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Does The Minimum Wage Affect Welfare Caseloads?

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Abstract

Although minimum wages are advocated as a policy that will help the poor, few studies have examined their effect on poor families. This paper uses variation in minimum wages across states and over time to estimate the impact of minimum wage legislation on AFDC caseloads, thus directly assessing whether minimum wages benefit a group they are intended to help. We find that the elasticity of the welfare caseload with respect to the minimum wage is 0.20. This suggests that minimum wages are not an efficient policy for facilitating the transition from welfare to work.

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The degree to which minimum wages affect employment has been of interest to economists and policy makers for many years. This interest has stemmed largely from a potential inconsistency between the intent of minimum wage laws and their theoretical effects. The goal of minimum wages is to improve individuals' abilities to support their families and avoid welfare (see, for example, Ellwood, 1988), but the textbook model of supply and demand predicts that such wage gains come at the expense of lower employment levels (Brown, 1988). In order for minimum wages to improve the overall well-being of families, the demand curve for low-skilled workers must be relatively inelastic.

Most research on minimum wages has focused on the relationship between minimum wage increases and employment rates. Examples from the literature include Baker, Benjamin and Stanger (1997), Brown, Gilroy and Kohen (1983), Card (1992a, 1992b), Card and Krueger (1994,1995), Card, Katz and Krueger (1994), Currie and Fallick (1996), Deere, Murphy and Welch (1995), Katz and Krueger (1992), Kim and Taylor (1995), Neumark and Wascher (1992, 1994), and Wellington (1991). In light of arguments in favor of minimum wage legislation (to help the poor), little attention has been paid to the impact of minimum wage increases on poor adults. While a few studies have recently emerged to fill this gap (e.g. Addison and Blackburn (1996), Connolly and Segal (1997), Neumark and Wascher (1997), Horrigan and Mincy (1993) and MaCurdy and O'Brien-Strain (1997)), to date there have been no studies estimating the impact of minimum wage legislation on potential welfare recipients.

A significant number of potential recipients could be affected by minimum wage policies; in 1996, for example, about 14% of unmarried female household heads with children had wages between the 1996 federal minimum wage and the 1997 federal minimum wage.[1] If higher minimum wages raise the earnings of low-skilled single mothers, then they may help to reduce the number of women who require welfare benefits in order to support their families. The minimum wage also may reduce welfare caseloads by increasing the return to work and, thus, encourage welfare recipients to enter the labor market. In 1992, the median Aid to Families with Dependent Children (AFDC) worker earned $4.50 an hour and the 25th percentile AFDC worker earned $3.45 per hour, which suggests that a significant fraction of working welfare recipients are willing to accept jobs for wages near the minimum.[2] However, if minimum wages reduce the demand for workers with limited skills, then they may unintentionally lead to increases in the number of women participating in welfare programs.

Recent passage of the Personal Responsibility and Work Opportunity Reconciliation Act of 1996 (PRWORA) intensifies the need to understand the relationship between work and welfare. PRWORA sets a limit on the maximum number of years an adult can receive welfare benefits at five years (although states may opt to impose a stricter time limit), and penalizes welfare recipients who do not find work (or an approved substitute for work) within two years of initiating welfare benefits. At nearly the same time as Congress passed PRWORA, the federal government increased the national minimum wage, and several states raised their minimum wages above the new federal level. The ostensible reason for these increases was to reduce individuals' need for welfare benefits by increasing the return to work. In fact, we have no idea whether higher minimum wages will help welfare recipients comply with the new welfare law.

Because existing estimates have typically been generated using different demographic groups, the relationship between minimum wages and employment cannot be relied on to predict how minimum wages will affect labor market outcomes and welfare participation among low-skilled single mothers. For example, many studies focus on teenagers, but single mothers are likely to have different labor supply and demand elasticities than teenagers. Teenagers' employment decisions are undoubtedly affected by the fact that they are primarily supported by their parents, whereas single mothers' labor supply decisions may be influenced by the availability and affordability of child care. On the demand side, low-skilled adults may be more desirable workers than teenagers because they have more work experience, are older, and are available for work over a wider range of hours.

In this study, we use variation in minimum wages across states and over time to identify their effect on the size of the state welfare caseload. Our empirical results indicate that, once state trends and a variety of other factors are accounted for, the elasticity of the welfare caseload with respect to the minimum wage is between 0.1 and 0.2. In other words, a 35 percent increase in the minimum wage (like the increase recently implemented in California) could lead to a 3 to 7 percent increase in the size of the welfare caseload, holding all else equal. These results suggest that minimum wages are not an efficient means of improving the financial independence of low-skilled single mothers, since the wage gains experienced by those who keep their jobs are counteracted by an increase in the welfare rolls. Policies like the Earned Income Tax Credit (EITC), which increases wages through the tax code without depressing the demand for low-skill labor, are likely to be more effective in facilitating the transition from welfare to work.

In the next section, we briefly describe the various mechanisms by which the minimum wage might be expected to affect the probability of becoming a welfare recipient and review the related empirical literature. In Section II, we lay out our empirical framework and discuss the data used to estimate the relationship between minimum wage levels and welfare caseloads. In Section III, we present our results, and in Section IV, we conclude with policy implications.

I. Background

Conventional neoclassical theory predicts a decline in the demand for low-skilled labor in response to an increase in the minimum wage. Since welfare spells are frequently precipitated by employment loss, a fall in job opportunities should lead to a rise in the number of welfare recipients.[3] Recent papers by the Council of Economic Advisors (1997): Blank (1997); Hoynes (1996); and Ziliak, Figlio, Davis and Connolly (1997) confirm that the probability of welfare participation increases when labor demand is low.

There are a number of reasons, however, that welfare participation rates might be unaffected when minimum wages increase. First, although the competitive model predicts a fall in employment when minimum wages increase, it has nothing to say about the magnitude of this decline. Empirical studies of the relationship between minimum wages and employment generally produce elasticity estimates ranging from 0 to -0.3. A few studies have even produced positive elasticity estimates. If the employment effects of minimum wage policies are negligible, then the elasticity of the welfare caseload with respect to the minimum wage also may be negligible.

A second reason that caseloads might be unresponsive to minimum wage increases is that many welfare recipients work in jobs that are “off the books” (see Edin and Lein (1997)). We find that more than 30% of working AFDC recipients in the 1993 CPS reported hourly wages that were more than 10% below the minimum.[4] The employment of women working such jobs is unlikely to be jeopardized by minimum wage increases, even if such policies reduce employment among other workers.[5]

If job opportunities facing potential welfare recipients are not substantially reduced when minimum wages rise, then such policies could lead to reductions in welfare caseloads. Those individuals previously working at or near the minimum wage almost certainly will experience an increase in earnings when the minimum wage is increased, and Bane and Ellwood (1994) find that up to 25% of AFDC exits are related to earnings increases. An increase in potential earnings also may induce some welfare recipients who otherwise would not look for work in order to become active members of the labor force. Studies of the Earned Income Tax Credit (see, for example, Dickert, Houser and Scholz (1995), and Eissa and Liebman (1996)) indicate that exogenous wage increases positively affect single mothers’ probabilities of working. The earnings generated from full-time work at the minimum wage typically hover close to the monthly AFDC guarantee (for a family of three) in the median state. As a result, fluctuations in the minimum wage could substantively affect the trade-off between work and welfare. Finally, if the labor market for low-skilled workers is not competitive, but is instead dominated by monopsonists, then increases in the minimum wage may lead to an increase in the demand for labor, which would subsequently reduce the welfare caseload.

