Minimum Wages and Employment: A Case Study of the Fast-Food Industry in New Jersey and Pennsylvania

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On April 1, 1992, New Jersey's minimum wage rose from $4.25 to $5.05 per hour. To evaluate the impact of the law we surveyed 410 fast-food restaurants in New Jersey and eastern Pennsylvania before and after the rise. Comparisons of employment growth at stores in New Jersey and Pennsylvania (where the minimum wage was constant) provide simple estimates of the effect of the higher minimum wage. We also compare employment changes at stores in New Jersey that were initially paying high wages (above $5) to the changes at lower-wage stores. We find no indication that the rise in the minimum wage reduced employment. (*JEL* J30, J23)

How do employers in a low-wage labor market respond to an increase in the minimum wage? The prediction from conventional economic theory is unambiguous: a rise in the minimum wage leads perfectly competitive employers to cut employment (George J. Stigler, 1946). Although studies in the 1970's based on aggregate teenage employment rates usually confirmed this prediction,[1] earlier studies based on comparisons of employment at affected and unaffected establishments often did not (e.g., Richard A. Lester, 1960, 1964). Several recent studies that rely on a similar comparative methodology have failed to detect a negative employment effect of higher minimum wages. Analyses of the 1990–1991 increases in the federal minimum wage (Lawrence F. Katz and Krueger, 1992; Card, 1992a) and of an earlier increase in the minimum wage in California (Card, 1992b) find no adverse employment impact. A study of minimum-wage floors in Britain (Stephen Machin and Alan Manning, 1994) reaches a similar conclusion.

This paper presents new evidence on the effect of minimum wages on establishment-level employment outcomes. We analyze the experiences of 410 fast-food restaurants in New Jersey and Pennsylvania following the increase in New Jersey's minimum wage from $4.25 to $5.05 per hour. Comparisons of employment, wages, and prices at stores in New Jersey and Pennsylvania before and after the rise offer a simple method for evaluating the effects of the minimum wage. Comparisons within New Jersey between initially high-wage stores (those paying more than the new minimum rate prior to its effective date) and other stores provide an alternative estimate of the impact of the new law.

In addition to the simplicity of our empirical methodology, several other features of

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[1]See Charles Brown et al. (1982, 1983) for surveys of this literature. A recent update (Allison J. Wellington, 1991) concludes that the employment effects of the minimum wage are negative but small: a 10-percent increase in the minimum is estimated to lower teenage employment rates by 0.06 percentage points.

the New Jersey law and our data set are also significant. First, the rise in the minimum wage occurred during a recession. The increase had been legislated two years earlier when the state economy was relatively healthy. By the time of the actual increase, the unemployment rate in New Jersey had risen substantially and last-minute political action almost succeeded in reducing the minimum-wage increase. It is unlikely that the effects of the higher minimum wage were obscured by a rising tide of general economic conditions.

Second, New Jersey is a relatively small state with an economy that is closely linked to nearby states. We believe that a control group of fast-food stores in eastern Pennsylvania forms a natural basis for comparison with the experiences of restaurants in New Jersey. Wage variation across stores in New Jersey, however, allows us to compare the experiences of high-wage and low-wage stores within New Jersey and to *test* the validity of the Pennsylvania control group. Moreover, since seasonal patterns of employment are similar in New Jersey and eastern Pennsylvania, as well as across high- and low-wage stores within New Jersey, our comparative methodology effectively "differences out" any seasonal employment effects.

Third, we successfully followed nearly 100 percent of stores from a first wave of interviews conducted just before the rise in the minimum wage (in February and March 1992) to a second wave conducted 7–8 months after (in November and December 1992). We have complete information on store closings and take account of employment changes at the closed stores in our analyses. We therefore measure the overall effect of the minimum wage on average employment, and not simply its effect on surviving establishments.

Our analysis of employment trends at stores that were open for business before the increase in the minimum wage ignores any potential effect of minimum wages on the rate of new store openings. To assess the likely magnitude of this effect we relate state-specific growth rates in the number of McDonald's fast-food outlets between 1986 and 1991 to measures of the relative minimum wage in each state.

I. The New Jersey Law

A bill signed into law in November 1989 raised the federal minimum wage from $3.35 per hour to $3.80 effective April 1, 1990, with a further increase to $4.25 per hour on April 1, 1991. In early 1990 the New Jersey legislature went one step further, enacting parallel increases in the state minimum wage for 1990 and 1991 and an increase to $5.05 per hour effective April 1, 1992. The scheduled 1992 increase gave New Jersey the highest state minimum wage in the country and was strongly opposed by business leaders in the state (see Bureau of National Affairs, *Daily Labor Report*, 5 May 1990).

In the two years between passage of the $5.05 minimum wage and its effective date, New Jersey's economy slipped into recession. Concerned with the potentially adverse impact of a higher minimum wage, the state legislature voted in March 1992 to phase in the 80-cent increase over two years. The vote fell just short of the margin required to override a gubernatorial veto, and the Governor allowed the $5.05 rate to go into effect on April 1 before vetoing the two-step legislation. Faced with the prospect of having to roll back wages for minimum-wage earners, the legislature dropped the issue. Despite a strong last-minute challenge, the $5.05 minimum rate took effect as originally planned.

II. Sample Design and Evaluation

Early in 1992 we decided to evaluate the impending increase in the New Jersey minimum wage by surveying fast-food restaurants in New Jersey and eastern Pennsylvania.[2] Our choice of the fast-food industry was driven by several factors. First, fast-food stores are a leading employer of low-wage workers: in 1987, franchised restaurants em-

[2]At the time we were uncertain whether the $5.05 rate would go into effect or be overridden.

TABLE 1—SAMPLE DESIGN AND RESPONSE RATES

	All	Stores in: NJ	PA
Wave 1, February 15–March 4, 1992:			
Number of stores in sample frame:[a]	473	364	109
Number of refusals:	63	33	30
Number interviewed:	410	331	79
Response rate (percentage):	86.7	90.9	72.5
Wave 2, November 5–December 31, 1992:			
Number of stores in sample frame:	410	331	79
Number closed:	6	5	1
Number under rennovation:	2	2	0
Number temporarily closed:[b]	2	2	0
Number of refusals:	1	1	0
Number interviewed:[c]	399	321	78

[a]Stores with working phone numbers only; 29 stores in original sample frame had disconnected phone numbers.
[b]Includes one store closed because of highway construction and one store closed because of a fire.
[c]Includes 371 phone interviews and 28 personal interviews of stores that refused an initial request for a phone interview.

ployed 25 percent of all workers in the restaurant industry (see U.S. Department of Commerce, 1990 table 13). Second, fast-food restaurants comply with minimum-wage regulations and would be expected to raise wages in response to a rise in the minimum wage. Third, the job requirements and products of fast-food restaurants are relatively homogeneous, making it easier to obtain reliable measures of employment, wages, and product prices. The absence of tips greatly simplifies the measurement of wages in the industry. Fourth, it is relatively easy to construct a sample frame of franchised restaurants. Finally, past experience (Katz and Krueger, 1992) suggested that fast-food restaurants have high response rates to telephone surveys.[3]

Based on these considerations we constructed a sample frame of fast-food restaurants in New Jersey and eastern Pennsylvania from the Burger King, KFC, Wendy's, and Roy Rogers chains.[4] The first wave of the survey was conducted by telephone in late February and early March 1992, a little over a month before the scheduled increase in New Jersey's minimum wage. The survey included questions on employment, starting wages, prices, and other store characteristics.[5]

Table 1 shows that 473 stores in our sample frame had working telephone numbers when we tried to reach them in February-March 1992. Restaurants were called as many as nine times to elicit a response. We obtained completed interviews (with some item nonresponse) from 410 of the restaurants, for an overall response rate of 87 percent. The response rate was higher in New Jersey (91 percent) than in Pennsylva-

[3]In a pilot survey Katz and Krueger (1992) obtained very low response rates from McDonald's restaurants. For this reason, McDonald's restaurants were excluded from Katz and Krueger's and our sample frames.

[4]The sample was derived from white-pages telephone listings for New Jersey and Pennsylvania as of February 1992.

[5]Copies of the questionnaires used in both waves of the survey are available from the authors upon request.

nia (72.5 percent) because our interviewer made fewer call-backs to nonrespondents in Pennsylvania.[6] In the analysis below we investigate possible biases associated with the degree of difficulty in obtaining the first-wave interview.

