 Journal of Business & Economic Statistics, April 1995, Vol. 13, No. 2

The Employment Effect in Retail Trade of California's 1988 Minimum Wage Increase

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In this article, we study the labor-market effect of California's 1988 minimum wage increase in the retail trade industry. Two different approaches to evaluating the minimum wage effects suggest that the textbook economic analysis of minimum wages pertains; employment growth in California's low-wage retail trade sector appears to have been tempered by the minimum wage increase.

Since the passage of the Fair Labor Standards Act of 1938, minimum wage laws have represented one of this country's most prominent labor-market intervention policies. Subsequent to the establishment of the Act, several amendments have raised the federal minimum. Most recently, after almost a decade with no changes in the minimum wage, the federal minimum increased from \$3.35 to \$3.80 on April 1, 1990, and one year later to \$4.25. There is wide support for further increases in the minimum wage, with many policymakers favoring large increases in the nominal minimum and, possibly, indexation of the minimum wage.

The enthusiasm held by many policymakers for increasing the minimum wage appears to be based on a notion that such an increase will raise the earnings of low-wage workers while having little effect on these workers' employment prospects. This view has been bolstered recently by a series of interesting and compelling empirical studies that call into question the conventional economic prediction that an increase in the minimum wage moves the labor market equilibrium backward along the demand curve, thereby reducing the employment of low-wage workers. Each of these studies suggests that the most recent round of increases in the minimum wage in the United States had little effect on employment among low-wage workers.

A study by Card (1992a) explored the consequences for teenage employment of California's minimum wage increase of July 1988. (The federal minimum wage remained constant through most the 1980s, but California was one of several states that adopted legislation that raised the state minimum to levels above the national standard.) Card compared teenage employment changes in California and in a group of southwestern and southern states from 1987 to 1989. Although the minimum increased in California during this period while remaining fixed in the other states, Card found no signs of a negative impact on California's teenage employment. In a second study, Card (1992b) exploited regional variation in wages across the United States, again finding little evidence that recent minimum wage increases reduced low-wage employment. Katz and Krueger (1992) and Card and Krueger (1993), in studies of the fast-food industry, also detected little change in employment due to increases in the U.S. minimum wage. A study in the British context by Machin and Manning (1994) also failed to confirm conventionally predicted employment effects of a minimum wage.

There have been studies finding evidence of a negative effect on employment of the minimum wage; Wellington (1991), Brown (1988), and Brown, Gilroy, and Kohen (1982) provided useful overviews. Still, the most recent wave of empirical studies is noteworthy both for its policy relevance and for its potential to shed light on labor-market fundamentals. Because this new research is important and because much of the previous literature is equivocal, we undertake in this article a new empirical evaluation of California's increase in the minimum wage from \$3.35 to \$4.25 in July of 1988. Our approach is to look for the systematic industry- and county-level variation in employment growth and wage changes that would be the effect of a minimum wage increase predicted by conventional theory.

Using County Business Patterns (CBP) data, we find evidence that does indeed appear consistent with conventional theory. First we show that in California's broadly defined retail trade sector, for the time period March 1988 to March 1989, those industries (e.g., department stores or women's clothing stores) in which the relative wage increase was most rapid had relatively slow employment growth. This pattern did not appear in California in years when the minimum wage was unchanged. Second, we note that within California there was substantial intercounty variation prior to the 1988 minimum wage increase. Retail trade wages in some counties were rather high, but in other counties retail trade employees were poorly paid. Conventional theory suggests that employment would be adversely affected in these latter counties. Our empirical results accord with this prediction.

The magnitude of the estimated effect of the minimum wage change on employment growth is similar for estimates based on interindustry variation and intercounty variation.

1. RETAIL TRADE IN CALIFORNIA AND THE UNITED STATES, 1986–1989

Our study focuses on employment in the retail trade industry. Employment in retail trade is of particular interest because wages paid by many employers in retail trade are low, and the minimum wage is therefore likely to present a binding constraint for much of this industry. Moreover, the retail trade industry is by far the largest employer of low-wage workers. Using data from the Current Population Survey (CPS), Card (1992a) estimated that in 1987 over 30% of workers in retail trade earned $3.35 or more, but less than $4.25. Moreover, these workers in retail trade represented nearly half of all workers in California who were paid $3.35–$4.24.

For our analysis, we rely primarily on data from the CBP issued by the U.S. Department of Commerce, Bureau of the Census (various years a) to track trends in employment and pay. These data indicate firms' payroll (based on the employers' quarterly federal tax returns) for the first quarter of the year and the firms' total employment for a pay period including March 12. These data are tabulated from universe files, so in principle sampling error is not an issue. (As we shall discuss later, however, nonsampling error may be a problem.) In the retail trade industry, trends in employment appear to be quite similar in California and in the rest of the United States over the period 1986 to 1989. Table 1 presents a summary. Employment grew slightly less rapidly in California in 1986–1987 and somewhat more rapidly in 1987–1988. The 1988–1989 employment trend, which should reflect any effect of the 1988 increase in California's minimum wage, shows employment growing somewhat more slowly in California than in the rest of the United States.