While we know of no study that has explicitly assessed the impact of the minimum wage on the probability of welfare participation, a few recent studies have estimated the relationship between minimum wages and other measures of well-being among the poor. Connolly and Segal (1997) use changes in the federal minimum wage in 1990 and 1991 to estimate its impact on the earnings of poor working families. They compare outcomes across states with different fractions of affected workers and find that, in states where the increase should have had the biggest impact, poor and near-poor families experienced significantly higher growth in household earnings. In contrast, Addison and Blackburn (1996) do not find any evidence of a relationship between minimum wages and poverty. In a similar but more detailed study, Neumark and Wascher (1997) use CPS surveys matched over consecutive years to examine whether changes in family income are associated with state-level changes in minimum wages. They conclude that minimum wages effectively increase the incomes of some poor families, but that employment-reducing effects of minimum wages hurt other poor and near-poor families.

Two papers use simulations to measure the effect of minimum wages on poor families. Horrigan and Mincy (1993) simulate the effects of the minimum wage on income inequality. Controlling for disemployment effects, incomplete coverage, and compliance, they find that family income inequality would have been no different if the minimum wage had kept pace with inflation during the 1980’s. MaCurdy and O’Brien-Strain (1997) simulate the benefits and costs of a recent change in California’s minimum wage, noting that the costs of minimum wages must either reduce the profits of employers or be passed on to consumers in the form of higher prices for goods and services. When they assume that the minimum wage has no effect on employment, they find that low-income households pay a disproportionate share of these higher costs.

While the aforementioned studies shed light on the effect of minimum wages on poor families, they do not directly assess the impact of minimum wages on the well-being of single mothers, or their probability of participating in welfare programs. The purpose of this study is to estimate the impact of minimum wage legislation on welfare caseloads, explicitly investigating whether the minimum wage benefits a particular group that it is intended to help.

II. Data and Empirical Model

In the previous section we discussed a number of reasons why we are unable to predict from theoretical models whether minimum wages will increase or decrease welfare caseloads. The answer to this question must be obtained empirically. We begin our analysis by estimating a reduced form model motivated by the related literature on minimum wages and teenage employment:

$$C_{st} = bMW_{st} + X_{st}d + \sum_{s=1}^{s=50} a_s + \sum_{t=1}^{t=T} g_t + \sum_{s=1}^{s=50} w_s t + e_{st} \tag{1}$$

where C_{st} is the logarithm of the average monthly per capita AFDC-Basic caseload in state *s* and year *t*; MW_{st} is the logarithm of the minimum wage in state *s* and year *t*; X_{st} is a set of control variables that vary across states and over time; a_s is a time-invariant, state-specific effect; g_t is a year specific effect (we also present specifications that include a quartic time trend); and $w_s t$ is a state-specific linear time trend. The sign of $\hat{b}$ will tell us whether supply or demand effects dominate the response to minimum wage increases.

We use the AFDC caseload as our dependent variable instead of the number of AFDC recipients for two reasons. First, many recipients are children: we wish to distinguish between the number of households that are affected (and thus the number of adult recipients whose labor market opportunities are altered) from differences in household fertility across states. Another reason for choosing the caseload variable is that, in the wake of welfare reform, this statistic has received more attention from policymakers. We use actual caseloads rather than participation probabilities estimated from an individual level data set, such as the CPS or PSID, because this choice will substantially reduce measurement error in the dependent variable, particularly as available microdata yield small samples of single mothers at the state level.[6]

As in the employment literature, the vector X_{st} includes a number of variables reflecting the state's economy in time period t. It is important to control for a state's economic conditions because these conditions will affect the demand for labor, which may, in turn, affect the number of women requiring AFDC for support. We include a more abundant set of economic variables than most employment studies, including the logarithm of the real value of average production wage, the log of gross state product, and the state unemployment rate. We also include two lagged values of the unemployment rate because a number of recent studies (e.g. Blank (1997), Council of Economic Advisors (1997), Ziliak, Figlio, Davis and Connolly (1997)) have found that caseload fluctuations are a function of employment opportunities in previous periods.

Blanchard and Katz (1992) present evidence that individuals migrate to states experiencing economic booms; those who migrate to take advantage of improved labor market conditions are unlikely to be welfare participants. In light of this, we increase the flexibility of our model by moving the denominator of C_{st} to the right side of the equation, so that log (population) is treated as an explanatory variable. Note, however, that because we are controlling for the number of women in the state at time *t*, this specification does not change our interpretation of b, the elasticity of the per capita caseload with respect to the minimum wage.

Our vector of control variables includes average production wages because the demand for minimum wage workers is likely to depend on their wage relative to the wages of other workers in the state. The production wage is not an ideal measure of the market-clearing wage for adults at risk of becoming welfare recipients because the majority of these adults are single women, and women are more likely to be employed in the service sector. This is the only average wage series that is available both over time and across states, however, and it should be highly correlated with wages in more relevant occupations. We shall examine the sensitivity of our results to alternative wage measures that we have calculated ourselves using the CPS.

Many minimum wage studies bring alternative wages into the regression model through the Kaitz index—the ratio of the state minimum wage to the state average production wage—adjusted to reflect the fraction of workers covered by minimum wage legislation. There are several reasons that we do not use the Kaitz index in this study. First, information on coverage under state minimum wage laws is not available.[7] Second, there is no reason to assume that the weights given to the minimum wage and coverage components of the Kaitz index accurately reflect their relative importance. The Kaitz index treats the level and coverage effects equally, but many researchers find the estimated effect of coverage on employment weak, and that the relative level of the minimum wage is more important.[8] [9] Finally, if states in which average wages are rising tend to be states in which employment is rising and caseloads are falling, then estimates based on the Kaitz index might be misinterpreted as minimum wage effects. To overcome this problem, while still acknowledging the fact that the demand for minimum wage workers will depend on their *relative* wages, we make use of the fact that $\log(\frac{MW_{st}}{AvgWage_{st}})$ can be written as $\log MW_{st} - \log AvgWage_{st}$ and include the terms separately.

X_{st} also includes several variables that partially account for differences in welfare take-up rates across states and over time. The probability of participating in a welfare program is affected by the level of benefits available (see Moffitt, 1992). Moreover, residents' preferences for benefit levels may be correlated with their tastes for minimum wages. Thus, we add to our equation measures of the generosity of each state's welfare benefits. We include the total value of AFDC benefits and Food Stamp benefits for a family of three with no income, and a dummy variable indicating whether the state maintains an AFDC-Unemployed Parent program. Prior to

1990, states had the option of providing AFDC-UP benefits; in 1990, 28 states did so. After 1990, state provision of these benefits became mandatory. Our equation also controls for the fraction of households headed by a single mother,[10] and the fraction of women between the ages of 15 and 55 who are white.

The model is estimated using state-level panel data spanning the years from 1983 to 1996.[11] Panel data allow us to control for both national time trends and static differences across states that might be correlated with both minimum wage levels and the size of the welfare caseload. An alternative would have been to pool together individual-level data for the same set of years, allowing us to test the hypothesis that supply and demand responses vary across different family types. However, since no policy variation occurs at the individual level, we believe that the detail lost by turning to aggregate data is compensated by the improved accuracy with which we can measure our dependent variable. Sample sizes in the most common individual level datasets are very small for a large number of states. Many employment studies are based on aggregate time-series or state-level panel data.[12]

We choose 1983-1996 as the years for our analysis because several of our control variables are not readily available prior to the early 1980's or after 1996. In addition, as shown in Table 1, minimum wages display virtually no variation across states prior to 1985, whereas after 1987 there is substantial variation in minimum wages both across states and within states over time. Within-state variation in the *real* minimum wage after 1987 is even larger than the nominal variation shown in Table 1. If data from earlier years were added to the sample, the identification of minimum wage effects would increasingly rest on time-series variation in the federal minimum wage. Most of this variation is eliminated when we include time dummies (or time-trends) in our equation.[13]

Appendix A provides details on the sources of our data and Table 2 presents their means and standard deviations. We use the CPI to adjust all monetary variables to 1994 dollars.[14] The data set contains 700 observations.

