POVERTY AND MACROECONOMIC PERFORMANCE ACROSS SPACE, RACE, AND FAMILY STRUCTURE*

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We examined the effects of macroeconomic performance and social policy on the extent and depth of poverty in America using state-level panel data from the 1981–2000 waves of the Current Population Survey. We found that a strong macroeconomy at both the state and national levels reduced both the number of families who were living in poverty and the severity of poverty. The magnitude and source of these antipoverty effects, however, were not uniform across family structures and racial groups or necessarily over time. While gains in the eradication of poverty, in general, were tempered by rising wage inequality, simulations indicated that female-headed families and families that were headed by black persons experienced substantial reductions in poverty in the 1990s largely because of the growth in median wages. An auxiliary time-series analysis suggests that the expansions in the federal Earned Income Tax Credit of the 1990s accounted for upward of 50% of the reduction in after-tax income deprivation.

Changes in the economic landscape over the past two decades substantially altered the opportunities for economic progress among low-income Americans. The economy underwent the deepest recession since the Great Depression, followed by the longest and third-longest expansions in the post–World War II period. Tax and welfare policies also changed dramatically, ranging from the Tax Reform Act of 1986 to the Personal Responsibility and Work Opportunity Reconciliation Act (PRWORA) of 1996. Many in the policy and research communities have speculated about the effects of these changes on the financial well-being of low-income families, but heretofore the evidence has been scarce (Lichter and Crowley 2002; Meyer and Sullivan 2001). In this article, we aim to narrow the gap in the literature by using state-level panel data over the 1980s and 1990s to examine the collective impacts of macroeconomic performance and social policy on a broad measure of family well-being: the extent and depth of poverty.

The link between poverty and economic growth has garnered a great deal of research attention over the years (Aaron 1967; Anderson 1964; Blank 1993; Blank and Blinder 1986; Blank and Card 1993; Cain 1998; Cutler and Katz 1991; Freeman 2002; Gottschalk and Danziger 1985; Haveman and Schwabish 2000; Iceland 2003). This research, which has typically relied on aggregate time-series data, has established a strong, inverse association between economic growth and poverty rates in the 1960s and 1970s. However, this link was apparently weakened during the 1980s when aggregate poverty rates did not fall nearly as much as expected when the economy expanded. This was a troubling development because a large segment of the population was left anchored to the bottom during this much-heralded expansion.

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The tempering of the relationship between economic activity and poverty rates in the 1980s has been attributed to the substantial rise in income inequality (Blank and Card 1993; Iceland 2003).[1] Because official U.S. poverty lines are fixed in real terms, poverty rates are determined by both the shape and the location of the income distribution. All else being constant, a rising median reduces poverty, whereas rising inequality exacerbates poverty. Thus, in the 1980s, it appears that the latter force may have partially offset the former. Although rising income inequality likely attenuated the reduction in poverty in the 1980s, the aggregate link between poverty and the economy may have been spuriously weakened because the aggregate data mask important heterogeneity in the macroeconomy at the subnational level. As an example, whereas the country as a whole prospered in the late 1980s, certain regions experienced fairly severe recessions (e.g., the oil-producing Rocky Mountain states).

Much of the recent research on disadvantaged populations has focused on the effect of wholesale changes in social policies in the 1990s on various outcomes, such as welfare caseloads, labor-force behavior, consumption, and earnings (e.g., Meyer and Rosenbaum 2000; Meyer and Sullivan 2001; Moffitt 1999; Schoeni and Blank 2000; Ziliak et al. 2000; Ziliak, Gundersen, and Figlio 2003). These changes in social policies included significant federal and state expansions in the Earned Income Tax Credit (EITC). For example, from 1986 to 2000, the maximum EITC credit increased over sevenfold, from \$500 to \$3,800. Along with this expansion, states were granted waivers from federal rules governing the Aid to Families with Dependent Children (AFDC) program and, in 1996, there was a federal overhaul of welfare as part of PRWORA. Collectively, the intention of the EITC expansions and welfare reform was to make work more attractive than the receipt of welfare.

In this article, we contribute both to the literature on the relationship between macroeconomic performance and poverty and to the literature on welfare reform. The models we used in our study differed from the standard models in the poverty-macroeconomy literature in three key ways. First, in lieu of aggregate time-series data, we exploited the substantial heterogeneity in poverty and economic activity across states and over time.[2] Specifically, we assembled a 20-year panel of states from the 1981–2000 waves of the Current Population Survey (CPS) which, coupled with labor-force data from the Bureau of Labor Statistics, improved our ability to identify the effects of the economy on poverty compared to time-series data. Second, rather than restrict our focus to only the poverty rate (i.e., the fraction of the population with incomes below the poverty threshold), we used a poverty index that permitted us to portray not only the extent of poverty but also the depth of poverty via the so-called squared poverty gap. Although the head count is an easily interpretable measure of poverty status, its use as a measure of impoverishment is limited because it treats all persons identically, whether their incomes are \$1 or \$7,000 below the poverty line. The squared poverty gap does not suffer from this limitation, and by using this measure, we could address the benefits of economic growth for those who are the lowest in the income distribution.

Third, the literature has typically defined poverty in terms of before-tax income. However, substantial changes in both federal and state income tax systems over the past two decades, notably expansions of the federal and state EITCs, suggest that after-tax income may paint a different portrait of poverty. The EITC was expanded at the federal level in 1986, 1990, and 1993 such that by 1996, it reached nearly the fourth decile of

1. We do not mean to imply that economic growth always is accompanied by increases in inequality. The twentieth century also witnessed periods of wage compression (Goldin and Margo 1992).

2. Others who used less-aggregated data to examine the issue of poverty and economic activity include Blank and Card (1993), who used data for the 1970s and 1980s by census region, and Freeman (2001), who used state-level data for 1989–1998, although the state-level analysis is not the primary focus of his paper.

the married-couple income distribution and approached 150% of the median for female-headed families (Ventry 2000). To capture these changes in the tax system, we used both gross and net-of-tax income measures.

Aside from a more precise identification of the business cycle, an additional advantage of state-level panel data is that they permit identification of the effect of state-level policy reforms on poverty, such as welfare reform, state-level EITC expansions, and state-level minimum-wage expansions. While the primary aim of welfare policy makers was to reduce the dependence on welfare and nonmarital fertility, presumably policy makers also wanted to reduce poverty. To ascertain how well this intention was met, we exploited the differential timing of the waivers to AFDC policies and the implementation dates under the new rules in PRWORA.[3] Alongside the state-level changes to the welfare system, by the end of the 1990s, nearly 20% of the states had refundable credits on top of the federal EITC. Likewise, the states' minimum wages increasingly deviated from the federal level, perhaps because the latter was fixed in nominal terms for much of the 1980s, such that upward of 30% of the states had minimum wages that were higher than the federal minimum over the past two decades.

Unlike most studies (for an exception, see Iceland 2003), we also conducted our analysis separately by family structure and race. This separation seems especially relevant insofar as the extent of poverty across types of households differs widely in the United States, and there is no reason to expect the macroeconomy to have the same effect on all categories. Female-headed families are of particular interest because of their high levels of poverty and because they have been the focus of the many recent changes in social policy. Likewise, families that are headed by black persons are of interest relative to those headed by white persons because several studies have found differential effects of the business cycle on the labor-market outcomes of blacks (Bound and Holzer 1993; Clark and Summers 1979; Ziliak, Wilson, and Stone 1999).

Our results indicate that a strong macroeconomy at both the state and national levels reduces not only the number of families with incomes below the poverty line, but also the severity of poverty. The magnitude and source of these antipoverty effects, however, are not uniform across family structures and racial groups or necessarily over time. Growth in the median wage is critical for progress against poverty among female-headed and black families. Overall, the gains in the eradication of poverty are tempered by rising wage inequality, although the extent of this offset is less pronounced among female-headed families and black families. We found limited evidence that (after-tax) poverty was lower among female-headed families and black families after the implementation of state-specific welfare reforms, both before and after passage of PRWORA in 1996. An auxiliary time-series analysis suggested that the expansions in the federal EITC in the 1990s accounted for upward of 50% of the reduction in after-tax income deprivation.

