WAGE INEQUALITY IN THE UNITED STATES DURING THE 1980s: RISING DISPERSION OR FALLING MINIMUM WAGE?*

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The magnitude of growth in "underlying" wage inequality in the United States during the 1980s is obscured by a concurrent decline in the federal minimum wage, which itself could cause an increase in observed wage inequality. This study uses regional variation in the relative level of the federal minimum wage to separately identify the impact of the minimum wage from nationwide growth in "latent" wage dispersion during the 1980s. The analysis suggests that the minimum wage can account for much of the rise in dispersion in the lower tail of the wage distribution, particularly for women.

I. INTRODUCTION

A striking feature of the United States labor market experience during the past twenty years has been the dramatic rise in earnings and wage inequality that occurred during the 1980s.[1] Past research has documented the various dimensions of this trend: the sharp rise in wage differences between more- and less-educated workers, the growing wage disparity between more- and less-experienced workers, and the rise in wage inequality within groups narrowly defined by age, education, and gender—so-called "within-group" inequality.[2] Researchers have proposed a number of explanations for these relative wage movements, most of which can be characterized as either a demand explanation (technological change, import competition), a supply story (immi-

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1. Levy and Murnane [1992] and Katz and Autor [1999] provide a survey of research on earnings and wage inequality. Gottschalk [1997] also provides a summary of the general trends and dimensions of wage inequality in the United States.

2. See Katz and Murphy [1992]; Juhn, Murphy, and Pierce [1993]; and Murphy and Welch [1992].

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gration of less-skilled workers), or an institutional factor (deunionization, the minimum wage).[3]

An equally striking development that accompanied dramatic increases in wage inequality in the United States during the 1980s was the longest sustained decline in the real value of the federal minimum wage in the previous four decades, as it reached its lowest point since the 1950s.[4] As shown below, decreases in the minimum wage tend to increase measured wage inequality, regardless of its effect on employment, earnings, or income. DiNardo, Fortin, and Lemieux [1996] demonstrate that the minimum wage is an important empirical feature of the observed distribution of wages during the 1980s. At the least, their findings suggest that the declining real value of the minimum wage significantly obscures the impact of factors such as relative supply and demand shifts or other institutional changes on the wage distribution throughout the decade. In this paper I attempt to answer the question: after accounting for the impact of the minimum wage, did "underlying" wage inequality in the United States rise during the 1980s? In particular, what was the net impact of all factors, other than the minimum wage, on the wage distribution?

It is difficult to answer these questions from an analysis of aggregate time series data because the steady increase in the dispersion of observed wages is closely paralleled by a steady decrease in the relative level of the minimum wage during the 1980s, making it virtually impossible to separate the effect of the minimum wage from a time trend.

Instead, this study utilizes the differential impact of the federal minimum wage across regions within the United States to directly estimate the contribution of the falling relative value of the minimum wage to increasing wage inequality in the lower half of the wage distribution since 1979. The interaction of state variation in overall wage levels and a uniform federal minimum wage generates cross-state variation in "effective" minimum wage levels. The assumption that this variation is not systematically related to the shape of the "underlying" wage distribution—the distribution of wage rates that would prevail in the absence of any minimum wage—permits the data to separately identify the

3. For a series of articles surveying the leading explanations for rising wage inequality, see Johnson [1997], Topel [1997], and Fortin and Lemieux [1997].

4. In 1996 dollars using the CPI-U. Nominal rates are from Bureau of the Census [1996].

impact of the minimum wage from nationwide growth in "latent" wage dispersion during the 1980s.

When regional wage distribution data constructed from the Current Population Survey are used, the resulting estimates of the impact of the minimum wage imply that after accounting for the minimum wage, average growth in wage dispersion in the lower tail of states' wage distributions during the 1980s is quite modest. In fact, the estimates for men, women, as well as the combined sample, imply that almost all of the growth in the wage gap between the tenth and fiftieth percentiles is attributable to the erosion of the real value of the federal minimum during the decade.

I provide some evidence suggesting that the identifying assumption may be inappropriate for the male sample of workers. On the other hand, the same kind of evidence lends support to the validity of the main identification strategy when analyzing women, and when considering the unconditional distribution of wages. For example, estimates from using variation in *legislated changes* in minimum wages (which arise from an interaction between the 1990–1991 federal increases and preexisting variability in state-specific minimum wage laws) are quite similar to those generated by cross-state variability in the relative minimum.

The results also imply that accounting for the minimum wage only moderately affects the magnitude of changes in wage differentials by gender, race, educational attainment, and experience, and leaves the broad trends unaffected. On the other hand, this paper's results imply that the minimum wage may account for as much as 80 percent of the growth in so-called "within-group" wage inequality during the 1980s. When examining men and women separately, overall growth in wage inequality during the 1980s remains substantial, due to the appreciable growth in dispersion in the upper tail of the wage distribution, where the minimum wage is unlikely to have an impact. Curiously, after accounting for the minimum wage, the *unconditional* distribution (without regard to gender) of wages appears to be surprisingly stable throughout the 1980s.

The paper is organized as follows. Section II provides a brief summary of the trends in wage inequality that serve to motivate the present study. Section III discusses how state variation in effective minimum wage levels can be used to separately identify the effect of a declining minimum wage from an underlying rise in wage dispersion. Section IV describes the empirical implementation and presents the results of the estimation. Section V examines the effect of the increases in the federal minimum wage on wage inequality from 1989 to 1991. Section VI uses the estimates of this study to adjust measures of between- and within-group wage inequality for the minimum wage effect. Finally, Section VII concludes with a discussion of the implications of the findings in this paper for future research.

II. Changes in the Wage Distribution in the 1980s

A. Data

The analysis in this paper utilizes microdata from the National Bureau of Economic Research Extracts of the Current Population Survey (CPS) Outgoing Rotation Group Earnings Files. These data consist of point-in-time measures of the hourly rate of pay, which make it appropriate for a study of the impact of a wage floor on the wage distribution. For workers not paid on an hourly basis, usual weekly earnings and usual weekly hours can be used to construct an average hourly wage. These data's large sample sizes (three times the size of a single month of data from the CPS) and annual frequency make them suitable for a detailed regional analysis throughout the decade of the 1980s. Details of the preprocessing of the data are reported in the Data Appendix.

The three parts of Figure I depict the (hours-weighted) kernel density estimate of the distribution of log-wages from these data for the years 1979 and 1989, for men, women, and all workers (men and women combined).[5] In each panel the two vertical lines denote the federal minimum wage levels in each year. The figure suggests that in 1979 the minimum wage had, to some extent, a supporting effect on wages; after its relative level's erosion during the decade, by 1989, the minimum wage appears to have much less of an influence on the shape of the distribution. The impression that the minimum wage was an important feature of the lower part of the wage distribution is particularly striking for the female wage distribution. Figure IC suggests the possibility that the decline in the relative value of the minimum wage over the decade was primarily responsible for changes in the unconditional wage distribution, since much of the change appears to occur in the lower tail.

5. The sample includes individuals aged sixteen and over. The horizontal axis measures the log-wage relative to the overall (men and women) median in each year. All kernel density estimates use the rule-of-thumb bandwidth $h = 0.79Rn^{-1/5}$, where R is the interquartile range [Silverman 1986].

This paper follows the wage inequality literature's usual practice of conducting formal analyses for men and women separately. But since changes in the lower tail appear to be a relatively more important component of the total change when examining the unconditional distribution of wages (without regard to gender), some attention will also be given to the findings for the entire sample. Examining the entire sample—the distribution of hourly pay for hours worked in the economy—allows some abstraction from issues that arise when men and women are thought to compete in the same labor market, when changes in the male and female distributions are to some extent interdependent.[6] Thus, this paper's estimates of changes in the unconditional distribution of wages—after parsing out the minimum wage effect—could be interpreted as the *net* impact of between-gender wage changes [O'Neill and Polachek 1993; Fortin and Lemieux 1996; Blau and Kahn 1997], and other "within-gender" forces such as skill-biased technological change [Davis and Haltiwanger 1991; Bound and Johnson 1992; Autor, Katz, and Krueger 1997; Doms, Dunne, and Troske 1997], international trade [Borjas and Ramey 1995; Feenstra and Hanson 1996], and more generally rising "skill premiums" stemming from an increase in the demand for skilled labor [Katz and Murphy 1992; Berman, Bound, and Griliches 1994].

The available time series data on the United States wage distribution also point to a close connection between the minimum wage and changes in inequality. Figure II plots changes in percentiles of the wage distribution, relative to the median, for the fifth, tenth, twenty-fifth, seventy-fifth, and ninetieth wage percentiles between 1973 and 1993.[7] These measures of wage dispersion, as well as the log difference between the federal minimum wage and the median wage (also shown in the figure) are normalized to be zero in 1979, which appears to be a turning point for several of the series.

Figure II shows that after rising modestly during the latter half of the 1970s, the minimum wage (relative to the median wage) fell almost 40 log points in the period between 1979 and 1989. The sharp rebound in 1990 and 1991 reflects the legislated rise in the federal minimum wage from $3.35 to $3.80 in April of

6. This issue is addressed in Fortin and Lemieux [1996] and Topel [1997].
7. Includes both men and women, all those aged 16 and over. The data for 1973–1978 are taken from the May CPS dual job supplements. These data contain a point-in-time wage measure comparable to that from the Outgoing Rotation Group Files for 1979–1993.

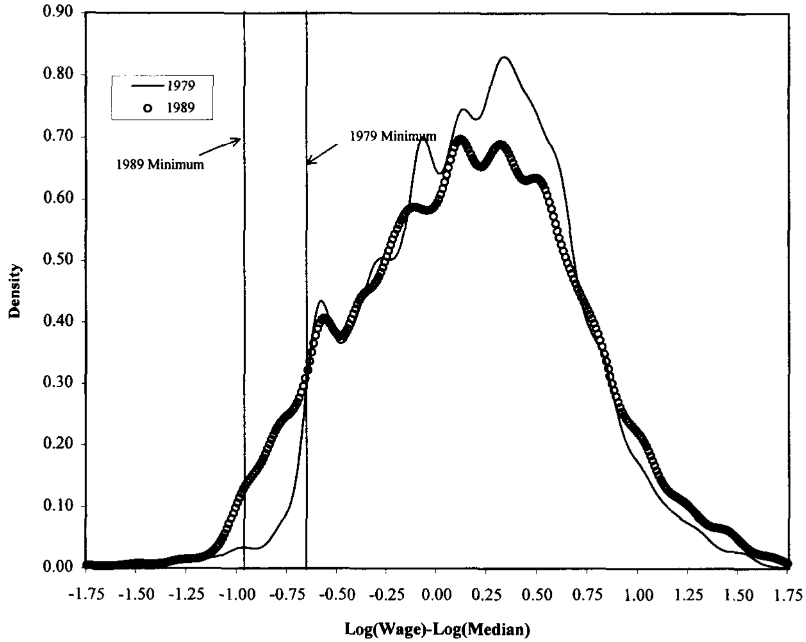

FIGURE Ia
Wage Distribution Density Estimates: Men, 1979–1989

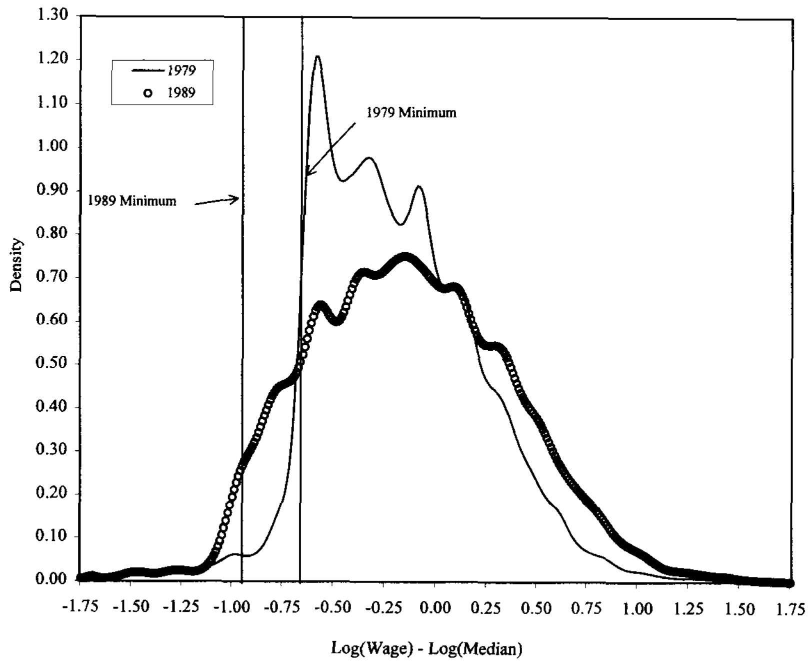

FIGURE Ib
Wage Distribution Density Estimates: Women, 1979–1989

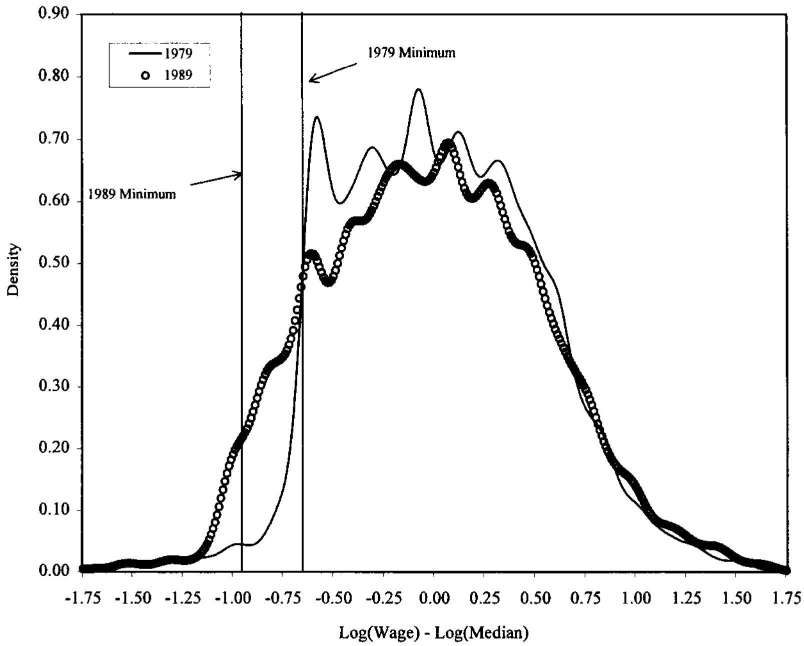

FIGURE Ic
Wage Distribution Density Estimates: Men and Women, 1979–1989

1990 and then to $4.25 in April of 1991. The figure illustrates that with the exception of the last few years of the 1980s and the period after 1991, movements in wage dispersion in the lower half of the wage distribution moved in close tandem with changes in the minimum wage. The fifth and tenth percentiles rise in the latter part of the 1970s, experience a sharp decline during much of the 1980s, and then rebound modestly in the early 1990s.[8]

Finally, Figure II shows that while changes in the seventy-fifth wage percentile have been small, there has been a steady rise in the ninetieth percentile throughout the 1973–1993 period. That the minimum wage is unlikely to have played any part in this particular trend underscores the limited inferences that can be made from an examination of the aggregate time series data. For example, there is nearly as much evidence here to suggest that the

8. It should be noted that trends in measures of dispersion in average hourly earnings as computed from the Census and the March CPS differ somewhat from that calculated from the May CPS during the 1970s. For example, the 90–10 differential for women falls by about 11 percent in the May CPS, but rises 3 percent in the March CPS. See Katz and Autor [1998] for a summary of differences across the data sets over the past few decades.