Table 2 presents similar statistics for reported pay per person in the retail trade industry. Unfortunately, the CBP data do not give us average wages by industry, so pay per person is constructed as payroll divided by employment. This is the "wage" in our article. As an alternative measure, we also computed the average wage for retail trade workers paid hourly from the CPS for a nine-month period, September 1987 to May 1988, that preceded the July 1988 minimum wage increase and for the same nine-month period for the following year, September 1988 to May 1989. CPS data show that in California the average wage increased 6.58% from the former to the latter period, whereas in the rest of the United States the average wage increased by 3.67%. These figures do not differ greatly from the corresponding statistics listed in the last column of Table 2.

The most notable feature of Table 2 is, of course, that retail trade wages grew much more rapidly in California than in the rest of the United States from the first quarter of 1988 to the first quarter of 1989. A natural explanation for this outcome is that the minimum wage increased in California, from $3.35 to $4.25 in July of 1988, but the wage floor remained constant in nearly all other states. In 43 states and in the District of Columbia, the minimum wage did not change at all during this period, and in states in which the minimum did increase, the increments were much smaller than California's $.90 increase. (The minimum increased during this period in Connecticut, Minnesota, and Washington by $.50, in Pennsylvania by $.35, and in Maine, New Hampshire, and Massachusetts by $.10.)

Table 1. Percent Change in Employment in Retail Trade

Location	1986–1987	1987–1988	1988–1989
California	4.46	2.95	2.48
U.S.	5.00	1.98	2.88
Cal.–U.S.	−.54	.97	−.40

Table 2. Percent Change in Per Person Pay in Retail Trade

Location	1986–1987	1987–1988	1988–1989
California	1.28	2.75	6.15
U.S.	.99	3.62	3.83
Cal.–U.S.	.29	−.87	2.32

On the basis of the data presented in Tables 1 and 2, it would appear that the minimum wage increase had little effect on employment in California's retail trade sector. Although wages in retail trade did increase in California relative to the rest of the United States for 1988–1989, presumably as a result of the new minimum wage law, retail trade employment growth was only slightly slower in California than in the rest of the United States.

This conclusion that the minimum wage increase had no substantial adverse effect on employment is predicated on the assumption that there were no other major changes to factors affecting retail trade employment in California. One can read the data in Tables 1 and 2 as the outcome of a "natural experiment" in which a treatment group, California, was subjected to a substantial minimum wage increase. This reading would be wrong, however, if California's retail trade sector was doing particularly well during the 1988 to 1989 period. In that case employment growth in California's retail trade sector from the first quarter of 1988 through 1989 may have been relatively strong *despite* negative employment effects of the minimum wage increase.

There is some evidence that reinforces this concern. The California State Board of Equalization (various years) indicated robust taxable retail trade in California during the 12 months following the minimum wage increase (August 12, 1988, to August 11, 1989); the retail sales volume during this period was 7.8% higher than during the previous 12 months. We do not have comparable data for the rest of the United States, but we can get an indication of retail sales strength from calendar year estimates listed by the U.S. Department of Commerce, Bureau of the Census (various years b) and California Department of Finance (various years). These indicate that sales in California were 9.4% stronger in calendar year 1989 than in 1988. The corresponding percent growth in sales for the rest of the United States was estimated to be just 4.9%.

Given the strong performance of California's retail trade sector from 1988 to 1989, we would naturally have expected employment growth to be much faster in California than in the rest of the United States for 1988–1989, all else being equal. It is therefore possible for data in Tables 1 and 2 to be consistent with a substantial negative employment effect of the minimum wage increase. In an attempt to resolve this issue we look beyond the aggregate wage and employment figures.

Table 3. Differences in Pay Growth and Employment Growth, California–U.S., 1988–1989

Industry categories	Pay per employee (1)	Employment (2)
General merchandise	6.86	−6.16
Eating and drinking	4.63	−1.24
Food stores	2.36	.37
Apparel and accessory	2.31	1.19
Building and garden supplies	.77	3.15
Furniture	.00	3.87
Auto dealers and service stations	−1.56	2.02

NOTE: Column (1) gives the percent of change in pay per person in California minus the percent of change in pay per person in the rest of the United States and column (2) is the difference in the percent of change in employment in California and the rest of the United States. These categories represent all retail trade industries except "Miscellaneous retail."