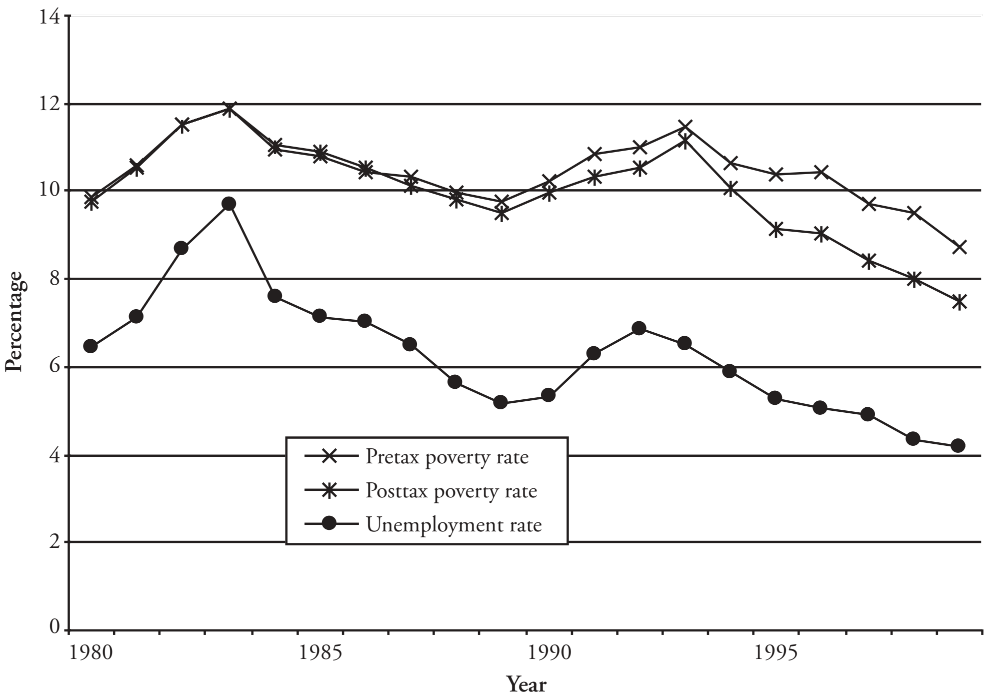
BACKGROUND AND EMPIRICAL MODEL

The adage "A rising tide lifts all boats" suggests that poverty should be countercyclical insofar as economic growth should reduce poverty. Evidence of such a countercyclical relationship seems apparent in Figure 1.[4] Family-level poverty (both before and after tax) surged in the early 1980s during the deep recession and then fell with the long expansion in the mid- and late 1980s. The unemployment rate rose with the 1990–1991 recession, peaking in 1992, while poverty continued its upward climb for another year.

3. Schoeni and Blank (2000) examined the impact of welfare reform on poverty among female-headed families, but their study did not focus on the broader links of macroeconomic performance and poverty. Meyer and Sullivan (2001) examined the impact of broad-based changes in taxes and social policy on the consumption expenditures among single mothers.

4. All data used in the ensuing figures are from the CPS, as is detailed later.

Figure 1. Poverty Rates and Unemployment Rates

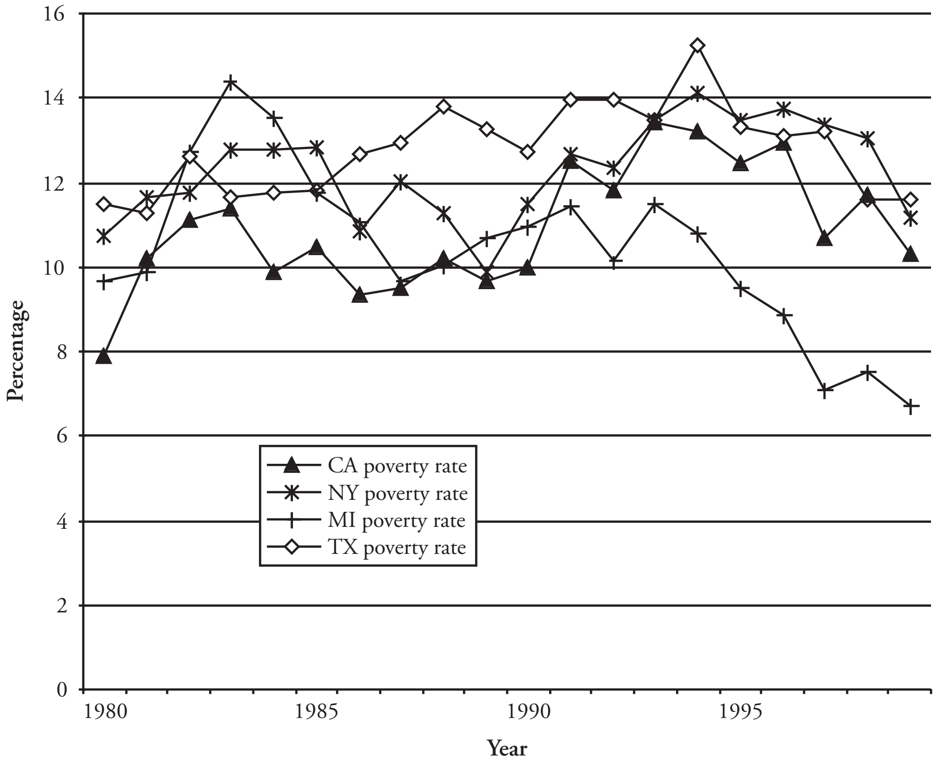


In the 1990s, the reduction in poverty paralleled the long decline in unemployment rates. Until the Tax Reform Act of 1986 (TRA86), poverty rates before and after taxes were coincident; however, after TRA86 was passed, and more dramatically after 1993, after-tax poverty rates diverged from their before-tax counterparts, suggesting a strong anti-poverty role for the EITC.[5]

The changes in poverty and unemployment rates in Figure 1, while instructive, mask important heterogeneity at the subnational level, heterogeneity that is likely to enable accurate identification of the relationship between poverty and the economy. For example, over the past two decades, average poverty rates have ranged from a high of 18.7% in Mississippi to a low of 5.3% in New Jersey, and average unemployment rates have ranged from 10.4% in West Virginia to 3.7% in Nebraska. More important for the purposes of our models, the differences in poverty and macroeconomic performance across states are not invariant over time. We highlight these trends in Figures 2 and 3. The figures depict the 20-year time series of poverty rates and unemployment rates of four large states from different regions of the country—California, Michigan, New York, and Texas. Figure 2 shows that Michigan experienced a rise in poverty in the early 1980s recession, a recession that had a severe impact on the auto industry. Following this recession, there was a

5. In the official poverty rates compiled by the U.S. Census Bureau, a before-tax measure of income is used. Consistent with this practice, we use *poverty* to refer to measures that use a before-tax measure of income and *after-tax poverty* to refer to measures that use after-tax measures of income.