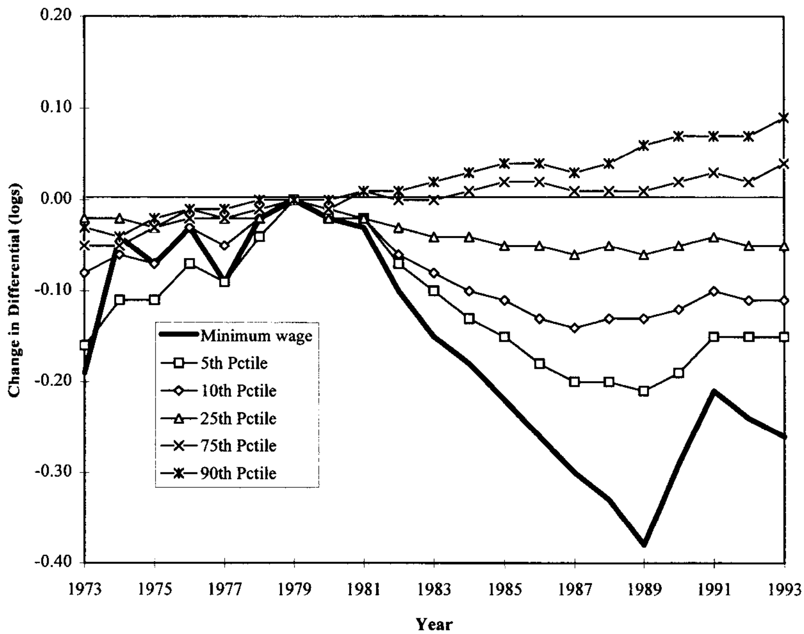

FIGURE II
Selected Percentiles of the Wage Distribution, Minimum Wage, Relative to the Median: 1973–1993
All series are normalized to be 0 in 1979.

minimum wage contributed to growth in the ninetieth percentile, as there is to suggest that it caused a decline in the tenth percentile. The effective decline in the minimum wage, and the growth in inequality in the lower half of the wage distribution, are both monotonic throughout the 1980s. As a result, no analysis of the aggregate time series data will be able to adequately distinguish a minimum wage effect from a rising trend in "latent" wage inequality—which I define here to be a rise in wage inequality that would have occurred in the absence of a minimum wage policy.

B. Regional Variation in the "Effective" Minimum Wage

The time-series pattern of wage dispersion in the lower tail is not uniform across regions in the United States. Figure III plots the average 10–50 log (wage) differential during 1979–1991 for two groups of states: the three highest-wage states (New Jersey,

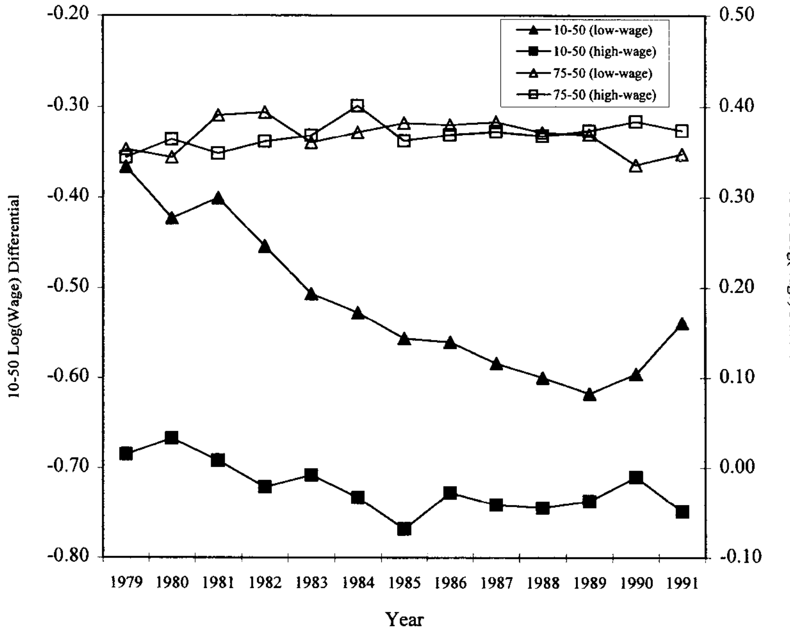

FIGURE III
10–50, 75–50 Log(Wage) Differentials: High- versus Low-Wage States, 1979–1991

Michigan, and Maryland) and the three lowest-wage states (South Dakota, Mississippi, and Arkansas).[9] The low-wage states, which initially begin with a more compressed lower tail than the high-wage states, experience an average increase in dispersion of about twenty log points throughout the 1980s, and a subsequent compression of about five log points during 1989–1991. By contrast, the 10–50 differential for the highest-wage states appears relatively stable during the entire 1979–1991 period.

This additional dimension of wage inequality and its relation to the relative level of the minimum wage across regions offer an opportunity to distinguish the minimum wage effect from a national trend in latent wage dispersion. In particular, the interaction between a uniform federal minimum wage and wide

9. These six states were selected from the state-level panel data set described in the Data Appendix. Average median wages for the entire period were computed. The three highest and three lowest overall averages were chosen, excluding all states for which there was at least one year where the state minimum exceeded the federal minimum wage rate. The differentials are for the entire sample aged sixteen or over.

variation in wage levels across states yields variability in "effective" minimum wage levels. The federal minimum wage highly impacts low-wage states, while minimally affecting high-wage states.

This is demonstrated in Figure IV, which plots (hours-weighted) kernel density estimates of log-wages in 1979 for three groups of states: high-, medium-, and low-wage states.[10] The horizontal scale measures the log-wage relative to each group's respective median wage. The vertical lines represent the federal minimum wage, also relative to each group's respective median. Figure IV gives the impression that low-wage states are indeed more highly impacted by the minimum wage than higher-wage states. It is this variation that is exploited in the present study.

Stratifying the wage distribution by a dimension other than geography would likely produce a similar result. And cross-sectional variation in effective minimum wages across educational groups, for example, could be used to estimate a minimum wage effect. However, in this study variation across states has a few important advantages over variation across other groups such as education, age, industry, or occupation.

First, on a conceptual level, variation in the relative minimum across states more closely mimics the variation that would be most ideal (but that is also unattainable) for estimating the impact of the minimum wage on the wage distribution. These ideal data would consist of a sample of independent realizations of the wage distribution of the United States labor market at one point in time, and exhibit variability in the minimum wage across realizations. Each state's labor market can be more readily considered a microcosm of the United States labor market than can any particular group of workers defined by education, age, or industrial or occupational sector.

Second, as described in detail in Section V, by the late-1980s there arose moderate variation in state-legislated minimum wage levels. This variability in state-specific minimum wages, combined with the 1990–1991 legislated increases in the federal minimum, generates an alternative source of variation from which the impact of the minimum wage can be identified.

A potential problem with utilizing the substantial regional

10. I rank each state by the mean log-wage. I choose the bottom seven (Arkansas, Mississippi, Vermont, Maine, South Carolina, South Dakota, and North Carolina), middle seven (Idaho, Oklahoma, Kentucky, Kansas, Virginia, Missouri, and Indiana), and highest seven (Oregon, Wyoming, California, Illinois, Maryland, Michigan, and Washington) for the three groups. The sample sizes for the low-, medium-, and high-wage states are 12,263, 16,190, 31,932, respectively.

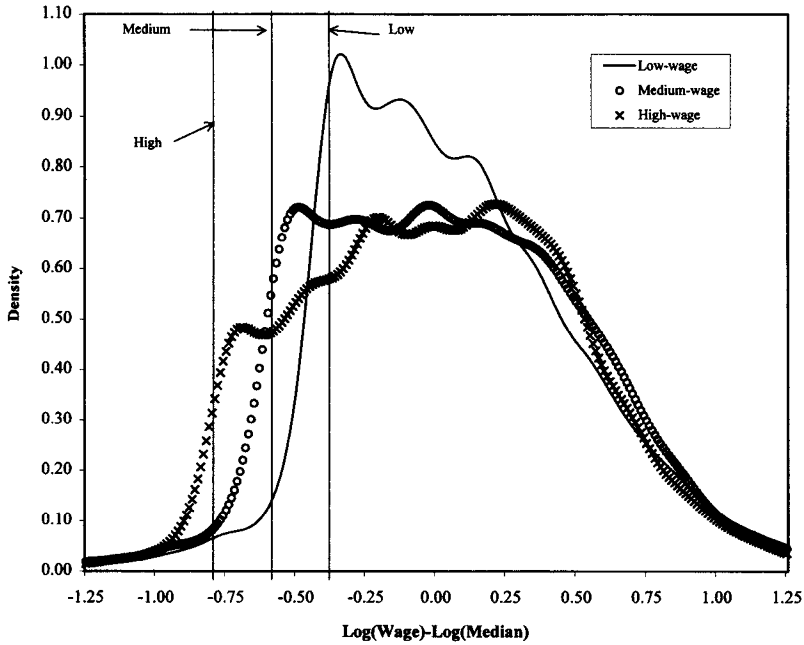

FIGURE IV
Wage Distribution Density Estimates:
Low-, Medium-, and High-Wage States, 1979

variation in wage levels, is that if high-wage states tend to possess greater latent wage dispersion, the use of the cross-sectional variation illustrated in Figure IV would lead to an exaggerated estimate of the impact of the minimum wage. A test of whether overall latent wage dispersion is *empirically* related to wage levels lies in being able to detect such a correlation where the minimum wage is not likely to be a factor, such as at the upper tail of the wage distribution. Figure III provides evidence to suggest that such a systematic relation is not apparent in these data. The average 75–50 differentials for the very highest- and lowest-wage states appear to be comparable throughout the 1979–1991 period.

Of course, the assumption that this property also holds for the lower tails of states' latent wage distributions is, strictly speaking, untestable. However, a significant, systematic relation between latent wage dispersion in the upper tail and overall wage levels (and hence the "effective" minimum wage) across regions would seem to considerably weaken our confidence in its validity. Hence, an examination of the relation between the relative minimum wage and measures of dispersion of the upper tail will provide a useful specification check for the formal empirical analyses presented below.

III. Identification of Changes in Latent Wage Inequality

A. Theoretical Relation between Wage Dispersion and the Minimum Wage

Below, I formally model the relation between the variability in the relative "bite" of the minimum wage and the observed wage distribution across regions in a way that allows separate identification of the average growth in latent wage inequality. The approach in this study is to use the various log(wage) percentile differentials to measure wage dispersion, and the log-differential between the logs of the minimum wage and some measure of the centrality (or location) of the distribution to measure the "effective minimum wage."[11] Below, I characterize the theoretical relation between these two quantities under three distinct scenarios. I denote the latent and observed pth log-wage percentile of state j by w_j^{p*} and w_j^p, respectively, and the log of the nominal minimum wage by *minwage.* Although I utilize other measures of centrality in the formal empirical analysis, I begin by considering w_j^{50}, the median wage. Also, for the sake of exposition, I assume that the shapes of the latent wage distribution in all states are strictly identical, up to location: $w_j^{p*} - w_j^{50*} = w_k^{p*} - w_k^{50*}\ \forall j,k$.[12]

Case 1, "Censoring": no spillovers, no disemployment. In this extreme case, the only effect of the minimum wage is to raise the wages of those initially making less than the minimum to exactly the level of the wage floor. Across states we expect to see the relation,

$$(1)\quad \begin{aligned} w_j^{10} - w_j^{50} &= w_j^{10*} - w_j^{50*} \quad \text{if } (minwage - w_j^{50}) < (w_j^{10*} - w_j^{50*}) \\ &= (minwage - w_j^{50}) \text{ otherwise.} \end{aligned}$$

11. In the remainder of the paper, I often refer to the effective minimum wage as the "relative minimum wage" or simply "the minimum wage," when dealing with a cross section of states.

