How Firms Construct Price Changes: Evidence from Restaurant Responses to Increased Minimum Wages

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We use price data underlying the Consumer Price Index to assess how restaurants, whose prices are generally quite sticky, respond to minimum wage increases. Aggregate prices rise, quickly, by amounts reflecting the increase in costs, and they rise more among fast food outlets and in low-wage locations. But restaurants do not construct price increases by raising all their prices by amounts reflecting the increase in wages. Instead, they raise only some prices, but by larger amounts. Prices at cluster points are less likely to be changed, and prices that were recently increased (decreased) are less (more) likely to be raised.

Key words: minimum wages, price pass-through, price stickiness.

How are industry-wide cost shocks, such as an increase in minimum wages, passed on to retail consumers? The question is more complicated than it may first appear. It first implies an aggregated question: do prices rise, and by how much? The aggregated question, concerning the extent of pass-through, occurs frequently in analyses of the transmission of changes in agricultural prices to wholesale and retail food prices, and in analyses of the transmission of tax, exchange rate, and materials price changes.

But the question also implies a finer issue of strategy and pricing: retailers sell many items, and can construct store-wide price increases by choosing both the set of items to undergo price increases as well as the size of the increases to be applied. The choices matter because retail prices are often quite sticky. They are changed periodically rather than continuously, and the periods between changes may be long. Prices and price changes also cluster at a few discrete values. Sticky retail prices may respond slowly to price shocks, and if prices and price changes are clustered, retail prices may not respond at all to small cost changes.

The finer question, of how price changes are constructed in a world of sticky prices, is more novel than the aggregated question. However, an emerging class of empirical research, based on microlevel price data, has placed a new emphasis on analyses of retail price formation. Recent studies assess the costs of changing retail food prices (Levy et al., 1997); the extent to which retail supermarket prices remain unchanged over time (Bils and Klenow, 2004; Hosken and Reiffen, 2004); and the unexpected reactions of retail food prices to demand shifts (MacDonald, 2000; Chevalier, Kashyap, and Rossi, 2003) and to rival entry (Ward et al., 2002).

We use microlevel price data to assess price rigidity in an industry, to analyze price transmission, and to investigate whether and how price rigidity affects price transmission. Specifically, we assess the response of restaurant prices to 1996 and 1997 increases in minimum wages, using item prices collected by the U.S. Bureau of Labor Statistics (BLS) for use in the Food Away from Home (FAFH) component of the Consumer Price Index (CPI). The transmission issue (the impact of minimum wage increases on product prices) is important in its own right. Since restaurants, particularly fast food establishments, are major employers of minimum wage workers, and since minimum wage labor constitutes an important share of industry costs, the industry is an ideal choice

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for analysis of that issue. But the data also allow us to extend recent pricing research to FAFH (prior research has focused on supermarket scanner data), and to look more closely at the links between item price rigidity and outlet price transmission, or what we call the "construction" of outlet price responses.[1]

Restaurant responses to minimum wage changes followed textbook expectations—prices rose, quickly, by amounts consistent with the modest costs imposed by minimum wage increases. And prices rose more where the wage increases raised costs more, in fast food outlets and in low-wage locations.

But the aggregate increases were constructed in interesting ways. Firms do not raise all prices by amounts reflecting the costs of minimum wage increases. Instead, they raised prices on a subset of items, but by larger amounts, a response suggesting item-specific fixed costs to changing price, or demand elasticities that vary across items. Prices at cluster points were less likely to be changed, and prices that were recently increased (decreased) were less (more) likely to be raised in response to a minimum wage increase.

In the next section, we distinguish this article from prior research on minimum wage effects. Then we describe our data source in more detail, and summarize the relevant elements of price stickiness. Later sections provide our evidence of a rapid and substantial response of retail prices to minimum wage changes, and of the complications, related to price stickiness, in how those retail price responses are constructed.

Prior Research on Minimum Wage Increases and Product Prices

There have been two recent studies of the impact of minimum wage increases on U.S. restaurant prices, and each used published BLS price indexes for selected metropolitan areas. Card and Krueger (1995) compared changes between 1989 and 1992 in FAFH indexes for 29 major metropolitan areas, a period that encompassed 1990 and 1991 federal minimum wage increases, and found that prices rose more in those cities with higher proportions of workers affected by minimum wage increases. With a small sample, their estimates were imprecise, with coefficient values and statistical significance that were each sensitive to model specification.[2]

Aaronson (2001) built a larger sample over a longer time span, and analyzed the impact of federal and state minimum wage increases on monthly variations in FAFH indexes for 27 major metropolitan areas from 1978 to 1986, and for 15 metro areas from 1986 to 1995 (numbers vary with changes in BLS programs). Minimum wage increases were associated with statistically significant price increases, with magnitudes approximating the likely cost effects of minimum wage changes. Most of the price responses occurred quickly, within a three-month window surrounding the minimum wage change. His findings were sensitive to the time period studied, with much smaller and less precisely estimated effects during 1983–95, a period of lower inflation and a smaller sample. Aaronson (2001) found nearly identical results for province-level Canadian price data over the same period.

The published indexes analyzed by Card and Krueger (1995) and by Aaronson (2001) aggregate price changes across many types of outlets in 15–29 major metropolitan areas. We use item price data drawn from 88 different metro and urban areas, a much wider range of locations than those covered by published indexes. With our sample, we can assess price responses over a wider range of product and labor market conditions and, by linking individual outlets and prices over time, can assess how changes in aggregate price indexes are constructed from price changes on specific items.[3]

[1] Hobijn, Ravenna, and Tambalotti (2004) investigate the construction of retail price changes with an analysis of the striking increases in restaurant prices upon introduction of the Euro. European Union (EU) and U.S. restaurant prices show similar patterns—few prices change each month, and they remain unchanged for long periods. Restaurants as a group do not usually synchronize price changes, but stagger them over time. But the Euro created a menu cost shock, consuming management time and necessitating changes in menus and signage. With currency-related menu costs to be incurred anyway, the marginal menu costs of price changes would be very small, and EU restaurants then synchronized price changes to coincide with currency changeover. The Euro created a one-time shift from staggered to synchronized price-setting among restaurants, leading in turn to a sharp spike in restaurant prices, with dips before and after the changeover (Lach and Tsiddon, 1996). Inflation in other EU products showed no apparent change, and EU countries that did not shift to the Euro saw no increase in restaurant inflation.

[2] In a related analysis, they surveyed prices for selected items at fast food outlets in New Jersey and Pennsylvania, and reported that prices increased in New Jersey outlets, but not in Pennsylvania's, after a 1992 increase in New Jersey's minimum wage.