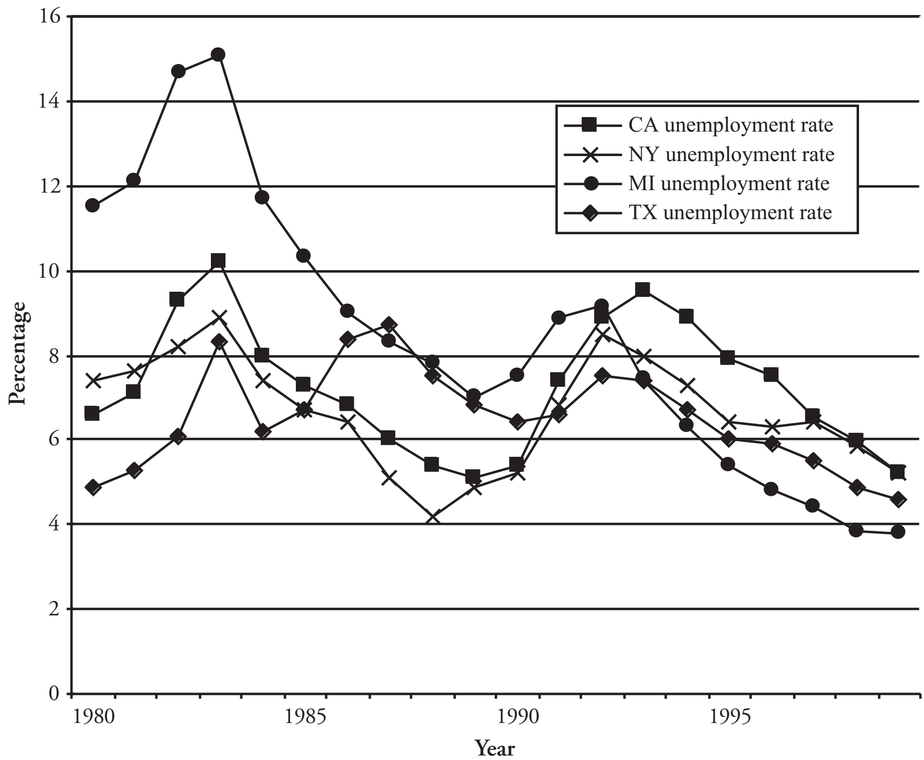
Figure 2. Spatial Differences in Poverty Rates



quick recovery in the mid-1980s and a further large decline in poverty in the mid-1990s. This decline seemed to track unemployment fairly well, as Figure 3 indicates. In contrast, Texas had a long secular increase in poverty throughout the 1980s until the mid-1990s. During this period, poverty appears not to have tracked unemployment as closely as in Michigan. Relative to the other states, California took an additional year to recover from the recession of the early 1990s, and, like New York, its poverty rate remained fairly high until the end of the decade.

Concomitant with the large changes in macroeconomic performance were dramatic changes in social policies, especially welfare reform. These changes have led many advocates for the poor to conclude that there may have been a sharp increase in poverty especially among female-headed families and families headed by black persons because both groups have had higher rates of participation in assistance programs. In fact, though, poverty rates across these groups actually declined in the mid- and late 1990s (see Figure 4). The declines in poverty rates were especially large for female-headed families and families headed by black persons. In 1992, 46.7% of the female-headed families were poor, but by 1999, this proportion had fallen over a quarter to 35.0%. For families headed by black persons, poverty rates fell from 29.7% to 20.0%, about a one-third decline. What is important is that federal tax policy reforms of the 1980s and 1990s

Figure 3. Spatial Differences in Unemployment Rates



reduced after-tax poverty rates among female heads of families more than among other groups, both in terms of timing and scope, highlighting the increasing role of the tax system as a stabilization and redistribution tool relative to income transfers (Kniesner and Ziliak 2002).

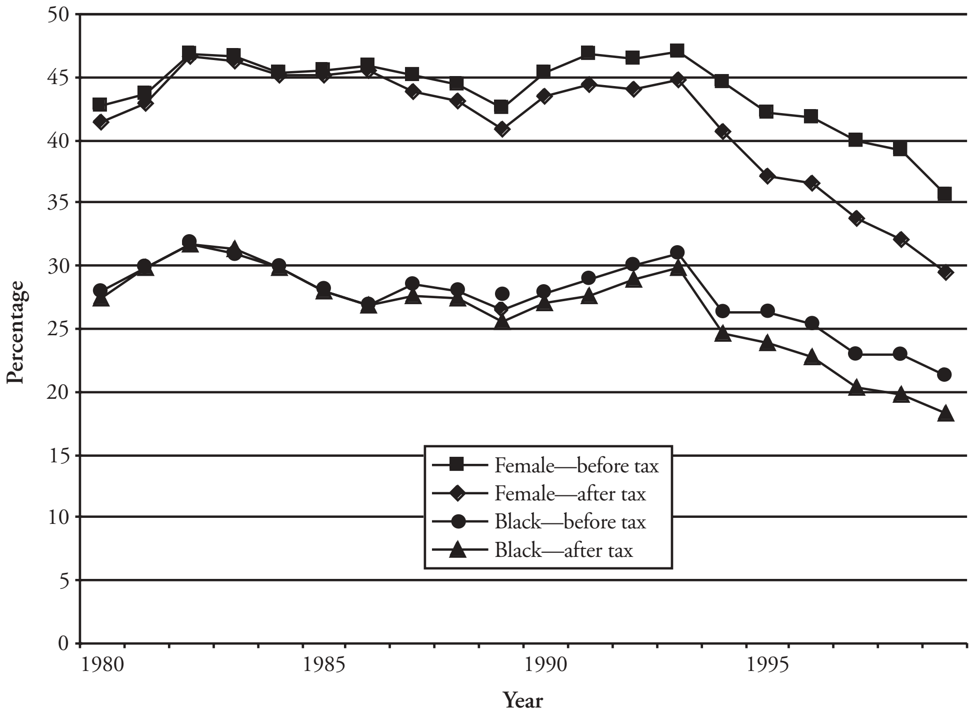
Empirical Model

Our objective was to relate a state-specific and time-varying measure of poverty to state-specific and time-varying indicators of macroeconomic performance and policy changes and simultaneously to control for other, less readily quantifiable, factors that affect poverty. This objective led us to the baseline econometric model for group *j* (*j* = all families, female-headed families, married-couple families, white families, black families) in state *s* (*s* = 1, . . . , 51) in time *t* (*t* = 1, . . . , 20):

$$\ln\left(P_{\alpha,t}^{j,s}\right) = \rho^j \ln\left(P_{\alpha,t-1}^{j,s}\right) + \sum_{k=1}^{K} \beta_k^j E_{k,t}^s + \sum_{m=1}^{M} \gamma_m^j R_{m,t}^s + \lambda_t^j + \mu^{j,s} + \varphi^{j,s} t + \varepsilon_t^{j,s}, \tag{1}$$

where P_α is the poverty measure, E_k reflects macroeconomic indicator k, R_m reflects public policy indicator m, λ_t captures aggregate time effects, μ^s and $\varphi^s t$ capture unobserved fixed and trending factors that are state specific, and ε_t^s is a random error term.

Figure 4. Poverty Rates for Female-Headed Families and Black Families



Of central import to this study is the measurement of poverty (i.e., the choice of α). In response to the deficiencies associated with the head-count measure, economists have constructed numerous axiomatically derived poverty measures (see, e.g., Atkinson 1987; Foster, Greer, and Thorbecke 1984; Sen 1976). Following Foster et al. (1984), consider the following class of poverty indices for a given group *j*:

$$P_{\alpha t}^{s} = \frac{1}{n_t^s} \sum_{q=1}^{Q_t^s} \left(\frac{z_t - y_t^q}{z_t} \right)^{\alpha}, \tag{2}$$

where n is the population, Q is the number of poor families, z is the family-size specific poverty threshold, and y is income. The index α, $0 \leq \alpha \leq \infty$, is known as the "poverty-aversion" index. As α increases, increasing weight is given to the poorest households in the state at time t. When $\alpha \rightarrow \infty$, all weight is given to the poorest family.