12. That the latent distributions are strictly identical across states is certainly false in practice, but abstracting from stochastic elements, which will be more formally introduced in Section IV, aids in describing the intuition behind the identification.

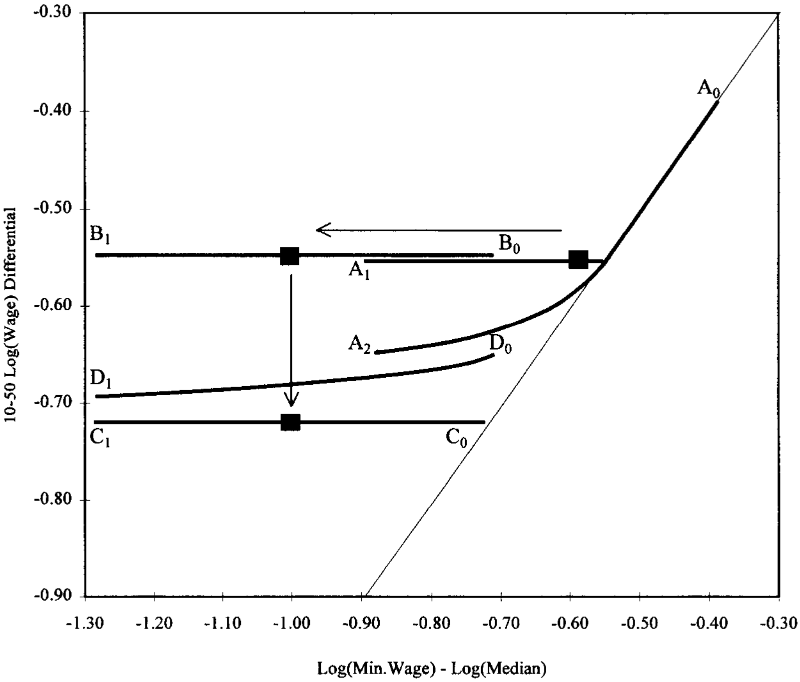

FIGURE V
Minimum Wage Decline and Growth in Latent Wage Dispersion

This relation across states is depicted in Figure V as line A_0A_1.[13] For states that possess "effective minimums" that are smaller than the latent 10–50 differential, there is no relation between the differential and the relative minimum (the flat portion of the line). For states where the relative minimum is above the latent 10–50 differential, the observed 10–50 differential is exactly equal to the differential between the minimum wage and the median (the portion coincident with the 45 degree line).

Case 2, "Spillovers," no disemployment. More generally, the minimum wage may have an effect on the pth percentile even if $(minwage - w_j^{50}) < (w_j^p - w_j^{50})$, so that we have

$$w_j^{10} - w_j^{50} = g(minwage - w_j^{50}), \tag{2}$$

13. I am assuming that the minimum wage is never higher than the latent median wage for any state.

where g will generally be an increasing function, as long as the "spillover" effect monotonically diminishes, the higher the wage percentile. An example of such a function is depicted as line A_0A_2 in Figure V. Note that, across states, a marginal rise in the effective minimum, even among states with minimums lower than the 10–50 differential, causes an increase in the tenth wage percentile relative to the median.

Case 3, "Truncation": no spillovers, full disemployment. In this case, the minimum wage has no impact on workers with latent wages already above the minimum, and causes job loss for all workers with latent wages below the minimum (i.e., truncation of the wage distribution); we will not observe the latter's wages. The loss of the sample in the lower tail mechanically leads to changes in the observed percentiles of the wage distribution. For example, if, for state j, $minwage = w_j^{10*}$, then this model implies that we will not observe wages for workers who would have earned w_j^{10*} (or less) in the absence of the minimum. As a result, the observed (posttruncation) w_j^{10} cannot equal w_j^{10*}; instead, it will equal the $10 + (0.10 \times 90) = 19$th percentile of the latent wage distribution. As shown in Lee [1998], under some regularity conditions for the shape of the latent wage distribution, this truncation model implies that $w_j^{10} - w_j^{50}$ will be an increasing function of $(minwage - w_j^{50})$; simulations of the shape of this relation using actual wage data show that it is generally increasing and exhibits curvature similar to that depicted by line A_0A_2 in Figure V.

In each of these three cases (and in hybrids), there is a nonlinear relation that is expected between the relative minimum wage measure and the observed 10–50 differential. As the relative minimum declines indefinitely, the relation asymptotes to a horizontal line, corresponding to the latent 10–50 differential that is common across all states. I forgo an attempt to empirically distinguish between disemployment (Case 3) and spillover effects (Case 2) of the minimum wage on the observed wage distribution. Instead, I simply note that in the arguably more realistic case of some disemployment and some spillover effects (a hybrid of Cases 2 and 3), any positive empirical relation between $w_j^{10} - w_j^{50}$ and $(minwage - w_j^{50})$ will be an overestimate of true spillover effects; some of it will be a statistical "illusion" that is a natural consequence of both employment loss and the fact that we only observe wages for those who are working.[14]

Consider first a significant decline in the relative value of the federal minimum wage. We expect a leftward shift in the locus representing the 50 states, from line A_0A_1 to B_0B_1 in Figure V, for example. In addition, if all of the states experienced an expansion in the latent 10–50 wage differential, this would be additionally reflected in a downward shift from line B_0B_1 to line C_0C_1. When these two events happen simultaneously, we will only observe data corresponding to lines A_0A_1 and C_0C_1. In this example, the rise in observed wage dispersion of the median state (which is denoted by the black square of each locus) is fully due to a rise in latent wage inequality rather than to the falling relative level of the minimum wage. Figure V also illustrates a contrasting example, where the states' relative minimums and the 10–50 differentials are represented by lines A_0A_2 (in the initial period) and D_0D_1 (the latter period). In this case, there is a much smaller increase in latent wage inequality, with most of the average rise in dispersion attributable to the decline in the federal minimum.

B. Wage Dispersion and the Minimum Wage across States

Figure VIa is the empirical analog of Figure V, constructed from a sample of states in 1979 and 1989.[15] The horizontal axis measures the federal minimum wage, relative to the state median wage, and the vertical axis measures the 10–50 log-wage differential. The two solid lines represent the fitted values of Ordinary Least Squares (OLS) regressions (weighted by the sample size of the state-year), one for each year, of $(w^{10} - w^{50})$ on $(minwage - w^{50})$ and its squared value.

The figure reveals a strong positive association between the relative positions of the tenth percentile and the minimum wage, particularly in 1979, when the tenth percentile is either at or

14. As discussed in Lee [1998], the full truncation model produces a small negative relation between upper tail measures of dispersion and the relative minimum, since the truncation effect diminishes higher up in the distribution. In the more realistic case of some disemployment and some spillovers, the relation is even weaker.

15. Data come from the state-level panel data set described in the Appendix. So that all the variation in the effective minimum wage comes from variation in the states' medians, I exclude states with legislated minimum wage rates higher than the federal minimum: Alaska for 1979, and Alaska, California, Connecticut, Hawaii, Maine, Minnesota, Massachusetts, New Hampshire, Pennsylvania, Rhode Island, Vermont, Washington, and Wisconsin for 1989. Robustness to differing samples is mentioned in Section IV.

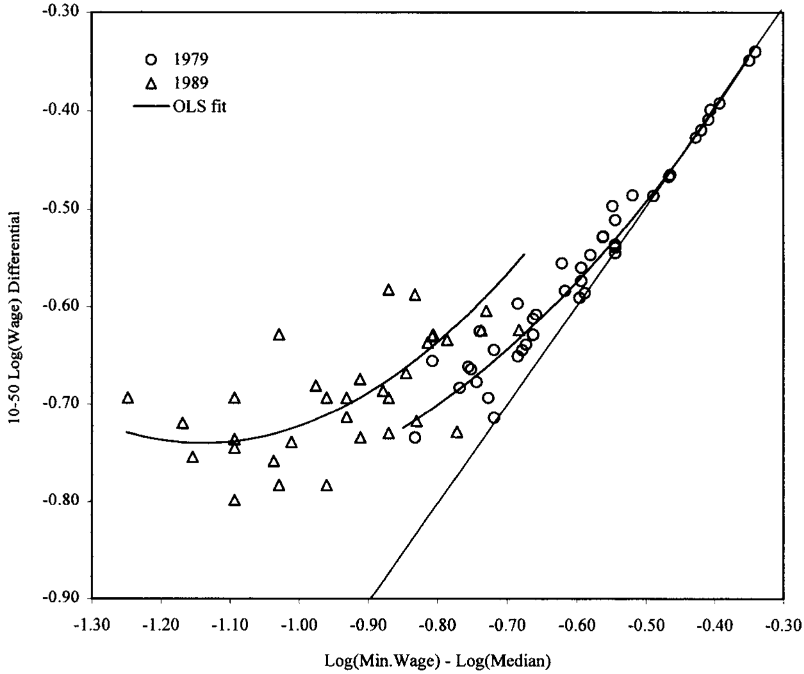

FIGURE VIa
10–50 Log (Wage) Differential, Relative Minimum Wage, across States: 1979, 1989

slightly above the minimum wage for a majority of the states. The regression lines, especially that for 1989, exhibit some nonlinearity, as the relation between the effective minimum wage and the 10–50 differential appears to flatten to the left. The estimated coefficients on the quadratic terms for both years are each statistically significant from zero at the 0.05 level.

Overall, there appears to be little evidence of a downward shift in the wage dispersion-minimum wage relation over time. In fact, an OLS regression using the data from both years, and including a year dummy variable, yields a coefficient of 0.062 (with a standard error of 0.017) on the 1989 dummy, implying that there is a slight *upward* shift in the relation—a *decrease* in latent wage dispersion as measured by the 10–50 wage differential. Most of the data points in 1989 are significantly above the 45 degree line. This implies that there was a real potential for the states'

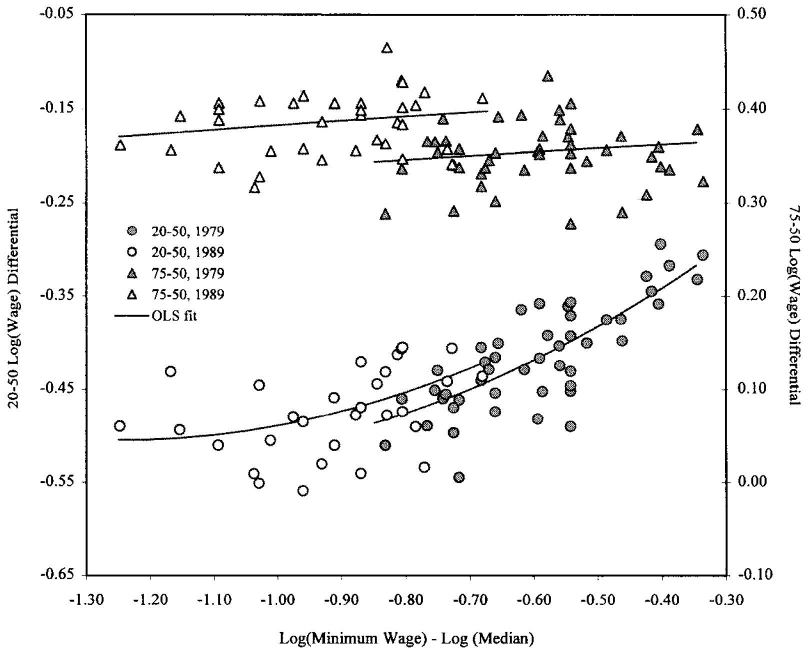

FIGURE VIb
20–50, 75–50 Log(Wage) Differentials, Relative Minimum Wage, across States: 1979, 1989

tenth percentiles to fall farther than they did. However, the 10–50 differentials fell by an amount that was no more—and, if anything, less—than what would be expected as a response to a uniform decline in the relative level of the minimum wage.

An important caveat to this empirical approach is that, in the presence of a nonlinear relation between the relative minimum and wage percentile differentials, the identification of a downward shift in the relation becomes difficult when the minimum wage is "too binding" across all states. For example, suppose that in 1979, all of the states' 10–50 lay exactly on the 45-degree line in Figure VIa. Even in the absence of any change in the relative minimum, we could expect an increase in latent wage dispersion to produce no change in the empirical relation; the states would continue to lie on the 45-degree line, and no downward shift could be detected. Nor could we detect a modest *upward* shift in the relation if the latent tenth percentile were initially far enough below the relative minimum for all states.[16]

The advantage of using percentile differentials as the dependent variable is that we can examine other wage percentiles, higher up in the wage distribution, where this problem does not exist.[17] For example, Figure VIb plots the analogous empirical relation to Figure VIa, for the 20–50 log-wage differential. Again, the figure reveals that after accounting for the declining federal minimum, there appears to be a slight *decrease* in wage dispersion.[18] Examining samples of workers with generally higher wages (older workers, men) (as is done in Section IV) similarly alleviates the potential identification problem described above.

Before presenting the formal empirical analysis, it is important to clarify the nature of what appears to be a "mechanical" relation between the dependent and independent variables, since the state median wage is used in the measure of dispersion, as well as to construct the relative minimum wage variable. It appears that the empirical relation between these two quantities may exaggerate the impact of the minimum wage. There are two distinct components to this possibility.

The first arises when sampling error is an important part of the total variability in state median wages. For example, even if no relation exists between the two variables, the fact that the sample median and the sample tenth percentile are not perfectly correlated implies there will be a spurious positive relation between the 10–50 differential and the relative minimum. However, the state-by-state sample sizes from the CPS Outgoing Rotation Group Files are reasonably large (on average about 3000 per state-year), and hence on average the sampling error for the typical state's median is about 0.01, implying that less than 1 percent of the variation in the relative minimum is due to sampling error.[19] Furthermore, this problem can be somewhat

16. Identification is also weakened when there is little "overlap" of the ranges of the effective minimum for the two years. If there is no overlap at all, identification relies heavily on functional form assumptions about the underlying shape of the minimum wage's impact on wage dispersion. In the formal empirical analysis, however, sufficient overlap is generated by using data from all intermediate years between 1979 and 1989.