Table 3 shows disaggregated data on wage and employment growth for the crucial 1988–1989 period. In particular, statistics similar to those in the bottom rows of Tables 1 and 2 are shown for two-digit Standard Industry Classification (SIC) categories within retail trade. This table indicates considerable variation by industry within the retail trade sector. In some industries, notably eating and drinking places and general merchandise stores, there were substantial increases in the wages in California relative to the United States, but in other industries the wage growth was actually slower in California than in the country as a whole. Similarly, there is substantial variation in the growth of employment in California compared to the national trend. Moreover, in industries in which wages grew rapidly in California relative to the national trend, employment appears to decline relative to the trend.

2. INTERINDUSTRY EMPLOYMENT EFFECTS OF THE MINIMUM WAGE

In the textbook labor-market model, wages and employment are jointly determined by supply and demand. It is therefore difficult to interpret coefficients estimated for a regression of the form

$$\Delta e_i = \beta_0 + \beta_1 \Delta w_i + u_i, \qquad (1)$$

where Δe_i and Δw_i are industry-level changes in employment and wages. Employment is related to wages, but any changes in employment are the result of shifts in the supply curve or the demand curve (or both), and Regression (1) does not distinguish between these movements.

Equation (1) can be interpreted as a demand curve if there is some exogenously generated variation in wage changes faced by firms across industries. One such exogenous shock is the minimum wage increase in California. If labor markets are competitive, industry wage changes that result from the minimum wage increase should induce movements along the demand curve for labor. We assume that this relationship between industry wage changes and employment growth can be represented by Equation (1), where we let e be the log of employment and w be the log of the wage, so that Δe and Δw are the approximate percent changes in employment and wages, respectively.

In particular, suppose that in California

$$\Delta e_{ci} = \alpha_c + \alpha_i + \beta_1 \Delta w_{ci} + u_{ci}, \qquad (2)$$

where the subscript ci indicates industry i in California. Note that in (2) the observed percent change in employment in industry i is the result not only of changes in the wage in the industry but also of California-specific and industry-specific shocks to the demand schedule, α_c and α_i, respectively.

If for the United States as a whole a similar demand curve pertains,

$$\Delta e_{us,i} = \alpha_{us} + \alpha_i + \beta_1 \Delta w_{us,i} + u_{ui}, \qquad (3)$$

we can subtract (3) from (2), giving

$$(\Delta e_c - \Delta e_{us})_i = \beta_0 + \beta_1 (\Delta w_c - \Delta w_{us})_i + u_i, \qquad (4)$$

where $\beta_0 = \alpha_c - \alpha_{us}$ and $u_i = u_{ci} - u_{ui}$.

We estimate Equation (4) using data from the CBP for the time period March 1988 to March 1989. Given that the minimum wage increased in California in July 1988, we would expect that much of the between-industry variation in wage growth in California relative to the rest of the country would be exogenously induced—that is, would be the result of the policy change.

Even though there is exogenously induced wage variation due to a minimum wage change, one might still be concerned that our estimated demand equation would be subject to simultaneous equation bias because industry employment changes are still influenced by both supply and demand. Fortunately, there is a natural way of dealing with this issue. For a time period when the minimum wage changes, some portion of the exogenous change in wages in California's industries relative to the United States, $\Delta w_{ci} - \Delta w_{us,i}$, may be systematically related to industry characteristics within California. These characteristics can then serve as instrumental variables (IV's). First, the wage increase within an industry will typically be negatively correlated with the initial wage level of the industry. That is, when the minimum wage changes between periods $t-1$ and t, we would expect the industry wage changes to be largest in industries that in period $t-2$ had low wages. Second, because smaller firms may have higher noncompliance rates (Ashenfelter and Smith 1979), industries with small firms will, all else being equal, tend to have smaller industry wage increases due to an increase in the wage floor. Alternatively, suppose that the variance in wages is larger in large firms. This means that, having conditioned on the mean, industries with large firms will tend to have a higher fraction of workers receiving wages near or at the minimum. In either of these cases, industry wage changes due to a minimum wage increase will be positively correlated with industry firm size (having conditioned on average industry wages).

More specifically, in estimating (4) we use two-stage least squares (2SLS) with instruments, $\overline{w}_i$, the log of the California industry mean wage for year $t-2$, and $\overline{s}_i$, the log of mean firm size in the industry for the year $t-2$. Thus for the 1988–1989 equation the instruments are taken from the year 1987. There are predetermined variables that are likely to be

Table 4. Regression Results for Dependent Variable ($\Delta e_c - \Delta e_{us}$)

	1985–1986		1986–1987		1987–1988		1988–1989	
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Intercept	.010	.003	−.010	−.017	.010	−.006	.028*	.027*
	(.008)	(.013)	(.008)	(.035)	(.009)	(.020)	(.005)	(.005)
Coefficient on	−.361*	−.945	−.317	3.870	−.092	−1.685	−.902*	−.879*
($\Delta W_c - \Delta W_{us}$)	(.123)	(.723)	(.154)	(4.935)	(.146)	(.989)	(.084)	(.133)
n	50	50	50	50	51	51	53	53
$\bar{R}^2$	.13	.01	.06	−.01	−.01	.04	.69	.45

NOTE: All regressions are weighted by employment in industry i in California in year $t-2$. Asymptotic standard errors are given in parentheses (*indicates sig. < .01). For the IV estimates, the instruments are log mean wage and log average firm size in industry i in year $t-2$. For 1985–1986 and 1986–1987 three industry categories are missing and for 1987–1988 two industry categories are missing. The employment of these missing industries is less than 2% of total retail industry employment in 1988.

correlated with the wage change induced by the increase in the mandated minimum wages.