III. Results

We began our analysis by imposing as few restrictions on the data as possible, and then systematically placed more structure on the model until we arrived at equation (1). When the minimum wage effect is identified using only time-series variation (including only state fixed effects) or only cross-sectional variation (including only year dummies or a quartic time trend), estimates of the minimum wage's effect are negative and sometimes statistically significant. In other words, states that tend to have high minimum wages also tend to have low caseloads, and years in which real minimum wages are relatively high correspond to years in which caseloads are relatively low.

We then placed more restrictions on the data by including both state and year fixed effects, as in Neumark and Wascher's (1992) employment study. The resulting minimum wage effect is about –0.16, and is statistically different from zero. On the other hand, if we account for unobservable state-specific factors by taking first differences (rather than by including indicator variables for each state), our estimate of the minimum wage effect is close to zero and not statistically significant. When we replace our year-specific effects with a quartic time trend the two estimation approaches continue to yield different results.

The state-fixed effect and first-differences estimation strategies both identify the impact of the minimum wage using variation in its real value within states over time, but the first-differences approach uses within-state variation across two years, whereas the inclusion of state-indicator variables (which is equivalent to mean-differencing) uses within-state variation across the entire length of the panel.[15] The difference in the estimates produced by these approaches is consistent with the existence of unobserved state factors that are not fixed over time.[16] We were able to reconcile the results of the two estimation approaches by including state-specific linear time trends in our regression model.[17] The estimates presented in Table 3 are, therefore, based on a specification that includes state dummies, year dummies (or quartic time trends), and linear state-specific trends. This is the same model described in equation (1). In this specification, the identification of our estimates comes from variation in minimum wages around a state-specific trend. Given that the real minimum wage typically trends downward over time and then increases abruptly with new legislation, the source of our identification seems appropriate.

Table 3 presents four different sets of results. All standard error estimates are corrected for heteroskedasticity and serial correlation using a procedure suggested by Newey and West (1987).[18] [19] In the first two columns, we report estimates based on equation (1), and in the second two columns we present estimates based on a model that replaces the year-specific dummy variables with a quartic time trend. The advantage of this first specification is that it allows us to implicitly control for all unobserved time-specific factors that are common across states. The disadvantage of this specification is that it wipes out our ability to use changes in the federal minimum wage as a source of identification. When we include both federal minimum wage and state minimum wage variables together in the model with a quartic time trend, the coefficient estimate on the state minimum wage variable is reduced by about one third, suggesting that variation in the federal minimum wage is providing some of the identification. As shown in Figure 1, predicted caseloads based on the models that include year-specific effects and the quartic trend are very similar. The quartic trend does a good job of tracing the caseload pattern over time, but still allows changes in federal minimum wage legislation to contribute to the identifying variation. We remain agnostic about whether the gains from this additional flexibility outweigh the potential omitted variables problems, and present both sets of results: they are very similar.

For each specification, we present both unweighted OLS estimates and weighted least squares estimates, where observations for each state are weighted by the population of women between the ages of 15 and 55. Card, Katz and Krueger (1994) and Krueger (1995) have shown that weighting can have a significant impact on estimated employment elasticities. Weighting is appropriate if one wants to estimate the degree to which minimum wage legislation will contribute to an overall rise or fall in U.S. welfare participation probabilities. California and New York, for example, together comprise almost 30% of the U.S. caseload, so caseload changes in those states will contribute a great deal to changes in the national caseload. On the other hand, big states and small states are contributing equal amounts of information to our analysis, so our WLS regressions may place too much emphasis on large states.

The estimates in Table 3 indicate that increasing the minimum wage by 10% will lead to a one- to two-percent increase in the size of the welfare caseload. To put this estimate into perspective, these results suggest that, in California, where minimum wages were recently increased from $4.25 to $5.75, we would expect welfare caseloads to rise by between 3 and 7%, holding everything else constant. Of course, PRWORA went into effect at approximately the same time that California's minimum wage rose: the new law places time limits on welfare recipiency and eliminates welfare as an entitlement, so TANF caseloads may be altered to a lesser degree.

Our results are consistent with a negative employment elasticity for low-skilled single mothers of between 1 and 2 %, provided that none of those who lost employment were already welfare recipients. While these elasticities are well within the range of researchers' estimates examining the relationship between minimum wage levels and teenage employment, it is important to note that if some job losers were already participating in the AFDC program, then the employment elasticity would need to be larger in order to generate a 1-2% increase in the size of the welfare caseload.

The other coefficient estimates presented in Table 3 are either of the expected sign or are indistinguishable from zero. Unsurprisingly, we find that good economic conditions are associated with lower caseloads, and that more generous AFDC benefits are associated with higher caseloads. Like previous studies of caseload dynamics (none of which have considered the minimum wage), we find that employment demand in previous periods affects the current caseload. Our estimated coefficients on these variables are similar in magnitude to Blank (1997). We also note that periods of high population growth (relative to trends) are associated with reductions in welfare caseloads. This is consistent with Blanchard and Katz (1992), who find evidence of state-to-state migration in response to changes in labor market conditions.

IV. Robustness of the Estimates

We carry forward specifications 2 and 3 (in Table 3) for the remainder of our analysis. These two specifications are chosen because they allow us to display the results from both year-specific and time-trend models, as well as results produced by weighted and unweighted regressions. In addition, the coefficient estimates produced by these specifications are in the mid-range of those shown in Table 3. In all of the regressions described below, specification 1 continues to yield the smallest estimates and specification 4 continues to produce the largest estimates. Minimum wage estimates are robust across all additional specifications, except that estimates produced by the unweighted model with year-specific effects lose some of their magnitude and statistical significance in certain models that include state specific business-cycle effects.

Alternative model specifications

Tables 4.a and 4.b show how our estimates of minimum wage effects change when we change the specification of the minimum wage, our controls for economic and other unobservable factors, or the time period of our estimation. First, we consider the possibility that our linear state-specific time trends are too restrictive. In the first column of each table, we present estimates that add quadratic state-specific time trends to our basic equation. The inclusion of these terms reduces the estimates of the minimum wage's effect somewhat, but they continue to be statistically significant. Next, we estimate the effect of the minimum wage *level* on welfare caseloads. Although it is convenient to think about minimum wage effects in terms of elasticities, one might argue that a logarithmic specification incorrectly implies that an increase in the minimum wage from $1 to $2 would have the same percentage effect on welfare caseloads as a wage increase from $3 to $6. Nonetheless, the estimate in column 2, which indicates that raising the minimum wage by $1 would increase welfare caseloads by 4%, is similar to our elasticity estimate.

Third, we examine what happens to our minimum wage estimates when the contemporaneous unemployment rate is eliminated from our regression equation. If the minimum wage increases the unemployment rate, and subsequently increases welfare caseloads, then including the unemployment rate in our regression will lead to biased estimates of the true minimum wage effect. In fact, we see that when we omit this variable from our regression equation, the estimated minimum wage coefficient barely changes.