The second wave of the survey was conducted in November and December 1992, about eight months after the minimum-wage increase. Only the 410 stores that responded in the first wave were contacted in the second round of interviews. We successfully interviewed 371 (90 percent) of these stores by phone in November 1992. Because of a concern that nonresponding restaurants might have closed, we hired an interviewer to drive to each of the 39 nonrespondents and determine whether the store was still open, and to conduct a personal interview if possible. The interviewer discovered that six restaurants were permanently closed, two were temporarily closed (one because of a fire, one because of road construction), and two were under renovation.[7] Of the 29 stores open for business, all but one granted a request for a personal interview. As a result, we have second-wave interview data for 99.8 percent of the restaurants that responded in the first wave of the survey, and information on closure status for 100 percent of the sample.

Table 2 presents the means for several key variables in our data set, averaged over the subset of nonmissing responses for each variable. In constructing the means, employment in wave 2 is set to 0 for the permanently closed stores but is treated as missing for the temporarily closed stores. (Full-time-equivalent [FTE] employment was calculated as the number of full-time workers [including managers] plus 0.5 times the number of part-time workers.)[8] Means are presented separately for stores in New Jersey and Pennsylvania, along with t statistics for the null hypothesis that the means are equal in the two states.

Rows 1a–e show the distribution of stores by chain and ownership status (company-owned versus franchisee-owned). The Burger King, Roy Rogers, and Wendy's stores in our sample have similar average food prices, store hours, and employment levels. The KFC stores are smaller and are open for fewer hours. They also offer a more expensive main course than stores in the other chains (chicken vs. hamburgers).

In wave 1, average employment was 23.3 full-time equivalent workers per store in Pennsylvania, compared with an average of 20.4 in New Jersey. Starting wages were very similar among stores in the two states, although the average price of a "full meal" (medium soda, small fries, and an entree) was significantly higher in New Jersey. There were no significant cross-state differences in average hours of operation, the fraction of full-time workers, or the prevalence of bonus programs to recruit new workers.[9]

The average starting wage at fast-food restaurants in New Jersey increased by 10 percent following the rise in the minimum wage. Further insight into this change is provided in Figure 1, which shows the distributions of starting wages in the two states before and after the rise. In wave 1, the distributions in New Jersey and Pennsylvania were very similar. By wave 2 virtually all

[6] Response rates per call-back were almost identical in the two states. Among New Jersey stores, 44.5 percent responded on the first call, and 72.0 percent responded after at most two call-backs. Among Pennsylvania stores 42.2 percent responded on the first call, and 71.6 percent responded after at most two call-backs.

[7] As of April 1993 the store closed because of road construction and one of the stores closed for renovation had reopened. The store closed by fire was open when our telephone interviewer called in November 1992 but refused the interview. By the time of the follow-up personal interview a mall fire had closed the store.

[8] We discuss the sensitivity of our results to alternative assumptions on the measurement of employment in Section III-C.

[9] These programs offer current employees a cash "bounty" for recruiting any new employee who stays on the job for a minimum period of time. Typical bounties are $50–$75. Recruiting programs that award the recruiter with an "employee of the month" designation or other noncash bonuses are excluded from our tabulations.

TABLE 2—MEANS OF KEY VARIABLES

Variable	Stores in: NJ	PA	t[a]
1. *Distribution of Store Types (percentages):*			
a. Burger King	41.1	44.3	−0.5
b. KFC	20.5	15.2	1.2
c. Roy Rogers	24.8	21.5	0.6
d. Wendy's	13.6	19.0	−1.1
e. Company-owned	34.1	35.4	−0.2
2. *Means in Wave 1:*			
a. FTE employment	20.4	23.3	−2.0
	(0.51)	(1.35)	
b. Percentage full-time employees	32.8	35.0	−0.7
	(1.3)	(2.7)	
c. Starting wage	4.61	4.63	−0.4
	(0.02)	(0.04)	
d. Wage = $4.25 (percentage)	30.5	32.9	−0.4
	(2.5)	(5.3)	
e. Price of full meal	3.35	3.04	4.0
	(0.04)	(0.07)	
f. Hours open (weekday)	14.4	14.5	−0.3
	(0.2)	(0.3)	
g. Recruiting bonus	23.6	29.1	−1.0
	(2.3)	(5.1)	
3. *Means in Wave 2:*			
a. FTE employment	21.0	21.2	−0.2
	(0.52)	(0.94)	
b. Percentage full-time employees	35.9	30.4	1.8
	(1.4)	(2.8)	
c. Starting wage	5.08	4.62	10.8
	(0.01)	(0.04)	
d. Wage = $4.25 (percentage)	0.0	25.3	—
		(4.9)	
e. Wage = $5.05 (percentage)	85.2	1.3	36.1
	(2.0)	(1.3)	
f. Price of full meal	3.41	3.03	5.0
	(0.04)	(0.07)	
g. Hours open (weekday)	14.4	14.7	−0.8
	(0.2)	(0.3)	
h. Recruiting bonus	20.3	23.4	−0.6
	(2.3)	(4.9)	

Notes: See text for definitions. Standard errors are given in parentheses.
[a]Test of equality of means in New Jersey and Pennsylvania.

restaurants in New Jersey that had been paying less than $5.05 per hour reported a starting wage equal to the new rate. Interestingly, the minimum-wage increase had no apparent "spillover" on higher-wage restaurants in the state: the mean percentage wage change for these stores was −3.1 percent.

Despite the increase in wages, full-time-equivalent employment *increased* in New Jersey relative to Pennsylvania. Whereas New Jersey stores were initially smaller, employment gains in New Jersey coupled with losses in Pennsylvania led to a small and statistically insignificant interstate