We considered two variants of the poverty-aversion index: $\alpha = 0$, which yields the head-count ratio, more commonly known as the poverty rate, and $\alpha = 2$, which is often referred to as the squared poverty gap.[6] The squared poverty-gap measure satisfies both

6. Strictly, when $\alpha = 2$, it is the sum of squared normalized poverty gaps, not the square of the poverty gap. In model estimation, one takes the natural log of each poverty index, which aids in reducing the influence of outliers and permits the interpretation of some of the coefficients as elasticities.

the monotonicity axiom (all else being equal, a reduction in the income of a poor family must increase the poverty measure) and the transfer axiom (all else being equal, a pure transfer of income from a poor family to any other less-poor family must raise the poverty measure). The head-count measure satisfies neither axiom; this is one of the reasons for our use of the squared poverty gap.[7]

Aside from satisfying certain axiomatic criteria, with the squared poverty-gap measure, we could identify the effect of the macroeconomy on the depth of poverty. For example, it is possible that economic growth "trickles down" to low-income families with incomes that are at or just below the poverty threshold, but has little effect on the very poor. Moreover, because the squared poverty gap assigns weights to families, depending on their income position (higher weights to lower-income families), we were able to assess a frequent criticism levied against welfare reform (Primus et al. 1999). To these critics, welfare reform was successful in moving those with incomes that were close to the poverty line out of poverty but had a much smaller impact on worse-off families. Insofar as states were assigned grant allocations that were based on the number of people who left welfare for work, they may have concentrated on members of households with incomes that were closer to the poverty line who are more likely to be job ready. The poverty rate cannot speak to this issue; the squared poverty gap can.

Turning now to the specification in Eq. (1), we included a lag in the dependent variable because poverty tends to persist at the household level (Sawhill 1988; Stevens 1999), and thus more aggregated measures are likely to respond sluggishly to changing economic conditions. A further advantage of the dynamic specification is its ability to distinguish between the short-run (i.e., β_k) and the long-run (i.e., β_k / (1 – ρ)) effects of the macroeconomy and social policies on poverty.[8] To portray the state of the macroeconomy, we used two common measures of the business cycle: state-level unemployment rates and per capita rates of growth in employment (hereafter employment growth rates). Both measures capture important features of labor-market opportunities among the disadvantaged. However, employment growth rates likely indicate the evolution of job vacancies over the business cycle and may better capture demand-side shocks to the labor market (Bartik 1991; Blanchard and Katz 1992). Its portrayal of demand-side shocks is one reason why the business-cycle dating committee at the National Bureau of Economic Research uses it as the primary indicator of economic activity (Hall et al. 2003).[9]

The unemployment rate and employment growth rate control for only a subset of state economic activity, so we also considered the impacts of state median wages and state wage inequality. These measures of inequality likely capture additional cyclical, as well as secular, changes in the economy. To portray these relationships, we used a quadratic in median wages and the ratio of the 80th to the 20th percentiles of wages.[10] We

7. In earlier versions, we also considered α = 1, which yields the so-called poverty gap. This measure satisfies the monotonicity axiom but not the transfer axiom. Because the results were qualitatively the same as those for the squared poverty gap shown later and because the transfer axiom was violated, we omitted our tabulation of the poverty-gap estimates for ease of presentation.

8. The dynamic fixed-effect model may suffer from the so-called Nickell (1981) bias; that is, the correlation between the lagged dependent variable and the fixed effects may bias the coefficient on the lagged dependent variable toward zero. However, this bias declines as the time-series dimension grows, and with T = 20, this bias is likely to be minimal. As a check, though, we estimated the dynamic model via instrumental variables, using a variety of instruments and techniques (i.e., fixed-effects instrumental variables and first-differences instrumental variables) with little change in the results.

9. We also considered other measures of macroeconomic activity, including the growth rate in the real gross state product (RGSP). After controlling for the unemployment rate and employment growth per capita, we found that RGSP was statistically and economically insignificant and thus dropped it from the analysis.

10. Instead of the 80–20 measure of inequality, we also ran models with the coefficient of variation with little change in the estimates. See Cowell (2000) for a detailed treatment of the measurement of inequality; Katz

used wages, rather than earnings, to avoid possible endogeneity with our poverty measures, since poverty rates and gaps are nonlinear transformations of earnings and nonlabor-market income.[11]

The first policy variable of our empirical model was the measurement of welfare reform. States began experimenting with their welfare programs in the early 1990s via waivers from federal regulations granted by the U.S. Department of Health and Human Services (DHSS). These waivers included time limits on the receipt of benefits; work requirements; and work incentives, such as higher earnings disregards and liquid-asset limits. The waivers were codified into federal legislation with the passage of PRWORA, which eliminated the AFDC program and replaced it with a state block-grant program known as Temporary Assistance for Needy Families (TANF). Under PRWORA, cash assistance is no longer an entitlement, and aid is subject to a federal lifetime limit of 60 months (or less, depending on a state's discretion). For our purposes, we delineated the welfare-reform era into two periods: "waivers" (1992–1996) and PRWORA (1996–1999). Moreover, for parsimony, we aggregated the state waiver programs into a single "any-waiver" indicator reflecting the fraction of the year that the program was in operation. On the basis of program-implementation dates assembled by DHHS, we could identify the effect of waivers by exploiting the fact that some states did not receive waivers, and those that did receive waivers implemented them at different times. Likewise, states implemented their TANF programs at different times over a two-year period from the fall of 1996 to the summer of 1998.[12] As with the macroeconomy, poverty may respond sluggishly to the implementation of welfare reform, so our dynamic model admits this possible protracted response.[13]

Before welfare waivers were implemented, the primary policy lever that states had was the level of AFDC benefits. The effect of welfare benefits on poverty, however, is ambiguous ex ante. All else being equal, a higher benefit level mechanically raises income and, as such, may move some recipients over the poverty threshold and would move all recipients higher in the income distribution. However, more generous transfers may have a negative behavioral effect on the supply of labor, thereby lowering earnings and reducing the antipoverty effectiveness of the benefits. To control for the consumption floor offered by welfare, we included the natural log of the sum of the maximum AFDC benefit level and the resulting allotment of food stamps for a three-person family. The coefficient in the regression model captured the net effect of these countervailing forces.

In the 1980s and 1990s, states increased their use of other programs to combat poverty. One such program is a state EITC program, which in most states is proportional to the federal credit (Dickert-Conlin and Houser 2002). We controlled for the state EITC by specifying the variable as the log difference of the state's maximum benefit level and the federal benefit level (equal to 0 for states with no supplemental EITC). Much like cash

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and Autor (1999) for an application of the 80–20 measure to U.S. earnings inequality; and Wu, Perloff, and Golan (2002) for a study of governmental policies on alternative measures of the income distribution.

11. To calculate the wage variable, we divided total annual earnings by the usual number of hours worked times the number of weeks worked in the past year.

12. We also considered models with disaggregated waiver variables, with little change in the key results. Some of these waiver policies, such as time limits on benefits or work requirements for the receipt of benefits, have obvious implications for current poverty, while others, such as responsibility-based waivers requiring that children receive regular health examinations, are less obvious. In results not presented, we found that time limits and earnings disregards decreased poverty, whereas work requirements increased poverty. Our specification in Eq. (1) captures the net effect of these different policies.

13. Bitler, Gelbach, and Hoynes (2003) argued that in models with unrestricted time effects, such as in Eq. (1), the PRWORA effect is identified but is best interpreted as a 1997 effect, since the bulk of states implemented welfare reform in that year. The waiver period in the mid-1990s offers a more robust identification of welfare reform's effects because of greater cross-state heterogeneity in the timing of the programs' implementation.

assistance, the effect of the EITC on after-tax poverty is ambiguous in theory. The EITC clearly raises after-tax income, when before-tax income is held constant, but because the labor supply is predicted to fall over certain ranges (the so-called stationary range and the phase-out range) in the presence of the credit, before-tax poverty may increase. Indeed, using various models, periods, and samples, previous work has found mixed results for the EITC as it pertains to changes in income and the supply of labor, two factors that affect the extent and depth of poverty. Some researchers have found positive impacts of the EITC on the supply of labor and earnings, whereas others have found no impact or, in some cases, negative effects (Eissa and Liebman 1996; Meyer and Rosenbaum 2000; Neumark and Wascher 2001). Our model permits a direct assessment of state EITCs on before- and after-tax poverty rates and gaps.