One concern elucidated in the minimum wage literature is that the timing of minimum wage increases corresponds to periods of economic growth, when employment levels would be expected to rise. If this hypothesis is true, then the coefficients presented in Table 3 will be biased downward, and can be thought of as a lower bound of the true minimum wage effect. Conversely, if minimum wages are legislated during downturns in the economy, then the estimates in Table 3 will be too high. In columns 4, 5, and 6 of Tables 4.a and 4.b, we allow for the possibility that state heterogeneity in business cycle responses is correlated with state minimum wage changes by adding state interactions with the unemployment rate, production wage and gross state product variables to our equation. The magnitude of the estimated minimum wage coefficient continues to lie between 0.1 and 0.2. Note, however, that in columns 1, 4, 5 and 6, the estimates are generally smaller than those produced by the basic equation in Table 3. This raises the possibility that, even though we include a much richer set of economic controls than do most studies of the minimum wage, our results (and the results of most employment studies) may still be biased upward by our inability to *completely* control for economic factors specific to the state in time period t. Unfortunately, we know of no minimum wage studies that have used a convincing instrument to evade this problem. Neumark and Wascher (1992) instrument the minimum wage with the average minimum wage in surrounding states, but in our model this instrument is either uncorrelated with the minimum wage or correlated with the dependent variable, depending on the specification we use. The potential endogeniety of the minimum wage deserves additional investigation.

We also conducted our analysis on a subsample of observations from the period 1987-1994 because we were concerned that our results prior to 1987 might be driven by the very few states (Alaska, Connecticut and Maine) that had minimum wages higher than the federal level, or, in the case of the time-trend specification, by movements in the federal minimum that coincided with other unobserved changes. As seen in column 7, our estimated minimum wage coefficient for this subsample continues to be statistically different from zero and in the 1-2% range.

Omitted variables

With the exception of the demographic variables included in equation (1), our empirical equation follows models that have been used to estimate the relationship between minimum wages and teenage employment rates. Like those studies, we control for a large number of economic variables, because we want to be as confident as possible that we are distinguishing between caseload changes resulting from the impact of minimum wages on the demand for low-skilled workers and arising from other economic conditions affecting the caseload. In this section, we consider the possibility that other factors correlated with minimum wage changes are driving our results. The inclusion of these additional variables is motivated by three recent studies of caseload dynamics by Blank (1997), the Council of Economic Advisors (1997) and Ziliak, Figlio, Davis and Connolly (1997).

All three studies explore the possible effect of welfare reform on caseload size, using information on state welfare waivers. Many states obtained waivers from federal rules governing the AFDC program prior to the passage of PRWORA in 1996, which allowed them to incorporate work incentives such as "making work pay", work requirements, limits on the length of time one could receive welfare, and enforcement of child support. With the exception of work incentives,[20] these waivers were expected to reduce welfare caseloads, and all three of the studies mentioned above find evidence that some of these waivers were positively correlated with caseload reductions. Since many waivers were initiated at the same time as an increasing number of states were establishing minimum wage levels above the federal minimum, it is important to consider whether their implementation is in any way responsible for the magnitude of our estimated minimum wage coefficient. In order to test the hypothesis that the imposition of welfare waivers is driving our results, we include four dummy variables indicating the presence of four categories of statewide waivers:[21] (1) those that require work; (2) those that provide incentives to work; (3) those that impose limits on the length of time an individual can receive welfare benefits; and (4) those that enforce child support rules.[22]

In addition to the welfare waivers, we also add two variables intended to capture changes in state Medicaid eligibility that occurred during the period we study. Prior to 1988, eligibility for Medicaid was explicitly tied to AFDC eligibility, and there is evidence that this link encouraged AFDC participation and reduced single mothers' probabilities of working (Yelowitz, 1995). Federal legislation passed in 1986 and 1987 severed this link by eliminating Medicaid eligibility criteria related to AFDC eligibility and increasing income thresholds for some children. States were given some flexibility in implementing these expansions. We control for these expansions because severing the relationship between AFDC and Medicaid may have changed the labor supply of low-income individuals who value health insurance (Yelowitz, 1995). In addition, states that implemented the expansions more rapidly and instituted broader eligibility criteria also might have provided more generous minimum wages during this period. The first variable, HIGHAGE, indicates the highest age for which children in low-income families are eligible for Medicaid benefits. It is equal to zero in years prior to the expansions and varies by state and year starting in 1988. The second variable, MAXINC, is the threshold level of income (as a percent of the poverty line) below which at least some children are eligible.[23] This variable also is equal to zero prior to 1988 and varies by state and year after that. Our expectation is that higher values of HIGHAGE and MAXINC will be associated with lower caseloads. Blank includes similar variables in her 1997 paper.

The first column of Tables 5.a and 5.b present estimates of equation (1) with the addition of the welfare waiver and Medicaid expansion variables described above. As in other studies, we find that the presence of welfare waivers is correlated with the size of the caseload. In particular, we find that child support programs reduce caseloads by about 6.4%. Ziliak et al. (1997) do not find evidence that child support programs reduce welfare caseloads, but their specification includes a much sparser set of control variables. Also in contrast to Ziliak et al., we find that time limits are associated with larger caseloads. The sign of this estimate is not intuitive, but it is consistent with Blank (1997), who includes a much richer set of control variables and concludes that the coefficients on her welfare waiver variables are correlated with other changes occurring at the same time. Martini and Wiseman (1997) also note that the apparent effect of welfare waivers may not be causal. In regressions not included in the table, we pooled all of the waivers together to include a single indicator variable equal to 1 if the state had been granted any waiver. Our estimated coefficients on this variable are about –0.026 (standard error estimate of 0.01) in each specification, which is very similar to Blank's estimated waiver coefficient. Of most relevance for this study, however, is that the inclusion of these variables does not alter the minimum wage estimate. If anything, the estimates of the effect of the minimum wage on caseloads become larger.

Although the caseload studies conducted by the Council of Economic Advisors and Ziliak, et al. include few control variables, Blank (1997) considers a number of other factors that might cause caseloads to grow. For example, in models that include state-specific time trends, she finds that the fraction of the population consisting of new immigrants helps to explain caseload dynamics. Borjas and Hilton (1996) have shown that, in the 1990s, immigrant households were 7 percentage points more likely than native households to participate in welfare programs; thus, large influxes of immigrants might increase welfare caseloads. It also is possible that immigrants crowd potential welfare recipients out of jobs, thus increasing the likelihood of welfare participation of non-immigrants. We examine these possibilities by rerunning our basic regression equation with the addition of data on the fraction of the population in each state that are newly admitted immigrants, and a one year lag of this variable.[24] The results are presented in column 2 of Tables 5.a and 5.b. The estimated minimum wage coefficient in this specification is virtually unchanged.

Next, we consider the possibility that our minimum wage estimates are biased upward because we have not sufficiently controlled for voter tastes. Suppose, for example, that state residents' preferences are such that they favor policies to help the poor. While our inclusion of state fixed effects and state-specific time trends should go a long way toward controlling for these unmeasurable tastes, there remains the possibility that fluctuations in preferences around the trend can explain our findings. We test this possibility in column 5 by adding to our regression a set of variables that measure the political atmosphere within the state. These include a dummy variable that is equal to 1 if there is Democratic control of both legislative houses *and* the governorship, a dummy variable equal to 1 if control of the legislative houses and/or governorship is divided between Democrats and Republicans (DIVIDED CONTROL), the ratio of per-capita state debt to per-capita income (DEBT/INCOME), a dummy variable equal to 1 if the governor does not have the power to use a line-item veto (NO LINE ITEM VETO), the percent of the upper legislative house who are Democrats (% UPPER DEMOCRATS), and the percent of the lower legislative house who are Democrats (% LOWER DEMOCRATS). These variables have been gathered for all states except Alaska, Hawaii and Nebraska, and are available from 1983-1994. Again, the inclusion of these variables has little effect on the minimum wage estimate.