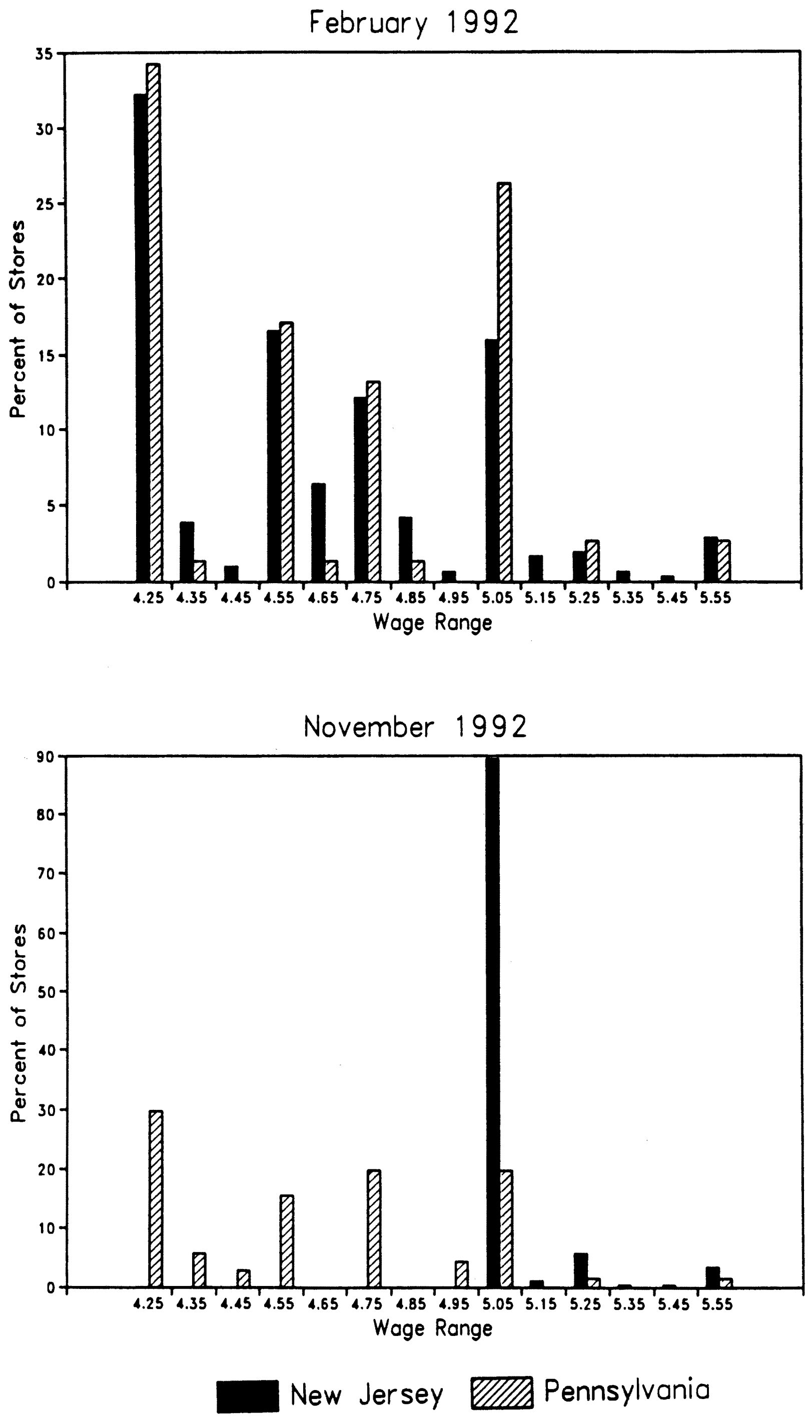

FIGURE 1. DISTRIBUTION OF STARTING WAGE RATES

difference in wave 2. Only two other variables show a relative change between waves 1 and 2: the fraction of full-time employees and the price of a meal. Both variables increased in New Jersey relative to Pennsylvania.

We can assess the reliability of our survey questionnaire by comparing the responses of 11 stores that were inadvertently interviewed twice in the first wave of the survey.[10] Assuming that measurement errors in the two interviews are independent of each other and independent of the true variable, the correlation between responses gives an estimate of the "reliability ratio" (the ratio of the variance of the signal to the combined variance of the signal and noise). The estimated reliability ratios are fairly high, ranging from 0.70 for full-time equivalent employment to 0.98 for the price of a meal.[11]

We have also checked whether stores with missing data for any key variables are different from restaurants with complete responses. We find that stores with missing data on employment, wages, or prices are similar in other respects to stores with complete data. There is a significant size differential associated with the likelihood of the store closing after wave 1. The six stores that closed were smaller than other stores (with an average employment of only 12.4 full-time-equivalent employees in wave 1).[12]

III. Employment Effects of the Minimum-Wage Increase

A. *Differences in Differences*

Table 3 summarizes the levels and changes in average employment per store in our survey. We present data by state in columns (i) and (ii), and for stores in New Jersey classified by whether the starting wage in wave 1 was exactly $4.25 per hour [column (iv)] between $4.26 and $4.99 per hour [column (v)] or $5.00 or more per hour [column (vi)]. We also show the differences in average employment between New Jersey and Pennsylvania stores [column (iii)] and between stores in the various wage ranges in New Jersey [columns (vii)–(viii)].

Row 3 of the table presents the changes in average employment between waves 1 and 2. These entries are simply the differences between the averages for the two waves (i.e., row 2 minus row 1). An alternative estimate of the change is presented in row 4: here we have computed the change in employment over the subsample of stores that reported valid employment data in both waves. We refer to this group of stores as the balanced subsample. Finally, row 5 presents the average change in employment in the balanced subsample, treating wave-2 employment at the four temporarily closed stores as zero, rather than as missing.

As noted in Table 2, New Jersey stores were initially smaller than their Pennsylvania counterparts but grew relative to Pennsylvania stores after the rise in the minimum wage. The relative gain (the "difference in differences" of the changes in employment) is 2.76 FTE employees (or 13 percent), with a t statistic of 2.03. Inspection of the averages in rows 4 and 5 shows that the relative change between New Jersey and Pennsylvania stores is virtually identical when the analysis is restricted to the balanced subsample, and it is only slightly smaller when wave-2 employment at the temporarily closed stores is treated as zero.

Within New Jersey, employment expanded at the low-wage stores (those paying $4.25 per hour in wave 1) and contracted at the high-wage stores (those paying $5.00 or more per hour). Indeed, the average change in employment at the high-wage stores (−2.16 FTE employees) is almost identical to the change among Pennsylvania stores (−2.28 FTE employees). Since high-wage stores in New Jersey should have been

[10] These restaurants were interviewed twice because their phone numbers appeared in more than one phone book, and neither the interviewer nor the respondent noticed that they were previously interviewed.

[11] Similar reliability ratios for very similar questions were obtained by Katz and Krueger (1992).

[12] A probit analysis of the probability of closure shows that the initial size of the store is a significant predictor of closure. The level of starting wages has a numerically small and statistically insignificant coefficient in the probit model.

largely unaffected by the new minimum wage, this comparison provides a specification test of the validity of the Pennsylvania control group. The test is clearly passed. Regardless of whether the affected stores are compared to stores in Pennsylvania or high-wage stores in New Jersey, the estimated employment effect of the minimum wage is similar.

The results in Table 3 suggest that employment contracted between February and November of 1992 at fast-food stores that were unaffected by the rise in the minimum wage (stores in Pennsylvania and stores in New Jersey paying \$5.00 per hour or more in wave 1). We suspect that the reason for this contraction was the continued worsening of the economies of the middle-Atlantic states during 1992.[13] Unemployment rates in New Jersey, Pennsylvania, and New York all trended upward between 1991 and 1993, with a larger increase in New Jersey than Pennsylvania during 1992. Since sales of franchised fast-food restaurants are procyclical, the rise in unemployment would be expected to lower fast-food employment in the absence of other factors.[14]

B. *Regression-Adjusted Models*

The comparisons in Table 3 make no allowance for other sources of variation in employment growth, such as differences across chains. These are incorporated in the estimates in Table 4. The entries in this table are regression coefficients from models of the form:

$$\text{(1a)} \qquad \Delta E_i = a + \mathbf{b}\mathbf{X}_i + c\mathrm{NJ}_i + \varepsilon_i$$

or

$$\text{(1b)} \qquad \Delta E_i = a' + \mathbf{b}'\mathbf{X}_i + c'\mathrm{GAP}_i + \varepsilon_i'$$

where ΔE_i is the change in employment from wave 1 to wave 2 at store i, $\mathbf{X}_i$ is a set of characteristics of store i, and NJ_i is a dummy variable that equals 1 for stores in New Jersey. GAP_i is an alternative measure of the impact of the minimum wage at store i based on the initial wage at that store (W_{1i}):

$$\begin{aligned} \mathrm{GAP}_i &= 0 && \text{for stores in Pennsylvania} \\ &= 0 && \text{for stores in New Jersey with } W_{1i} \geq \$5.05 \\ &= (5.05 - W_{1i})/W_{1i} && \text{for other stores in New Jersey.} \end{aligned}$$

GAP_i is the proportional increase in wages at store i necessary to meet the new minimum rate. Variation in GAP_i reflects both the New Jersey–Pennsylvania contrast and differences within New Jersey based on reported starting wages in wave 1. Indeed, the value of GAP_i is a strong predictor of the actual proportional wage change between waves 1 and 2 ($R^2 = 0.75$), and conditional on GAP_i there is no difference in wage behavior between stores in New Jersey and Pennsylvania.[15]

The estimate in column (i) of Table 4 is directly comparable to the simple difference-in-differences of employment changes in column (iv), row 4 of Table 3. The discrepancy between the two estimates is due to the restricted sample in Table 4. In Table 4 and the remaining tables in this section we restrict our analysis to the set of stores with available employment *and* wage data in both waves of the

[13] An alternative possibility is that seasonal factors produce higher employment at fast-food restaurants in February and March than in November and December. An analysis of national employment data for food preparation and service workers, however, shows higher average employment in the fourth quarter than in the first quarter.

[14] To investigate the cyclicality of fast-food restaurant sales we regressed the year-to-year change in U.S. sales of the McDonald's restaurant chain from 1976–1991 on the corresponding change in the unemployment rate. The regression results show that a 1-percentage-point increase in the unemployment rate reduces sales by \$257 million, with a t statistic of 3.0.

[15] A regression of the proportional wage change between waves 1 and 2 on GAP_i has a coefficient of 1.03.