17. An examination of Appendix 2 provides a sense of how close the relative minimums are to the various wage percentiles across states.

18. The coefficient on the 1989 year dummy in a (weighted by observation per state-year) pooled regression of the 20–50 differential on the relative minimum and its square is 0.0226 with a standard error of (0.0123).

19. A similar calculation yields that sampling variability is about 10 percent of the within-state variance.

alleviated by using an alternative measure of centrality to create the relative minimum wage variable, which I utilize in Section IV.

Second, apart from the sampling issue, even the absence of the minimum wage, there may be relatively greater variability in the median than in percentiles of the lower or upper tails. For example, if there is relatively little variability in the tenth latent wage percentiles across states, then on average a state with a large median wage will have both a lower relative minimum, as well as a larger gap between the tenth percentile and the median, independently of any minimum wage effect. It should be noted that this possibility is one specific example of a violation of the identifying assumption: that, on average, wage levels are uncorrelated with latent wage dispersion. On the other hand, if we could be assured that locational amenities, for example, generated compensating state wage differentials that were constant across all of each state's percentiles wage distribution, and that the difference in the median was an adequate measure of the state wage differential, then no relation would exist between any of the latent percentile differentials and the relative minimum wage, even though the median would be used to construct both measures.

More generally, some scenarios (including the hypothesis described above—greater variability across states in the middle of the distribution than in the tails) that posit a systematic relation between latent wage dispersion and the location of the distribution, as measured by the median, will have the observable implication of a significant relation between upper tail measures of dispersion and the median. Figure VIb, which plots the 75–50 differential against the relative minimum for this subsample, illustrates that such a correlation is weak in these data.[20] This finding is at odds with the notion that high-wage states inherently possess uniformly greater latent wage dispersion.

In order to be consistent with the pattern of data shown in Figure VI, alternative stories positing a correlation between latent wage dispersion and the median wage must be able to explain (1) why only the upper tail measure of dispersion appears uncorrelated with median, and (2) why the empirical relation for the lower tail is much stronger in 1979 than in 1989, diminishing over the course of the 1980s.

20. The coefficient on the relative minimum variable in a pooled regression of the 75–50 differential on a year dummy and the minimum is 0.046 with a standard error of 0.028.

IV. Estimation of Changes in Latent Wage Inequality: 1979–1988

A. Estimating Equation

In the presence of stochastic variation in latent wage dispersion across states, the main identifying assumption used in this study can be motivated as follows. Suppose that each state j at time t has a *latent* log-wage distribution that can be characterized by the cumulative distribution function,

$$F_t((w - \mu_{jt})/\sigma_{jt}), \tag{3}$$

where μ_{jt} and σ_{jt} are centrality and scale parameters, respectively. $F_t(\cdot)$, the "shape" of the latent wage distribution in year t, is constant across regions. The latent log-wage percentile p of state j at time t is thus

$$w_{jt}^{p*} = \mu_{jt} + \sigma_{jt}F_t^{-1}(p), \tag{4}$$

where the asterisk emphasizes that this is the latent wage percentile. Normalize the average $\bar{\sigma}_t = E[\sigma_{jt}|t]$ to be constant over time, $\bar{\sigma}_t = 1$. Then the quantity of interest is the change in the dispersion (particularly, in the lower tail) of $F_t(\cdot)$ over time.

The main identifying assumption in this analysis is that conditional on t, σ_{jt} is independent of μ_{jt}: conditional on the year, the centrality measure of the state wage distribution is not systematically correlated with latent wage dispersion across states. If the observed median wage w_{jt}^{50} is an adequate measure of μ_{jt} (i.e., $F_t^{-1}(50) = 0$), then for any given year and percentile p of the latent wage distribution,

$$\begin{aligned}(5)\quad & \operatorname{cov}[(w_{jt}^{p*} - w_{jt}^{50}), (minwage_t - w_{jt}^{50})|t] \\ & = \operatorname{cov}[\sigma_{jt} \cdot F_t^{-1}(p), minwage_t - \mu_{jt}|t], = 0,\end{aligned}$$

where $minwage_t$ is the log of the federal minimum wage in year t. Under the above assumptions a positive association between the relative minimum and the *observed* 10–50 differential reflects the impact of the minimum on the lower tail of the wage distribution, and a significant association between the observed 75–50 differential and the relative minimum, for example, constitutes evidence that the identifying assumptions are violated.[21]

21. If $F_t^{-1}(50) \neq 0$, $\operatorname{cov}[(w_{jt}^{p*} - w_{jt}^{50}), (minwage - w_{jt}^{50})] = -F_t^{-1}(50)[F_t^{-1}(p) - F_t^{-1}(50)] \operatorname{var}[\sigma_{jt}]$. Thus, in this model a positive correlation between the upper tail differential (like the 90–50 difference) and the median-deflated minimum wage, for

When the federal minimum wage is absent in years 0 and 1, we can estimate the average change in the latent 10–50 differential, $[F_1^{-1}(10) - F_0^{-1}(10)]$, by the sample analog of $E[w_{j1}^{10} - w_{j1}^{50}] - E[w_{j0}^{10} - w_{j0}^{50}]$. But when the federal minimum is "binding," $[F_1^{-1}(10) - F_0^{-1}(10)]$ is estimated as a downward "shift" in line A_0A_1 to C_0C_1 in Figure V, for example; if we knew a priori that the minimum wage had a straightforward "censoring" effect (Case 1), we could estimate $[F_1^{-1}(10) - F_0^{-1}(10)]$ by the sample analog of $E[w_{j1}^{10} - w_{j1}^{50} | w_{j1}^{10} - w_{j1}^{50} < mw_{j1}] - E[w_{j0}^{10} - w_{j0}^{50} | w_{j0}^{10} - w_{j0}^{50} < mw_{j0}]$, where $mw_{jt} = (minwage_t - w_{jt}^{50})$.

Since the exact form of the minimum wage effect is unknown (it is perhaps a hybrid of Case 2 and 3), I fit the state-year data to the parameterization,

$$E[w_{jt}^{10} - w_{jt}^{50} | mw_{jt}, t] = \frac{mw_{jt} - \overline{\alpha}_t}{1 - e^{B(mw_{jt} - \overline{\alpha}_t)}} + \overline{\alpha}_t, \tag{6}$$

where $B < 0$ is a "curvature" parameter. This function asymptotes to $E[w_{jt}^{10} - w_{jt}^{50} | mw_{jt}, t] = mw_{jt}$ as mw_{jt} gets large. More importantly, it also asymptotes to $E[w_{jt}^{10} - w_{jt}^{50} | mw_{jt}, t] = \overline{\alpha}_t$, as mw_{jt} vanishes, which means that in this parameterization, the estimated $\overline{\alpha}_t$ is an estimate of the average latent 10–50 differential in year t, or $F_t^{-1}(10)$. Note that the above parameterization captures the basic features of the various functions depicted in Figure V.[22] In particular, as $B \to -\infty$, the function becomes the "censoring" case (Case 1).

An examination of Figure VI suggests that a reasonable, simple alternative to the parameterization in equation (6) is the estimation of the equation,

$$w_{jt}^{10} - w_{jt}^{50} = \overline{\alpha}_t + \overline{\beta} \cdot mw_{jt} + \overline{\gamma} \cdot mw_{jt}^2 + u_{jt}, \tag{7}$$

where the approximation error u_{jt} is assumed to be orthogonal to mw_{jt} and its square. This is simply the regression of the 10–50 log-wage differential on the relative minimum (and its square), and year dummies, using the panel of states throughout the 1980s. Again, changes in the $\overline{\alpha}_t$'s represent changes in nationwide latent wage inequality (as measured by the 10–50 log-wage differential).

example, necessarily implies a negative correlation between the latent lower tail differential (such as the 10–50) and the relative minimum. This equation shows that the most appropriate deflator of the minimum is the best measure of centrality (a percentile q such that $F_t^{-1}(q) = 0$).

22. I am grateful to an anonymous referee for suggesting this parameterization.

B. Results: 1979–1988

Panel A of Table I reports the least squares estimates of equations (7) and (6) using the panel data set of 50 states across ten years from 1979–1988 for the entire sample of earners (both men and women), as described in the Appendix. In the estimation, mw is defined as log(max[state minimum wage, federal minimum wage]) $- w^{50}$. Column (1) shows that the dependent variable $(w^{10} - w^{50})$ declines significantly throughout the decade. Including a linear term in mw, column (2) shows that the estimated slope, 0.509, is highly significant, implying that a 1 percent fall in the minimum wage leads to a 0.5 percent fall in the tenth percentile, relative to the median. The empirical fit of the regression increases substantially, the R^2 rising from 0.245 to 0.755.

Most importantly, the coefficients on the year dummies all rise toward zero. Whereas the average 10–50 differential is -0.107 in 1985 (relative to 1979), the average change after accounting for a linear term in mw is virtually zero for that same year. In fact, there is an increase of 4.7 log points in the relative position of the tenth percentile—a *decrease* in latent wage dispersion—between 1979 and 1988, after the inclusion of mw.

Column (3) includes a quadratic term, and demonstrates that the nonlinear aspect of the empirical relation is important.[23] In this specification, none of the year coefficients are negative, and the coefficient on 1988 is statistically different from zero, at *positive* 0.043. Column (4) demonstrates that the trend toward compression in the 10–50 differential is even more pronounced when equation (6) is estimated by Nonlinear Least Squares, although the standard errors of the year coefficients rise appreciably.

To alleviate, at least to some extent, the potential spurious correlation induced by sampling error (since the sample median is used to construct both the 10–50 differential and mw), I examine an alternative estimate of centrality, $\overline{w}^t$, which is the trimmed mean (where the bottom 30 percent and top 30 percent of the sample for each state-year are eliminated) as a deflator for the minimum wage. Column (5), which uses $\widetilde{mw}_{jt} = (minwage - \overline{w}^t_{jt})$, shows that the coefficients are quite similar to those in column (3). Although the coefficients in column (3) and column (5) are nearly identical for this sample and specification, for the remainder of the paper I utilize $\widetilde{mw}_{jt}$ because sampling error in the median may be more serious in other specifications (such as a within-state

23. Cubic and quartic terms have negligible effects on the results.

estimator, where the population variability in the locational shift of the distribution is diminished).[24]

One should note that in this sample period, several of the states legislated their own minimum wage rates to be higher than the federal rate. The analysis thus far also implicitly assumes that the nominal *state-legislated* minimum wage is not systematically related to latent wage dispersion across states. To examine how legislative endogeneity may be affecting the results, I repeated the estimation for the subsample of states for which the state minimum wage did not exceed the federal rate at any time during the 1979–1988 period. Although not reported here, the results are quite similar, suggesting that even if state-legislated minimum wages were endogenously determined for these excluded states, the effect is not large enough to noticeably influence the basic results of Panel A of Table I.[25]

Panel B of Table I reports results from using the specification of column (5) in Panel A for various subsamples, based on gender and age. For each subsample, the table reports (1) the ten-year annualized trend of the 10–50 differential (based on a regression of the year coefficients on a linear trend), (2) the estimated annualized trend when $\widetilde{mw}$ and its square are included, (3) the coefficients on the linear and quadratic terms of $\widetilde{mw}$, (4) the derivative implied by the coefficients evaluated at the overall mean of $\widetilde{mw}$ for the 1979–1988 period, and (5) the R^2.

As might be expected, the table shows that among workers with generally higher wages (men and older workers) the relative minimum has a more modest impact on the 10–50 differential. The implied derivative is as high as 0.597 for all women earners aged 16 and over, to as low as 0.182 for men aged 25–64. Overall, the table shows that for almost every sample, most of the trend in the 10–50 differential disappears when $\widetilde{mw}$ is included. For women the trend in latent wage dispersion cannot be statistically distinguished from zero, and for the entire sample the estimated trend in latent wage dispersion is small, but in the opposite direction of the observed trend. An important exception is the finding for men, aged 25–64, where much of the growth in

24. The root mean squared error of a regression on mw on $\widetilde{mw}$ is 0.011 (the R^2 is 0.996).

25. This table is reported in Lee [1998], which also includes an analysis where the procedure of DiNardo, Fortin, and Lemieux [1996] is used to adjust for differences in observable covariates across state-year observations. After adjusting for state differences in demographic, industrial, and occupational composition in this way, the year coefficients analogous to those in Table I, Panel A, appear to be virtually unaltered. See Lee [1998] for more details and discussion.