Next, we check the robustness of our minimum wage estimates to the inclusion of alternative wage controls. Blank includes the log of wages at various percentiles of the wage distribution in her caseload regressions. We replace production wages with the log of the 10th percentile and the log of the 25th percentile wage in the CPS Monthly Outgoing Rotation Group files. Since these measures (particularly the 10th percentile wage) are more likely to be directly affected by minimum wage legislation than the average production wage, this is not our preferred specification, but we include it for purposes of comparison. The estimated minimum wage coefficient is unaffected by the inclusion of these alternative wage measures, but, unlike Blank, we find little evidence that higher "alternative" wages are correlated with lower caseloads. The difference appears to result from the fact that we include additional controls for economic conditions.

Finally, we approximately replicate Blank's specifications by including all of the variables added in columns 1-4 into a single regression. The estimates produced by the analysis of this more inclusive model are shown in the 5th column of Table 5. In the 6th column we include a similar specification, replacing the 10th and 25th percentile wages with the production wage. Even when all of these variables are included in the regression model, the estimated minimum wage effect remains similar to the estimates in Table 3.

We also have run two additional regressions that are not included in the tables. First, we ran specifications in which our dependent variable was replaced with the logarithm of the AFDC-UP caseload. Although the estimated coefficients were typically positive, they were not statistically different from zero. This leaves open the possibility that AFDC-UP recipients are affected differently by minimum wage increases, which would not be surprising since the eligibility criteria for AFDC-UP participation are different than criteria for participation in the AFDC-Basic program. In a second analysis, we estimated the relationship between minimum wage increases and the probability of employment among unmarried female heads of household, using the March *Current Population Survey*. Although our point estimates tended to be negative, the confidence intervals around them were quite large: we could not reject the possibility that the true relationship was positive. Small sample sizes may be partially responsible for the lack of precision in our estimates—for example, many of the state-year cells in our data have fewer than 35 observations. Furthermore, many of the women in our sample will be working for wages that are much higher than the minimum wage.

V. Conclusion

This paper has investigated the effect of minimum wages on welfare caseloads using state level panel data. Our results suggest that minimum wage levels have a substantial positive effect on the size of the AFDC caseload, and that, therefore, minimum wages will not help many low-income families achieve self-sufficiency. A ten-percent increase in the minimum wage is estimated to increase welfare caseloads between 1 and 2 percent. We find that the inclusion of other variables that influence the evolution of caseloads over time, such as state-specific welfare reforms and changing political preferences, do not weaken our results.

Increases in minimum wages might lead to rising welfare caseloads for several reasons. First, as suggested by the classical economic model, the imposition of minimum wages can result in a reduction in the number of job vacancies. As jobs become less available, low-skilled workers are less able to find work and are thus more likely to apply for AFDC benefits. Second, increasing minimum wages might cause more workers to enter the labor market. If the new labor market entrants possess more human capital than do likely welfare recipients, then potential welfare recipients may be "crowded out" of jobs. Lang and Kahn (1998) find that minimum wage increases shift food-service jobs from adults to teenagers and students, which provides support for this possibility. Either job loss or crowding-out could also lead to hours reductions rather than total job loss, and it may be that the earnings decline resulting from a reduction in hours of work is large enough to increase welfare participation. Our data do not enable us to distinguish between these possibilities, but we hope to explore them in future research.

Our results suggest that increases in the minimum wage may impose a real cost on government — the cost associated with larger welfare caseloads. For example, our estimated minimum wage elasticity suggests that the one-dollar minimum wage increase (from $5.15 to $6.15) recently proposed by President Clinton could increase welfare caseloads by about 130,000 families and cost state and federal governments close to 1 billion dollars a year.

Because higher production costs resulting from minimum wages are at least partly borne by consumers, the cost to society of minimum wage policies will be higher than the cost resulting from the increase in welfare payments alone. This additional cost might be considered acceptable, however, if minimum wage policies improved outcomes among the poor. Our findings suggest that the cost to society is not balanced by reductions in the number of families depending on welfare. These indirect costs should be considered when comparing the cost of the minimum wage to the cost of policies with similar goals, such as the Earned Income Tax Credit (see Burkhauser, Couch, and Glenn, 1996).

Our policy conclusions come with a number of caveats. First of all, it is important to remember that low-skilled adults who maintain their employment levels are likely to benefit from the higher wages mandated by minimum wage legislation. On balance, however, our finding that welfare caseloads increase when minimum wages rise suggests that the number of families whose earnings rise enough to help them move off of welfare is exceeded by the number whose employment-based income declines. Second, if minimum wages are increased in response to rising caseloads or poor economic conditions, then our policy variable will be endogenous and our estimates may be biased upward. We have tried to eliminate potential omitted variables bias by including as many control variables as possible, but if these controls are not complete then our study will be subject to the same criticism as the teenage employment studies. We know of no minimum wage studies that have used a convincing instrument to get around this problem. Finally, it is important to remember that research on the impact of minimum wages on teenage employment has produced elasticity estimates that are all over the map, and that these estimates vary according to the data years analyzed, the source of the identifying variation, and the empirical strategies used to eliminate fixed effects. We have considered a number of alternative specifications in our analysis, but, as in the employment literature, an alternative approach may yield different results.

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Appendix A

The following list documents the sources for the independent variables included in our regression analyses.

AFDC Caseloads	Social Security Bulletin, *Annual Statistical Supplement*, various years.
Minimum Wage	U.S. Department of Labor, Division of State Standards Programs, Wage and Hour Division, Employment Standards Administration.
Production Wages	U.S. Department of Labor, Bureau of Labor Statistics: *State and Area Current Statistics-Most Requested Series* website.
Unemployment Rate	U.S. Department of Labor, Bureau of Labor Statistics: *State and Area Current Statistics-Most Requested Series* website.
State Population	U.S. Census Bureau: *State Population Estimates* website.
Gross State Product	U.S. Department of Commerce: *Bureau of Economic Analysis* website.
AFDC+Foodstamp Benefit Levels	U.S. House of Representatives: *Background Materials and Data on Programs Within the Jurisdiction of the Committee on Ways and Means*, Washington, D.C., various years.
Existence of AFDC-UP program	Congressional Research Service, The Library of Congress, 1987. *State Use of Aid to Families with Dependent Children-Unemployed Parent (AFDC-UP) Program: An Overview*, 87-969 EPW & U.S. Department of Health and Human Services, Social Security Administration, 1992. *Characteristics and State Plans for Aid to Families with Dependent Children.*

HighAge MaxInc	National Governors' Association, "State Coverage Of Pregnant Women and Children," NGS Center for Policy Research, various years, and Intergovernmental Health Policy Project, "Medicaid And Indigent Care: An Overview of State Legislative Activity," 1992 and 1994 (George Washington University).
Welfare Waivers	Ziliak, James P., David N. Figlio, Elizabeth E. Davis, and Laura S. Connolly, "Accounting for the Decline in AFDC Caseloads: Welfare Reform or Economic Growth?" University of Oregon, unpublished mimeo.
Immigration	Immigration and Naturalization Service, *Statistical Yearbooks*, 1988 and 1996. Washington, D.C.: NTIS.
Percent of households headed by single women with children	Current Population Survey Monthly Outgoing Rotation Groups, 1980-1996.