[3] Bils and Klenow (2004) use the same underlying data source, item prices observed as inputs to CPI indexes over 1995–97. They used product-level measures of price stickiness calculated for them by the BLS to assess the incidence of price stickiness across a broad range of retail industries, and to draw implications for macroeconomic theories of price-setting, while we accessed detailed outlet datafiles at BLS.

Restaurant Price Data

We base our analysis on a deep and richly detailed dataset of items and prices. We describe our data source below, show why restaurant prices can be described as "sticky," and introduce some important empirical regularities that affect our analysis.

Sample Source and Construction

We use prices sampled over a three-year period (January 1995 through December 1997) to construct the CPI for FAFH. BLS field personnel collected prices for nearly 7,500 food items at over 1,000 different outlets. Outlets were drawn from eighty-eight primary sampling units (PSUs), which included seventy-six Metropolitan Statistical Areas and twelve other areas representing the urban nonmetro United States. During this period, PSUs were assigned to one of the three reporting cycles: outlets in the five largest PSUs 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, and randomly assigned outlets in the five largest PSUs to odd or even two-month cycles.[4]

Each outlet has a BLS "type of business" code. In "limited service" (LS) outlets, meals are served for on- or off-premises consumption and patrons typically place orders and pay at the counter before they eat. In "full service" (FS) outlets, wait-service is provided, food is sold primarily for on-premises consumption, orders are taken while patrons are seated at a table, booth, or counter, and patrons typically pay after eating. FS outlets account for about half of all price quotes in our sample, and LS outlets account for about 29%, with the remainder collected from many other outlet types, such as department stores, supermarkets, convenience stores, gas stations, and vending machines. The classification is useful because LS outlets employ higher proportions of teenage and unskilled workers and hence should be more sensitive to minimum wage changes.[5]

Within an outlet, specific items (usually seven or eight) 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 (e.g., 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 BLS introduced a complete outlet and item resampling in January 1998, we only use data through December 1997. 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). 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, and we do not compare prices where BLS has made an item substitution.[6]

Restaurant Prices Change Infrequently

Prior studies find that prices can remain fixed for long periods (Carlton, 1986; Cecchetti, 1986; Kashyap, 1995; Hosken and Reiffen, 2004). That pattern should hold in restaurants, where firms review and change prices periodically rather than continuously (Bils and Klenow, 2004; Hobijn, Ravenna, and Tambalotti, 2004). Figure 1 shows, for items enumerated in any month, the share whose prices remained unchanged from two months before. There is remarkable stability—on average, 87.4% of FS prices remained unchanged in any period. Aside from sharp changes immediately after the 1996 and 1997 federal minimum wage increases, LS meal prices are also stable. Across all outlet types and comparison periods, 86.6% of prices remained unchanged.

Table 1 adds more detail, comparing price changes in periods with and without minimum wage increases. Among LS outlets, only 11.4% of prices are increased, on average,

[4] Counts are based on all priced items in Fall, 1996. A more complete description of outlet and item selection procedures can be found in the BLS *Handbook of Methods* (U.S. Bureau of Labor Statistics, 1992).

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

[6] Firms could respond to a minimum wage increase by reducing quality instead of raising price, which would likely be reflected in the substitution of a new items for old in BLS pricing samples. However, in that case, we ought to observe an increased rate of item substitutions just after federal minimum wage increases, when most sample outlets faced minimum wage changes. Instead, the incidence of item substitutions is no greater in those periods than during other bimonthly periods in our 1995–97 time span.

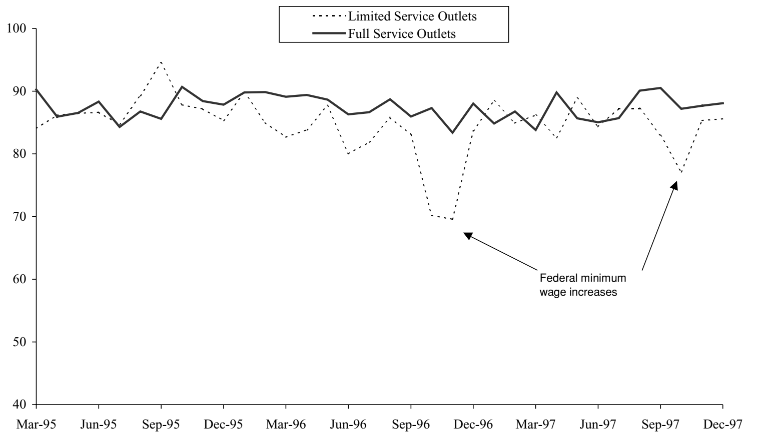

Figure 1. Percentage of items with unchanged price, bimonthly periods, 1995–97

Table 1. The Incidence and Distribution of Bimonthly Price Changes, 1995–97

	Outlet and Month Classification			
Outlet Types:	Limited Service Outlets		Full Service Outlets	
Periods with Minimum Wage:	Unchanged	Increasing	Unchanged	Increasing
Item price increases				
Percentage of items with increase	11.4%	22.6%	10.8%	11.2%
Mean price increase	5.6%	5.1%	4.8%	4.9%
Distribution of increases	Percentage of items with price increase			
0–2%	25.0	22.0	32.7	26.2
2–6%	47.4	52.6	40.4	46.4
6–10%	14.5	13.5	16.3	17.8
>10%	13.1	12.0	10.8	9.6
	100.0	100.0	100.0	100.0
Item price cuts				
Percentage of items with cut	2.9%	2.5%	1.7%	1.7%
Mean price cut	7.6%	7.5%	7.5%	9.3%
Distribution of cuts	Percentage of items with price cut			
0–2%	30.0	24.5	25.3	21.4
2–6%	29.9	30.6	32.0	23.2
6–10%	11.9	17.4	18.7	22.3
>10%	28.3	27.6	23.9	33.0
	100.0	100.0	100.0	100.0

Note: "Periods with Minimum Wage Increasing" include any bimonthly period with an increase in the minimum wage effective in an outlet's state.

in bimonthly periods with no minimum wage increase, but the fraction jumps to 22.6% in periods with minimum wage increases. FS outlets show a much smaller increase in incidence, from 10.8% to 11.2%. The data show no large change in the incidence of price cuts, in either outlet type.