TABLE I, PANEL A
OLS/NLLS ESTIMATES: 10–50 DIFFERENTIAL ON MINIMUM WAGE, 1979–1988

Variable	(1)	(2)	(3)	(4)[a]	(5)
Year dummies					
1980	−0.001	0.004	0.006	0.027	0.004
	(0.007)	(0.004)	(0.003)	(0.026)	(0.004)
1981	−0.010	−0.002	0.001	0.004	0.002
	(0.007)	(0.005)	(0.004)	(0.027)	(0.005)
1982	−0.040	0.001	0.010	0.035	0.012
	(0.007)	(0.007)	(0.004)	(0.024)	(0.005)
1983	−0.065	−0.002	0.010	0.035	0.009
	(0.007)	(0.007)	(0.005)	(0.024)	(0.005)
1984	−0.090	−0.006	0.005	0.020	0.007
	(0.007)	(0.009)	(0.007)	(0.023)	(0.007)
1985	−0.107	−0.001	0.007	0.022	0.008
	(0.007)	(0.010)	(0.008)	(0.023)	(0.008)
1986	−0.111	0.014	0.018	0.038	0.019
	(0.009)	(0.012)	(0.009)	(0.023)	(0.010)
1987	−0.115	0.025	0.025	0.045	0.025
	(0.008)	(0.013)	(0.011)	(0.023)	(0.011)
1988	−0.105	0.047	0.043	0.071	0.043
	(0.009)	(0.016)	(0.012)	(0.023)	(0.013)
(Min.-median)	—	0.509	1.418	−9.395	—
		(0.051)	(0.152)	(0.547)	
(Min.-median)^2	—	—	0.589	—	—
			(0.114)		
(Min.-trim. mean)	—	—	—	—	1.494
					(0.155)
(Min.-trim. mean)^2	—	—	—	—	0.634
					(0.116)
Trend in year coef.[b]	−0.014	0.002	0.003	0.006	0.003
	(0.001)	(0.002)	(0.001)	(0.003)	(0.001)
R^2	0.245	0.755	0.806	—	0.791

N = 500. Data are constructed from the CPS Merged Outgoing Rotation Earnings Files. See Data Appendix for details. Regressions are the 10–50 differential on the relative minimum wage. In columns (2) and (3) relative minimum is max [fed. min., state min.] − median; in column (5) it is max [fed. min., state min.] − (trimmed mean), where the top 30 and bottom 30 percent of the sample are taking the mean. Weighted by observations per state-year. Standard errors (in parentheses) are heteroskedasticity-consistent and consistent with unrestricted autocorrelation within-state. Excluded year is 1979.

a. Estimated by Nonlinear Least Squares. Parameterization described in text.

b. From a regression of the estimated year coefficients on a linear trend.

inequality remains after the inclusion of $\widetilde{mw}$. For the rest of the analysis, so that results can be readily compared with previous research utilizing the CPS Outgoing Rotation Group data [DiNardo, Fortin, and Lemieux 1996], I focus on workers aged 18–64.

In principle, the panel nature of the state-level data set provides an alternative to cross-sectional identifying variation. *Within-state* variability in $\widetilde{mw}$ can also be exploited to identify the

TABLE I, PANEL B
OLS ESTIMATES: 10–50 DIFFERENTIAL ON MINIMUM WAGE, BY GENDER AND AGE GROUPS, 1979–1988

	Trend in 10–50 dif.	OLS Estimates: Trend in yr. coef.	(Minimum-trim. mean)	(Minimum-trim. mean)^2	Derivative at average	R^2
All						
16+	−0.014	0.003	1.494	0.634	0.515	0.791
	(0.001)	(0.001)	(0.155)	(0.116)	(0.038)	
18–64	−0.014	0.002	1.396	0.588	0.469	0.741
	(0.001)	(0.001)	(0.149)	(0.112)	(0.042)	
25–64	−0.008	0.003	0.999	0.391	0.309	0.564
	(0.001)	(0.001)	(0.196)	(0.124)	(0.036)	
Men						
16+	−0.013	0.000	1.250	0.450	0.407	0.583
	(0.001)	(0.002)	(0.219)	(0.134)	(0.049)	
18–64	−0.013	−0.002	1.036	0.353	0.364	0.488
	(0.001)	(0.002)	(0.279)	(0.165)	(0.058)	
25–64	−0.011	−0.005	0.197	0.007	0.182	0.221
	(0.001)	(0.003)	(0.348)	(0.188)	(0.086)	
Women						
16+	−0.027	−0.001	1.331	0.639	0.597	0.893
	(0.001)	(0.002)	(0.091)	(0.103)	(0.044)	
18–64	−0.025	−0.001	1.294	0.623	0.561	0.868
	(0.001)	(0.002)	(0.088)	(0.103)	(0.048)	
25–64	−0.020	−0.001	1.119	0.528	0.424	0.793
	(0.001)	(0.002)	(0.087)	(0.088)	(0.045)	

N = 500. Data are constructed from the CPS Merged Outgoing Rotation Earnings Files. See Data Appendix for details. Specification in column (5) in Panel A of Table I is used. Relative minimum wage is max [fed. min., state min.] − (trimmed mean) where the top 30 and bottom 30 percent of the sample is trimmed before taking the mean. Weighted by observations per state-year. Derivative is computed at the average value of the relative minimum wage variable for each regression. Trend is the average annual change in the year coefficients. Standard errors (in parentheses) are heteroskedasticity-consistent and consistent with unrestricted autocorrelation within-state.

impact of the minimum wage on the wage distribution. In particular, suppose that instead of $\overline{\alpha}_t$ in equation (7) we have

$$\overline{\alpha}_{jt} = \overline{\alpha}_j + \overline{\alpha}_t. \tag{8}$$

The validity of this "fixed effects" approach relies on the assumption that, *within-state,* wage levels are uncorrelated with latent wage dispersion after taking out year effects. Hence, a consistent sensitivity of latent wage inequality to regional economic booms or downturns is likely to contaminate the true impact of $\widetilde{mw}$ on $(w_j^p - w_j^{50})$. A disadvantage of using this variation is that most of the variability in $\widetilde{mw}$ (after taking out time effects) is generated by cross-state as opposed to within-state variability. The root mean

TABLE II
OLS ESTIMATES: UPPER TAIL DIFFERENTIALS ON MINIMUM WAGE, 1979–1988

	Pooled cross section			With state dummies		
Differential	All	Men	Women	All	Men	Women
60–50	−0.019	0.049	−0.033	0.008	0.029	−0.011
	(0.011)	(0.017)	(0.010)	(0.036)	(0.024)	(0.027)
70–50	0.022	0.124	−0.035	0.069	0.119	0.013
	(0.018)	(0.029)	(0.019)	(0.041)	(0.039)	(0.033)
80–50	0.087	0.207	−0.016	0.113	0.256	0.085
	(0.029)	(0.040)	(0.029)	(0.053)	(0.045)	(0.040)
90–50	0.162	0.299	0.017	0.280	0.438	0.131
	(0.043)	(0.048)	(0.036)	(0.061)	(0.065)	(0.059)

N = 500. Data are constructed from the CPS Merged Outgoing Rotation Earnings Files. See Data Appendix for details. Workers aged 18–64. Weighted by observations per state-year. Entries are the estimated coefficients on the relative minimum wage variable in regressions of each upper tail differential on max [fed. min., state min.] minus the (top 30 and bottom 30 percent) trimmed mean, year (and state) dummies. Standard errors (in parentheses) are heteroskedasticity-consistent and consistent with unrestricted autocorrelation within-state.

squared error from a regression of $\widetilde{mw}$ on year dummies for the full sample is 0.125; adding state dummies to this regression lowers it to 0.031, implying that only $(0.031/0.125)^2 = 0.062$ of the residual variance (after taking out time effects) in $\widetilde{mw}$ is due to "within-state" variability.[26] Thus, we might expect identification in a state fixed effects specification to be considerably weaker than in a pooled cross-section specification. In addition, the reduced identifying variation resulting from eliminating the "permanent" state effects may magnify biases due to misspecification, in the same way biases stemming from measurement error in the independent variable are magnified when true variation in the independent variable is reduced.

Perhaps the best available indication of which specification is more appropriate is given by the correlation between $\widetilde{mw}$ and measures of wage dispersion in the upper tail. A specification that yields a significant relation between $\widetilde{mw}$ and dispersion in the upper tail, where we are reasonably confident that the minimum has no effect, might be viewed with more suspicion when considering the empirical relation between $\widetilde{mw}$ and dispersion in the lower tail.

Table II reports the coefficient on $\widetilde{mw}$ in OLS regressions of the 60–, 70–, 80–, and 90–50 wage differentials on $\widetilde{mw}$ and a set of

26. A relatively small fraction of the variation in $\widetilde{mw}$ is due to variability in the nominal state-specific minimum wage. The addition of these nominal rates to the pooled cross-section and fixed effects specifications yield root mean squared errors of 0.121 and 0.032.

year dummies, with and without state dummies, for the three samples. Focusing on the 60–, 70–, and 80–50 differentials, two patterns emerge. First, it appears that the slope coefficients are considerably smaller for women or for the combined sample than for men alone. For women the pooled cross-section estimates for the upper percentiles range between −0.035 and −0.016 with standard errors of about 0.01 or 0.02. For the combined sample the pooled cross-section coefficients for the 60– and 70–50 differentials cannot be statistically rejected from zero, but the 80–50 estimates are statistically different from zero for both specifications. The comparable estimates for men are much larger in absolute magnitude; for example, the 70–50 pooled cross-section estimate is 0.124 and as high as 0.256 for the 80–50 fixed effects estimate.

Second, the absolute values of the coefficients tend to be larger in the state fixed effects specification. The 70–50 coefficient for the entire sample rises from 0.022 to 0.069, and the 80–50 coefficient rises from −0.016 to 0.085 for women, and from 0.207 to 0.256 for men.

Although I include results for the 90–50 differential for completeness, the coefficient for men and the entire sample should be interpreted with caution, because of the increasing importance of the weekly earnings topcode in the latter part of the 1980s. This is because a nontrivial and nominal topcode will naturally cause a positive association between the relative minimum and the 90–50 differential, since high-wage states will have their ninetieth percentiles artificially lowered to a greater extent than low-wage states.[27]

If the empirical relations between $\widetilde{mw}$ and the upper tail differentials are considered a valid specification check, Table II seems to suggest that when examining the empirical results for the lower tails, we can be most confident in the results for women in the pooled cross section, more cautious about the results for the entire sample, and considerably more suspicious for the results for men. In the same way, the state fixed effects might be viewed with more caution relative to the pooled cross-sectional results.

27. In order to assess the magnitude of this problem, I use the 1989 data (when the topcode changed to an irrelevantly higher level), and artificially topcode the weekly earnings level to where the pre-1989 level is (adjusted by the CPI), for each year. I compare regression coefficients of the 90–50 differential in the actual data for 1989, with that from utilizing the simulated data in a similar regression (that additionally includes year dummies). For the entire sample the coefficients are 0.038 and 0.101 for the actual and simulated data, respectively, suggesting that about 0.06 of the coefficient in the pooled cross section for the entire sample may be attributable to the topcode. For men the coefficients are 0.190 and 0.2623, and for women they are −0.003 and −0.002.

TABLE III
OLS ESTIMATES: LOWER TAIL DIFFERENTIALS ON MINIMUM WAGE, 1979–1988

	All			Men			Women		
Differential	Trend (dif.)	Trend (coef.)	Derivative	Trend (dif.)	Trend (coef.)	Derivative	Trend (dif.)	Trend (coef.)	Derivative
				Pooled cross section					
10–50	−0.014	0.002	0.469	−0.013	−0.002	0.364	−0.025	−0.001	0.561
	(0.001)	(0.001)	(0.042)	(0.001)	(0.002)	(0.058)	(0.001)	(0.002)	(0.048)
20–50	−0.006	0.003	0.262	−0.011	−0.005	0.184	−0.013	−0.002	0.268
	(0.001)	(0.001)	(0.036)	(0.001)	(0.002)	(0.059)	(0.001)	(0.001)	(0.031)
30–50	−0.004	0.000	0.134	−0.005	−0.004	0.058	−0.008	−0.002	0.134
	(0.001)	(0.001)	(0.025)	(0.001)	(0.001)	(0.037)	(0.000)	(0.001)	(0.019)
40–50	−0.002	0.000	0.057	−0.003	−0.003	0.004	−0.003	−0.001	0.056
	(0.001)	(0.001)	(0.013)	(0.000)	(0.001)	(0.017)	(0.000)	(0.001)	(0.012)
				With state dummies					
10–50	−0.015	−0.006	0.248	−0.014	−0.012	0.080	−0.026	−0.012	0.324
	(0.001)	(0.002)	(0.071)	(0.001)	(0.003)	(0.087)	(0.001)	(0.003)	(0.084)
20–50	−0.007	−0.003	0.107	−0.011	−0.011	0.028	−0.014	−0.006	0.190
	(0.001)	(0.002)	(0.066)	(0.001)	(0.002)	(0.067)	(0.001)	(0.003)	(0.066)
30–50	−0.005	−0.002	0.082	−0.006	−0.007	−0.037	−0.008	−0.005	0.064
	(0.001)	(0.001)	(0.033)	(0.001)	(0.001)	(0.038)	(0.000)	(0.002)	(0.040)
40–50	−0.002	0.000	0.044	−0.003	−0.004	−0.029	−0.003	−0.001	0.049
	(0.001)	(0.001)	(0.032)	(0.000)	(0.001)	(0.035)	(0.000)	(0.001)	(0.037)

N = 500. Data are constructed from the CPS Merged Outgoing Rotation Earnings Files. See Data Appendix for details. Workers aged 18–64. Weighted by observations per state-year. Specification of column (5) of Panel A, Table I is used. The columns for each sample show 1) the trend in the percentile differential, 2) the trend in the year coefficients when (a quadratic of) the relative minimum (max [fed. min, state min.] − trimmed (top 30 and bottom 30 percent) mean) is included, and 3) the estimated slope evaluated at the overall mean of the relative minimum for 1979–1988. Lower panel includes state dummy variables. Standard errors (in parentheses) are heteroskedasticity-consistent and consistent with unrestricted autocorrelation within-state.

Table III reports, for the three samples and two specifications, (1) the annualized trend in the average 10–, 20–, 30–, and 40–50 differentials, (2) the trend when $\widetilde{mw}$ and its square is included, and (3) the estimated derivative evaluated at the overall mean of $\widetilde{mw}$ throughout 1979–1988. For women in the pooled cross section the minimum wage appears to explain roughly 75 percent to virtually all of the growth in lower tail percentile differentials.[28] For the pooled cross-section specification using the entire sample, all of the expansion in the lower tail differentials can be accounted for by $\widetilde{mw}$.