Table 1

State Minimum Wage Levels: 1980-1996[1]

	1980	1981	1982	1983	1984	1985	1985	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996
Federal Min. Wage	3.1	3.35	3.35	3.35	3.35	3.35	3.35	3.35	3.35	3.35	3.35	3.8	4.25	4.25	4.25	4.25	4.25
AL	3.6	3.85	3.85	3.85	3.85	3.85	3.85	3.85	3.85	3.85	3.85	4.75	4.75	4.75	4.75	4.75	4.75
AK																	
AZ																	
AR																	
CA										4.25	4.25						
CO																	
CT	3.12	3.37	3.37	3.37	3.37	3.37	3.37	3.37	3.75	4.25	4.25	4.25	4.27	4.27	4.27	4.27	4.27
DE	3.14	3.48	3.62	3.82	3.85	3.86	4.16	4.33	4.33	4.33	4.85	4.85	5.45	5.45		5.25	5.25
DC																	
FL																	
GA																	
HI									3.85	3.85	3.85	3.85		5.25	5.25	5.25	5.25
ID											3.85	4.25	4.65	4.65	4.65	4.65	4.65
IL																	
IN																	
IA																	
KS																	
KY																	
LA																	
ME								3.55	3.65	3.75	3.75						4.75
MD																	
MA						3.45	3.55	3.65	3.65	3.75	3.85	3.85					
MI																	
MN									3.55	3.85	3.95	4.25					
MS																	
MO																	
MT																	
NE																	
NV											3.40						
NH																	
NJ								3.45	3.55	3.65	3.75	3.85					
NM														5.05	5.05	5.05	5.05
NY																	
NC																	
ND																	
OH																	
OK																	
OR											4.25	4.75	4.75	4.75	4.75	4.75	4.75
PA										3.70	3.70	4.25	4.45	4.45	4.45		
RI								3.55	3.65	4.00	4.25	4.25				4.45	4.45
SC																	
SD																	
TN																	
TX																	
UT																	
VT																	
VA								3.45	3.55	3.65	3.75	3.85				4.75	4.75
WA										3.85	4.25	4.25			4.9	4.9	4.9
WV																	
WI																	
WY																	
# of States	3	3	3	3	3	4	4	8	10	13	16	13	6	8	8	10	11

Exceeding Fed. Min. [2]

[1] As of January 1, in the given year.
[2] Alaska's minimum wage is always fifty cents above the federal minimum.

Table 2

Summary Statistics

	Mean (Standard Deviation)		
Log (Per Capita Caseload)	10.58 (1.184)	% New Immigrant	0.002 (0.002)
Log (Minimum Wage)	1.480 (0.075)	% New Immigrant$_{t-1}$	0.002 (0.002)
Log (Average Production Wage)	2.506 (0.126)	State has any Welfare Waiver	0.121 (0.327)
Unemployment Rate	6.320 (2.052)	State has Work Incentive	0.080 (0.271)
Unemployment Rate$_{t-1}$	6.612 (2.052)	State has Time Limits	0.047 (0.212)
Unemployment Rate$_{t-2}$	6.762 (2.052)	State has Child Support	0.090 (0.286)
Log (Gross State Product)	-3.733 (0.198)	State has Work Requirement Requirement	0.080 (0.241)
Employment Growth	0.005 (0.008)	Democratic Control of Both Houses & Gvshp	0.319 (0.467)
Log (Population)	14.94 (1.005)	Divided Control	0.365 (0.482)
Fraction of the Population White	0.805 (0.136)	(Debt/Income)	82.76 (51.83)
Fraction of Households Female Headed	0.117 (0.031)	No Line Item Veto	0.135 (0.342)
Maximum Benefit: AFDC+FS Family of Three	699.2 (151.3)	% Upper Democrats	60.05 (18.21)
State Provides AFDC-UP	0.720 (0.449)	% Lower Democrats	58.83 (17.58)
HighAge	5.954 (6.173)		
MaxInc	102.3 (81.00)		
Sample Size	700		

Table 3

Basic Regression Estimates of the Effect of the Minimum Wage on Welfare Caseloads

	Time Dummies		Quartic Time Trend	
	Unweighted	Weighted	Unweighted	Weighted
	(1)	(2)	(3)	(4)
Minimum wage	0.124	0.184	0.184	0.236
	(0.070)	(0.053)	(0.065)	(0.056)
Log (Average Production Wage)	-0.362	0.087	-0.284	0.286
	(0.252)	(0.249)	(0.248)	(0.235)
Unemployment Rate	0.006	0.004	0.008	0.003
	(0.004)	(0.004)	(0.003)	(0.004)
Unemployment Rate t-1	0.011	0.004	0.012	0.008
	(0.004)	(0.003)	(0.003)	(0.003)
Unemployment Rate t-2	0.012	0.018	0.011	0.013
	(0.004)	(0.004)	(0.003)	(0.003)
Log(Gross State Product)	-0.483	-0.601	-0.507	-0.629
	(0.107)	(0.110)	(0.102)	(0.099)
Log(Population)	-1.358	-1.661	-1.457	-1.846
	(0.541)	(0.352)	(0.482)	(0.292)
Fraction of Population White	-2.973	0.625	-2.570	1.051
	(1.998)	(2.624)	(1.938)	(2.655)
Fraction of Households Female Headed	0.416	0.591	0.190	0.148
	(0.219)	(0.199)	(0.157)	(0.139)
Maximum Benefit For AFDC+FS Family of 3	0.001	0.001	0.001	0.001
	(0.0002)	(0.0002)	(0.0002)	(0.0002)
State Provides AFDC-UP	0.053	0.089	0.061	0.095
	(0.019)	(0.024)	(0.017)	(0.021)
State*Time Trends	yes	yes	yes	yes

Table 4.a

Additional Estimates of Minimum Wage Effects on Welfare Caseloads Weighted and with Time Indicator Variables

	Quadratic State Trends	Use MWage Level	Eliminate Urate	State-Specific Business Cycles? State* Urate	State* Prod Wg	State* GSP	Sample= 1987-96
	(1)	(2)	(3)	(5)	(6)	(7)	(8)
Minimum Wage	0.118	0.041	0.178	0.145	0.247	0.083	0.111
	(0.055)	(0.012)	(0.054)	(0.059)	(0.065)	(0.051)	(0.050)
Log (Average Production Wage)	0.119	0.086	0.126	0.218	2.353	0.160	0.258
	(0.185)	(0.249)	(0.253)	(0.264)	(1.815)	(0.235)	(0.253)
Unemployment Rate	0.006	0.004		-0.011	0.004	0.005	0.008
	(0.004)	(0.004)		(0.006)	(0.007)	(0.004)	(0.004)
Unemployment Rate t-1	0.010	0.004	0.006	0.006	0.002	0.007	0.003
	(0.003)	(0.003)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
Unemployment Rate t-2	0.012	0.018	0.017	0.012	0.017	0.014	0.021
	(0.004)	(0.004)	(0.005)	(0.004)	(0.008)	(0.004)	(0.005)
Log(Gross State Product)	-0.169	-0.601	-0.646	-0.586	-0.741	-0.206	-0.183
	(0.114)	(0.110)	(0.098)	(0.113)	(0.128)	(0.181)	(0.152)
Log(Population)	-2.454	-1.656	-1.600	-1.837	-1.720	-1.574	-1.711
	(0.451)	(0.355)	(0.349)	(0.348)	(0.333)	(0.352)	(0.353)
Fraction of Population White	10.143	0.602	0.695	1.634	0.325	3.116	8.706
	(2.995)	(2.622)	(2.596)	(2.366)	(2.360)	(2.857)	(2.601)
Fraction of Households Female Headed	0.322	0.591	0.606	0.321	0.473	0.201	0.255
	(0.197)	(0.200)	(0.202)	(0.171)	(0.284)	(0.183)	(0.201)
Maximum Benefit For AFDC+FS Family of 3	0.001	0.001	0.001	0.001	0.001	0.001	0.001
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0003)	(0.0001)	(0.0002)
State Provides AFDC-UP	0.099	0.089	0.087	0.049	0.059	0.064	0.101
	(0.025)	(0.024)	(0.023)	(0.013)	(0.030)	(0.017)	(0.023)
State*Time Trends	yes	yes	yes	yes	yes	yes	yes