TABLE 3—AVERAGE EMPLOYMENT PER STORE BEFORE AND AFTER THE RISE IN NEW JERSEY MINIMUM WAGE

	Stores by state			Stores in New Jersey[a]			Differences within NJ[b]	
Variable	PA (i)	NJ (ii)	Difference, NJ − PA (iii)	Wage = \$4.25 (iv)	Wage = \$4.26–\$4.99 (v)	Wage ≥ \$5.00 (vi)	Low–high (vii)	Midrange–high (viii)
1. FTE employment before, all available observations	23.33 (1.35)	20.44 (0.51)	−2.89 (1.44)	19.56 (0.77)	20.08 (0.84)	22.25 (1.14)	−2.69 (1.37)	−2.17 (1.41)
2. FTE employment after, all available observations	21.17 (0.94)	21.03 (0.52)	−0.14 (1.07)	20.88 (1.01)	20.96 (0.76)	20.21 (1.03)	0.67 (1.44)	0.75 (1.27)
3. Change in mean FTE employment	−2.16 (1.25)	0.59 (0.54)	2.76 (1.36)	1.32 (0.95)	0.87 (0.84)	−2.04 (1.14)	3.36 (1.48)	2.91 (1.41)
4. Change in mean FTE employment, balanced sample of stores[c]	−2.28 (1.25)	0.47 (0.48)	2.75 (1.34)	1.21 (0.82)	0.71 (0.69)	−2.16 (1.01)	3.36 (1.30)	2.87 (1.22)
5. Change in mean FTE employment, setting FTE at temporarily closed stores to 0[d]	−2.28 (1.25)	0.23 (0.49)	2.51 (1.35)	0.90 (0.87)	0.49 (0.69)	−2.39 (1.02)	3.29 (1.34)	2.88 (1.23)

Notes: Standard errors are shown in parentheses. The sample consists of all stores with available data on employment. FTE (full-time-equivalent) employment counts each part-time worker as half a full-time worker. Employment at six closed stores is set to zero. Employment at four temporarily closed stores is treated as missing.

[a]Stores in New Jersey were classified by whether starting wage in wave 1 equals \$4.25 per hour ($N = 101$), is between \$4.26 and \$4.99 per hour ($N = 140$), or is \$5.00 per hour or higher ($N = 73$).

[b]Difference in employment between low-wage (\$4.25 per hour) and high-wage (≥ \$5.00 per hour) stores; and difference in employment between midrange (\$4.26–\$4.99 per hour) and high-wage stores.

[c]Subset of stores with available employment data in wave 1 and wave 2.

[d]In this row only, wave-2 employment at four temporarily closed stores is set to 0. Employment changes are based on the subset of stores with available employment data in wave 1 and wave 2.

TABLE 4—REDUCED-FORM MODELS FOR CHANGE IN EMPLOYMENT

	Model				
Independent variable	(i)	(ii)	(iii)	(iv)	(v)
1. New Jersey dummy	2.33 (1.19)	2.30 (1.20)	—	—	—
2. Initial wage gap[a]	—	—	15.65 (6.08)	14.92 (6.21)	11.91 (7.39)
3. Controls for chain and ownership[b]	no	yes	no	yes	yes
4. Controls for region[c]	no	no	no	no	yes
5. Standard error of regression	8.79	8.78	8.76	8.76	8.75
6. Probability value for controls[d]	—	0.34	—	0.44	0.40

Notes: Standard errors are given in parentheses. The sample consists of 357 stores with available data on employment and starting wages in waves 1 and 2. The dependent variable in all models is change in FTE employment. The mean and standard deviation of the dependent variable are −0.237 and 8.825, respectively. All models include an unrestricted constant (not reported).

[a]Proportional increase in starting wage necessary to raise starting wage to new minimum rate. For stores in Pennsylvania the wage gap is 0.

[b]Three dummy variables for chain type and whether or not the store is company-owned are included.

[c]Dummy variables for two regions of New Jersey and two regions of eastern Pennsylvania are included.

[d]Probability value of joint F test for exclusion of all control variables.

survey. This restriction results in a slightly smaller estimate of the relative increase in employment in New Jersey.

The model in column (ii) introduces a set of four control variables: dummies for three of the chains and another dummy for company-owned stores. As shown by the probability values in row 6, these covariates add little to the model and have no effect on the size of the estimated New Jersey dummy.

The specifications in columns (iii)–(v) use the GAP variable to measure the effect of the minimum wage. This variable gives a slightly better fit than the simple New Jersey dummy, although its implications for the New Jersey–Pennsylvania comparison are similar. The mean value of GAP_i among New Jersey stores is 0.11. Thus the estimate in column (iii) implies a 1.72 increase in FTE employment in New Jersey relative to Pennsylvania.

Since GAP_i varies within New Jersey, it is possible to add both GAP_i and NJ_i to the employment model. The estimated coefficient of the New Jersey dummy then provides a test of the Pennsylvania control group. When we estimate these models, the coefficient of the New Jersey dummy is insignificant (with t ratios of 0.3–0.7), implying that inferences about the effect of the minimum wage are similar whether the comparison is made across states or across stores in New Jersey with higher and lower initial wages.

An even stronger test is provided in column (v), where we have added dummies representing three regions of New Jersey (North, Central, and South) and two regions of eastern Pennsylvania (Allentown-Easton and the northern suburbs of Philadelphia). These dummies control for any region-specific demand shocks and identify the effect of the minimum wage by comparing employment changes at higher- and lower-wage stores within the same region of New Jersey. The probability value in row 6 shows no evidence of regional components in employment growth. The addition of the region dummies attenuates the GAP coefficient and raises its standard error, however, making it no longer possible to reject the null hypothesis of a zero employment effect of the minimum wage. One explanation for this attenuation is the presence of measurement error in the starting wage. Even if employment growth has no regional component, the addition of region dummies will lead to some attenuation of the estimated GAP coefficient if some of the true variation in GAP is explained by region. Indeed, calculations based on the estimated reliability of the GAP variable (from the set of 11 double interviews) suggest that the fall in the estimated GAP coefficient from column (iv) to column (v) is just equal to the expected change attributable to measurement error.[16]

We have also estimated the models in Table 4 using as a dependent variable the proportional change in employment at each store.[17] The estimated coefficients of the New Jersey dummy and the GAP variable are uniformly positive in these models but insignificantly different from 0 at conventional levels. The implied employment effects of the minimum wage are also smaller when the dependent variable is expressed in proportional terms. For example, the GAP coefficient in column (iii) of Table 4 implies that the increase in minimum wages raised employment at New Jersey stores that were initially paying $4.25 per hour by 14 percent. The estimated GAP coefficient from a corresponding proportional model implies an effect of only 7 percent. The difference is attributable to heterogeneity in the effect of the minimum wage at larger and smaller stores. Weighted versions of the proportional-change models (using initial employment as a weight) give rise to wage elastici-

[16]In a regression model without other controls the expected attenuation of the GAP coefficient due to measurement error is the reliability ratio of GAP (γ_0), which we estimate at 0.70. The expected attenuation factor when region dummies are added to the model is $\gamma_1 = (\gamma_0 - R^2)/(1 - R^2)$, where R^2 is the R-square statistic of a regression of GAP on region effects (equal to 0.30). Thus, we expect the estimated GAP coefficient to fall by a factor of $\gamma_1/\gamma_0 = 0.8$ when region dummies are added to a regression model.

[17]These specifications are reported in table 4 of Card and Krueger (1993).

ties similar to the elasticities implied by the estimates in Table 4 (see below).

C. *Specification Tests*

The results in Tables 3 and 4 seem to contradict the standard prediction that a rise in the minimum wage will reduce employment. Table 5 presents some alternative specifications that probe the robustness of this conclusion. For completeness, we report estimates of models for the change in employment [columns (i) and (ii)] and estimates of models for the proportional change in employment [columns (iii) and (iv)].[18] The first row of the table reproduces the "base specification" from columns (ii) and (iv) of Table 4. (Note that these models include chain dummies and a dummy for company-owned stores). Row 2 presents an alternative set of estimates when we set wave-2 employment at the temporarily closed stores to 0 (expanding our sample size by 4). This change has a small attenuating effect on the coefficient of the New Jersey dummy (since all four stores are in New Jersey) but less effect on the GAP coefficient (since the size of GAP is uncorrelated with the probability of a temporary closure within New Jersey).

Rows 3–5 present estimation results using alternative measures of full-time-equivalent employment. In row 3, employment is redefined to exclude management employees. This change has no effect relative to the base specification. In rows 4 and 5, we include managers in FTE employment but reweight part-time workers as either 40 percent or 60 percent of full-time workers (instead of 50 percent).[19] These changes have little effect on the models for the level of employment but yield slightly smaller point estimates in the proportional-employment-change models.