Consider another measure of the duration of prices, using LS items whose prices increased just after the federal minimum wage increase in October 1996. Ten months later, 56% of the new prices remained unchanged. The pattern was close to that for items that did not change price just after the minimum wage increase: 49% of those prices remained unchanged ten months later. A similar pattern holds broadly across all possible ten-month spans in our sample: 48% of LS and 55% of FS prices remain unchanged. Taken together the duration measures suggest that the half-life of a restaurant

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 increase?	No	Yes	No	Yes
Limited service outlets				
Outlet-periods	1,852	287	1,070	162
Number of price increases	Share of outlet-periods (%)			
0	77.3	58.9	68.2	59.2
1–2	11.0	12.1	17.3	19.0
3–5	6.5	14.1	9.8	10.3
6	1.2	4.5	2.1	1.7
7	4.0	10.4	1.7	5.2
8			1.0	4.6
Full service outlets				
Outlet-periods	2,965	528	1,508	286
Number of price increases	Share of outlet-periods (%)			
0	80.4	79.6	80.7	78.7
1–2	8.1	8.5	8.5	7.3
3–5	5.9	6.5	4.5	7.0
6	1.9	0.8	1.1	2.1
7	3.7	4.7	2.1	1.8
8			3.2	3.2

Note: "Minimum Wage Increases" include any bimonthly period with an increase in the minimum wage effective in an outlet's state.

price in this low-inflation period was about ten months.[7]

The Distribution of Price Changes

Table 1 also contains information on the size distribution of price changes. Mean LS increases, conditional on a price rise, are slightly higher than FS increases. Mean increases change little in months after minimum wage increases, and stay very stable over time, within a range of 4–6%.[8]

Price changes peak (cluster near the mean), compared to a normal distribution. Among LS outlets, about a quarter of all price increases are under 2%, and about three-quarters fall below 6%. Similar magnitudes obtain among FS outlets, with nearly three-quarters of all increases under 6%.[9]

Although price cuts are rare, they can be large. Over a quarter of price cuts exceed 10%, and they frequently exceed 20%. Because many price cuts reflect sales, and hence are of large size but limited duration, we may need to explicitly account for prior price changes in our later modeling of minimum wage effects.[10]

But Price Changes Are Not Synchronized

Although outlets change prices infrequently, they do not synchronize the price changes that they do make. That is, outlets do not all change prices at the same time, as automobile dealerships or fashion retailers do at the start of a model year (if they did, figure 1 would show sharp spikes at regular intervals).

When they do change prices, outlets usually only change some. Table 2 details the incidence of price increases among the large majority of

[7] If 93.3% of prices were unchanged each month (86.6% bimonthly), and if the probability of change was independent of the length of time that a price was unchanged, then in ten months 50% of prices ($1 - 0.933^{10}$) would be unchanged. With the ten-month survival rates so close to 50%, the probability of price change does appear to be independent of the length of time that it had been fixed. Bils and Klenow (2004) reached similar conclusions regarding price durations in the industry.

[8] 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 LS increases. FS price increases do show a statistically significant shift, driven by the lower incidence of 0–2% increases when minimum wages are raised.

[9] Excess kurtosis, a measure of peakedness, was 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.

[10] Recall that BLS aims to price full meals, so that an item might be a "combo meal" at a limited service restaurant, and such meals are sometimes offered at sale prices for limited periods. Our later analyses show that a large price cut is tied to a sharply increased probability of a next period price increase, and that increase is likely to be large.

sample outlets that have seven or eight item price quotes. The table sorts incidence by outlet type and by whether the minimum wage was increased during the bimonthly price comparison period.

Most outlets in most periods change no sample prices. For example, among outlets with seven price quotes in periods with no minimum wage change, 77.3% of LS outlets and 80.4% of FS outlets change no sample prices at all. Among those outlets that increase any sampled prices, just under half raised prices of one or two items, while less than a fifth increased all sample prices. During periods of minimum wage increases, there is a substantial decline in the fraction of LS outlets that keep all sample prices unchanged, and a very small decline in the share of FS outlets with unchanged sample prices. When minimum wages increase, many more LS sample outlets raise all sample prices; but most raise more, not all, prices. Outlets raising prices in the wake of common cost increases appear to select a set of items to increase, meaning that the construction of a price increase is a strategic choice.

Prices Cluster at Price Points

One in eight LS item prices ends in 99, while over 30% end in 9, fractions that are twelve and three times those that would appear by random chance if all endings were equally likely. Few FS prices end in 99, but one in five end in 00, 25, or 95 (almost seven times random chance). Price changes are also strikingly clustered: one-quarter of all LS price changes are 10 cents, while 5, 10, 20, and 30 cents account for half. Those four, plus 25 cents, 50 cents, and a dollar, account for half of all FS price changes.

Kashyap (1995), Blinder et al. (1998), and Hall, Walsh, and Yates (2000) identify such psychological price points as important elements in retailer pricing decisions, suggesting a discontinuous threshold at which buyers are unusually sensitive to price increases. At price points, we should see longer quote durations and price changes, when they occur, that are larger. If outlets put off price changes until costs change to a price changing point, then price increases will lag cost increases. If price points matter to consumers, outlets may be less likely to change prices of items at price points, and instead select other items.

To summarize, restaurants change item prices infrequently, and prices and price changes cluster at common values. Price increases tend to range closely around a mean percentage value that varies little over time. When restaurants increase prices, they are likely to increase a subset, not all. Given these observed elements of price stickiness, one might easily anticipate that the modest cost changes associated with minimum wage increases might not be passed through to retail prices at all. Instead, we find below that retail prices respond rapidly to minimum wage increases, in magnitudes consistent with the increased costs, but that price stickiness plays a role in determining which items bear the price increase.

Modeling the Effects of Minimum Wage Increases on Prices

A broad literature, useful for thinking about how minimum wage increases affect product prices, analyzes the impact of changes in taxes, exchange rates, or intermediate goods prices on product prices. A textbook result is that, under perfect competition, no input substitution, and constant marginal costs, an input price increase will be fully passed through to the product price, and the percentage increase in the product price will equal the percentage change in the input price times the input's share in costs.

When these restrictive assumptions are relaxed, the pass-through to product prices depends on a variety of factors, including the magnitude of the demand elasticity, the convexity of demand, the labor supply elasticity, the elasticity of marginal cost with respect to output, the degree to which inputs can be substituted for one another, and the nature of competition. Given the variety of factors that can impact pass-through, it is not surprising that product prices can sometimes increase less, and sometimes more, than predicted under the full pass-through case.