Table 4.b

Additional Estimates of Minimum Wage Effects on Welfare Caseloads
Unweighted and with Quartic Time Trend

	Quadratic State Trends	Use MWage Level	Eliminate Urate	State-Specific Business Cycles? State* Urate	State* Prod Wg	State* GSP	Sample= 1987-96
	(1)	(2)	(3)	(5)	(6)	(7)	(8)
Minimum Wage	0.120	0.040	0.199	0.125	0.197	0.113	0.123
	(0.060)	(0.015)	(0.065)	(0.053)	(0.058)	(0.059)	(0.049)
Log (Average Production Wage)	-0.021	-0.281	-0.218	-0.251	1.991	-0.165	-0.075
	(0.182)	(0.248)	(0.241)	(0.257)	(0.838)	(0.233)	(0.265)
Unemployment Rate	0.008	0.008		-0.005	0.009	0.007	0.011
	(0.003)	(0.004)		(0.006)	(0.003)	(0.004)	(0.004)
Unemployment Rate t-1	0.012	0.012	0.015	0.013	0.012	0.013	0.007
	(0.002)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Unemployment Rate t-2	0.008	0.011	0.010	0.008	0.013	0.010	0.017
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.004)
Log(Gross State Product)	-0.343	-0.507	-0.601	-0.456	-0.541	-0.222	-0.291
	(0.082)	(0.103)	(0.084)	(0.101)	(0.114)	(0.195)	(0.104)
Log(Population)	-1.937	-1.453	-1.413	-1.538	-1.292	-1.080	-1.171
	(0.466)	(0.482)	(0.474)	(0.477)	(0.462)	(0.531)	(0.410)
Fraction of Population White	4.801	-2.620	-2.437	-1.550	-3.211	-2.648	2.911
	(2.871)	(1.940)	(1.962)	(1.857)	(2.019)	(1.899)	(2.136)
Fraction of Households Female Headed	-0.089	0.183	0.216	0.038	0.006	-0.156	-0.010
	(0.116)	(0.157)	(0.162)	(0.143)	(0.126)	(0.129)	(0.158)
Maximum Benefit For AFDC+FS Family of 3	0.001	0.001	0.001	0.001	0.001	0.001	0.001
	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
State Provides AFDC-UP	0.074	0.062	0.061	0.050	0.046	0.052	0.083
	(0.016)	(0.017)	(0.017)	(0.013)	(0.015)	(0.014)	(0.018)
State*Time Trends	yes	yes	yes	yes	yes	yes	yes

Table 5.a

Additional Estimates of Minimum Wage Effects on Welfare Caseloads Weighted and with Quartic Time Trend

	Add Waivers	Add % Immigrants	Add Political Variables	Alternative Wage Measures	Blank Model & Log(MAXBEN)	Blank Model Prod. Wages & Log(MAXBEN)
	(2)	(1)	(3)	(4)	(5)	(6)
Minimum Wage	0.230	0.151	0.153	0.154		0.191
	(0.045)	(0.074)	(0.054)	(0.047)		(0.058)
Log (Average Production Wage)	0.109	0.032	-0.194			-0.361
	(0.257)	(0.246)	(0.204)			(0.202)
Log(10th Wage percentile)				0.179		
				(0.089)		
Log(25th Wage percentile)				-0.136		
				(0.071)		
Unemployment Rate	0.004	0.004	0.005	0.004		0.009
	(0.004)	(0.004)	(0.004)	(0.004)		(0.004)
Unemployment Rate t-1	0.005	0.005	0.007	0.005		0.011
	(0.003)	(0.003)	(0.003)	(0.003)		(0.003)
Unemployment Rate t-2	0.017	0.017	0.014	0.018		0.015
	(0.004)	(0.004)	(0.004)	(0.004)		(0.004)
Log(Gross State Product)	-0.584	-0.611	-0.347	-0.587		
	(0.113)	(0.112)	(0.126)	(0.105)		
Log(Population)	-1.819	-1.778	-1.569	-1.672		-1.669
	(0.357)	(0.369)	(0.406)	(0.344)		(0.402)
Fraction of Population White	0.326	0.492	-12.059	0.612		-14.829
	(2.347)	(2.572)	(3.264)	(2.435)		(3.266)
Fraction of Households Female Headed	0.656	0.579	0.572	0.549		0.633
	(0.198)	(0.203)	(0.208)	(0.207)		(0.212)
Maximum Benefit For AFDC+FS Family of 3	0.001	0.001	0.001	0.001		0.712
	(0.0002)	(0.0002)	(0.0002)	(0.0002)		(0.118)
State Provides AFDC-UP	0.090	0.094	0.055	0.086		0.057
	(0.022)	(0.024)	(0.017)	(0.024)		(0.017)
High Age	0.001					0.002
	(0.002)					(0.001)
Max Inc	-0.0001					-0.0001
	(0.0001)					(0.0001)
% Immig		0.202				0.015
		(0.492)				(0.397)
% Immig at t-1		2.775				2.302
		(1.294)				(1.224)
State has Work Requirement	0.013					-0.024
	(0.014)					(0.018)
State has Work Incentive	0.001					0.023
	(0.017)					(0.026)
State has Time Limits	0.033					0.073
	(0.024)					(0.027)
State has Child Support	-0.066					-0.046
	(0.017)					(0.020)
Democratic Control of Both Houses & Gvshp*100			-0.319			-0.498
			(1.078)			(1.084)
Divided Control			-0.021			-0.023
			(0.010)			(0.010)
(Debt/Income)*100			0.142			0.130
			(0.032)			(0.029)
No Line Item Veto*100			0.0467			1.376
			(2.346)			(2.797)
% Upper Democrats*100			0.0097			0.0065
			(0.0488)			(0.0471)
% Lower Democrats*100			0.002			0.001
			(0.001)			(0.001)
State*Time Trends	yes	yes	yes	yes	yes	yes

Table 5.b

Additional Estimates of Minimum Wage Effects on Welfare Caseloads
Unweighted and with Quartic Time Trend