In row 6 we present estimates obtained from a subsample that excludes 35 stores in towns along the New Jersey shore. The exclusion of these stores, which may have a different seasonal pattern than other stores in our sample, leads to slightly larger minimum-wage effects. A similar finding emerges in row 7 when we add a set of dummy variables that indicate the week of the wave-2 interview.[20]

As noted earlier, we made an extra effort to obtain responses from New Jersey stores in the first wave of our survey. The fraction of stores called three or more times to obtain an interview was higher in New Jersey than in Pennsylvania. To check the sensitivity of our results to this sampling feature, we reestimated our models on a subsample that excludes any stores that were called back more than twice. The results, in row 8, are very similar to the base specification.

Row 9 presents weighted estimation results for the proportional-employment-change models, using as weights the initial levels of employment in each store. Since the proportional change in average employment is an employment-weighted average of the proportional changes at each store, a weighted version of the proportional-change model should give rise to elasticities that are similar to the implied elasticities arising from the levels models. Consistent with this expectation, the weighted estimates are larger than the unweighted estimates, and significantly different from 0 at conventional levels. The weighted estimate of the New Jersey dummy (0.13) implies a 13-percent relative increase in New Jersey employment —the same proportional employment effect implied by the simple difference-in-differences in Table 3. Similarly, the weighted estimate of the GAP coefficient in the proportional-change model (0.81) is close to

[18]The proportional change in employment is defined as the change in employment divided by the average level of employment in waves 1 and 2. This results in very similar coefficients but smaller standard errors than the alternative of dividing by wave-1 employment. For closed stores we set the proportional change in employment to -1.

[19]Analysis of the 1991 Current Population Survey reveals that part-time workers in the restaurant industry work about 46 percent as many hours as full-time workers. Katz and Krueger (1992) report that the ratio of part-time workers' hours to full-time workers' hours in the fast-food industry is 0.57.

[20]We also added dummies for the interview dates for the wave-1 survey, but these were insignificant and did not change the estimated minimum-wage effects.

TABLE 5—SPECIFICATION TESTS OF REDUCED-FORM EMPLOYMENT MODELS

	Change in employment		Proportional change in employment	
Specification	NJ dummy (i)	Gap measure (ii)	NJ dummy (iii)	Gap measure (iv)
1. Base specification	2.30 (1.19)	14.92 (6.21)	0.05 (0.05)	0.34 (0.26)
2. Treat four temporarily closed stores as permanently closed[a]	2.20 (1.21)	14.42 (6.31)	0.04 (0.05)	0.34 (0.27)
3. Exclude managers in employment count[b]	2.34 (1.17)	14.69 (6.05)	0.05 (0.07)	0.28 (0.34)
4. Weight part-time as 0.4 × full-time[c]	2.34 (1.20)	15.23 (6.23)	0.06 (0.06)	0.30 (0.33)
5. Weight part-time as 0.6 × full-time[d]	2.27 (1.21)	14.60 (6.26)	0.04 (0.06)	0.17 (0.29)
6. Exclude stores in NJ shore area[e]	2.58 (1.19)	16.88 (6.36)	0.06 (0.05)	0.42 (0.27)
7. Add controls for wave-2 interview date[f]	2.27 (1.20)	15.79 (6.24)	0.05 (0.05)	0.40 (0.26)
8. Exclude stores called more than twice in wave 1[g]	2.41 (1.28)	14.08 (7.11)	0.05 (0.05)	0.31 (0.29)
9. Weight by initial employment[h]	—	—	0.13 (0.05)	0.81 (0.26)
10. Stores in towns around Newark[i]	—	33.75 (16.75)	—	0.90 (0.74)
11. Stores in towns around Camden[j]	—	10.91 (14.09)	—	0.21 (0.70)
12. Pennsylvania stores only[k]	—	−0.30 (22.00)	—	−0.33 (0.74)

Notes: Standard errors are given in parentheses. Entries represent estimated coefficient of New Jersey dummy [columns (i) and (iii)] or initial wage gap [columns (ii) and (iv)] in regression models for the change in employment or the percentage change in employment. All models also include chain dummies and an indicator for company-owned stores.

[a] Wave-2 employment at four temporarily closed stores is set to 0 (rather than missing).

[b] Full-time equivalent employment excludes managers and assistant managers.

[c] Full-time equivalent employment equals number of managers, assistant managers, and full-time nonmanagement workers, plus 0.4 times the number of part-time nonmanagement workers.

[d] Full-time equivalent employment equals number of managers, assistant managers, and full-time nonmanagement workers, plus 0.6 times the number of part-time nonmanagement workers.

[e] Sample excludes 35 stores located in towns along the New Jersey shore.

[f] Models include three dummy variables identifying week of wave-2 interview in November-December 1992.

[g] Sample excludes 70 stores (69 in New Jersey) that were contacted three or more times before obtaining the wave-1 interview.

[h] Regression model is estimated by weighted least squares, using employment in wave 1 as a weight.

[i] Subsample of 51 stores in towns around Newark.

[j] Subsample of 54 stores in town around Camden.

[k] Subsample of Pennsylvania stores only. Wage gap is defined as percentage increase in starting wage necessary to raise starting wage to $5.05.

the implied elasticity of employment with respect to wages from the basic levels specification in row 1, column (ii).[21] These findings suggest that the proportional effect of the rise in the minimum wage was concentrated among larger stores.

One explanation for our finding that a rise in the minimum wage has a positive employment effect is that unobserved demand shocks within New Jersey outweighed the negative employment effect of the minimum wage. To address this possibility, rows 10 and 11 present estimation results based on subsamples of stores in two narrowly defined areas: towns around Newark (row 10) and towns around Camden (row 11). In each case the sample area is identified by the first three digits of the store's zip code.[22] Within both areas the change in employment is positively correlated with the GAP variable, although in neither case is the effect statistically significant. To the extent that fast-food product market conditions are constant within local areas, these results suggest that our findings are not driven by unobserved demand shocks. Our analysis of price changes (reported below) also supports this conclusion.

A final specification check is presented in row 12 of Table 5. In this row we exclude stores in New Jersey and (incorrectly) define the GAP variable for Pennsylvania stores as the proportional increase in wages necessary to raise the wage to $5.05 per hour. In principle the size of the wage gap for stores in Pennsylvania should have no systematic relation with employment growth. In practice, this is the case. There is no indication that the wage gap is spuriously related to employment growth.

We have also investigated whether the first-differenced specification used in our employment models is appropriate. A first-differenced model implies that the *level* of employment in period t is related to the lagged level of employment with a coefficient of 1. If short-run employment fluctuations are smoothed, however, the true coefficient of lagged employment may be less than 1. Imposing the assumption of a unit coefficient may then lead to biases. To test the first-differenced specification we reestimated models for the change in employment including wave-1 employment as an additional explanatory variable. To overcome any mechanical correlation between base-period employment and the change in employment (attributable to measurement error) we instrumented wave-1 employment with the number of cash registers in the store in wave 1 and the number of registers in the store that were open at 11:00 A.M. In all of the specifications the coefficient of wave-1 employment is close to zero. For example, in a specification including the GAP variable and ownership and chain dummies, the coefficient of wave-1 employment is 0.04, with a standard error of 0.24. We conclude that the first-differenced specification is appropriate.

D. *Full-Time and Part-Time Substitution*

Our analysis so far has concentrated on full-time-equivalent employment and ignored possible changes in the distribution of full- and part-time workers. An increase in the minimum wage could lead to an increase in full-time employment relative to part-time employment for at least two reasons. First, in a conventional model one would expect a minimum-wage increase to induce employers to substitute skilled workers and capital for minimum-wage workers. Full-time workers in fast-food restaurants are typically older and may well possess higher skills than part-time workers. Thus, a conventional model predicts that stores may respond to an increase in the minimum wage by increasing the proportion of full-time workers. Nevertheless, 81 percent of restaurants paid full-time and part-time

[21] Assuming average employment of 20.4 in New Jersey, the 14.92 GAP coefficient in row 1, column (ii) implies an employment elasticity of 0.73.