Furthermore, the minimum wage literature has provided an even starker alternative to consider: do prices rise at all after a cost shock? Spurred by Card and Krueger's (1995) surprising finding that employment barely budges, perhaps even goes up, in response to an increase in the minimum wage, a variety of models (e.g., Burdett and Mortensen, 1998; Bhaskar and To, 1999; Manning, 2003) were developed that incorporate monopsony (or monopsonistic competition) in low-wage labor markets and that have such employment implications. Aaronson and French (2005) and Aaronson, French, and MacDonald

(2004) formally quantify the implications of monopsony-like behavior on prices within a structural model of labor demand, showing that prices need not rise, and may even fall, in response to a minimum wage change.[11]

Moreover, the models outlined above all assume that prices are flexible, with no cost of changing prices. But with sticky prices, perhaps due to substantive costs of changing product prices, product prices may not respond at all to small cost changes. There is some evidence (Bils and Klenow, 2004; Hobijn, Ravenna, and Tambalotti, 2004) that restaurants might be particularly susceptible to such concerns. Consequently, the impact of a minimum wage hike on product prices is very much an empirical issue.

We have a rich set of microdata, consisting of price observations gathered at bimonthly intervals for meals sold at specific outlets, to use in evaluating the relationship between minimum wage changes and output price changes. A common drawback to microdata, and one that holds here, is that the precisely detailed units of observation cannot be cleanly matched to other data on other factors that may affect product pricing, like other input prices, technological relationships, and elements of product demand. As a result, there is a legitimate concern that measured impacts of minimum wage changes on prices may reflect the influence of omitted variables.

That issue would be a major concern if we only observed a once and for all change in minimum wages—it might then be difficult to separate the effects of changes in minimum wages from other unmeasured changes occurring at the same time. But we actually observe a finely graduated set of minimum wage changes affecting different outlets at different times, which allows us to effectively identify the effects of such changes on prices. First, states may have minimum wages, and the effective minimum wage in a state is the higher of the federal or state minimum. As a result, a given change in the federal minimum wage can lead to varying changes in effective minimum wages across states, depending on the existing pattern of state minimums. Second, one type of outlet—LS restaurants—employs more minimum wage workers than other outlet types; hence, minimum wage changes should have much greater impacts on LS costs and prices. Third, minimum wages do not always exceed market wages, and minimum wage increases should then have bigger impacts on restaurant costs in some local labor markets than in others. We detail those elements of our identification strategy below.

Measuring Effective Changes in Minimum Wages

On August 20, 1996, President Clinton signed a bill raising the federal minimum wage in two stages: an October 1, 1996 increase from \$4.25 an hour to \$4.75 an hour, and a second increase eleven months later, on September 1, 1997, to \$5.15.

Table 3 shows how effective minimum wage changes varied across states between 1995 and 1997. Consider the October 1996 federal increase. Outlets in 39 states faced the full 11.8% increase implied by the new federal law. But six states (Hawaii, Massachusetts, New Jersey, Oregon, Vermont, and Washington) already imposed minimum wages that were at or above the new federal levels, and therefore faced no change in the effective minimum wage. State minimums in Delaware, Iowa, and Rhode Island fell between the old and new federal minimums, and outlets in those states faced effective minimum wage increases that were less than the federal increase. Alaska, Connecticut, and the District of Columbia maintain minimum wages at a constant amount above the federal level, so federal increases trigger contemporaneous increases in those jurisdictions; but since they started at higher levels, their 1996 percentage increases were smaller than the federal increase.

A similar pattern held at the time of the 1997 federal increase (8.4%)—five states faced no effective change in minimum wages, while seven faced smaller percentage changes than the federal increase (table 3). Finally, six states (California, Delaware, Massachusetts, Oregon, Rhode Island, and Vermont) changed their state minimums at least once during 1995–97 (ten changes altogether), and outlets in those states faced effective minimum wage

[11] Under monopsony, increasing the minimum wage can cause employment to fall, rise, or have no effect depending upon how high the minimum wage is set. Since a monopsonist faces the market labor supply curve, it must raise market wages to attract additional workers, and (absent wage discrimination) will face a marginal factor cost of labor that exceeds the wage paid by the monopsonist. But a minimum wage rule can alter the monopsonist's marginal factor cost, and therefore its employment calculation. If set between the market wage and the old marginal factor cost, a minimum wage will induce the monopsonist to expand employment, compared to an environment with no minimum wage. In this circumstance, increased employment will increase output and, with downward sloping product demand curves, output prices will fall. This result holds under very general product market structures and production functions.

Table 3. Minimum Wages, by Jurisdiction, during 1995–97

Jurisdiction	Initial Minimum January, 1995	Months When New Minimums Were Initiated 1/96	4/96	10/96	1/97	3/97	8/97	9/97
		Minimum wage (dollars per hour)						
United States	4.25			4.75				5.15
States								
AK	4.75			5.25				5.65
CA	4.25			4.75		5.00		5.15
CT	4.27			4.77				5.18
DC	5.25			5.75				6.15
DE	4.25		4.65	4.75	5.00			5.15
HA	5.25							
IA	4.65			4.75				5.15
MA	4.25	4.75			5.25			
NJ	5.05							5.15
OR	4.75				5.50			
RI	4.45			4.75	5.15			
VT	4.50	4.75			5.00		5.25	
WA	4.90							5.15

Source: *Monthly Labor Review*, Annual Surveys of State Labor Legislation, January issues.
Note: States that are not separately listed here had the federal (United States) minimum wage.

increases at times that differed from other states.

We capture changes in effective minimum wages with MW_{it}, the percentage increase in the minimum wage effective in state i during period t. If the minimum wage effective in a state increased by 10% on October 1, then MW_{it} will equal 10 for August–October and September–November price comparisons, and 0 for other bimonthly comparisons.

Minimum wage effects should also vary between LS and FS outlets. Tips are important for many FS workers, and changes in nontipped minimum wages also represent effective minimum wage changes for tipped employees.[12] But because wages and tip earnings at FS outlets usually exceed effective minimum wages, FS employees are far less likely to be affected by minimum wage increases.

A third factor should create further local variation in price responses. Market wages vary across geographic areas; where prevailing low-skill wages exceed minimum wages, minimum wage increases should have little effect on market wages, and hence little effect on costs and prices. Where minimums exceed market wages for low-skill workers, 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 prices in low-wage areas.

With outlet locations, we can link outlets to related geographic information. We use Current Population Survey data to summarize 1996 hourly wage distributions in the outlet's PSU, and use the 20th percentile of the PSU's hourly wage distribution (WAGE20) as our measure of low-skill wages in a metro area.[13] In our analysis, we allow the effect of MW_{it} to vary across locations with different values of WAGE20.