Ordinary least squares (OLS) and IV estimates of Equation (4) are presented in Tables 4 and 5. These results were produced using data for 3–4-digit industries within the retail trade sector. In Table 4, a weighted regression procedure is used, with weights given by California's employment in the 3–4-digit industry for the year $t-2$. Table 5 provides unweighted results that indicate that the results are not sensitive to the weighting scheme employed. We were not able to include all 3–4-digit retail trade industries in our regressions because there was some shuffling of the SIC codes between 1987 and 1988. Fortunately, though, industries we did include represent 97.8% of total employment in retail trade.

As a simple specification test, we estimated Equation (4) for the time frame of interest, 1988–1989 and also for several other years when the minimum wage laws did not change. Consider first the OLS estimates, presented in the odd numbered columns of Tables 4 and 5. As expected, for the years 1985–1986, 1986–1987, and 1987–1988, the regression results show little of interest; few coefficients are significantly different from 0, and the typical value of the R^2 is low. For 1988–1989, however, a different picture emerges: In those retail trade industries in which the wages increased most rapidly, relative employment declined.

Our OLS estimates of the elasticity of demand for labor in retail trade are in the −.8 to −.9 range. By way of comparison, Coterill's (1975) estimates of wage elasticities for various retail trade industries for years between 1948 and 1963 were generally between −.4 and −1.1, and Zucker's (1973) estimates of the elasticity of demand in the U.S. low-wage manufacturing sector were close to −1.0. It is worth mentioning that the wage change on the right side of our demand equation is *not* the change in the minimum wage. Rather, it is the change in average wage cost per worker, which is considerably smaller; the absolute value of the elasticity of employment with respect to the *minimum wage* itself is obviously much smaller than .9.

The results we present accord with standard theory and appear to fall in the general range of previous estimates. We cannot, however, ignore the possibility that the negative elasticity we estimate using OLS is due to the way we have attempted to deal with shortcomings in our data. In particular, we do not have in our data set the actual wages firms are paying. Instead the "wage" we use is the average pay per worker, calculated as average weekly payroll during the quarter divided by employment in the last week of the quarter. That is, for any particular industry, we use as our measured log wage in time t, $w_t^m = \ln(P_t/E_t^m) = p_t - e_t^m$. To see the implication of this, consider an industry in which the "true" wage is constant during the quarter, but our measure of log employment for the last week of the quarter, e_t^m, varies randomly around its mean, e_t. In particular, suppose that $e_t^m = e_t + v_t$, where v_t is the percentage deviation in our measure of employment from its "true" quarterly mean e_t. (This random variation could be due to employment fluctuations within an industry during the quarter or could be due to measurement error.) In this case, our measure of the log wage for time t, w_t^m, will also be error ridden because $w_t^m = p_t - e_t^m = w_t - v_t$. This in turn implies that, for each of the industries in our data,

Table 5. Regression Results (Unweighted) for Dependent Variable ($\Delta e_c - \Delta e_{us}$)

	1985–1986		1986–1987		1987–1988		1988–1989	
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Intercept	.001	.049	.065	−.089	−.009	−.034	.031*	.029*
	(.019)	(.059)	(.058)	(.205)	(.021)	(.040)	(.007)	(.009)
Coefficient on	−.117	1.076	−.365	6.600	−.276	.379	−.825*	−.898**
($\Delta W_c - \Delta W_{us}$)	(.186)	(1.332)	(.611)	(7.875)	(.207)	(.887)	(.115)	(.391)
n	51	51	51	51	53	53	53	53
$\bar{R}^2$	−.01	−.01	−.01	−.01	.01	−.02	.49	.08

NOTE: Asymptotic standard errors are given in parentheses (* indicates sig. < .01 and ** indicates sig. < .05). For the IV estimates, the instruments are log mean wage and log average firm size in industry i in year $t-2$. For 1985–1986 and 1986–1987, two industry categories are missing. The employment of these missing industries is 1.64% of total retail industry employment in 1988.