Our final policy variable was whether states chose to implement a minimum wage in excess of the federal minimum wage. If they did, we used the log difference between the state and federal minimums. The minimum wage has two possible effects on poverty. In states with relatively higher minimum wages, the incomes of employed workers may be higher, thereby lowering the extent of poverty. Conversely, states with higher minimum wages may have fewer employed workers because of the higher wage floor. The contentious empirical and theoretical debate over whether either or both of these effects is important is an ongoing one (e.g., Card and Krueger 1995; Neumark and Wascher 2001, 2002). Another possible role for the minimum wage is its moderating effect on the extent of inequality. Lee (1999) found that the falling real minimum wage was an important contributor to the rise of inequality in the 1980s. So, insofar as increasing inequality leads to increasing poverty, the minimum wage may also affect poverty through this indirect route. However, because we controlled for inequality in our model, we were able to identify the direct effect of the state minimum wage on poverty.

Identification Issues

In Eq. (1), we assumed that ε_t^s is an independent identically distributed (IID) random error, and that with the inclusion of the lagged dependent variable on the right-hand side, the model is "dynamically complete" in the sense that further lags are redundant (Wooldridge 2002: chap. 13.8).[14] Conditional on the observed regressors, aggregate time effects (λ_t), state fixed effects (μ^s), and state trends ($\varphi^s t$), the assumption of an IID random error implies that (1) economic conditions vary exogenously across states and (2) there are no remaining systematic differences among states' policy choices such that reforms to welfare can be viewed as exogenous to a given state's poverty rate. Under these assumptions, ordinary least squares (OLS) provides consistent estimates of the model parameters in Eq. (1).

In their study of the evolution of state labor markets, Blanchard and Katz (1992) made a strong case in favor of the first assumption (just presented). They characterized states as producing different bundles of goods for sale on the national market under constant returns to scale technology with infinite long-run mobility of workers and firms. Because of product differentiation, states face different shocks to the demand for labor and thus state-specific fluctuations. Under these assumptions and controls for permanent and trending differences across states, our model permits the identification of indicators of states' business cycles, such as growth rates in unemployment and employment on poverty.

Some researchers have suggested that states' applications for waivers from federal rules governing cash assistance programs were an endogenous function of the states'

14. In results not presented, we tested the assumption of dynamic completeness and found no evidence in favor of including further lags of the dependent variable or of including lags in economic conditions or policy variables. This finding is at odds with some of the recent literature on welfare caseloads (e.g., Ziliak et al. 2000) and highlights the fact that poverty and welfare caseloads respond differently to the business cycle.

caseloads (see, e.g., Martini and Wiseman 1997). If this possibility is correct, and if it also applied to concerns over states' poverty rates, then the second assumption may not be valid. To our knowledge, no one has conducted a complete statistical analysis of this potential endogeneity. However, a limited set of tests in Ziliak et al. (2000) cast doubt on the endogeneity of waivers to AFDC caseloads. In the year leading up to the application for a waiver, Ziliak et al. found no statistical difference in the caseloads of states that applied for waivers and states that did not. Likewise, in a study on the effects of welfare reform on household savings, Hurst and Ziliak (2002) found no evidence that states that expanded the limits of liquid assets for program eligibility had systematically higher or lower saving rates prior to welfare reform.[15]

Role of Aggregate Economic and Policy Effects

We have been relatively silent on the role of aggregate economic activity and aggregate policy reforms on state-level poverty. The vector of time indicators (λ_t) is intended to control for the aggregate influences that are common to all states and households in a given period but that vary over time. Examples include shocks to aggregate labor markets by innovations in workplace technology; shocks to aggregate growth, perhaps emanating from oil-price shocks; and national reforms of the tax and transfer systems. Most prominent among the latter are the expansions of the federal EITC in the 1990s and the 1996 welfare reform. We attempted to capture some of these aggregate influences by means of an auxiliary regression.[16]

Specifically, let $\hat{\lambda}_t$ be the vector of estimated time effects. These effects can be viewed as an aggregate poverty rate (and squared poverty gap) net of idiosyncratic state differences. We then used this vector as the dependent variable in an auxiliary time-series regression on aggregate economic activity, trends, and policy reforms, or

$$\hat{\lambda} = \omega_1 AggUn_t + \omega_2 \%\Delta RGDP_t + \omega_3 t + \omega_4\left(t \times p90\right) + \omega_5\left(t \times p93\right) + \omega_4\left(t \times p96\right) + u_t, \quad (3)$$

where $AggUn_t$ represents the aggregate unemployment rate in a given year, $\%\Delta RGDP_t$ represents the growth rate in the real gross domestic product (GDP), t reflects an aggregate trend in poverty, $p90$ is an indicator that equals 1 for the 1990–1992 period (inclusive), $p93$ is an indicator for the 1993–1995 period, $p96$ is an indicator for the 1996–1999 period, and u_t is a random error that is heteroskedastic by construction (Saxonhouse 1976).

The specification in Eq. (3) portrays the effect of the aggregate business cycle and aggregate trends on poverty. Since they are common in the time-series literature on poverty, we included the unemployment rate and growth rate in the real GDP as proxies for the business cycle. To capture reforms of national policies, we permitted breaks in the aggregate trend. Specifically, the breaks were intended to capture the 1990 expansion in the federal EITC ($p90$), the phase-in period of the 1993 expansion ($p93$), and the period in which the EITC was fully phased in ($p96$). To the extent that there is an aggregate effect of welfare reform on state-level poverty (e.g., via national media coverage and statements by national politicians), interpretation of the latter two breaks in the trend is confounded because the welfare-waiver period overlapped with the phase-in period of the 1993 expansion, and PRWORA overlapped with the post-1996 period. However, because

15. In results not presented, we estimated a model that treated the welfare waiver as endogenous to poverty rates. For instruments, we used time-varying state political factors, such as whether both houses in a state legislature were under Democrat or Republican control, and an indicator of whether the governor was a Democrat. Although these variables were good predictors of waiver status, the results for these instrumental variables did not differ significantly from the OLS results presented here.

16. Solon, Barsky, and Parker (1994) used a similar two-step model to identify the effect of the aggregate unemployment rate on the wages of men in the Panel Study of Income Dynamics.

we conducted our analysis with both before- and after-tax poverty, we could examine the difference in estimated coefficients across models to isolate the effect of the federal EITC on after-tax poverty. We recognize the pitfalls of identification off time-series data (which is why we used state panel data for our primary empirical analysis), but we believe that this two-step method is informative about the separate effects of local and national activity and policy on poverty. Interpretation of the aggregate effects, though, merits caution.

DATA

The data we used to estimate the impact of macroeconomic performance and welfare reform on poverty came primarily from the 1981–2000 waves (1980–1999 calendar years) of the CPS. For each wave, we uniquely identified each state, and within each state, we obtained data on the total income and wages for all families, as well as by family structure (female-headed families with children or married-couple families with children) and by race of the family head (white and black).[17] The measure of before-tax income we used is the same as that used in the Census Bureau's official poverty measure, which includes labor-market earnings, governmental and nongovernmental cash transfers, and interest and dividends. In-kind transfers and capital gains are omitted from this definition. After-tax income is then calculated by netting out the family's federal income tax liability[18] and adding in the EITC.[19] Given gross and net family income, we constructed the poverty rate and the squared poverty gap. The average hourly wage for the family head was found by taking the ratio of labor-market earnings to annual hours of work for the sample of working persons. The poverty thresholds are adjusted annually by the Consumer Price Index–All Urban Consumers, and the wage data are deflated with the same price index.[20]

Although it is possible to construct state-specific unemployment rates and growth rates in per capita employment directly from the CPS, to improve measurement, we turned to the Local Area Unemployment Statistics, produced by the Bureau of Labor Statistics (BLS) for data on employment and unemployment, and to the Census Bureau for data on state populations.[21] For data on the labor force, the BLS uses information from the CPS, the Current Employment Statistics survey, and state-specific data on Unemployment Insurance claims to refine state-level estimates of unemployment and employment over those obtained from the CPS alone. For the data on state populations, the Census Bureau imputes annual population figures from the decennial census on the basis of births, deaths, and domestic and international migration.