We evaluate lead and lagged responses in our analysis, for three reasons. First, the nature of BLS price collection could create lags in observed price changes. BLS enumerators collect prices in three-week-long collection periods during the first 22 days of a month. Suppose that an outlet is visited in the first week of October, and suppose that, in response to an October 1 minimum wage increase, it raises

[12] Federal law sets a separate cash minimum for tipped employees ($2.13), but requires that tips plus cash wages must at least equal the nontipped employee minimum ($5.15 in September, 1997); the same requirement holds for state laws.

[13] Wage data for the twelve nonmetro PSUs are drawn from the nonmetro parts of the outlet's state. CPS codes are unavailable for nine MSAs, so sample sizes decline when area wage data are included in the analysis. Employment studies use other measures of labor market exposure to minimum wage increases (Card and Krueger, 1995). We also evaluate Card and Krueger's measure, the fraction of area workers earning the minimum wage (PCMIN). WAGE20 gives better results, but we also believe that it is a better measure for this sample. The smaller cities in our sample have smaller CPS worker samples (earlier studies relied on samples of major metro area), providing less reliable estimates of PCMIN.

price on October 10. The price increase will not be observed until the next enumerator visit, in the first week of December. Thus, the very quick actual response will only be observed with a one-period lag. Second, actual price responses could occur with a lag, if restaurants only review and change prices at a few discrete intervals during the year. Third, restaurants could change some prices before an anticipated cost increase. Consider the timing of the 1996–97 federal increase. Businesses knew of the 1996 increase just 2–4 months prior to implementation, while they knew of the 1997 increase, specified in the 1996 bill, 12–13 months before implementation.[14] Hence, outlets could have changed prices in response to the October 1996 increase during August or September price reviews, and they could have responded to the September 1997 increase well before it was implemented.

Aggregate Response: The Effects of Minimum Wage Increases on Prices

Our empirical analysis assesses whether prices at retail outlets increased by more following upon a minimum wage increase, and more still among outlets dependent on minimum wage labor and in low-wage locations, than they did in months when there was no minimum wage increase. Accordingly, our basic statistical model is specified as follows:[15]

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

$P_{kj,t}$ is the price of the item k at outlet j in month t. The dependent variable is the percentage change in price over a bimonthly period. PPI is the bimonthly percentage change in the Producer Price Index for Processed Foods, a measure of input price shocks faced by sample outlets. We include contemporaneous as well as one- and two-period lagged values. We also include dummy variables for the type of meal (MEALTYPE)—BLS codes identify dinner, lunch, snack, and breakfast/brunch items (snacks, mostly offered through unusual outlet types, were dropped).

It is important to control for pricing dynamics. Some prices fall sharply during sales (table 1), and rebound after the promotion. Similarly, some increases are retracted later, because the original increase reflected temporary cost increases or because rivals did not match an increase. To capture that process, we use the vector IP, with measures of the item's recent pricing history. Specifically, IPUP is the percentage increase in an item's price in a previous period (zero if the price did not increase). Similarly, IPDOWN is the percentage decrease in price in a previous period (zero if there was no decrease). We enter one-, two-, and three-period lags for each.[16]

MW_{it} measures effective minimum wage increases facing outlets in state i during period t. In addition to contemporaneous values, we also use lead and lag values, MW_{it-1} and MW_{it+1}, to assess price effects one period before and one period after 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.

Table 4 presents the basic analysis, using all outlets. Because price quotes from the same outlet are unlikely to be statistically independent, all standard error calculations account for quote clustering, using Huber–White robust estimation techniques. The adjustment matters—standard errors are two to three times greater than in OLS.

Item price dynamics matter: prior price cuts lead to current period price increases, and past increases lead to current cuts. All estimated IP coefficients were highly significant, although the effects are small; full reversion to prior prices implies absolute coefficient values of 1, while these fall well below 1 and usually below 0.1.

We multiply all reported minimum wage coefficients and standard errors by 10 to save space (results should then be read as the effects of 10% minimum wage increases). In column 1 of table 4, the minimum wage effect is positive and highly significant; outlet prices rise by 0.33% for 10% increases in the minimum

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[14] The 1996 increase was unexpected, and passage could not have been predicted until shortly before the House of Representatives vote on May 23 (Weisman, 1996). Even then, the final timing did not become clear until adoption of the conference report on August 2. The unexpected nature of the 1996–97 increase suggests that endogeneity was not an issue—that the wage increase was not a predictable response to price increases. Aaronson (2001) tested for such endogeneity in 1978–95, when it was more likely to occur, and found no evidence.

[15] The basic statistical relationship is discussed and derived in Aaronson and French (2005) and Aaronson, French, and MacDonald (2004).

[16] Adding lags reduces sample size. Three lags (six months) were always statistically significant, while a fourth period was not. If we exclude the lagged prices, estimated standard errors on the minimum wage coefficients increase, although the coefficients remain statistically significant. Exclusion of the lags has modest impacts on the size of the coefficient estimates on minimum wage variables.

Table 4. Magnitude and Timing of Price Responses to Minimum Wage Increase

Variable	Description	Coefficients (Standard Errors) (1)	(2)
Intercept		0.379 (0.026)	0.315 (0.032)
PPI_t	% increase in producer price index for processed foods	0.069 (0.033)	0.070 (0.030)
PPI_{t-1}	% increase in producer price index for processed foods, one-period lag	−0.042 (0.040)	−0.040 (0.038)
PPI_{t-2}	% increase in producer price index for processed foods, two-period lag	0.046 (0.031)	0.046 (0.030)
MTYPE2	Dummy variable = 1 if dinner item	0.006 (0.029)	0.007 (0.021)
MTYPE3	Dummy variable = 1 if other item (lunch was null)	−0.075 (0.061)	−0.077 (0.063)
$IPUP_{t-1}$	% increase in item price, one-period lag; zero if no increase	−0.089 (0.013)	−0.090 (0.013)
$IPUP_{t-2}$	% increase in item price, two-period lag; zero if no increase	−0.070 (0.012)	−0.070 (0.012)
$IPUP_{t-3}$	% increase in item price, three-period lag; zero if no increase	−0.036 (0.010)	−0.036 (0.009)
$IPDOWN_{t-1}$	% decrease in item price, one-period lag; zero if no decrease	0.388 (0.045)	0.389 (0.045)
$IPDOWN_{t-2}$	% decrease in item price, two-period lag; zero if no decrease	0.096 (0.023)	0.096 (0.023)
$IPDOWN_{t-3}$	% 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 (∗10)	0.331 (0.065)	0.406 (0.063)
$MW_{i,t-1}$	% minimum wage change, state i, period $t-1$ (∗10)		0.207 (0.006)
$MW_{i,t+1}$	% minimum wage change, state i, period $t+1$ (∗10)		0.115 (0.007)
R^2		0.065	0.066
N		68,887	68,887

wage. Column 2 adds lead and lag effects for minimum wages. Each effect is positive; the lag effect is highly significant, 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 combined effect is more than double that in equation (1), a pattern that closely matches that found by Aaronson (2001) with aggregated BLS data for 1978–95. With monthly data, 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 contemporaneous, lead, and lag periods, at two months each, sum to six months.