[22] The "070" three-digit zip-code area (around Newark) and the "080" three-digit zip-code area (around Camden) have by far the largest numbers of stores among three-digit zip-code areas in New Jersey, and together they account for 36 percent of New Jersey stores in our sample.

TABLE 6—EFFECTS OF MINIMUM-WAGE INCREASE ON OTHER OUTCOMES

Outcome measure	Mean change in outcome: NJ (i)	PA (ii)	NJ − PA (iii)	Regression of change in outcome variable on: NJ dummy (iv)	Wage gap[a] (v)	Wage gap[b] (vi)
Store Characteristics:						
1. Fraction full-time workers[c] (percentage)	2.64	−4.65	7.29	7.30	33.64	20.28
	(1.71)	(3.80)	(4.17)	(3.96)	(20.95)	(24.34)
2. Number of hours open per weekday	−0.00	0.11	−0.11	−0.11	−0.24	0.04
	(0.06)	(0.08)	(0.10)	(0.12)	(0.65)	(0.76)
3. Number of cash registers	−0.04	0.13	−0.17	−0.18	−0.31	0.29
	(0.04)	(0.10)	(0.11)	(0.10)	(0.53)	(0.62)
4. Number of cash registers open at 11:00 A.M.	−0.03	−0.20	0.17	0.17	0.15	−0.47
	(0.05)	(0.08)	(0.10)	(0.12)	(0.62)	(0.74)
Employee Meal Programs:						
5. Low-price meal program (percentage)	−4.67	−1.28	−3.39	−2.01	−30.31	−33.15
	(2.65)	(3.86)	(4.68)	(5.63)	(29.80)	(35.04)
6. Free meal program (percentage)	8.41	6.41	2.00	0.49	29.90	36.91
	(2.17)	(3.33)	(3.97)	(4.50)	(23.75)	(27.90)
7. Combination of low-price and free meals (percentage)	−4.04	−5.13	1.09	1.20	−11.87	−19.19
	(1.98)	(3.11)	(3.69)	(4.32)	(22.87)	(26.81)
Wage Profile:						
8. Time to first raise (weeks)	3.77	1.26	2.51	2.21	4.02	−5.10
	(0.89)	(1.97)	(2.16)	(2.03)	(10.81)	(12.74)
9. Usual amount of first raise (cents)	−0.01	−0.02	0.01	0.01	0.03	0.03
	(0.01)	(0.02)	(0.02)	(0.02)	(0.11)	(0.11)
10. Slope of wage profile (percent per week)	−0.10	−0.11	0.01	0.01	−0.09	−0.08
	(0.04)	(0.09)	(0.10)	(0.10)	(0.56)	(0.57)

Notes: Entries in columns (i) and (ii) represent mean changes in the outcome variable indicated by the row heading for stores with available data on the outcome in waves 1 and 2. Entries in columns (iv)–(vi) represent estimated regression coefficients of indicated variable (NJ dummy or initial wage gap) in models for the change in the outcome variable. Regression models include chain dummies and an indicator for company-owned stores.

[a]The wage gap is the proportional increase in starting wage necessary to raise the wage to the new minimum rate. For stores in Pennsylvania, the wage gap is zero.

[b]Models in column (vi) include dummies for two regions of New Jersey and two regions of eastern Pennsylvania.

[c]Fraction of part-time employees in total full-time-equivalent employment.

workers exactly the same starting wage in wave 1 of our survey.[23] This suggests either that full-time workers have the same skills as part-time workers or that equity concerns lead restaurants to pay equal wages for unequally productive workers. If full-time workers are more productive (but equally paid), there may be a second reason for stores to substitute full-time workers for part-time workers; namely, a minimum-wage increase enables the industry to attract more full-time workers, and stores would naturally want to hire a greater proportion of full-time workers if they are more productive.

Row 1 of Table 6 presents the mean changes in the proportion of full-time work-

[23]In the other 19 percent of stores, full-time workers are paid more, typically 10 percent more.

ers in New Jersey and Pennsylvania between waves 1 and 2 of our survey, and coefficient estimates from regressions of the change in the proportion of full-time workers on the wage-gap variable, chain dummies, a company-ownership dummy, and region dummies [in column (vi)]. The results are ambiguous. The fraction of full-time workers increased in New Jersey relative to Pennsylvania by 7.3 percent (t ratio = 1.84), but regressions on the wage-gap variable show no significant shift in the fraction of full-time workers.[24]

E. *Other Employment-Related Measures*

Rows 2–4 of Table 6 present results for other outcome variables that we expect to be related to the level of restaurant employment. In particular, we examine whether the rise in the minimum wage is associated with a change in the number of hours a restaurant is open on a weekday, the number of cash registers in the restaurant, and the number of cash registers typically in operation in the restaurant at 11:00 A.M. Consistent with our employment results, none of these variables shows a statistically significant decline in New Jersey relative to Pennsylvania. Similarly, regressions including the gap variable provide no evidence that the minimum-wage increase led to a systematic change in any of these variables [see columns (v) and (vi)].

IV. Nonwage Offsets

One explanation of our finding that a rise in the minimum wage does not lower employment is that restaurants can offset the effect of the minimum wage by reducing nonwage compensation. For example, if workers value fringe benefits and wages equally, employers can simply reduce the level of fringe benefits by the amount of the minimum-wage increase, leaving their employment costs unchanged. The main fringe benefits for fast-food employees are free and reduced-price meals. In the first wave of our survey about 19 percent of fast-food restaurants offered workers free meals, 72 percent offered reduced-price meals, and 9 percent offered a combination of both free and reduced-price meals. Low-price meals are an obvious fringe benefit to cut if the minimum-wage increase forces restaurants to pay higher wages.

Rows 5 and 6 of Table 6 present estimates of the effect of the minimum-wage increase on the incidence of free meals and reduced-price meals. The proportion of restaurants offering reduced-price meals fell in both New Jersey and Pennsylvania after the minimum wage increased, with a somewhat greater decline in New Jersey. Contrary to an offset story, however, the reduction in reduced-price meal programs was accompanied by an *increase* in the fraction of stores offering free meals. Relative to stores in Pennsylvania, New Jersey employers actually shifted toward more generous fringe benefits (i.e., free meals rather than reduced-price meals). However, the relative shift is not statistically significant.

We continue to find a statistically insignificant effect of the minimum-wage increase on the likelihood of receiving free or reduced-price meals in columns (v) and (vi), where we report coefficient estimates of the GAP variable from regression models for the change in the incidence of these programs. The results provide no evidence that employers offset the minimum-wage increase by reducing free or reduced-price meals.

Another possibility is that employers responded to the increase in the minimum wage by reducing on-the-job training and flattening the tenure–wage profile (see Jacob Mincer and Linda Leighton, 1981). Indeed, one manager told our interviewer in wave 1 that her workers were forgoing ordinary scheduled raises because the minimum wage was about to rise, and this would provide a raise for all her workers. To determine whether this phenomenon occurred more generally, we analyzed store managers' responses to questions on the amount

[24] Within New Jersey, the fraction of full-time employees increased about as quickly at stores with higher and lower wages in wave 1.

of time before a normal wage increase and the usual amount of such raises. In rows 8 and 9 we report the average changes between waves 1 and 2 for these two variables, as well as regression coefficients from models that include the wage-gap variable.[25] Although the average time to the first pay raise increased by 2.5 weeks in New Jersey relative to Pennsylvania, the increase is not statistically significant. Furthermore, there is only a trivial difference in the relative change in the amount of the first pay increment between New Jersey and Pennsylvania stores.

Finally, we examined a related variable: the "slope" of the wage profile, which we measure by the ratio of the typical first raise to the amount of time until the first raise is given. As shown in row 10, the slope of the wage profile flattened in both New Jersey and Pennsylvania, with no significant relative difference between states. The change in the slope is also uncorrelated with the GAP variable. In summary, we can find no indication that New Jersey employers changed either their fringe benefits or their wage profiles to offset the rise in the minimum wage.[26]

V. Price Effects of the Minimum-Wage Increase

A final issue we examine is the effect of the minimum wage on the prices of meals at fast-food restaurants. A competitive model of the fast-food industry implies that an increase in the minimum wage will lead to an increase in product prices. If we assume constant returns to scale in the industry, the increase in price should be proportional to the share of minimum-wage labor in total factor cost. The average restaurant in New Jersey initially paid about half its workers less than the new minimum wage. If wages rose by roughly 15 percent for these workers, and if labor's share of total costs is 30 percent, we would expect prices to rise by about 2.2 percent ($=0.15\times0.5\times0.3$) due to the minimum-wage rise.[27]

In each wave of our survey we asked managers for the prices of three standard items: a medium soda, a small order of french fries, and a main course. The main course was a basic hamburger at Burger King, Roy Rogers, and Wendy's restaurants, and two pieces of chicken at KFC stores. We define "full meal" price as the after-tax price of a medium soda, a small order of french fries, and a main course.