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17. We focused on the family, which the CPS defines as a group of two or more persons who are related by birth, marriage, or adoption and who live together. We limited our sample to families with positive incomes, which is likely to reduce measurement error if some low-income households underreport income in the CPS, a finding by Edin and Lein (1997) in other contexts. Moreover, there is a discontinuity between households with negative and zero incomes and other households with incomes that are below the poverty line insofar as households with negative and zero incomes are much less likely to evidence correlates with poverty, such as participation in food-assistance programs (Wemmerus and Porter 1996).

18. To calculate federal income tax liability, the Census Bureau uses data from the American Housing Survey, the Income Survey Development Program, and the Internal Revenue Service to create simulations of taxes paid, number of tax-filing units, adjusted gross income, and other tax characteristics into the CPS. *The State Tax Handbook* from the Commerce Clearing House is also used as a source of information for data on taxes.

19. Meyer and Sullivan (2003) argued that the CPS captures only about 75% of EITC receipts, which suggests a likely understatement of the actual magnitude of the effect of the program on poverty. Moreover, take-up rates were increasing over the period of our study. With the assumption of a 100% take-up rate for the full period, our measured effect of the EITC is also diminished in the state-level analyses.

20. Poverty thresholds are available on-line at http://www.census.gov/hhes/poverty/threshld.html

21. Labor-force data from the BLS are available on-line at http://www.bls.gov/top20.html#LAUS, and data on state populations are available on-line at http://www.census.gov/population/www/estimates/statepop.html

There may be concerns about possible measurement error plaguing annual state-specific estimates of poverty, wages, and inequality in the CPS. This measurement error may have biased our estimates of Eq. (1), especially for subpopulations, such as female-headed or black families, whose sample sizes are limited in some smaller states. (This is less of an issue for the state-level employment, unemployment, and population figures because of the rigorous refining they receive from the BLS and Census Bureau.) To alleviate these concerns, we constructed three-year moving averages of all variables (except for the welfare-reform dummy variables), resulting in the loss of the two end points for each state's time series.[22] By using three-year moving averages, we used the same method used by the Census Bureau in its annual reports on the extent of poverty in the United States (Dalaker and Proctor 2000). In addition, for families headed by black persons, we limited our sample to state-years with more than 30 observations.

Our use of three-year moving averages necessitated a change to our lag structure of Eq. (1). Instead of the $(t-1)$ lagged dependent variable in Eq. (1), we had to use the $(t-2)$ value. To see this change, note that at time t, the three-year moving average for $P(t)$ is $\overline{P}_t = (P_{t+1} + P_t + P_{t-1})/3$; at time $(t-1)$, $\overline{P}_{t-1} = (P_t + P_{t-1} + P_{t-2})/3$; and at time $(t-2)$, $\overline{P}_{t-2} = (P_{t-1} + P_{t-2} + P_{t-3})/3$. Under the assumption that ε_t is IID, then both $\overline{P}_t$ and $\overline{P}_{t-1}$ are endogenous to ε_t (because of the presence of P_t), but $\overline{P}_{t-2}$ is only predetermined. Hence, using the lag dependent variable at $(t-2)$ still permits consistent estimation via OLS.[23]

RESULTS

We next investigated more formally the impact of macroeconomic performance and social policies on the (log) poverty rate and the (log) squared poverty gap using both before-tax and after-tax income. Each regression was based on three-year moving averages; was weighted by the number of families in each state-year-group cell; and controlled for year effects, state fixed effects, and state-specific trends. Moreover, each of the ensuing four tables contains two panels: an upper panel for the results of the primary model from Eq. (1), based on state panel data, and a lower panel for the results of the auxiliary model, based on the aggregate time series from Eq. (3).

The Macroeconomy, Social Policy, and the Extent of Poverty Across Families

Table 1 presents weighted OLS estimates of the models of before-tax poverty rates for all families and across family structure and race. With the exception of black families, the unemployment rate has a positive and statistically significant effect on before-tax poverty rates. For example, in the case of all families, a one-percentage-point decrease in the unemployment rate leads to a 4.5% decline in the short-run poverty rate; for female-headed families, the response is a 3.2% decline; and for married-couple families, it is a 5.6% decline. Because of the small coefficient on the lagged dependent variable across the poverty measures, the long-run effects are only slightly higher. Growth in employment per capita also demonstrates a strong countercyclical effect on poverty; across all families, a one-percentage-point increase in the growth rate of employment per capita leads to a 1.4% decline in poverty rates. Again, however, there are differences across family structure and race in the antipoverty effectiveness of rates of growth in employment insofar as poverty rates among female-headed families and black families are fairly nonresponsive to such growth.

22. In the descriptive analysis presented in Figures 1–4, we used annual data because the data were either aggregated across all states or came from relatively large states. Each state-year was weighted by the number of observations in the CPS for the respective cell.

23. Note that we took the log of the three-year average, not the average of the logs.

Table 1. Estimates of the Impact of Macroeconomic Performance and Social Policies on Before-Tax Poverty Rates

Variable	All Families	Female-Headed Families	Married-Couple Families	White Families	Black Families
Poverty (t – 2)	0.136	0.074	0.074	0.136	0.156
	(0.040)	(0.040)	(0.038)	(0.048)	(0.063)
Unemployment Rate	0.045	0.032	0.056	0.048	0.013
	(0.005)	(0.005)	(0.007)	(0.006)	(0.016)
Growth in Employment per Capita	–0.014	–0.003	–0.013	–0.012	–0.008
	(0.004)	(0.005)	(0.007)	(0.005)	(0.011)
Median Wage	–0.060	–0.174	0.024	–0.062	–0.335
	(0.053)	(0.043)	(0.076)	(0.058)	(0.056)
Median Wage, Squared	–0.000	0.006	–0.004	–0.000	0.012
	(0.002)	(0.002)	(0.002)	(0.002)	(0.003)
Ratio of 80th to 20th Wages	0.268	0.028	0.317	0.306	–0.025
	(0.032)	(0.012)	(0.053)	(0.040)	(0.011)
Log of State-Federal EITC	0.035	0.023	0.044	–0.005	0.052
	(0.017)	(0.015)	(0.030)	(0.021)	(0.030)
Log of State-Federal Minimum Wage	–0.027	–0.015	–0.030	–0.040	–0.056
	(0.011)	(0.011)	(0.025)	(0.010)	(0.025)
Pre-PRWORA Waiver	0.006	–0.022	0.067	0.020	0.000
	(0.014)	(0.018)	(0.027)	(0.020)	(0.035)
Post-PRWORA Waiver	0.009	–0.020	0.059	–0.010	–0.017
	(0.026)	(0.031)	(0.058)	(0.034)	(0.062)
Log Max AFDC/FSP Benefit	0.371	0.543	0.446	0.374	0.877
	(0.160)	(0.194)	(0.256)	(0.191)	(0.572)
Impact on Year Fixed Effects					
Aggregate unemployment	0.014	0.013	0.023	0.015	0.009
	(0.003)	(0.003)	(0.006)	(0.002)	(0.009)
Real GDP growth	–0.007	–0.008	–0.004	–0.007	0.002
	(0.003)	(0.002)	(0.007)	(0.003)	(0.007)
Trend	–0.013	–0.006	–0.036	–0.014	–0.026
	(0.002)	(0.002)	(0.004)	(0.001)	(0.005)
Year trend × post 1990	–0.033	–0.022	–0.004	–0.033	0.035
	(0.014)	(0.012)	(0.032)	(0.013)	(0.032)
Year trend × post 1992	–0.056	–0.059	–0.059	–0.052	0.001
	(0.020)	(0.020)	(0.038)	(0.011)	(0.053)
Year trend × post 1995	–0.093	–0.087	–0.122	–0.056	–0.088
	(0.028)	(0.024)	(0.051)	(0.016)	(0.073)