The Effect of Outlet Type on Price Responses

We examine how minimum wage effects vary across outlet types in table 5. The table reports minimum wage effects, but the models retain all other explanatory variables from table 4. Estimated minimum wage effects are robust to the inclusion or exclusion of the other explanatory variables, and the pattern of coefficients of the other explanatory variables changes little as the model changes to capture different minimum wage effects.[17]

Column 1 repeats the minimum wage coefficients from table 4, column 2, for comparison.

[17] We also estimated models with fixed chain and location (PSU) effects. Minimum wage coefficients were unaffected by the inclusion of fixed effects and the fixed effects themselves added almost nothing to the model's fit.

Table 5. Effects of Outlet Type on the Price Response to a Minimum Wage Increase

Outlet Type	(1) All	(2) FS	(3) LS	(4) LS
MW_{it}	0.406	0.191	0.805	0.937
	(0.063)	(0.086)	(0.125)	(0.134)
$MW_{i,t-1}$	0.207	0.212		0.305
	(0.006)	(0.083)		(0.119)
$MW_{i,t+1}$	0.115	−0.062		0.319
	(0.007)	(0.075)		(0.142)
N	68,887	35,759	21,064	21,064
R^2	0.066	0.017	0.151	0.152

Note: Each regression also includes other variables listed in table 4. FS refers to full service outlets, while LS refers to limited service outlets. $MW_{i,t-1}$ and $MW_{i,t+1}$ refer to one period before and one period after minimum wage changes. All standard error estimates are adjusted for error clustering within outlets.

Column 2 reports the coefficients for FS outlets only, while columns 3 and 4 report results for LS outlets. FS minimum wage effects are statistically significant but very small. LS effects are much larger. The coefficient of MW_{it} in column 3, positive and highly significant, suggests that LS prices rise by 0.8% in response to contemporaneous 10% increases in minimum wages. In column 4, we add lead and lag effects to the LS equation; the effects are positive, large, and statistically significant, and their inclusion raises the contemporaneous estimate.

Including lead and lag effects, LS prices rise by 1.56% in response to a 10% minimum wage increase, more than double the all-outlet estimate of 0.73%. FS price effects are much lower, 0.34% including the negative but not significant lead effect. Our estimated all-outlet magnitude corresponds closely to that found by Aaronson (2001), who reported a 0.72% price increase in the United States in response to a 10% increase in the minimum wage (and 0.74% for Canada, using Statistics Canada Province-level price data). He also used American Chamber of Commerce data to show that reported prices for McDonald's hamburgers and KFC chicken rose by 1.5–1.6% in response to 10% minimum wage increases, but reported Pizza Hut pizza prices did not change significantly. Those findings mirror ours: if they were to be surveyed for the CPI, McDonald's and KFC outlets would be LS outlets, which had a large response to the minimum wage increases, while Pizza Hut outlets would be FS (a very small response). Hence the magnitude of response, in the aggregate and by outlet types, appears to have remained consistent over twenty years, since Aaronson's (2001) analysis covers a period, 1978–95, which just precedes our 1995–97 span.[18]

Our magnitudes are also consistent with the likely cost-effects of minimum wage increases. On average, payroll accounts for 25% of sales in LS restaurants (U.S. Census Bureau, 2000). If LS outlet wages rose by 10% in response to a 10% minimum wage increase, and if there were no substitution between labor and other factors of production, we then would expect LS costs to rise by 2.5% under the base case defined above. But since many LS workers earn wages above the minimum wage, and higher-paid workers account for a disproportionate share of payroll, the actual increase under full pass-through should be less than 2.5%. Without further information on the share of minimum wage workers in payroll, spillovers of minimum wage increases to other wages, and substitution in demand and in production, we cannot identify the likely cost-effects of minimum wage increases. Nevertheless our estimate of a 1.56% price increase across all LS outlets within a six-month window suggests that restaurant prices respond rapidly to cost shocks, and by amounts that represent substantial pass-through of cost increases to retail consumers, in spite of widespread evidence of price stickiness.

How Location Affects Price Responses

Price responses vary systematically by location. Columns 1–3 in table 6 introduce an interaction 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 should be less strongly affected by changes in the minimum wage.[19]

In the all-outlet sample (column 1) the main MW coefficient remains positive and highly significant, while the interaction term is negative and statistically significant ($t = 2.07$). Minimum wage increases have larger effects on prices in low-wage areas. Among FS outlets (column 2), the coefficients of MW and the

[18] Recall that Aaronson (2001) reported that his results were much weaker (losing statistical significance) for the 1983–95 subperiod. Our findings suggest that the weakness followed from the reduction in sample size (from twenty-seven to fifteen metro areas) in the period, and not from any change in responsiveness.

[19] In unreported regressions, we entered WAGE20 separately and in interaction with MW. Coefficients of WAGE20 were always small and never statistically significant, unsurprising since we have no good reason to expect larger price increases in low-wage areas, absent a change in the minimum wage. Coefficients of the interaction term shrank and became only marginally significant when we included WAGE20 in the model.

Table 6. Effects of Location on the Price Response to a Minimum Wage Increase

Outlet Type	(1) All	(2) FS	(3) LS	(4) LS	(5) All	(6) FS	(7) LS
MW_{it}	1.449	1.048	2.698	0.613	1.453	1.040	2.709
	(0.531)	(0.692)	(0.883)	(0.272)	(0.544)	(0.710)	(0.900)
$MW_{i,t-1}$	0.200	0.227	0.263	0.265	0.225	0.249	0.295
	(0.068)	(0.088)	(0.121)	(0.121)	(0.067)	(0.087)	(0.120)
$MW_{i,t+1}$	0.124	−0.041	0.296	0.297	0.077	−0.082	0.243
	(0.070)	(0.080)	(0.155)	(0.155)	(0.067)	(0.078)	(0.149)
$MW_{it} * WAGE20$	−0.162	−0.129	−0.278				
	(0.078)	(0.099)	(0.133)				
$MW_{it} * PCMIN$				0.048			
				(0.039)			
$MW_{it} * RWAGE20$					−0.757	−0.595	−1.304
					(0.374)	(0.480)	(0.622)
N	61,716	32,822	18,024	18,024	61,716	32,822	18,024
R^2	0.068	0.018	0.164	0.164	0.070	0.019	0.171

Note: Each regression also includes other variables listed in table 4. FS refers to full service outlets, while LS refers to limited service outlets. $MW_{i,t-1}$ and $MW_{i,t+1}$ refer to one period before and one period after effective minimum wage changes. WAGE20 is the wage at the 20th percentile of an area's hourly wage distribution, PCMIN is the estimated share of an area's workers earning the minimum wage, and RWAGE20 is WAGE20, divided by the effective minimum wage in an outlet's state. All standard error estimates are adjusted for error clustering within outlets.

interaction term are of the expected sign but are only marginally significant.