28. The result for the 10–50 differential for women should be interpreted with the caveat that the average difference between the minimum and the tenth percentile is about two log points 1979 (see Data Appendix), meaning that identification of the growth in the latent 10–50 is more tenuous, as mentioned in subsection III.B.

In light of Table II the results for the state fixed effects specification, as well as for men, might be interpreted with considerable caution, but I include them in Table III for completeness. For the entire sample and for women, the fixed effects estimates imply that the minimum wage can account for about half of the rise in inequality. For men the results are mixed: for example, the pooled cross-section estimate implies that about 55 percent of the fall in the 20–50 differential is due to the declining minimum, while the corresponding fixed effects estimate suggests 0 percent.

V. Wage Compression and the Minimum Wage: 1989–1991

After a decade of constancy the nominal federal minimum wage rate rose, effective in April 1990, from $3.35 to $3.80, and again in April of 1991, from $3.80 to $4.25. Prior to this time, several of the states (see Appendix 3) had already legislated their own minimum wage rates that were higher than the forthcoming rise in the federal rate. As a result, the newly imposed federal minimum generated variability in *changes* in the effective minimum wage level across states.

The source of this variability stands in sharp contrast to that which generated variation in the effective minimum wage in the previous analysis. It affords an opportunity to estimate the impact of the minimum wage on the wage distribution under an alternative set of identifying assumptions. Instead of relying on the assumption that the measures of centrality of states' wage distributions (used to "deflate" the federal minimum rate) are uncorrelated with latent wage dispersion, the analysis below adopts the identifying assumption that the binding (relative) minimum rate (the maximum of the federal and the state-specific rate) is uncorrelated with *changes* in latent wage inequality across states.

Card and Krueger [1995] use this latter approach to examine the impact of the minimum wage on the wage distribution. We can compare the estimates using the approach of Card and Krueger with those utilizing the procedure of Section IV, for the same time period (1989–1991).[29] We can further compare these estimates with those using the data for the 1979–1988 period. Finding consistency among these various estimates would seem to lend some support to the findings of the previous section.

29. Card and Krueger [1995] utilize a different minimum wage measure: the fraction of workers in each state in 1989 who earn more than the 1989 rate and less than the 1991 rate.

TABLE IV
MEANS OF 10–50 DIFFERENTIAL AND MINIMUM WAGE, BY STATE MINIMUM WAGE STATUS, 1989–1991

	Minimum wage		10–50 differential	
	State > fed	State <= fed	State > fed	State <= fed
All				
1989 (Level)	−0.959	−0.978	−0.662	−0.683
	(0.031)	(0.030)	(0.040)	(0.013)
1990–1989 (Change)	−0.009	0.085	−0.015	0.019
	(0.007)	(0.004)	(0.015)	(0.006)
1991–1989 (Change)	0.016	0.167	−0.015	0.032
	(0.026)	(0.004)	(0.016)	(0.010)
Men				
1989 (Level)	−1.088	−1.118	−0.724	−0.723
	(0.030)	(0.032)	(0.064)	(0.018)
1990–1989 (Change)	0.001	0.089	0.008	0.012
	(0.008)	(0.004)	(0.014)	(0.009)
1991–1989 (Change)	0.035	0.176	0.006	0.016
	(0.021)	(0.004)	(0.019)	(0.009)
Women				
1989 (Level)	−0.802	−0.813	−0.557	−0.583
	(0.033)	(0.028)	(0.028)	(0.009)
1990–1989 (Change)	−0.024	0.081	−0.018	0.006
	(0.006)	(0.005)	(0.014)	(0.010)
1991–1989 (Change)	−0.013	0.154	−0.021	0.033
	(0.029)	(0.005)	(0.020)	(0.010)
N	12	38	12	38

Data are constructed from the CPS Merged Outgoing Rotation Earnings Files. See Data Appendix for details. Workers aged 18–64. The relative minimum here is the max [fed. min, state min.] minus the (top 30 and bottom 30 percent) trimmed mean. "State > fed" means that the state had a minimum wage that was higher than the federal minimum in 1989. "State <= fed" means that the state had a minimum wage that was equal to or less than (or had no state minimum wage provision) the federal minimum in 1989. Standard errors (in parentheses) are heteroskedasticity-consistent and consistent with unrestricted autocorrelation within-state.

Table IV gives a sense of the magnitude of the average changes in both the minimum wage and the 10–50 differential for two groups of states: the states that in 1989 had minimum wages greater than the federal minimum, and all other states.[30] For men, women, and the entire sample, the group with state minimums equal to or less than the federal rate in 1989 experience an average increase in $\widetilde{mw}$ of about 8 or 9 percent by 1990, and an additional 7 percent in 1991. By contrast, the group of states

30. The states with higher minimums were California, Connecticut, Hawaii, Maine, Minnesota, Massachusetts, New Hampshire, Pennsylvania, Rhode Island, Vermont, Washington, and Wisconsin. Alaska was excluded because its state minimum is fixed, relative to the federal rate.

beginning with higher state-specific minimums experience a very small change in $\widetilde{mw}$.

Table IV illustrates that the two groups' trends in the 10–50 log-wage differential also diverge for women, as well as for the entire sample. For these samples, the group of states that experienced the 16 percent average growth in $\widetilde{mw}$ witnessed an average compression in their 10–50 differentials of about 3 percent, whereas the remaining states experienced a 1 or 2 percent expansion in the 10–50 gap. The divergent pattern is not found in the results for men, for whom the difference between the initial minimum and the tenth percentile in 1989 was greatest, at about 38 log points.

Utilizing legislation-induced changes in the minimum wage is formally implemented by estimating the state fixed effects specification that was used in the lower panel of Table III (using equations (7) and (8)). Within-state variation in $\widetilde{mw}$ during 1989–1991 is much smaller than the total cross-sectional variability, as was the case for the 1979–1988 data. For the combined sample in the 1989–1991 period, a regression of $\widetilde{mw}$ on time dummies yields a root mean squared error of 0.137. Adding state dummies to this regression reduces it to 0.040, implying that $(0.040^2/0.137^2) = 0.085$ of the residual variability (after taking out time effects) in $\widetilde{mw}$ is within-state, and that identification will be relatively weaker than in the pooled cross-section specification. However, virtually all of this within-state variation in $\widetilde{mw}$ is generated by changes in the legislated minimum rate, as opposed to transitory, within-state changes in the centrality measures of the states' wage distributions. Adding the nominal legislated rate to the above regression brings the root mean squared error from 0.040 to 0.013, implying that $((0.040)^2 - 0.013^2)/(0.040^2) = 0.894$ of the "within-state" variance in $\widetilde{mw}$ is indeed driven by legislation-induced changes for the 1989–1991 period; as noted in Section IV, the comparable figure for the 1979–1988 data is close to zero.

Table V presents results for the 1989–1991 period using the 10–50 differential as the dependent variable. For both women and the combined sample, the coefficients on both the linear and quadratic terms in the fixed effects specification are quite similar to the corresponding estimates in the pooled cross section. The implied derivatives (and standard errors) evaluated at the 1989–1991 mean of $\widetilde{mw}$, are 0.262 (0.096) and 0.349 (0.072) in the state fixed effects specification, for the combined sample and women, respectively. The corresponding derivatives are 0.320 (0.073) and

TABLE V
OLS ESTIMATES: 10–50 DIFFERENTIAL ON MINIMUM WAGE, 1989–1991

	Pooled cross section		With state dummies	
All				
Year dummies				
1990	0.010	−0.007	0.013	−0.002
	(0.007)	(0.007)	(0.008)	(0.010)
1991	0.020	−0.023	0.023	−0.013
	(0.010)	(0.013)	(0.011)	(0.015)
(Min.-trim. mean)	—	1.598	—	1.401
		(0.527)		(0.341)
(Min-trim. mean)^2	—	0.702	—	0.625
		(0.283)		(0.191)
Men				
Year dummies				
1990	0.011	−0.003	0.015	0.014
	(0.007)	(0.011)	(0.010)	(0.014)
1991	0.013	−0.024	0.017	0.012
	(0.008)	(0.013)	(0.009)	(0.020)
(Min.-trim. mean)	—	1.545	—	0.713
		(0.767)		(0.434)
(Min.-trim. mean)^2	—	0.614	—	0.326
		(0.366)		(0.195)
Women				
Year dummies				
1990	0.000	−0.016	0.002	−0.017
	(0.009)	(0.009)	(0.009)	(0.009)
1991	0.019	−0.019	0.020	−0.023
	(0.011)	(0.010)	(0.012)	(0.011)
(Min.-trim. mean)	—	1.532	—	1.472
		(0.299)		(0.297)
(Min.-trim. mean)^2	—	0.811	—	0.744
		(0.205)		(0.188)

N = 150. Data are constructed from the CPS Merged Outgoing Rotation Earnings Files. See Data Appendix for details. Workers aged 18–64. Relative minimum variable is the max [fed. min., state min.] minus the (top 30 and bottom 30 percent) trimmed mean wage. Weighted by observations per state-year. Standard errors (in parentheses) are heteroskedasticity-consistent and consistent with unrestricted autocorrelation within-state.

0.308 (0.060) for the pooled cross-section specification. Furthermore, evaluated at the 1979–1988 mean of $\widetilde{mw}$, the pooled cross-section (and fixed effects) estimates of the derivative are 0.482 (0.415) and 0.578 (0.597), for the combined sample and women, respectively. These are quite comparable to the estimated derivatives from the pooled cross-section estimates reported in Table III (0.469 and 0.561), which used the 1979–1988 data.

By contrast, the two specifications' coefficients on $\widetilde{mw}$ and its square are in less agreement for the sample of men. The implied derivatives (and standard errors) are 0.033 (0.0120) and 0.270 (0.068) for the fixed effects and pooled cross-section specifications, respectively. This finding provides some reason to lessen our confidence in the results for men. Examination of the empirical relation between the upper tail percentile differentials and $\widetilde{mw}$ provides some further reason to place more confidence in the results for the combined sample, and women, relative to the results for men. For example, a regression of the 80–50 differential on $\widetilde{mw}$ leads to a coefficient (standard error) of 0.094 (0.035), −0.003 (0.027), and 0.040 (0.026) for men, women, and the combined sample, respectively. A similar qualitative pattern exists for the various other measures of dispersion in the upper tail.[31]

Finally, note that a comparison of the year coefficients in regressions, with and without $\widetilde{mw}$, reveals that for women, and the combined sample, the modest compression in the observed 10–50 differential of about two log points across states belies what appears to be an *expansion* of the *latent* 10–50 gap of about one to two log points, during the 1989–1991 period.

VI. Implications for Between- and Within-Group Wage Inequality

The analysis to this point has focused on the extent to which the minimum wage has contributed to growth in broad measures of wage inequality—the various percentile differentials. Since much of the empirical literature has focused on changes in between-group wage differentials (based on gender, education, age), it is instructive to consider the potential impact of the minimum wage on these wage differentials, as implied by the findings in the previous sections. Equally informative is the implied effect of the minimum wage on changes in so-called "within-group" inequality, which constitutes a large portion of the total growth in wage dispersion in recent decades [Juhn, Murphy, and Pierce 1993].

A simple way to produce these calculations is given by

31. Using the 90–50 differential yields 0.124, 0.033, and 0.048 for men, women, and the combined sample, respectively; these are the largest (in absolute value) among regressions using the various upper tail percentile differentials (i.e., 60–50, 70–50, 80–50, and 90–50). The largest (in absolute value) for analogous fixed effects estimates using the upper tail was −0.108 (with a standard error of 0.095) for the 90–50 differential for the combined sample; all other estimates (using any of the samples, or upper tail measures) were smaller in absolute value and could not be statistically rejected from zero.

DiNardo, Fortin, and Lemieux [1996], who simulate a counterfactual wage density for 1988, adjusting the minimum wage to its 1979 level, in real terms. I follow this general approach; but instead of making specific assumptions about how the minimum wage affects the lower tail of the distribution, I simply employ the estimates of the previous section on the empirical relation between the minimum and the lower tail percentile differentials across states.

I do so in the following way. First, I determine for each observation in the microdata (that was used to create the state-level panel data set), its percentile in the overall wage distribution, on a state-by-state basis. Then according to this relative position, I add or subtract an amount to the log-wage of the worker according to estimates of $\bar{\beta}$ and $\bar{\gamma}$ (from equation (7)) and the simulated change in $\widetilde{mw}$. So, for example, in order to simulate what individuals at the pth percentile of their state's wage distribution in 1989 would earn in the face of a minimum wage at its 1979 relative level, I add the amount

$$(9)\qquad \Delta^p_{j,89} = \hat{\bar{\beta}}^p \cdot (\widetilde{mw}_{j,79} - \widetilde{mw}_{j,89}) + \hat{\bar{\gamma}}_p \cdot (\widetilde{mw}^2_{j,79} - \widetilde{mw}^2_{j,89})$$

to the worker's actual log-wage in 1989, where p denotes the within-state percentile, j the state, $\hat{\bar{\beta}}^p$ and $\hat{\bar{\gamma}}^p$ the estimated coefficients from the regression described by column (5) of Table I, Panel A, $\widetilde{mw}_{j,89}$ the actual $\widetilde{mw}$ for state j in 1989, and $\widetilde{mw}_{j,79} = \widetilde{mw}_{j,89} - \log(3.35) + \log(2.90) + 0.50$ denotes the hypothetical (1979) relative level of the federal minimum wage for state j.[32]

Although the separate estimates $\hat{\bar{\beta}}^p$ and $\hat{\bar{\gamma}}^p$ for men, women, and the combined sample can be applied to each respective sample, the accumulation of evidence to this point suggests that we might place the most confidence in the estimates from the pooled cross-section specification for women, and for the combined sample. And in particular, it suggests that the results for men be viewed with considerable suspicion. Thus, in the simulation exercise I adjust the wages of the entire sample of earners, using the pooled cross-sectional estimates of $\hat{\bar{\beta}}^p$ and $\hat{\bar{\gamma}}^p$ from the entire sample, but report the various between- and within-group wage dispersion measures by men and women separately.[33] It is impor-

32. The nominal minimum rate was \$2.90 in 1979, \$3.35 in 1989, and the change in the national median log(wage) for all workers was 0.50.