Notes: Robust standard errors are in parentheses. All regressions, based on three-year moving averages of CPS data from 1980–1999 for all 50 states and the District of Columbia, are weighted by the number of families in each state and control for year effects, state-specific fixed effects, and state-specific trends. For all families, the average cell size is 792.565, and the interquartile range is 289; for female-headed families, the mean cell size is 78.016, and the interquartile range is 30; for married couples, the mean cell size is 304.618, and the interquartile range is 139.5; for white families, the mean cell size is 690.228, and the interquartile range is 251.5; and for black families, the mean cell size is 139.28, and interquartile range is 111

A different story emerges with the more secular measures of the macroeconomy. For example, the median wage has neither a sizable economic impact nor a statistically significant impact on poverty across all families, married-couple families, or white families. With regard to all families, a possible reason for the lack of effect of the median wage is that if someone worked 40 hours a week for an entire year, the median wage ($13.64) would place him or her comfortably above the poverty line. Even the lowest median wage in a state during this period ($9.84) would place a person's income above the poverty line.[24] In contrast, growth in real median wages leads to substantial reductions in poverty among female-headed and black families, probably because of the lower median wages of these two demographic groups. When they work 40 hours a week, the median wage places members of these families at or below the poverty line.

If increases in median real wages are accompanied by increases in wage inequality, the estimates in Table 1 indicate that, in general, the benefits of economic growth are tempered by rising inequality. However, this dampening effect is much smaller for families that are headed by single mothers and is weakly beneficial for black families. The smaller effect for these groups suggests that wages for high-wage women and high-wage blacks (in the respective female-headed and black-family wage distributions) were still low enough relative to the poverty threshold in the 1980s and 1990s that increases at the high end relative to the lower tail of the distribution do not significantly lessen the benefits of wage growth.

We now turn to the variables that depict cross-state and time-series variation in social policies in Table 1. Families who reside in states with supplemental EITCs have, all else being equal, slightly higher before-tax poverty rates. (Because we used a before-tax measure, the monetary value of the EITC is not reflected in our income measure.) This can occur if the EITC encourages some workers to reduce their hours of labor-market work (Eissa and Hoynes 1999). Note, though, that this effect is small, since the elasticities range from zero to 0.05. Our results indicate that conditional on wage inequality, the higher the state minimum wage relative to the national level, the lower the head count. This finding is consistent with Lee's (1999) finding of increased inequality in the 1980s because of the declining minimum wage. Again, like the state EITC, these effects are small, since the largest elasticity found among black families suggests that a 10% increase in the state minimum wage lowers poverty by only 0.5%.

Welfare reform, as proxied by the indicators for pre-PRWORA welfare waivers and post-PRWORA TANF implementation, had little impact, positive or negative, on before-tax poverty rates. Among married couples, poverty rates rose upward of 6% both before and after PRWORA (although the latter is statistically insignificant). This is a surprising finding with no ready explanation, since less than 10% of the welfare caseload is composed of two-parent families. The positive correlation between welfare reform and poverty for married couples may reflect changes in the composition of married-couple families if some of the new rules that states adopted led more single women with children to marry. Indeed, since the implementation of welfare waivers, there has been an increase in marriage among teenage mothers, a decline or leveling off of divorce rates, and a decrease in unmarried childbearing, especially among black families (Lichter and Crowley 2002; "Sharp Increase in Marriages of Teenagers" 2002). Whether these trends have simply coincided with welfare reform or were caused by it requires further investigation. Finally, among welfare policies that are under (partial) state control, a 10% increase in

24. In an earlier version of the analysis, we used a quadratic in median earnings instead of the median wage and found a sizable and significant antipoverty role of earnings for all groups. Because of concerns over potential endogeneity of earnings with the poverty rate, we used wages in the current version. However, the results for earnings highlight the relative heterogeneity and cyclicality of hours compared to wages.

the maximum cash and food-stamp benefit for a three-person family leads to 4% to 9% higher poverty rates. The time-series literature tended to find no effect of transfers on poverty, and in some cases slightly positive effects (Blank 1993). This perverse effect, while surprising, does not appear to be robust to alternative specifications. Specifically, in results not presented, we found that the positive relation decreased to zero when the welfare-reform period was dropped from the analysis (but other parameter estimates were little changed).

Our model in Eq. (1) allowed us to portray the state-level factors we think are especially important to an improved understanding of the determinants of poverty. As we argued earlier, changes in macroeconomic conditions and social policies at the national level presumably also have an impact. In the lower panel of Table 1, we reported the estimates from Eq. (3). These estimates verify our assertion that aggregate macroeconomic conditions also matter at the local level, although the effects are not as pronounced. Indeed, a lower aggregate unemployment rate leads to lower state poverty rates, but the effect is about one third the magnitude of a change in unemployment at the state level. Likewise, growth in the real GDP weakly translates into lower poverty rates for most demographic groups.

The results of the auxiliary regression also confirm the downward trend in before-tax poverty rates over the past two decades. This trend was not constant, however, especially in the mid- to late 1990s. For example, from 1993 to 1996, poverty rates for female-headed families fell 6% faster than they did in the 1980s and 9% faster after 1996. Although the time-series analysis did not permit us to pinpoint the source of the decline (e.g., welfare reform, federal EITC expansions, or some other source), the results that are based on after-tax poverty rates, presented next, point to the EITC.

Accounting for Income Taxes

In Table 2, we present results that are parallel to those presented in Table 1 except that the dependent variable is now based on after-tax income. Overall, the results that are based on the state panel data from Eq. (1) changed little after tax liabilities were netted out and EITC credits were added. There are a few exceptions, however. Among female-headed families, the effect of PRWORA on after-tax poverty is about double the effect of before-tax poverty, although it remains statistically insignificant at usual levels of significance. More notable are changes among black families. First, the increase in the persistence of poverty from a before-tax to an after-tax measure tends to be substantially more pronounced among black families. This finding suggests the possible existence of racial differences in tax filing status (single parenthood is more heavily concentrated among this group) and in tax deductions (say, lower rates of itemizing because of lower rates of owner-occupied housing) that tend to perpetuate after-tax poverty among black families. Second, after accounting for taxes, we identified a strong antipoverty role for the macroeconomy, as measured by both unemployment rates and employment growth. This effect for black families is now more akin to the effects for other groups.

Particularly notable in the lower panel of Table 2 is that most of the coefficients on the 1990s trend-break variables are larger in absolute value than in the before-tax case. After-tax poverty tended to fall 1% to 2% faster than before-tax poverty between 1990–1992 and 1993–1995, and upward of 8% faster after 1995. This finding clearly points to a positive role for the federal EITC in eradicating after-tax poverty, although the actual magnitude is difficult to quantify. If the EITC entices nonworkers to enter the labor force or current workers to increase their hours of work, then it can be expected to reduce before-tax poverty rates, as is found in Table 1. Absent any other influences, this finding would imply that after 1995, the EITC lowered poverty by 9% more than in the 1980s, on average, and an additional 5% after taxes. This is an upper-bound estimate because there were other influences on the decline in poverty, including welfare reform.

