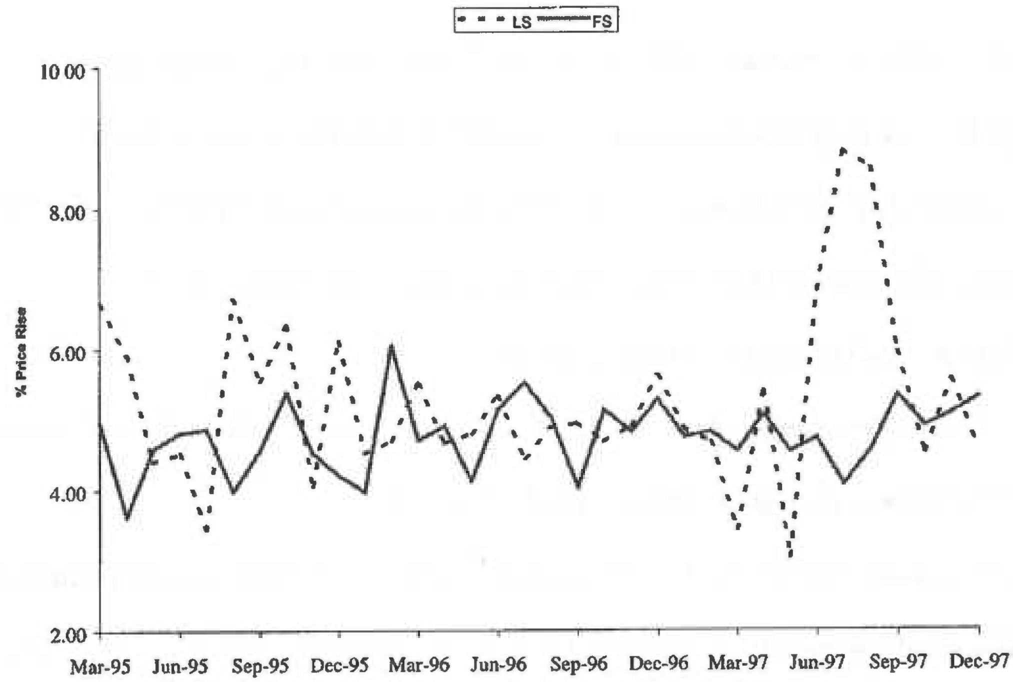


Figure 4: Mean Price Increases, Full and Limited Service Outlets



[1] Price responses are critical for measuring the welfare implications of minimum wage legislation.

[2] They estimated price increases of 3.3% for a 10% increase in minimum wages, suggesting that low wage labor accounted for one third of restaurant costs, if cost increases were fully passed through to prices. But total payroll accounts for less than 30% of restaurant sales, according to Retail Census data. Low wage labor more likely accounts for one sixth of costs (Aaronson, 2000).

[3] The calculations, presented in Aaronson (2000), are based on the published monthly and bimonthly Food Away from Home indexes for 27 U.S. cities.

[4] To maintain consistency in our price analyses, we randomly assigned outlets in the five largest PSUs to odd or even two month cycles. Counts are based on all items for which prices were obtained in October or November, 1996, just after the October 1, 1996 increase in the Federal minimum wage. A more complete description of the sample's construction, and particularly of outlet and item selection procedures, can be found in the BLS *Handbook of Methods*.

[5] BLS replaced an old ordering with these type of business codes in July, 1996, and began to report Food Away from Home price indexes for type of business groupings in January, 1998. Businesses surveyed early in our sample period were assigned these codes retroactively.

[6] It's possible that firms could respond to a minimum wage increase by reducing quality, instead of raising price. We don't have measures of quality, but our dataset notes whether an item is the same as the item priced in the previous month, or whether BLS has had to

substitute another item. The data show no evidence of any increased incidence of item substitutions following minimum wage increases.

[7] The measures are consistent in the following sense: if 93.3 percent of prices were unchanged each month (implied by 86.6 percent bimonthly stability), and if the probability of change was independent of the length of time that the price had been unchanged, then 10 months later we would expect 50 percent of prices ($1-0.933^{10}$) to be unchanged. With the ten month survival rates so close to 50 percent, the probability of price change does appear to be independent of the length of time that it had been fixed.