33. I use estimates from regressing each of the 49 lower tail differentials (1–50 to 49–50) on a quadratic in $\widetilde{mw}$, using data from the entire time period, 1979–1991, weighting by the number of observations in each state-year cell.

tant to note that the validity of the resultant "counterfactual" wage differentials rests upon the assumption that the minimum wage affects worker types (irrespective of their education, gender, age, etc.) equally, *conditional on the worker's wage level.*

Panel A of Table VI reports various comparisons between actual wage differentials (by race, education, and experience), and corresponding differentials from the adjusted microdata for male earners aged 18–64. The mean wages in each group are expressed as relative to the combined-sample median wage. The first two columns present the differentials computed from the actual data, the third reports the counterfactual of applying the 1979 relative level of the minimum wage to the 1989 distribution, and the last two columns show the counterfactual differentials resulting from simulating the impact of the 1989 relative minimum on the 1979 and 1989 distributions.[34]

An examination of the first and third columns of the table reveals that the main beneficiary of this hypothetical minimum wage increase are less-experienced high school dropouts, as their average log-wage rises about eight log points due to the adjustment of the 1989 data. The lower part of the table, which reports unadjusted and adjusted changes in the differentials, shows that *changes* in race, experience, and educational differentials are only moderately attenuated by an adjustment for the minimum wage. It shows that about 14 percent of the increase in the college-high school wage differential for those with 1–10 years of experience is due to the falling minimum wage throughout this period. One notes that this is comparable to the 13 percent calculation provided by DiNardo, Fortin, and Lemieux [1996]. The alternative adjusted changes in the fourth and fifth columns give very similar results.

Panel A of Table VI also reports the estimated slope coefficient on years of completed schooling in a wage equation that additionally includes a quartic in potential experience. The adjustment in the third column implies that about $(0.004)/(0.088 - 0.061) = 0.148$ of the change in the return to schooling is attributable to the minimum wage.

The adjustment has a much greater impact on the standard deviation of the residuals from the wage regression, a measure of "within-group" wage inequality. RMSE is the root mean squared error of a wage regression that includes a complete interaction of

34. The simulation for 1989 makes an adjustment only for workers in states with minimum wages higher than the federal rate.

TABLE VI, PANEL A
WAGE DIFFERENTIALS, ACTUAL AND ADJUSTED FOR MINIMUM WAGE: MEN, 1979–1989

	Actual		Adjusted		
Year	1979	1989	(1979 min.) 1989	(1989 min.) 1979	(1989 min.) 1989
All men	0.188	0.127	0.151	0.166	0.125
Standard deviation	0.459	0.534	0.501	0.494	0.538
Differentials					
10–50	−0.657	−0.781	−0.654	−0.739	−0.781
25–50	−0.324	−0.395	−0.375	−0.341	−0.398
75–50	0.288	0.351	0.351	0.288	0.351
90–50	0.561	0.670	0.670	0.561	0.670
White	0.207	0.149	0.171	0.186	0.146
Black	−0.015	−0.095	−0.055	−0.051	−0.095
1–10 yrs. exp.	0.052	−0.072	−0.035	0.021	−0.075
HSDO	−0.226	−0.450	−0.374	−0.292	−0.459
HS	−0.010	−0.219	−0.174	−0.043	−0.223
COL	0.269	0.291	0.305	0.254	0.290
21–30 yrs. exp.	0.329	0.299	0.313	0.315	0.298
HSDO	0.067	−0.141	−0.105	0.041	−0.146
HS	0.303	0.191	0.205	0.291	0.191
COL	0.642	0.673	0.678	0.637	0.673
Return to education	0.061	0.088	0.084	0.064	0.089
RMSE	0.393	0.432	0.405	0.422	0.434
Residuals					
90–10	0.995	1.092	1.021	1.060	1.099
50–10	0.517	0.572	0.508	0.578	0.579
90–50	0.478	0.520	0.513	0.482	0.520
	Change from 1979 actual			Change from 1979 (at 1989 minimum)	
Black–white	—	−0.022	−0.004	—	−0.004
(21–30 yrs)–(1–10 yrs.)	—	0.095	0.071	—	0.080
COL-HS, (1–10 yrs.)	—	0.232	0.200	—	0.216
COL-HSDO, (1–10 yrs.)	—	0.246	0.184	—	0.203
COL-HS, (21–30 yrs.)	—	0.144	0.135	—	0.136

Computed from the CPS Merged outgoing Rotation Earnings Files microdata. Data from 1989 are from 4/89–3/90. Ages 18–64. For variable construction see Data Appendix. All average wage levels are expressed as relative to the overall median wage. The first two columns use the actual data for 1979 and 1989. The third column is 1989 data adjusted to the 1979 relative minimum, and fourth and fifth columns are the 1979 and 1989 differentials adjusted to the 1989 relative minimum (for all states). The adjustment procedure is described in the text. Return to Education is the coefficient on education in a regression of log (wage) on education, a quartic in potential experience (age-educ-6) and a black dummy. RMSE is the root mean squared error from an hours-weighted regression of log (wage) on a set of fully interacted educational categories (4) and single year experience dummies. 90–10, 50–10, and 90–50 are various percentile differentials of the residuals in that regression.

TABLE VI, PANEL B
WAGE DIFFERENTIALS, ACTUAL AND ADJUSTED FOR MINIMUM WAGE: WOMEN, 1979–1989

	Actual		Adjusted		
Year	1979	1989	(1979 min.) 1989	(1989 min.) 1979	(1989 min.) 1989
All women	−0.177	−0.136	−0.093	−0.235	−0.140
Standard deviation	0.377	0.478	0.430	0.433	0.481
Differentials					
10–50	−0.389	−0.575	−0.447	−0.543	−0.588
25–50	−0.246	−0.323	−0.274	−0.313	−0.329
75–50	0.305	0.363	0.346	0.329	0.363
90–50	0.560	0.669	0.652	0.589	0.669
White	−0.170	−0.125	−0.083	−0.227	−0.129
Black	−0.227	−0.222	−0.165	−0.296	−0.223
1–10 yrs. exp.	−0.197	−0.199	−0.147	−0.256	−0.202
HSDO	−0.443	−0.636	−0.519	−0.556	−0.647
HS	−0.294	−0.404	−0.332	−0.367	−0.409
COL	0.053	0.150	0.165	0.028	0.148
21–30 yrs. exp.	−0.132	−0.070	−0.035	−0.183	−0.073
HSDO	−0.330	−0.451	−0.375	−0.415	−0.460
HS	−0.170	−0.186	−0.142	−0.222	−0.189
COL	0.201	0.295	0.306	0.183	0.294
Return to education	0.066	0.097	0.088	0.076	0.098
RMSE	0.333	0.399	0.359	0.383	0.401
Residuals					
90–10	0.836	1.008	0.895	0.976	1.012
50–10	0.359	0.485	0.395	0.464	0.489
90–50	0.478	0.522	0.500	0.512	0.523
	Change from 1979 actual			Change from 1979 (at 1989 minimum)	
Black–white	—	−0.041	−0.025	—	−0.026
(21–30 yrs)–(1–10 yrs.)	—	0.063	0.048	—	0.056
COL-HS, (1–10 yrs.)	—	0.207	0.150	—	0.162
COL-HSDO, (1–10 yrs.)	—	0.290	0.188	—	0.211
COL-HS, (21–30 yrs.)	—	0.110	0.077	—	0.079

Computed from the CPS Merged outgoing Rotation Earnings Files microdata. Data from 1989 are from 4/89–3/90. Ages 18–64. For variable construction see Data Appendix. All average wage levels are expressed as relative to the overall median wage. The first two columns use the actual data for 1979 and 1989. The third column is 1989 data adjusted to the 1979 relative minimum, and fourth and fifth columns are the 1979 and 1989 differentials adjusted to the 1989 relative minimum (for all states). The adjustment procedure is described in the text. Return to education is the coefficient on education in a regression of log (wage) on education, a quartic in potential experience (age-educ-6) and a black dummy. RMSE is the root mean squared error from an hours-weighted regression of log (wage) on a set of fully interacted educational categories (four) and single year experience dummies. 90–10, 50–10, and 90–50 are various percentile differentials of the residuals in that regression.

the four educational groups (<12, 12, 13–15, 16+) and dummies for each single year of experience. Rather than rising from 0.393 to 0.432 (9.9 percent rise) from 1979 to 1989, it rises to 0.405 (3.1 percent), accounting for the minimum wage. The following three lines, "90–10," "50–10," and "90–50," report the corresponding percentile differentials of these residuals. It reveals that the adjustment makes a difference mostly in the lower tail of the distribution of residuals.

Panel B of Table VI reports the results for women. A comparison of the first and fourth columns suggest that less-experienced, high school dropout females would have been paid, on average, about 11 percent less in 1979 if it had not been for the supporting effect of the minimum wage. Although the minimum wage appears to be a proportionately greater component of changes in differentials by education, relative to men, two-thirds of the observed changes remain after the adjustment. The adjusted college-high school wage differential continues to rise significantly, by about fifteen log points for the least experienced, and the college-dropout differential rises about 19 percent.

A comparison of the unadjusted and adjusted measures of the residual dispersion suggests that between two-thirds and three-quarters of the approximately 20 percent rise in within-group inequality could be attributable to the declining minimum wage during the 1980s. This estimate, and that for men, are substantially higher than the lower bound estimates given by DiNardo, Fortin, and Lemieux [1996], who find that the minimum wage accounts for 24.2 and 34.0 percent of the rise in residual dispersion, for men and women, respectively.[35]

I note that while the estimates here suggest that much of the growth in inequality in the lower tail can be attributed to the minimum wage, substantial growth in dispersion in the upper tail remains during the 1980s. As shown in Panels A and B of Table VI, the 90–50 log-wage differential expands by about ten or eleven log points for both men and women. For men the minimum wage appears to account for about 70 (or more) percent of the growth of the 50–10 differential, but about 75 percent of the widening 50–25 gap still remains, implying that an analysis

35. DiNardo, Fortin, and Lemieux [1996] adopt three assumptions to produce their lower bound estimates: (1) no disemployment effect, (2) no spillovers, and (3) the shape of the density below the minimum remains the same (even if the fraction below changes). They note that these assumptions work to understate the impact of the minimum. The estimates in their study as well as the present analysis depart somewhat from the analysis of Bernard and Jensen [1998], who find virtually no role for the minimum wage in explaining differences in residual wage inequality, using cross state differences from Census data.

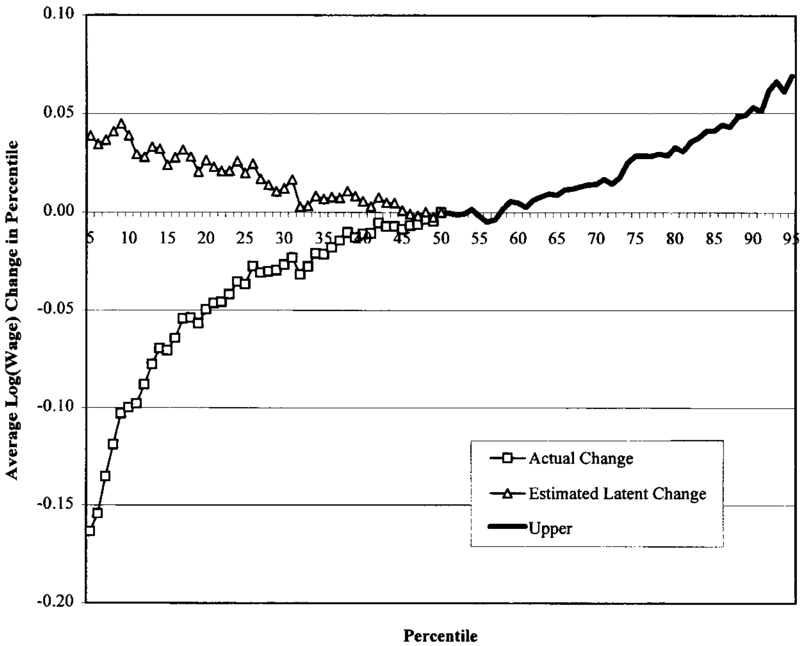

FIGURE VII
Summary of Changes in Actual and Estimated Latent Log (Wage) Differentials, 1979–1989, by Percentile: All Workers, 18–64

which ignores the minimum wage would not lead to a serious overstatement of increasing inequality, *if the focus is on the upper 75 percent of the distribution.* It should also be noted, however, that the same cannot be said for women. As the upper part of Panel B reveals, the minimum appears to have had such an important impact that about 50 to 60 percent of the thirteen log point rise in the 75–25 differential is explained by the minimum.

Curiously, accounting for the minimum wage has a considerably more striking impact on the impression of overall inequality for the unconditional distribution of wages (men and women combined). As a graphical summary of the core specification provided in Table I, Figure VII plots the actual average change (1979 to 1989) for each (from the 5–50 to the 95–50) percentile differential; these are essentially the "1989" coefficients in column (1) of Table I for each differential.[36] I also plot the 1989 year

36. Here, I append the data from 1989–1991 to the data (1979–1988) used for Table I.

coefficients from the core specification (column (5), Table I) which includes $\widetilde{mw}$ for each percentile up to the median. It shows that, after accounting for the minimum wage, the modest increase in dispersion in the upper tail of the overall wage distribution appears to be offset by an equally modest *compression* in the lower tail, so as to keep measures of the latent 90–10 or 75–25 differential relatively constant over time. This finding is roughly consistent with Teulings [1998], who in a recent alternative analysis of the impact of the minimum wage, uses a flexible parameterization of the wage distribution across time and regions within the United States, and concludes that the minimum wage can fully explain the growth in inequality of the unconditional distribution of wages during the 1980s.