$\Delta e_i^m = \Delta e_i + \Delta v_i$ and $\Delta w_i^m = \Delta w_i - \Delta v_i$. Even without simultaneity bias, a consequence of this particular form of measurement error is inconsistency in the OLS estimator $\widehat{\beta}_1$ from Equation (4). Specifically,

$$\text{plim}(\widehat{\beta}_1) = \frac{\beta_1}{1+r} - \frac{r}{1+r}, \qquad (5)$$

where r is the ratio, $\text{var}(\Delta v_c - \Delta v_{us})/\text{var}(\Delta w_c - \Delta w_{us})$. If the true parameter β_1 is 0, then the probability limit of our estimator goes to -1 as r gets big—for example, if measured wage changes are all noise. This is troubling in our context because our estimated elasticity of labor demand for California's retail trade industry is indeed not so far from -1.

Having made this point, we believe that there are two important pieces of evidence that argue against its importance to our analysis. First, we note that if the "division bias" described in the preceding paragraph is serious, it will induce a negative correlation between wage and employment changes for all years, not just years when the minimum wage changes. Equation (5) implies that the built-in-negative correlation between measured employment and wage changes is likely to be more severe in years when the minimum wage remains constant because a minimum wage increase is a source of true variation in relative wages. As in clear from Tables 4 and 5, however, there is little systematic relationship between observed wage and employment growth in years when the minimum wage is unchanged.

Second, and more importantly, the IV estimates of the wage elasticity [in column (8) of Tables 4 and 5] are very similar to OLS counterparts. As discussed previously, we use as instruments, $\overline{w}_i$, the log of the California industry mean wage for year $t-2$, and $\overline{s}_i$, the log of mean firm size in the industry for the year $t-2$. This industry log mean wage is the log of pay per employee, and log firm size is the log of number of employees per firm, both of which were calculated from the yearly averages reported by the Bureau of Labor Statistics (BLS). We use $\overline{w}_i$ and $\overline{s}_i$ for year $t-2$ instead of year $t-1$ because of concern over the possibility that there may be measurement error that would be common to both the CBP and BLS data. (Such error would mean that our instruments would be correlated automatically with Δw_i, but also with the error term of our OLS equation, because that error term may include components due to mismeasurement of the employment variables.)

For the IV approach to be consistent, we not only need our instruments to be correlated with $(\Delta w_c - \Delta w_{us})_i$, as they clearly are, but the instruments must also be orthogonal to the error term in our regression equation (4). Because the model is overidentified (i.e., because we use two instruments, not one), we can implement a conventional 2SLS overidentification test. Let b_0 and b_1 be the parameter estimates of (4) using 2SLS. Then the proposed test statistic, based on the Lagrange multiplier principle, is nR^2, where n is sample size and R^2 is the uncentered R^2 in the regression of the residuals $\widehat{u}_i = (\Delta e_c - \Delta e_{us})_i - b_0 - b_1(\Delta w_c - \Delta w_{us})_i$ on $\overline{w}_i$ and $\overline{s}_i$. This test statistic has an asymptotic distribution chi-squared with 1 df. In our case the calculated test statistic was found to be .25; quite clearly we cannot reject the overidentification restriction.

For a subset of industries we can construct instruments from an alternative data source—the CPS—as a means of evaluating the robustness of our approach. Our reason for finding new instruments is due to the concern about measurement error, as outlined previously. A potential problem is that we take our instruments $\overline{w}_i$ and $\overline{s}_i$ from the BLS, which in principle draws its data from the same sources as the CBP data used in the regressions. Thus, if there is serial correlation in the measurement error (i.e., if the errors in reported industry employment in 1988–1989 are correlated with errors in reported employment from 1987), there is a possibility that these instruments are correlated with the dependent variable through this mechanism. CPS-based instruments, however, should be free of such correlations.

To construct appropriate instruments, we first calculate the mean wage for each industry, but now using individual-level data from the CPS rather than pay per person from the BLS. In principle, we might have calculated the mean industry wages, $\overline{w}_i$, using 1988 data. Because CPS samples of workers within specific retail trade industries in California are small, we used CPS outgoing rotation group files from January 1984 to June 1988. There are 33 retail trade industries reported in the CPS, of which we matched 25 with the 1987 SIC codes for the CBP. We did not match all of the categories because in a few cases there were changes in the SIC codes between 1987 and 1988 and because we did not construct the instruments for any industry for which we had fewer than 20 CPS observations. In all, we used 5,114 of the 5,205 available observations on retail trade employees in our CPS sample in constructing instruments. For the individuals in our sample, the "hourly wage" is simply the reported hourly wage, or, when this is unavailable, the reported weekly earnings divided by reported usual weekly hours. We find that CPS industry mean wages constructed in this fashion are highly correlated with BLS pay per person; the correlation coefficient is .97.

Next we calculate a new firm size variable by dividing the BLS industry payroll by the industry mean wage calculated from the CPS. This gives us a measure of industry employment (in hours) that does not rely on the possibly error-ridden industry employment figures reported by the BLS. We divide this latter figure by the number of firms in the industry, which then gives us a measure of firm size in the industry. We find that per-firm industry employment measured in this fashion is highly correlated with employment per firm as given by the BLS; the correlation coefficient of the two variables is over .99.