Table 7 presents reduced-form estimates of the effect of the minimum-wage increase on prices. The dependent variable in these models is the change in the logarithm of the price of a full meal at each store. The key independent variable is either a dummy indicating whether the store is located in New Jersey or the proportional wage increase required to meet the minimum wage (the GAP variable defined above).

The estimated New Jersey dummy in column (i) shows that after-tax meal prices rose 3.2-percent faster in New Jersey than in Pennsylvania between February and November 1992.[28] The effect is slightly larger controlling for chain and company-ownership [see column (ii)]. Since the New Jersey sales tax rate fell by 1 percentage point between the waves of our survey, these estimates suggest that pretax prices rose 4-percent faster as a result of the

[25] In wave 1, the average time to a first wage increase was 18.9 weeks, and the average amount of the first increase was $0.21 per hour.

[26] Katz and Krueger (1992) report that a significant fraction of fast-food stores in Texas responded to an increase in the minimum wage by raising wages for workers who were initially earning more than the new minimum rate. Our results on the slope of the tenure profile are consistent with their findings.

[27] According to the McDonald's Corporation 1991 *Annual Report*, payroll and benefits are 31.3 percent of operating costs at company-owned stores. This calculation is only approximate because minimum-wage workers make up less than half of payroll even though they are about half of workers, and because a rise in the minimum wage causes some employers to increase the pay of other higher-wage workers in order to maintain relative pay differentials.

[28] The effect is attributable to a 2.0-percent increase in prices in New Jersey and a 1.0-percent decrease in prices in Pennsylvania.

TABLE 7—REDUCED-FORM MODELS FOR CHANGE IN THE PRICE OF A FULL MEAL

Independent variable	Dependent variable: change in the log price of a full meal (i)	(ii)	(iii)	(iv)	(v)
1. New Jersey dummy	0.033 (0.014)	0.037 (0.014)	—	—	—
2. Initial wage gap[a]	—	—	0.077 (0.075)	0.146 (0.074)	0.063 (0.089)
3. Controls for chain and[b] ownership	no	yes	no	yes	yes
4. Controls for region[c]	no	no	no	no	yes
5. Standard error of regression	0.101	0.097	0.102	0.098	0.097

Notes: Standard errors are given in parentheses. Entries are estimated regression coefficients for models fit to the change in the log price of a full meal (entrée, medium soda, small fries). The sample contains 315 stores with valid data on prices, wages, and employment for waves 1 and 2. The mean and standard deviation of the dependent variable are 0.0173 and 0.1017, respectively.

[a]Proportional increase in starting wage necessary to raise the wage to the new minimum-wage rate. For stores in Pennsylvania the wage gap is 0.

[b]Three dummy variables for chain type and whether or not the store is company-owned are included.

[c]Dummy variables for two regions of New Jersey and two regions of eastern Pennsylvania are included.

minimum-wage increase in New Jersey—slightly more than the increase needed to pass through the cost increase caused by the minimum-wage hike.

The pattern of price changes *within* New Jersey is less consistent with a simple "pass-through" view of minimum-wage cost increases. In fact, meal prices rose at approximately the same rate at stores in New Jersey with differing levels of initial wages. Inspection of the estimated GAP coefficients in column (v) of Table 7 confirms that within regions of New Jersey, the GAP variable is statistically insignificant.

In sum, these results provide mixed evidence that higher minimum wages result in higher fast-food prices. The strongest evidence emerges from a comparison of New Jersey and Pennsylvania stores. The magnitude of the price increase is consistent with predictions from a conventional model of a competitive industry. On the other hand, we find no evidence that prices rose faster among stores in New Jersey that were most affected by the rise in the minimum wage.

One potential explanation for the latter finding is that stores in New Jersey compete in the same product market. As a result, restaurants that are most affected by the minimum wage are unable to increase their product prices faster than their competitors. In contrast, stores in New Jersey and Pennsylvania are in separate product markets, enabling prices to rise in New Jersey relative to Pennsylvania when overall costs rise in New Jersey. Note that this explanation seems to rule out the possibility that store-specific demand shocks can account for the anomalous rise in employment at stores in New Jersey with lower initial wages.

VI. Store Openings

An important potential effect of higher minimum wages is to discourage the opening of new businesses. Although our sample design allows us to estimate the effect of the minimum wage on *existing* restaurants in New Jersey, we cannot address the effect of the higher minimum wage on potential entrants.[29] To assess the likely size of such an effect, we used national restaurant directories for the McDonald's restaurant chain to compare the numbers of operating restaurants and the numbers of newly opened restaurants in different states over the 1986–1991 period. Many states raised their minimum wages during this period. In addition, the federal minimum wage increased in the early 1990's from \$3.35 to \$4.25, with differing effects in different states depending on the level of wages in the state. These policies create an opportunity to measure the impact of minimum-wage laws on store opening rates across states.

The results of our analysis are presented in Table 8. We regressed the growth rate in the number of McDonald's stores in each state on two alternative measures of the minimum wage in the state and a set of other control variables (population growth and the change in the state unemployment rate). The first minimum-wage measure is the fraction of workers in the state's retail trade industry in 1986 whose wages fell between the existing federal minimum wage in 1986 (\$3.35 per hour) and the effective minimum wage in the state in April 1990 (the maximum of the federal minimum wage and the state minimum wages as of April 1990).[30] The second is the ratio of the state's effective minimum wage in 1990 to the average hourly wage of retail trade workers in the state in 1986. Both of these measures are designed to gauge the degree of upward wage pressure exerted by state or federal minimum-wage changes between 1986 and 1990.

The results provide no evidence that higher minimum-wage rates (relative to the retail-trade wages in a state) exert a negative effect on either the net number of restaurants or the rate of new openings. To the contrary, all the estimates show *positive* effects of higher minimum wages on the number of operating or newly opened stores, although many of the point estimates are insignificantly different from zero. While this evidence is limited, we conclude that the effects of minimum wages on fast-food store opening rates are probably small.

VII. Broader Evidence on Employment Changes in New Jersey

Our establishment-level analysis suggests that the rise in the minimum wage in New Jersey may have increased employment in the fast-food industry. Is this just an anomaly associated with our particular sample, or a phenomenon unique to the fast-food industry? Data from the monthly Current Population Survey (CPS) allow us to compare state-wide employment trends in New Jersey and the surrounding states, providing a check on the interpretation of our findings. Using monthly CPS files for 1991 and 1992, we computed employment–population rates for teenagers and adults (age 25 and older) for New Jersey, Pennsylvania, New York, and the entire United States. Since the New Jersey minimum wage rose on April 1, 1992, we computed the employment rates for April–December of both 1991 and 1992. The relative changes in employment in New Jersey and the surrounding states then give an indication of the effect of the new law.

A comparison of changes in adult employment rates show that the New Jersey labor market fared slightly worse over the 1991–1992 period than either the U.S. labor market as a whole or labor markets in Pennsylvania or New York (see Card and Krueger, 1993 table 9).[31] Among teenagers, however, the situation was reversed. In New Jersey, teenage employment rates fell by 0.7 percent from 1991 to 1992. In New York,

[29]Direct inquiries to the chains in our sample revealed that Wendy's opened two stores in New Jersey in 1992 and one store in Pennsylvania. The other chains were unwilling to provide information on new openings.

[30]We used the 1986 Current Population Survey (merged monthly file) to construct the minimum-wage variables. State minimum-wage rates in 1990 were obtained from the Bureau of National Affairs *Labor Relations Reporter Wages and Hours Manual* (undated).

[31]The employment rate of individuals age 25 and older fell by 2.6 percent in New Jersey between 1991 and 1992, while it rose by 0.3 percent in Pennsylvania, and fell by 0.2 percent in the United States as a whole.