	Add Waivers	Add % Immigrants	Add Political Variables	Alternative Wage Measures	Blank Model & Log(MAXBEN)	Blank Model Prod. Wages & Log(MAXBEN)
	(2)	(1)	(3)	(4)	(5)	(6)
Minimum Wage	0.218 (0.061)	0.178 (0.067)	0.239 (0.066)	0.132 (0.063)	0.266 (0.062)	0.255 (0.070)
Log (Average Production Wage)	-0.245 (0.238)	-0.289 (0.247)	-0.367 (0.252)			-0.405 (0.229)
Log(10th Wage percentile)				0.002 (0.127)	-0.144 (0.115)	
Log(25th Wage percentile)				-0.058 (0.086)	-0.136 (0.092)	
Unemployment Rate	0.007 (0.004)	0.008 (0.004)	0.008 (0.004)	0.005 (0.004)	0.012 (0.004)	0.013 (0.004)
Unemployment Rate t-1	0.012 (0.003)	0.012 (0.003)	0.012 (0.002)	0.011 (0.003)	0.011 (0.002)	0.012 (0.003)
Unemployment Rate t-2	0.010 (0.003)	0.011 (0.003)	0.008 (0.003)	0.012 (0.004)	0.008 (0.003)	0.009 (0.003)
Log(Gross State Product)	-0.532 (0.100)	-0.509 (0.100)	-0.331 (0.134)	-0.529 (0.108)		
Log(Population)	-1.561 (0.503)	-1.448 (0.485)	-1.498 (0.534)	-1.427 (0.561)	-1.515 (0.596)	-1.595 (0.551)
Fraction of Population White	-2.167 (1.951)	-2.595 (1.952)	-6.467 (3.573)	-2.479 (1.868)	-7.783 (3.489)	-7.308 (3.651)
Fraction of Households Female Headed	0.181 (0.137)	0.201 (0.161)	0.228 (0.168)	0.391 (0.217)	0.265 (0.158)	0.293 (0.152)
Maximum Benefit For AFDC+FS Family of 3	0.001 (0.0002)	0.001 (0.0002)	0.001 (0.0002)	0.001 (0.0002)	0.572 (0.147)	0.618 (0.150)
State Provides AFDC-UP	0.057 (0.017)	0.061 (0.017)	0.064 (0.019)	0.057 (0.020)	0.065 (0.015)	0.062 (0.018)
High Age	0.002 (0.001)				0.003 (0.002)	0.003 (0.002)
Max Inc	0.0001 (0.0002)				0.00005 (0.0002)	0.0001 (0.0002)
% Immig		-0.419 (0.557)			0.006 (0.536)	-0.099 (0.515)
% Immig at t-1		2.402 (1.909)			0.910 (1.732)	1.656 (1.571)
State has Work Requirement	0.001 (0.017)				-0.045 (0.019)	-0.039 (0.019)
State has Work Incentive	-0.010 (0.017)				0.009 (0.023)	0.003 (0.023)
State has Time Limits	0.032 (0.021)				0.063 (0.024)	0.062 (0.024)
State has Child Support	-0.055 (0.016)				-0.038 (0.015)	-0.035 (0.015)
Democratic Control of Both Houses & Gvshp*100			0.772 (1.366)		0.810 (1.330)	0.353 (1.359)
Divided Control*100			-0.265 (1.181)		-0.627 (1.101)	-0.829 (1.153)
(Debt/Income)*100			0.110 (0.036)		0.124 (0.0346)	0.129 (0.035)
No Line Item Veto*100			-0.0907 (3.448)		0.555 (3.501)	1.403 (3.460)
% Upper Democrats*100			-0.00021 (0.0500)		0.00725 (0.0481)	-0.0075 (0.0493)
% Lower Democrats*100			0.145 (0.063)		0.097 (0.060)	0.126 (0.063)
State*Time Trends	yes	yes	yes	yes	yes	yes

[1] This number was calculated using data from the 1996 Current Population Survey Monthly Outgoing Rotation Groups. 13.7% of unmarried female household heads with children earned between \$4.25 and \$5.15 per hour. An additional 3.1% report wages lower than the 1996 federal minimum wage of \$4.25.

[2] Calculated from the March 1993 Current Population Survey.

[3] Blank (1989) finds that one third of AFDC spells begin when a female head experiences a decline in earnings and about one fifth of spells end because of earnings increases. Bane and Ellwood (1983), find that 12% of spell beginnings can be attributed to earnings declines and about one third of exits can be attributed towards earnings increases. In more recent work (Bane and Ellwood, 1994) these numbers are revised downward to 7% and 25%.

[4] This number may result in part from measurement error.

[5] It is possible, however, that increases in the minimum wage indirectly increase wages in the uncovered (or underground) sector because employers in the uncovered sector must compete for employees who now can seek minimum wage jobs.

[6] While measurement error in the dependent variable will not bias coefficient estimates, it does reduce the precision with which they are estimated.

[7] It is not clear that we would want to use a Kaitz index as our measure of minimum wages even if information on coverage were available. The Kaitz index confounds the effect of changes in minimum wage levels with changes in the percent of workers who are covered. We would, however, like to use information on the percent of workers covered to check that our minimum wage estimates are robust by including in our regression an interaction between *percent covered* and MW_{st}. In states with more covered workers, minimum wage effects should be bigger.

[8] See Brown, Gilroy and Kohen (1983), Gramlich (1976), Lovell (1972), Moore (1971), and Wachter and Kim (1982).

[9] Wellington (1991) provides an explanation for this asymmetry, suggested by Gary Solon: if some industries and establishments are uncovered but compete with the covered sector (and, therefore, pay close to the minimum wage) then, when coverage is expanded into these industries, we would not expect to find that an increase in coverage will have a large impact on the employment of low wage workers. However, an increase in the minimum wage level will affect both the covered sector and the uncovered sector.

[10] We use the CPS Outgoing Rotation Group files to estimate the fraction of households headed by a single mother. Prior 1990, the ORG data do not provide enough demographic information to allow us to distinguish between children and other related persons in the household. For these years, we follow Blank (1997), and use the fraction of households headed by a single woman and including other related persons under age 18 in the household as a proxy for single mother households.

[11] We were unable to include the District of Columbia in our analysis because we were unable to obtain a measure of production wages for the district.

[12] For example, see Baker, Benjamin and Stanger (1996), and Card, Katz and Krueger (1994) and Neumark and Wascher (1992).

[13] David Ellwood has pointed out to us that raw correlations between the federal minimum wage and the size of the national welfare caseload are much larger between 1980 and 1994 than in previous years. If our estimates were primarily identified from time series variation in the federal minimum wage, then our choice of years would surely be suspect. Since we include time dummies (or quartic time trends) in our analysis, virtually all of our identification is coming from cross-sectional variation in state minimum wage changes over time.

[14] The inclusion of year dummies makes this superfluous, but we believe that it aids in the interpretation of our estimates.

[15] See Baker, Benjamin, and Stanger (1999) for a thorough discussion of the importance of this distinction.

[16] To see this consider the following: suppose that the unobserved state component in year *t* is $F_{st} = rF_{st-1}$, where $0 < r < 1$. Then when first differences are computed over a single year, not all of the unobserved state component will be swept away. In fact, $(r-1)F_{st-1}$ will remain. When differences are computed over longer time periods, even less of the unobserved state component is eliminated: $(r^T - 1)F_{st-T}$ remains. Because LSDV (Mean Differencing) uses differences computed over a longer time period, on average, less of the unobserved state component will be eliminated.

[17] Of the four specifications we present in this paper (see below), the only pair of point estimates that are not particularly close are those produced by the unweighted specification that includes year-effects. Using the first-differences estimation strategy produces a coefficient estimate of 0.032, whereas using the strategy that includes state-specific dummy variables produces an estimate of 0.124. The confidence intervals around each estimate include the other estimate.

[18] The Newey-West method produces consistent standard errors as N and T approach infinity; our sample, by comparison, is relatively small. However, OLS and White standard error estimates are very similar.

[19] Our model does not provide intuition regarding the appropriate number of autocovariances to be used in the autocorrelation weighting matrix, so we computed the Newey-West standard error estimates using values of m, the number of autocovariances, ranging from 4 to 8. Our results did not change when we used different values of m. The results reported in the paper set m=6.

[20] Work incentives might increase welfare caseloads if they include provisions such as raising the amount of income that AFDC recipients can earn before they begin to lose benefits.

[21] Some states implemented waivers that only affected small portions of the state. We identify the presence of a waiver only if it was implemented throughout the state.

[22] These data were obtained from the U.S. Department of Health and Human Services and are included in Ziliak, Figlio, Davis and Connolly (1997).

[23] This threshold varies within a state/year cell depending on the age of the child. We have chosen to use the maximum threshold for *any* child in the state/year cell.

[24] These data are reported in the INS's Statistical Yearbooks. The number of immigrants to each state is measured as the number of legally admitted immigrants who intend to reside in that state. These figures omit undocumented immigrants. Measurement error will arise from differences in the intended residence of an immigrant and the actual residence of that immigrant.