Finally, we construct a new instrument using the CPS—the fraction of the labor force in retail trade industries in California that was directly "affected" by the minimum wage. This variable, F_i, is the fraction of workers in industry i who, prior to the minimum wage increase, earned wages at or above \$3.35 but less than \$4.25. We define a worker as "affected" if the nominal hourly wage is equal to or greater than \$3.35 but less than \$4.25 in real terms, using June 1988 as the base and using the Consumer Price Index to adjust to real terms.

Table 6. Regression Results Using Alternative Instruments 1988–1989 for Dependent Variable ($\Delta e_c - \Delta e_{us}$)

	OLS (1)	IV (2)	IV (3)	IV (4)
Intercept	.027*	.030*	.030*	.029*
	(.006)	(.006)	(.006)	(.006)
Coefficient on	−.869*	−.946*	−.962*	−.916*
($\Delta W_c - \Delta W_{us}$)	(.113)	(.142)	(.146)	(.137)
n	25	25	25	25
$\bar{R}^2$	.71	.64	.64	.64

NOTE: All regressions are weighted by employment in industry i in California in 1987. Asymptotic standard errors are given in parentheses (* indicates sig. <.01). For the IV estimates, the instruments are (2) log mean wage and log average firm size calculated from the BLS, (3) log mean wage calculated from the CPS and log average firm size estimated from the CPS and the BLS, (4) log of the fraction of workers below minimum wage and log average firm size estimated from the CPS and the BLS.

The "total" number of workers is just the number whose wages are at the minimum or above. For each industry we calculate F_i, the ratio of "affected" to "total." Overall, we find that using our definition for "fraction affected," about 28% of California's workers in retail trade were affected by the minimum wage increase.

In Table 6 we present first the OLS estimates of Equation (4) for the 25 observations for which we were able to construct the CPS-based instruments. The estimates reported as IV(2) use $\bar{w}_i$ and $\bar{s}_i$ as calculated from the BLS as instrument (as in Table 4, but now with the smaller sample). IV (3) shows the results of using the instruments $\bar{w}_i$ and $\bar{s}_i$, but now calculated from the CPS, not the BLS. IV(4) is similar to IV(3) but replaces $\bar{w}_i$ with the log of F_i. In each regression we used a weighted procedure as described previously. The test statistic for the Lagrange multiplier test of the overidentifying restrictions for the three IV procedures (2), (3), and (4) reported in Table 5 are .17, .25, and 1.82, respectively. In none of the cases do we reject the restrictions.

Results in Table 6 are quite similar to those presented in Tables 4 and 5. In particular, these estimates show that, within retail trade, cross-industry variation in wage changes induced by the minimum wage increase is negatively correlated with industry employment changes.

Another way of seeing how the minimum wage change affected industry wages and employment for the 1988–1989 period is to look at the reduced-form equations underlying our IV procedures. These are presented in Table 7. Columns (1) and (2) show regressions in which ($\Delta w_c - \Delta w_{us}$) is the dependent variable. In column (1) the wage-change equation shows the expected negative coefficient on the mean hourly wage of workers in the industry (as calculated from the CPS), and also a positive coefficient on mean firm size in the industry (as calculated using wages from the CPS and total payroll from the BLS). As discussed previously, this latter outcome would be expected if small firms have lower rates of compliance to minimum wage laws, or if the variance of wages is typically higher in large firms than in small firms that pay the same average wage. Similarly, column (2) shows a positive coefficient for $\ln(F_i)$ and a positive coefficient for $\bar{s}_i$.

Columns (3) and (4) give reduced-form estimates for an equation specifying ($\Delta e_c - \Delta e_{us}$) as the dependent variable.

Table 7. Reduced-Form Regression Results: CPS Matched Industry 1988–1989

	Dependent variable ($\Delta W_c - \Delta W_{us}$)		Dependent variable ($\Delta e_c - \Delta e_{us}$)	
	(1)	(2)	(3)	(4)
Intercept	.072	−.031	.027	.093*
	(.036)	(.019)	(.038)	(.021)
$\bar{W}_i$	−.080*	—	.050**	—
	(.020)		(.020)	
$\ln(F_i)$	—	.029*	—	−.017**
		(.006)		(.007)
$\bar{s}_i$	.028*	.028*	−.032*	−.031*
	(.005)	(.005)	(.006)	(.005)
n	25	25	25	25
$\bar{R}^2$	.59	.66	.57	.58

NOTE: All regressions are weighted by employment in industry i in California in year 1987. Asymptotic standard errors are given in parentheses (* indicates sig. <.01 and ** indicates sig. <.05). The instruments are log mean wage calculated from the CPS and log average firm size estimated from the CPS and BLS for (1) and (3) and log of the fraction of workers below minimum wage and log average firm size estimated from the CPS and the BLS for (2) and (4).