TABLE 8—ESTIMATED EFFECT OF MINIMUM WAGES ON NUMBERS OF MCDONALD'S RESTAURANTS, 1986–1991

	Dependent variable: proportional increase in number of stores				Dependent variable: (number of newly opened stores) ÷ (number in 1986)			
Independent variable	(i)	(ii)	(iii)	(iv)	(v)	(vi)	(vii)	(viii)
Minimum-Wage Variable:								
1. Fraction of retail workers in affected wage range 1986[a]	0.33 (0.20)	—	0.13 (0.19)	—	0.37 (0.22)	—	0.16 (0.21)	—
2. (State minimum wage in 1991) ÷ (average retail wage in 1986)[b]	—	0.38 (0.22)	—	0.47 (0.22)	—	0.47 (0.23)	—	0.56 (0.24)
Other Control Variables:								
3. Proportional growth in population, 1986–1991	—	—	0.88 (0.23)	1.03 (0.23)	—	—	0.86 (0.25)	1.04 (0.25)
4. Change in unemployment rates, 1986–1991	—	—	−1.78 (0.62)	−1.40 (0.61)	—	—	−1.85 (0.68)	−1.40 (0.65)
5. Standard error of regression	0.083	0.083	0.071	0.068	0.088	0.088	0.077	0.073

Notes: Standard errors are shown in parentheses. The sample contains 51 state-level observations (including the District of Columbia) on the number of McDonald's restaurants open in 1986 and 1991. The dependent variable in columns (i)–(iv) is the proportional increase in the number of restaurants open. The mean and standard deviation are 0.246 and 0.085, respectively. The dependent variable in columns (v)–(viii) is the ratio of the number of new stores opened between 1986 and 1991 to the number open in 1986. The mean and standard deviation are 0.293 and 0.091, respectively. All regressions are weighted by the state population in 1986.

[a]Fraction of all workers in retail trade in the state in 1986 earning an hourly wage between $3.35 per hour and the "effective" state minimum wage in 1990 (i.e., the maximum of the federal minimum wage in 1990 ($3.80) and the state minimum wage as of April 1, 1990).

[b]Maximum of state and federal minimum wage as of April 1, 1990, divided by the average hourly wage of workers in retail trade in the state in 1986.

Pennsylvania, and the United States as a whole, teenage employment rates dropped faster. Relative to teenagers in Pennsylvania, for example, the teenage employment rate in New Jersey rose by 2.0 percentage points. While this point estimate is consistent with our findings for the fast-food industry, the standard error is too large (3.2 percent) to allow any confident assessment.

VIII. Interpretation

As in the earlier study by Katz and Krueger (1992), our empirical findings on the effects of the New Jersey minimum wage are inconsistent with the predictions of a conventional competitive model of the fast-food industry. Our employment results are consistent with several alternative models, although none of these models can also explain the apparent rise in fast-food prices in New Jersey. In this section we briefly summarize the predictions of the standard model and some simple alternatives, and we highlight the difficulties posed by our findings.

A. *Standard Competitive Model*

A standard competitive model predicts that establishment-level employment will fall if the wage is exogenously raised. For an entire industry, total employment is predicted to fall, and product price is predicted to rise in response to an increase in a binding minimum wage. Estimates from the time-series literature on minimum-wage effects can be used to get a rough idea of the elasticity of low-wage employment to the minimum wage. The surveys by Brown et al. (1982, 1983) conclude that a 10-percent increase in the coverage-adjusted minimum wage will reduce teenage employment rates by 1–3 percent. Since this effect is for *all* teenagers, and not just those employed in low-wage industries, it is surely a lower bound on the magnitude of the effect for fast-food workers. The 18-percent increase in the New Jersey minimum wage is therefore predicted to reduce employment at fast-food stores by 0.4–1.0 employees per store. Our empirical results clearly reject the upper range of these estimates, although we cannot reject a small negative effect in some of our specifications.

A possible defense of the competitive model is that unobserved demand shocks affected certain stores in New Jersey—specifically, those stores that were initially paying wages less than $5.00 per hour. However, such localized demand shocks should also affect product prices. (In fact, in a competitive model, product demand shocks work through a rise in prices.) Although lower-wage stores in New Jersey had relative employment gains, they did not have relative price increases. Furthermore, our analysis of employment changes in two major suburban areas (around Newark and Camden) reveals that, even within local areas, employment rose faster at the stores that had to increase wages the most because of the new minimum wage.

B. *Alternative Models*

An alternative to the conventional competitive model is one in which firms are price-takers in the product market but have some degree of market power in the labor market. If fast-food stores face an upward-sloping labor-supply schedule, a rise in the minimum wage can potentially increase employment at affected firms and in the industry as a whole.[32]

This same basic insight emerges from an equilibrium search model in which firms post wages and employees search among posted offers (see Dale T. Mortensen, 1988). Kenneth Burdett and Mortensen (1989) derive the equilibrium wage distribution for a noncooperative wage-search/wage-posting model and show that the imposition of a binding minimum wage can increase both wages and employment relative to the initial equilibrium. Furthermore, their model predicts that the minimum wage will increase employment the most at firms that initially paid the lowest wages.

Although monopsonistic and job-search models provide a potential explanation for the observed employment effects of the New Jersey minimum wage, they cannot explain the observed price effects. In these models, industry prices should have fallen in New Jersey relative to Pennsylvania, and at low-wage stores in New Jersey relative to high-wage stores in New Jersey. Neither prediction is confirmed: indeed, prices rose faster in New Jersey than in Pennsylvania, although at about the same rate at high- and low-wage stores in New Jersey. Another puzzle for equilibrium search models is the absence of wage increases at firms that were initially paying $5.05 or more per hour.

The strict link between the employment and price effects of a rise in the minimum wage may be broken if fast-food stores can vary the quality of service (e.g., the length of the queue at peak hours, or the cleanliness of stores). Another possibility is that stores altered the *relative* prices of their various menu items. Comparisons of price changes for the three items in our survey show slight declines (-1.5 percent) in the price of french fries and soda in New Jersey relative to Pennsylvania, coupled with a relative increase (8 percent) in entrée prices. These limited data suggest a possible role for relative price changes within the fast-food industry following the rise in the minimum wage.

One way to test a monopsony model is to identify stores that were initially "supply-constrained" in the labor market and test for employment gains at these stores relative to other stores. A potential indicator of market power is the use of recruitment bonuses. As we noted in Table 2, about 25 percent of stores in wave 1 were offering cash bonuses to employees who helped find a new worker. We compared employment

[32] Daniel G. Sullivan (1989) and Michael R Ransom (1993) present empirical results for nurses and university teachers that suggest monopsony-like behavior of employers.

changes at New Jersey stores that were offering recruitment bonuses in wave 1, and also interacted the GAP variable with a dummy for recruitment bonuses in several employment-change models. We do not find faster (or slower) employment growth at the New Jersey stores that were initially using recruitment bonuses, or any evidence that the GAP variable had a larger effect for stores that were using bonuses.

IX. Conclusions

Contrary to the central prediction of the textbook model of the minimum wage, but consistent with a number of recent studies based on cross-sectional time-series comparisons of affected and unaffected markets or employers, we find no evidence that the rise in New Jersey's minimum wage reduced employment at fast-food restaurants in the state. Regardless of whether we compare stores in New Jersey that were affected by the \$5.05 minimum to stores in eastern Pennsylvania (where the minimum wage was constant at \$4.25 per hour) or to stores in New Jersey that were initially paying \$5.00 per hour or more (and were largely unaffected by the new law), we find that the increase in the minimum wage increased employment. We present a wide variety of alternative specifications to probe the robustness of this conclusion. None of the alternatives shows a negative employment effect. We also check our findings for the fast-food industry by comparing changes in teenage employment rates in New Jersey, Pennsylvania, and New York in the year following the increase in the minimum wage. Again, these results point toward a relative *increase* in employment of low-wage workers in New Jersey. We also find no evidence that minimum-wage increases negatively affect the number of McDonald's outlets opened in a state.

Finally, we find that prices of fast-food meals increased in New Jersey relative to Pennsylvania, suggesting that much of the burden of the minimum-wage rise was passed on to consumers. Within New Jersey, however, we find *no* evidence that prices increased more in stores that were most affected by the minimum-wage rise. Taken as a whole, these findings are difficult to explain with the standard competitive model or with models in which employers face supply constraints (e.g., monopsony or equilibrium search models).

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