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

[9] By comparison, Kashyap (1995) reports that more than 20 percent of the price increases in his study of catalog prices are less than 3 percent. His study looks at changes over six month intervals rather than the two month intervals that we report.

[10] We applied Kolmogorov-Smirnov D-tests for differences in the price distributions in table 1. We found no significant differences in price decreases in months with minimum wage increases compared to other months, and no significant differences in the distribution of increases among LS outlets. FS outlets with price increases do show a statistically significant shift, driven by the higher incidence of 0-2% increases in months without minimum wage increases.

[11] Price points may be connected to coordination failures. In markets with a limited number of sellers, prices are strategic complements. A seller may be reluctant to raise a price unless they're sure that other sellers will also raise price. That uncertainty may be particularly pronounced for items with price points, if consumers are particularly sensitive to changes in those prices.

[12] Firms clearly had more foreknowledge of the 1997 increase. The 1996 increase could not have been predicted until shortly before the House of Representatives vote on May 23, after a week of legislative maneuvering that almost consigned the bill to defeat (Weisman, 1996; Rubin, 1996). Even then, the final timing didn't become clear until adoption of the conference report on August 2. Businesses therefore knew of the 1996 increase 2-4 months prior to implementation, while they knew of the 1997 increase, which was specified in the 1996 bill, 12-13 months before implementation.

[13] During 1995-97, 12 states had minimum wages above the Federal standard. We summarize state-specific minimum wage changes with information from the *Monthly Labor Review's* annual (January issue) surveys of state labor legislation. If the effective minimum wage in a state increased by 10% on October 1, then MW will equal 0.10 for August-October price comparisons and for September-November price comparisons (and 0.0 for other comparisons).

[14] For the 12 non-metro urban areas, we use wage data for the non-metro parts of the outlet's state. CPS codes are unavailable for 9 MSAs, so sample sizes decline when area wage data are included in the analysis.

[15] Adding lags reduces sample size. Three period lags (up to 6 months, with two month periods) were always statistically significant, while a fourth period was not.

[16] This pattern closely matches that found by Aaronson using aggregated data for 1978-95. He found a sharp price spike at the minimum wage increase, statistically significant individual month effects at the one month lead and lag, and a full effect captured in a window of six months surrounding the minimum wage increase (three before and three after). Our two month contemporaneous, lead, and lag periods sum to six months.

[17] Estimated minimum wage effects are quite robust to the inclusion or exclusion of the other explanatory variables, and the pattern of coefficients on the other explanatory variables changes little as the model changes to capture different minimum wage effects. As a final check, we estimated models with fixed firm and PSU effects. Key minimum wage coefficients were unaffected by the inclusion of fixed effects, and the fixed effects themselves added almost nothing to the model's fit.

[18] BLS enumerators collect prices in 3 roughly week-long collection periods during the first 22 days of a month. Outlets are visited during the same collection period each month, although not necessarily on the same day each month. Therefore, prices observed at "one period lags" after the October 1, 1996 minimum wage increase include, at the near minimum, comparisons made for the first week of December compared to the first week of October and, at the far maximum, comparisons for the third week of January compared to the third week of November. One period lead effects include, at the near minimum, comparisons of the third week of September to the third week of July and, at the far maximum, comparisons of the first week of August to the first week of June. In short, some "lead" and "lagged" price changes could have been made quite close to the date of the minimum wage change, as little as 10 days before or after.

[19] While our findings are consistent with full pass-through, they are also consistent with less than full pass-through, coinciding with spillovers.

[20] We also looked at outlet price increases two periods before a minimum wage increase, but adding another period provided no new information for the analysis. The results suggest the following sequence. Outlets that reviewed prices two periods before a minimum wage increase did not then review again in anticipation of the minimum wage change. Those that did raise price in anticipation of the minimum wage increase (one period before) had only modest contemporaneous price increases, while those without anticipatory price increases before the minimum wage change raised prices substantially.

[21] We looked at other LS price points. About 3 percent of prices end in 98 cents, and over half end in the digits 8 or 9, but additional price point measures were not nearly significant, once we entered PP99.

[22] Price point effects were also important in unreported FS outlet models. There, prices that ended in 25, 50, or 95 cents were much less likely to be increased, and were changed by more when they were changed.

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