These show that, having conditioned on firm size in the industry, employment growth over this period was most rapid in industries with higher initial wages. Similarly, we find the anticipated negative coefficient on the measure of an industry's workers having initial wages lower than the new minimum wage.

3. INTERCOUNTY EMPLOYMENT EFFECTS OF THE MINIMUM WAGE

We turn next to an alternative way of looking for the effect of the minimum wage increase on retail trade employment in California. Prior to the 1988 minimum wage increase there was substantial variation across California's counties in wages workers earned in the retail trade industry. As in Section 2, our argument is that the legislated change in the minimum wage induced exogenous variation in wages and this variation can be exploited to estimate the demand curve for labor in a low-wage sector. One reason for examining county data instead of industry data is that this procedure relies on an alternative source of variation and thus provides an opportunity for examining robustness of the results of the industry-level analysis. Another advantage of the county analysis is that in estimating the demand curve we can condition our labor-demand equation on county-level retail sales trends, estimated using sales-tax data.

The central idea is to treat the demand for labor in a county as a derived demand function—dependent on the county demand for "output" in the retail trade industry. Consider San Francisco and Sutter Counties, for instance. San Francisco County in the first quarter of 1988 had the highest average wages in retail trade of any county in California. Here the average wage grew by only 1.7% from the first quarter of 1988 to first quarter of 1989, suggesting that the minimum wage increase had little bite. County retail trade employment grew rapidly, by 8.0%. In contrast, Sutter County, an initially lower-wage county, had a 9.2% increase in average retail trade wages and a change in retail trade employment of −2.1%. That the employment trends may have been influenced by the wage trends in this case is made more plausible when we note that these counties had comparable strength in retail sales growth from the 12 months preceding the minimum wage increase to the subsequent 12-month period; Sutter County taxable sales increased by 7.2%, while San Francisco County taxable sales increased by 6.9%.

In more formal terms, suppose that growth in demand for retail trade services in county j, $\Delta \ln Y_j$, is determined by secular local economic trends, which in turn are largely unrelated to income of workers in retail trade. Suppose further that production in retail trade entails a constant elasticity of substitution production process. Then, as discussed by Hammermesh (1993), the derived demand function for labor can be written

$$\Delta \ln E_j = a_0 + \sigma \Delta \ln W_j + a_1 \Delta \ln Y_j, \qquad (6)$$

where σ is the elasticity of substitution between the services of capital and labor and a_1 is a parameter that equals 1 under constant returns to scale. We use this specification as the basis of the regressions presented later.

We do not directly observe the variable $\Delta \ln Y_j$ for our counties, but we do not have measures of a closely related variable—the percent change in retail sales in the county, $\Delta r_j = \Delta \ln R_j$. We thus specify and estimate a regression of the form

$$\Delta e_j = \alpha_0 + \alpha_1 \Delta w_j + \alpha_2 \Delta r_j + u_j, \qquad (7)$$

where Δe_j is the percent change in the county's retail sector employment, Δw_j is the percent wage change, and Δr_j is our measure of retail output growth.

Estimates of Equation (7) will of course be nonsensical in a typical year when there is no change in the minimum. Employment changes depend in general on local supply and demand, and any correlation between the wage change and employment change is the result of the interaction of these two. For a year when there is a mandated increase in the minimum wage, though, we have exogenously induced cross-county variation in the wage growth. For such a time period we can estimate our derived demand equation. As discussed in Section 2, we will want to use instruments for Δw_j in estimating (7); appropriate instruments are $\overline{w}_j$, the average wage in the county, and $\overline{s}_j$, the average firm size.

Table 8 presents estimates of the key Equation (7) for the years 1985–1986, 1986–1987, 1987–1988, and 1988–1989. As in our industry regressions, data on retail trade employment and pay per person are from the CBP and are from March of year $t-1$ to March of year t. For our measure of retail sales growth, Δr_j, we use August-to-August county retail trade data based on sales-tax information from the California State Board of Equalization. For example, in 1985–1986 regression, we use the percent change in taxable retail sales in the county from the 12-month period, August 1984 through August 1985, to August 1985 through August 1986. As instruments in the regression for years $t-1$ to t, we use average wages and firm size for the year $t-2$, as calculated from the CBP data. (We do not take our instruments from the BLS in this case, because these data are not tabulated at the county level.) CBP data were available to us for all but one (the smallest) of the 58 counties in California.

The results for 1985–1986, 1986–1987, and 1987–1988 are as anticipated: IV estimates of the coefficient on Δw_j are not significantly different from 0. Results are strikingly different for 1988–1989. Having conditioned on retail sales growth, we find a strong negative correlation between county wage changes and employment changes.

Because the regressions we present are weighted by the retail employment in the county (for year $t-2$), readers may wonder about the extent to which the results are being driven by a few large counties. If we use unweighted 2SLS instead, though, we find qualitatively similar results for 1988–1989. Finally, using the Lagrange multiplier test for overidentification, we find that $nR^2 = .006$. Given that this test statistic is asymptotically distributed chi-squared with 1 df, we clearly cannot reject our specification.