In column 3 we assess the LS sample. As in the all-outlet sample, price responses are larger in low-wage areas—the WAGE20 coefficient is negative, large, and statistically significant. With WAGE20 set equal to \$5.50, predicted prices would rise by 1.83% in response to a 10% minimum wage increase, while the rise would be 1.31% where WAGE20 was \$7.37 (\$7.37 and \$5.50 are the midpoints of the top and bottom quartiles of WAGE20).[20]

In column 4, we replace WAGE20 with an alternative, the share of workers earning the minimum wage (PCMIN), a measure used in Card and Krueger's (1995) analysis (see footnote 12). The sign of PCMIN is positive, as it should be if the cost effects are greater in low-wage areas. But the coefficient is not statistically significant.

We tried one more alternative in columns 5–7. RWAGE20 is the ratio of WAGE20 to the effective minimum wage faced by an outlet. It measures the gap between our indicator of low-skill market wages (WAGE20) and the effective minimum, and ranges from 1.1 to 1.7. The results are almost identical to those obtained with WAGE20, with a slightly better fit to the data. Again, areas with relatively high wages realize price increases, in response to a 10% minimum wage increase, that are half a percentage point lower than areas with relatively low wages (again using the top and bottom quartile midpoints as high and low).

How Are Price Increases Constructed?

We now turn to the issue of how restaurants construct price changes. Prices at LS outlets rise by nearly 1.6% in response to 10% increases in minimum wages. Restaurants can arrange that increase in many ways. They could raise all prices by 1.6%, or they could raise fewer prices by greater amounts. If they choose the latter, they must decide which item prices to increase.

Mean LS price increases, conditional on increasing price, were 5.1% in periods following a minimum wage increase and 5.6% in other periods (table 1), certainly no evidence that price increases are larger after minimum wage increases. More formal regression analysis supports that view. When we ran our price models with the samples restricted to items with price increases, the coefficients of the minimum wage variables were positive, but very small and not significant. Minimum wage increases have virtually no effect on the magnitude of price increases.

It appears that outlets construct store-wide price increases by raising more prices, not by

[20] We also explored interactions between WAGE20 and the lead and lag MW terms. Lagged effects have the expected sign but were not quite significant, while lead effects lost all power, suggesting that there was no interaction and that inclusion simply created multicollinearity with the main effect.

Table 7. Logit Model Coefficients: Probability of Price Increase, Limited Service Outlets

	Coefficients (Standard Errors)				
Variables	(1)	(2)	(3)	(4)	(5)
$IPUP_{t-1}$	−0.0867	−0.0910	−0.0960	−0.0939	−0.0944
	(0.0148)	(0.0150)	(0.0180)	(0.0123)	(0.0175)
$IPUP_{t-2}$	−0.0432	−0.0459	−0.0403	−0.0406	−0.0401
	(0.0156)	(0.0157)	(0.0178)	(0.0102)	(0.0177)
$IPUP_{t-3}$	−0.0225	−0.0206	−0.0291	−0.0295	−0.0290
	(0.0127)	(0.0125)	(0.0140)	(0.0103)	(0.0140)
$IPDOWN_{t-1}$	0.1072	0.1120	0.1199	0.1205	0.1197
	(0.0120)	(0.0126)	(0.0145)	(0.0092)	(0.0145)
$IPDOWN_{t-2}$	0.0893	0.0910	0.0903	0.0891	0.0892
	(0.0135)	(0.0138)	(0.0140)	(0.0110)	(0.0138)
$IPDOWN_{t-3}$	0.0514	0.0562	0.0552	0.0550	0.0550
	(0.0172)	(0.0180)	(0.0201)	(0.0100)	(0.0200)
REVIEW	−0.2377	−0.2551	−0.1614	−0.2292	−0.1600
	(0.1066)	(0.1106)	(0.1406)	(0.0549)	(0.1358)
PP99	−0.4882	−0.4804	−0.4605	−0.4762	−0.4637
	(0.1123)	(0.1160)	(0.1398)	(0.0818)	(0.1394)
$MW_{i,t}$		0.1040	0.2744	0.2680	0.2714
		(0.0118)	(0.0794)	(0.0365)	(0.0792)
$MW_{i,t-1}$		0.0095	0.0094		
		(0.0135)	(0.0152)		
$MW_{i,t+1}$		0.0174	0.0115		
		(0.0149)	(0.0166)		
$MW_{i,t} * WAGE20$			−0.0257	−0.0263	−0.0257
			(0.0123)	(0.0057)	(0.0123)
$MW_{i,t} * REVIEW$			−0.0308		−0.0309
			(0.0321)		(0.0316)
$MW_{i,t} * PP99$			−0.0058		−0.0054
			(0.0285)		(0.0285)
Constant	−1.7468	−1.9745	−2.0341	−1.9856	−2.0024
	(0.0621)	(0.0739)	(0.0843)	(0.0416)	(0.0761)

Note: Dependent variable is dummy variable equal to 1 if price increased in period. Sample sizes are 21,064 for columns 1 and 2 and 18,024 for columns 3–5.

raising prices more. We explore that issue in more detail, by assessing the question of which prices get increased. We choose a framework that mirrors that used for the analysis of price changes, except that the focus is on the incidence of price increases. We use a logit model, with the dichotomous dependent variable, UP, set equal to 1 for items at outlet *j* whose price was increased in period *t*:

$$(2) \qquad UP_{kj,t} = f(\text{PPI, IP, MEALTYPE, MW, REVIEW}_{j,t-1}, \text{PP99}_{kj,t-1}).$$

Our base model includes the vector IP used in the prior models to capture price dynamics—that is, separate one-, two-, and three-period lags for item price increases and decreases. We also entered (unreported in the tables that follow) dummy variables for meal type and measure of changes in the PPI for food products, with two lags.