VII. Summary and Conclusions

Two principal findings emerge from the above empirical analysis. First, estimates that exploit cross-state variation in the "effective" minimum wage imply that a great majority of the observed growth in inequality in the lower tail of the distribution is attributable to the erosion of the real value of the federal minimum wage rate during the 1980s. The estimates imply that the falling relative level of the minimum wage can explain from 70 to 100 percent of the growth in inequality in the lower tail of the female wage distribution, and for men, about 70 percent and 25 percent of the growth in the 50–10 and 50–25 differentials, respectively. For both men and women, analyzed separately, much growth in inequality remains (mostly in the upper tail of the distribution) during this period. Curiously, however, after accounting for the minimum wage, the unconditional distribution (men and women combined) exhibits a *compression* in the lower tail (50–10 differential) that is of the same magnitude as the modest expansion in the upper tail (90–50 differential) of the distribution.

Second, the magnitudes of this study's estimates imply that ignoring the declining real value of the minimum wage leads to only a moderate exaggeration of the growth in between-group wage inequality; the broad trends in educational and experience differentials are unaffected. But ignoring the minimum wage leads to a substantial overstatement of the growth in residual, "within-group" wage inequality during the 1980s; the minimum wage may explain between 60 to 80 percent of the rise in this component of wage inequality.

The credibility of the estimates provided here depend on the validity of the assumption that latent wage inequality is uncorrelated with measures of the centrality of states' log-wage distributions (and hence with the relative minimums). I find evidence that supports this notion for women, and to a lesser extent, for the entire sample of workers. Also, I find evidence that casts doubt on this notion for the sample of male workers.

For example, for the sample of women there are negligible correlations between measures of wage dispersion in the upper tail and the relative minimum. This is at odds with the notion that an inherent relation between the overall scale and location of the states' distributions is causing an upward bias of the estimate of the minimum wage's impact across states. An analogous analysis provides a weaker case for the validity of the cross-sectional results for the combined sample. The weakest case is for the analysis of the male sample, where the correlations between the relative minimum and the several measures of upper tail wage dispersion are all large and statistically significant.

Second, I provide estimates generated by a qualitatively different source of variability in changes in the effective minimum—that which is driven by an interaction between preexisting variability in state-legislated wage floors and the imposition of a higher, binding federal minimum in the early 1990s. For women as well as the combined sample, the resulting estimates are quite similar to those resulting from the earlier cross-sectional analysis. By contrast, the divergence of the two estimates when analyzing men in isolation provides reason to suspect the main results for the male sample.

The findings in this paper point have a few implications for future research on wage and earnings inequality. First, suppose that we take the findings here at face value—that a minimum wage that kept pace with inflation would have resulted in a relatively unchanging unconditional distribution of wages *for hours worked in the economy* during the 1980s. Given the suggestion of the previous literature that the minimum wage played a small role in the substantial rise in household income inequality, explaining this rise would seem necessarily to involve a *change* in the nature of the *mapping* between hourly wages and household income.[37]

This intermediate link includes how the distributions of wages and hours worked map into weekly earnings, and in turn,

37. See Brown's [1999] recent survey of the literature on the minimum wage for a discussion of these issues.

how weekly earnings and the distribution of weeks worked in the year map into annual earnings. It also includes issues of how individual earners come together to form households. And of course, the magnitude of disemployment effects due to the minimum wage—especially at a time such as the early 1980s, when the minimum wage was relatively high—is an integral part of this intermediate link. Since this study has focused exclusively on the *observed* wage distribution without accounting for these issues, its findings, taken alone, cannot be used to adequately assess the impact of the minimum wage on the distribution of economic welfare more generally.

Appendix 1: Data Appendix

State Wage Data

All computations (except for the pre-1979 data shown in Figure II) were derived from the National Bureau of Economic Research Extract of the Current Population Survey Merged Outgoing Rotation Group Earnings Files. Unless otherwise noted, the sample includes all individuals aged sixteen or over. The sample excludes the self-employed, unemployed, and those not in the labor force. In order to create a wage measure consistent over the entire time period, I construct the wage from "unedited" variables of the hourly rate of pay, usual weekly earnings, and usual weekly hours worked, since the imputation procedure for those with missing earnings variables seems to have changed significantly between 1988 and 1989. In order to have a nonmissing wage, an individual must report either (1) the hourly rate of pay or (2) usual weekly earnings *and* usual weekly hours worked. In cases where an individual had both values, the higher of the two was used. Note that since I use "unedited" wage variables, all individuals with "imputed" wages are excluded from the analysis. Note also that no exclusions of observations were made due to implausible or "extreme" wage values. A comparable sample selection and variable construction procedure was used for the May CPS data for Figure II. Resulting sizes per year (for workers aged 18–64, for example) for the Outgoing Rotation Groups ranged from about 130,000 in 1979 to 140,000 in 1989.

Using these microdata, a panel of the 50 states' wage percentiles was calculated for the 1979–1991 period. Three main panels were constructed: that for men only, women only, and that for the combined sample. In addition, separate panels were constructed for the various age-groups used in the paper: 16 and over, 18–64, and 25–64. Wage percentiles were calculated using the sample weight multiplied by usual hours worked. In order to make the wage data correspond to the timing of minimum wage legislation, data for "1989" correspond to data from 4/89 to 3/90. Similarly, "1990" and "1991" contains wage data from 4/90–3/91 and 4/91–3/92, respectively. Other state-year observations use data from January to December.

New state-specific minimum wage laws often came into effect midyear, and hence the following exceptions were made. Wisconsin in 1989 actually represents data from 7/89 to 3/90. Vermont: 1986, 7/85–6/86; 1987, 7/86–6/87; 1988, 7/87–6/88; 1989, 7/89–3/90. Massachusetts: 1986, 7/85–6/86; 1987, 7/86–6/87; 1988, 7/87–6/88. Rhode Island: 1986, 7/85–6/86; 1987 7/86–6/87; 1988, 7/87–6/88; 1989 8/89–3/90. Connecticut: 1987, 10/86–9/87; 1988 10/87–9/88. California: 1988, 7/87–6/88. Oregon: 1989, 1/89–8/89.

Appendix 2 gives a summary of the (sample-size-weighted) means of selected wage percentile differentials, and the log(max-[state minimum wage rate, federal rate]) − median wage for state panel data. The average sample size for the state-year cell is about 2800, with a median of about 2000.

Minimum Wage Legislation by State

Information on state-specific minimum wage legislation was compiled from the January issues of the *Monthly Labor Review* [Bureau of Labor Statistics], which contain an annual summary of the previous year's proposed and enacted state labor legislation. Details on exemption rules, and coverage criteria varied across states, but it was possible to isolate a "basic adult rate" for each state. The District of Columbia was excluded from the sample of states because frequent changes in sector-specific rates, and coverage rules made it difficult to isolate a consistently broad adult rate. For states that did not enact any minimum wage legislation throughout the 1979–1993 period, a summary of the minimum wage laws across states was available from the *Daily Labor Report* [Bureau of National Affairs 1990], which was also used to corroborate the information for the rest of the states.

The states' minimum wage legislation can be broadly categorized into four groups: (1) those with no minimum wage provision, (2) those with legislated minimum rates below, (3) equal, and (4) higher than the basic federal rate. Appendix 3 provides a graphical summary of the evolution of minimum wage laws in the 50 states during the 1979–1993 period.

APPENDIX 2: RELATIVE MINIMUM WAGE, LOWER PERCENTILE DIFFERENTIALS: MEANS FROM STATE-LEVEL DATA, 1979–1991

	All				Men				Women			
	Min–50	Min–TM	10–50	20–50	Min–50	Min–TM	10–50	20–50	Min–50	Min–TM	10–50	20–50
1979	−0.651	−0.653	−0.578	−0.418	−0.836	−0.832	−0.637	−0.415	−0.405	−0.410	−0.384	−0.295
1980	−0.662	−0.659	−0.584	−0.422	−0.847	−0.841	−0.654	−0.430	−0.423	−0.426	−0.398	−0.296
1981	−0.667	−0.668	−0.592	−0.427	−0.854	−0.847	−0.680	−0.450	−0.441	−0.443	−0.415	−0.314
1982	−0.729	−0.729	−0.625	−0.440	−0.915	−0.904	−0.713	−0.475	−0.512	−0.517	−0.465	−0.337
1983	−0.767	−0.767	−0.644	−0.447	−0.942	−0.934	−0.722	−0.484	−0.565	−0.568	−0.500	−0.354
1984	−0.813	−0.812	−0.664	−0.460	−0.982	−0.975	−0.723	−0.478	−0.612	−0.615	−0.528	−0.370
1985	−0.855	−0.854	−0.675	−0.466	−1.020	−1.010	−0.737	−0.487	−0.660	−0.661	−0.556	−0.385
1986	−0.894	−0.890	−0.681	−0.469	−1.050	−1.045	−0.731	−0.491	−0.703	−0.705	−0.569	−0.395
1987	−0.919	−0.919	−0.688	−0.464	−1.070	−1.066	−0.719	−0.488	−0.740	−0.742	−0.579	−0.401
1988	−0.943	−0.942	−0.667	−0.453	−1.092	−1.087	−0.717	−0.480	−0.771	−0.770	−0.578	−0.399
1989	−0.976	−0.973	−0.678	−0.468	−1.120	−1.111	−0.724	−0.489	−0.810	−0.810	−0.577	−0.400
1990	−0.916	−0.913	−0.668	−0.466	−1.049	−1.044	−0.713	−0.484	−0.753	−0.756	−0.577	−0.399
1991	−0.845	−0.846	−0.658	−0.459	−0.974	−0.971	−0.711	−0.479	−0.693	−0.699	−0.558	−0.393

N = 50 for each year. Workers aged 18–64. Min–50 is max [federal minimum, state minimum] minus the median wage of the state. Min–TM is max [federal minimum, state minimum] minus the (top 30 and bottom 30 percent) trimmed mean wage of the state.

Appendix 3: Minimum Wage Legislation Status, by State: 1979–1993

	1979	1980	1981	1982	1993	1984	1985	1986	1987	1988	1989	1990	1991	1992	1993
Alabama	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
Alaska	■	■	■	■	■	■	■	■	■	■	■	■	■	■	■
Arizona	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
Arkansas	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	◍	◍	◍	◍
California	●	●	●	●	●	●	●	●	●	■	■	■	■	●	●
Colorado	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍
Connecticut	●	●	●	●	●	●	●	●	■	■	■	■	■	●	●
Delaware	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	●	●	●
Florida	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
Georgia	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍
Hawaii	●	●	●	●	●	●	●	●	●	■	■	■	●	■	■
Idaho	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●
Illinois	◍	◍	◍	◍	◍	◍	●	●	●	●	●	●	●	●	●
Indiana	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	◍	◍	◍
Iowa	○	○	○	○	○	○	○	○	○	○	●	■	■	●	●
Kansas	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍
Kentucky	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	●	●	●	●
Louisiana	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
Maine	●	●	●	●	●	●	■	■	■	■	■	■	■	●	
Maryland	●	●	●	●	●	●	●	●	●	●	●	●	●	●	●
Massachusetts	●	●	●	●	●	●	●	■	■	■	■	◍	◍	◍	◍
Michigan	●	●	●	●	●	●	●	●	●	●	●	●	◍	◍	◍
Minnesota	◍	◍	◍	●	●	●	●	●	●	■	■	■	■	●	●
Mississippi	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
Missouri	○	○	○	○	○	○	○	○	○	○	○	●	●	●	●
Montana	○	○	○	○	○	○	○	●	●	●	●	●	●	●	●
Nebraska	○	○	○	○	○	○	○	○	●	●	●	●	●	●	●
Nevada	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	●	●	●
New Hampshire	●	●	●	●	●	●	●	●	■	■	■	■	●	●	●
New Jersey	●	●	●	●	●	●	●	●	●	●	●	●	●	■	■
New Mexico	○	○	●	●	●	●	●	●	●	●	●	●	◍	◍	
New York	●	●	●	●	●	●	●	●	●	●	●	●	●	●	●
North Carolina	◍	◍	◍	◍	●	●	●	●	●	●	●	●	◍	◍	●
North Dakota	○	○	○	○	○	○	○	○	○	○	●	●	●	●	●
Ohio	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	●
Oklahoma	◍	●	◍	◍	●	●	●	●	●	●	●	●	●	●	●
Oregon	◍	◍	◍	◍	◍	◍	◍	●	●	●	■	■	■	■	■
Pennsylvania	●	●	●	●	●	●	●	●	●	●	■	●	●	●	●
Rhode Island	◍	◍	◍	◍	◍	◍	◍	■	■	■	■	■	■	■	■
South Carolina	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
South Dakota	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	●	●
Tennessee	○	○	○	○	○	○	○	○	○	○	○	○	○	○	○
Texas	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	◍	◍	◍
Utah	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●
Vermont	●	●	●	●	●	●	●	■	■	■	■	■	■	●	●
Virginia	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●
Washington	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	■	■	●	●	●
West Virginia	◍	◍	◍	◍	◍	◍	◍	◍	●	●	●	●	●	●	●
Wisconsin	◍	◍	◍	◍	◍	◍	◍	◍	●	●	■	●	●	●	●
Wyoming	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍	◍

Compiled from the *Monthly Labor Review,* January issues,1979–1994. ○—no state provision. ◍—state minimum lower than federal. ●—state minimum equal to federal. ■—state minimum higher than federal. In years when the law changes midyear, the "higher" status is given.

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