Reduced-form estimates for equations underlying our 2SLS procedure are presented in Table 9. Estimates in the second column show a positive correlation between county employment growth and county growth in retail sales. Furthermore, having conditioned on the change in retail sales, we find that employment growth is slowest in counties with low initial average wages and large average firm sizes.

We thus find that California's 1988 minimum wage increase looks nearly the same from two different angles: In Section 2, we show that in retail trade industries in which the

Table 8. Regression Results for Dependent Variable: Δe in County j

	1985–1986		1986–1987		1987–1988		1988–1989	
	OLS (1)	IV (2)	OLS (3)	IV (4)	OLS (5)	IV (6)	OLS (7)	IV (8)
Intercept	.022	.037	.000	−.037	−.003	−.001	.017	.017
	(.006)	(.019)	(.009)	(.031)	(.012)	(.014)	(.008)	(.012)
Coefficient	−.039	1.849	.028	1.661	−.016	−.114	−.735*	−.734*
on ΔW_j	(.118)	(1.576)	(.162)	(1.225)	(.111)	(.360)	(.103)	(.221)
Coefficient	.608*	−.122	.784*	1.042*	.461*	.472*	.689*	.689*
on Δr_j	(.121)	(.666)	(.134)	(.294)	(.160)	(.165)	(.093)	(.106)
n	57	57	57	57	57	57	57	57
$\overline{R}^2$	.32	.07	.37	.18	.10	.10	.59	.42

NOTE: All regressions are weighted by employment in county j in year $t-2$. Asymptotic standard errors are given in parentheses (*indicates sig. <.01). For the IV estimates, the instruments are log mean wage and log average firm size in county j in year $t-2$.

Table 9. Reduced-Form Regression Results: County Analysis 1988–1989

	Dependent variable ΔW in county j	Dependent variable Δe in county j
Intercept	.046	.001
	(.044)	(.049)
$\overline{W}_j$	−.111*	.078**
	(.030)	(.034)
$\overline{s}_j$	.051*	−.043**
	(.018)	(.021)
Δr_j	.057	.635*
	(.120)	(.136)
n	57	57
$\overline{R}^2$	.24	.27

NOTE: All regressions are weighted by employment in county j in year 1987. Asymptotic standard errors are given in parentheses (* indicates sig. <.01 and ** indicates sig. <.05). The instruments are log mean wage, log average firm size, and change of sales in county j in year 1987.

wage grew rapidly in California as a result of the increase in the minimum wage, employment was adversely affected. In this section, we find that in counties where retail trade wages increased due to the minimum wage boost, county employment growth in retail trade was tempered. Moreover, our IV estimate based on county-level data [see column (8) of Table 8] suggests an employment elasticity, −.7, not greatly different from the elasticities presented in Section 2.

4. CONCLUSION

One of the most direct and compelling applications of conventional theory to public policy is the prediction that an increase in the minimum wage reduces employment in competitive markets.

Because California's large July 1988 minimum wage increase represents an excellent opportunity to empirically evaluate the effects of minimum wages and because of Card's (1992a) findings regarding this minimum wage increase, we have undertaken a further investigation of the effects of the minimum wage increase in the low-wage retail trade industry. We first show that the increase in wages in California's retail trade (from the first quarter of 1988 to the first quarter of 1989) was more than two full percentage points higher than for the rest of the United States, due, we believe, to the increase in California's minimum wage. We then show that within industries in retail trade there is a strong negative correlation between the relative wage changes and employment growth. IV estimates suggest an elasticity of employment with respect to wages in the neighborhood of −.9. We also explore the effect of the minimum wage increase by examining intercounty variation in the wage change in retail trade and the corresponding employment change. Again we find a strong negative correlation, with a wage elasticity for retail trade employment of about −.7.

These results appear to differ from some recently published work. As we mention in the introduction, a series of careful empirical studies have documented cases in which minimum wage increases have not resulted in reduced employment. There is also a growing theoretical literature providing explanations for such outcomes [including Burdett and Mortenson (in press), Lang and Dickens (1992), and Rebitzer and Taylor (in press)]. Still, our reading of the 1988 increase in California's minimum wage supports the more conventional view about the effects of the minimum wage for a very large sector of low-wage employment, retail trade. Indeed, we are in the fortunate position of having found two natural ways of evaluating the effect of a minimum wage increase, and both appear to reject the null hypothesis that the minimum wage increase had no effect on employment.

ACKNOWLEDGMENTS

Taylor gratefully acknowledges research support from the Institute for Research on Poverty, University of Wisconsin-Madison. We thank David Card, John Engberg, Seth Sanders, and the editor and associate editor of this journal for helpful comments.

[Received August 1993. Revised September 1994.]

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