The model includes two dummy variables intended to capture aspects of price stickiness. One, PP99, is set equal to 1 for items whose price ends in 99 cents; we expect that items with prices at such "price points" are less likely to be changed. Second, we define the variable REVIEW, and set it equal to 1 if the outlet changed price for *any* sampled item in the previous period (thus providing evidence of a price review).

We initially analyzed a base model, with no controls for minimum wage increases (table 7, column 1). Next we introduced MW_{it}, our measure of current-month minimum wage increases, as well as one period lead and lag MW measures in column 2. Finally, in columns 3–5, we interacted MW_{it} with WAGE20, PP99, and

Table 8. Predicted Probabilities That an Item's Price Will Increase

Events	(1)	(2)	(3)	(4)	(5)
10% Item price cut last period?	No	No	No	No	Yes
5% Item price raise last period?	No	No	No	Yes	No
Outlet price review last period?	No	Yes	No	Yes	No
P1 ends in 99 cents?	No	No	Yes	No	No
Base probability of price increase:	12.1	9.8	7.9	6.4	31.4
After 10% minimum wage increase:					
With mean WAGE20	26.3	22.1	18.1	15.0	54.3
In high-wage area	22.3	18.6	13.1	12.4	49.0
In low-wage area	32.0	27.2	22.6	18.9	61.1

Note: Predicted probabilities from logit model of price increase (table 7, column 4). Mean WAGE20 is \$6.55 an hour, while WAGE20 is set at \$5.50 in a low-wage area and \$7.37 in a high-wage area.

REVIEW, to identify any changes in the pattern of price response during minimum wage increases.

Several robust patterns stand out. First, consider pricing history. All IPDOWN and IPUP coefficients are statistically significant, with the expected signs. That is, items with recent price cuts are more likely to see current increases, and items with recent increases are less likely to show current increases.

Second, the two "price stickiness" indicators also matter (column 2). REVIEW is negative and statistically significant—an item's price is less likely to be increased if its outlet raised other prices in the prior period, suggesting that outlets hold regular price reviews at discrete times, and are unlikely to change prices outside of those reviews. The coefficient of PP99 is negative, large, and statistically significant—items with prices ending in 99 cents are less likely to experience price increases.

Third, the coefficient of MW_{it} is positive and statistically significant—price increases are more likely when minimum wages are increased. The incidence of minimum wage effects varies with low-skill wages, as in earlier equations—the interaction between MW_{it} and WAGE20 is negative and significant.[21]

The logit model provides some striking implications. We use the results of column 4, table 7, to estimate price increase probabilities under alternative scenarios in table 8. We first set a base probability of an increase to 12.1%, which is predicted on the basis of the intercept in equation 4 (setting all independent variables at zero—no recent item price changes, no recent review, no minimum wage change, and price not at a pricing point). A 10% minimum wage increase has a large contemporaneous effect, raising the probability of a price increase to 26.3%, for outlets in areas where low-skill wages are at the sample mean. Prevailing area wages condition the effect; the probability rises to 32% in low-wage areas and falls to 22% in high-wage areas.

Base price increase probabilities decline noticeably if an outlet had recently reviewed price, if an item's price had been increased in the prior period, or if the current price ended in 99 cents. Minimum wage increases raise the likelihood that those items will see price increases, but such items are still much less likely to be used to construct a price increase than other items after a minimum wage increase.

In contrast, items whose prices had recently been cut are far more likely to be chosen for a price increase in response to a minimum wage increase (note that we use a 5% increase but a 10% cut, asymmetric but representative magnitudes). The base probability of a price increase for such items is 31.4% (sharply higher than any other base probability), but the probability jumps to over 50% in most areas after a minimum wage increase, and to over 60% in low-wage areas—that is, the marginal effect of minimum wages on price increase probabilities for such items approaches 30 percentage points, compared to 5.2 percentage points for items at price points in high-wage areas.

In logit models, the marginal effect of a coefficient on the probability depends on values of other variables, and is largest when the base probability is at 50%. For that reason, the logit specification drives the patterns reported in table 8, particularly the variations in the marginal effects of minimum wage increases. But we tested the specification extensively against a variety of alternatives. We

[21] We use WAGE20 here instead of RWAGE20 for simplicity; results are nearly identical with each, and WAGE20 is more straightforward to apply.

found that coefficients of measures of price stickiness (PP99 and REVIEW) showed no substantive or significant change during minimum wage increases. Moreover, estimated coefficients of the IPUP and IPDOWN variables were insensitive to period. The results from our best specification suggest that minimum wage increases have simple proportional impacts on estimated log odds ratios for price increases, a finding that does yield some striking implications for the pattern of price increase probabilities.

Conclusion

Price stickiness characterizes restaurant prices: there are long price durations and infrequent changes, while prices and price changes cluster at recognizable price points. Such patterns lead to a reasonable prior expectation that there might be little transmission of cost shocks to retail prices.

Yet prices respond quickly to minimum wage increases. Most of the observed response occurs in the two-month period just after a minimum wage increase, while the rest occurs in the next two months or in the two months just before. Prices rose far more in LS (fast food) restaurants, which rely more on low-wage labor, and prices rose more in low-wage than in high-wage areas. The rise in costs attendant upon the minimum-wage increase, while small, was a widely publicized and permanent shock. Such shocks may pass through more fully and quickly to prices, as firms may more easily coordinate price changes without inducing increased consumer search among alternatives (Blinder et al., 1998; Hall, Walsh, and Yates, 2000; Levy, Bergen, and Dutta, 2002).

Pricing features matter when we turn to the issue of *how* restaurants construct a price increase. They generally do not raise all prices in response to a general cost shock. They also do not change the size of typical price increases, which cluster at a few absolute values and around a limited range of percentage values. Instead, they raise prices on subsets of items, by amounts that in the aggregate reflect the cost change. These patterns suggest that restaurants may face item-specific costs of changing prices (Levy et al., 1997) and that they may also perceive costs to departing from price change points—the typical and limited set of changes that they apply to items.

As to which items to choose, two striking features appeared. First, restaurants showed a marked reluctance to change prices from psychological pricing points. Specifically, outlets were reluctant to raise price on items whose prices ended in 99 cents, and they were particularly reluctant to use them to effect general price increases in response to minimum wage hikes. Second, there was a clear linkage between current and recent price changes. Outlets were far more likely to choose sale items, whose prices had recently been cut, as vehicles for constructing a response to minimum wage increases, and noticeably less likely to choose items whose prices had recently been increased. Psychological elements appear to play an important role in the construction of a response to a small, publicized, and widespread cost shock, even as the aggregate response itself appears to follow a straightforward economic framework.

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