Table 2. Estimates of the Impact of Macroeconomic Performance and Social Policies on After-Tax Poverty Rates

Variable	All Families	Female-Headed Families	Married-Couple Families	White Families	Black Families
Poverty ($t-2$)	0.127	0.073	0.105	0.153	0.251
	(0.040)	(0.044)	(0.041)	(0.044)	(0.040)
Unemployment Rate	0.044	0.035	0.044	0.044	0.046
	(0.005)	(0.006)	(0.007)	(0.006)	(0.005)
Growth in Employment per Capita	-0.015	–0.004	–0.018	–0.015	–0.011
	(0.004)	(0.005)	(0.007)	(0.005)	(0.005)
Median Wage	–0.070	–0.164	0.081	–0.017	–0.007
	(0.053)	(0.048)	(0.085)	(0.056)	(0.007)
Median Wage, Squared	0.000	0.006	–0.005	–0.002	–0.000
	(0.002)	(0.003)	(0.003)	(0.002)	(0.000)
Ratio of 80th to 20th Wages	0.251	0.049	0.310	0.302	–0.005
	(0.033)	(0.013)	(0.054)	(0.042)	(0.004)
Log of State-Federal EITC	0.031	0.024	0.004	–0.014	0.068
	(0.017)	(0.019)	(0.037)	(0.022)	(0.015)
Log of State-Federal Minimum Wage	–0.025	–0.018	–0.034	–0.037	–0.022
	(0.011)	(0.012)	(0.025)	(0.010)	(0.013)
Pre-PRWORA Waiver	0.014	–0.021	0.077	0.025	0.007
	(0.017)	(0.020)	(0.032)	(0.023)	(0.019)
Post-PRWORA Waiver	0.023	–0.052	0.112	–0.002	–0.025
	(0.028)	(0.034)	(0.073)	(0.039)	(0.026)
Log Max AFDC/FSP Benefit	0.378	0.340	0.454	0.367	0.249
	(0.162)	(0.202)	(0.274)	(0.195)	(0.173)
Impact on Year Fixed Effects					
Aggregate unemployment	0.022	0.018	0.039	0.024	0.011
	(0.005)	(0.005)	(0.011)	(0.004)	(0.005)
Real GDP growth	–0.010	–0.011	–0.008	–0.011	–0.005
	(0.004)	(0.003)	(0.008)	(0.003)	(0.003)
Trend	–0.018	–0.009	–0.054	–0.021	–0.009
	(0.003)	(0.003)	(0.007)	(0.003)	(0.003)
Year trend × post 1990	–0.041	–0.031	0.004	–0.042	0.012
	(0.018)	(0.017)	(0.042)	(0.015)	(0.016)
Year trend × post 1992	–0.067	–0.083	–0.038	–0.064	0.003
	(0.035)	(0.033)	(0.072)	(0.029)	(0.033)
Year trend × post 1995	–0.147	–0.136	–0.204	–0.109	–0.080
	(0.044)	(0.037)	(0.094)	(0.037)	(0.042)

Notes: Robust standard errors are in parentheses. All regressions, based on three-year moving averages of CPS data from 1980–1999 for all 50 states and the District of Columbia, are weighted by the number of families in each state and control for year effects, state-specific fixed effects, and state-specific trends. For all families, the average cell size is 792.565, and the interquartile range is 289; for female-headed families, the mean cell size is 78.016, and the interquartile range is 30; for married couples, the mean cell size is 304.618, and the interquartile range is 139.5; for white families, the mean cell size is 690.228, and the interquartile range is 251.5; and for black families, the mean cell size is 139.28, and interquartile range is 111.

The Macroeconomy, Social Policy, and the Depth of Poverty Across Families

In Table 3, we extend our previous analysis by estimating the effect of macroeconomic conditions and social policies on the depth of before-tax poverty, as measured by the squared poverty gap. Similar to the poverty-rate results in Table 1, the business cycle has a strong effect on the squared poverty gaps—a one-percentage-point decline in the unemployment rate leads to declines in the squared poverty gap from a low of 3.4% for poverty among female-headed families to a high of 5.2% for poverty among white families. While the business cycle did not affect before-tax poverty rates among black families, it improved the economic status of black families with incomes below the poverty line, especially those whose incomes were far below the poverty line. Likewise, the increase in median wages lowered the depth of poverty for both female-headed families and black families, with only limited evidence that economic progress was hindered by increases in inequality for these groups. This weak effect of wage inequality is contrary to the results for the other demographic groups under consideration.

State supplemental EITCs have no effect on the depth of before-tax poverty. In light of the results presented in Table 1, this finding suggests that the supplements most likely affected the behavior of those with incomes near the poverty threshold. State-mandated minimum wages, on the other hand, improved the economic position of the disadvantaged, particularly among female-headed and white families.

Pre-PRWORA waivers had a strong antipoverty role among female-headed and black families during the waiver period, reducing the squared poverty gaps by 12% among black families and 4% among female-headed families. The waivers adopted by states carried both carrots and sticks; the carrots often included higher earnings disregards and higher asset limits, while the sticks often entailed time limits on benefits and benefit sanctions. Collectively, these reforms trickled down into the lower tail of the income distribution to reduce the depths of poverty. On the other hand, during the PRWORA period, the depth of poverty among all families increased by 7% (the result is statistically significant at the 10% level)—primarily as a result of the 18% increase for married couples. As before, this variable probably captures a shift in composition among married-couple families and is not necessarily a causal channel. This shift in composition is also evident in the pre-PRWORA waivers, when married families saw an increase in the squared poverty gap of 13% in states with waivers.

Last, contrary to the results for poverty rates, more generous welfare benefits appear to improve the financial well-being of families who are the most likely to qualify for cash assistance—female-headed families. One possible explanation for the different results across poverty rates and squared poverty gaps is that low-benefit states have more compressed income-to-benefit schedules that may allow persons who are ex ante closer to the poverty line to escape poverty in low-benefit states. This more compressed scale, however, would not lead to larger decreases in poverty in low-benefit states. For all families, though, higher welfare benefits still led to increases in the depth of poverty, again the latter being driven by married-couple families.

The lower panel of Table 3 reveals that aggregate business cycle and economic growth do, in fact, "lift all boats" insofar as the squared poverty gaps fall with either a decline in aggregate unemployment or growth in the real GDP or both, although the benefits are more pronounced among married-couple families and families that are headed by white persons. Similar to the trends in poverty rates, there was a marked break in the trend of squared poverty gaps during the 1990s, particularly among female-headed and black families. In the mid-1990s, the trend fell between 8% and 11% among these groups and then an additional 14% and 24%, respectively, in the late 1990s. The period was clearly remarkable in the scope of families who were affected by the economic expansion and social policy reforms.

Table 3. Estimates of the Impact of Macroeconomic Performance and Social Policies on Before-Tax Squared Poverty Gaps

Variable	All Families	Female-Headed Families	Married-Couple Families	White Families	Black Families
Poverty ($t-2$)	0.128	–0.011	0.104	0.107	0.126
	(0.039)	(0.039)	(0.041)	(0.052)	(0.068)
Unemployment Rate	0.044	0.034	0.051	0.052	0.042
	(0.007)	(0.006)	(0.014)	(0.008)	(0.014)
Growth in Employment per Capita	–0.018	–0.006	–0.027	–0.017	–0.021
	(0.006)	(0.006)	(0.012)	(0.008)	(0.013)
Median Wage	–0.016	–0.166	0.326	–0.013	–0.350
	(0.074)	(0.060)	(0.101)	(0.068)	(0.060)
Median Wage, Squared	–0.001	0.007	–0.012	–0.001	0.012
	(0.002)	(0.003)	(0.003)	(0.002)	(0.003)
Ratio of 80th to 20th Wages	0.323	0.080	0.388	0.404	0.020
	(0.049)	(0.018)	(0.072)	(0.056)	(0.018)
Log of State-Federal EITC	0.024	0.003	0.004	–0.024	0.024
	(0.024)	(0.025)	(0.051)	(0.028)	(0.046)
Log of State-Federal Minimum Wage	–0.036	–0.025	–0.008	–0.053	–0.007
	(0.013)	(0.013)	(0.027)	(0.015)	(0.023)
Pre-PRWORA Waiver	0.027	–0.043	0.133	0.067	–0.124
	(0.025)	(0.026)	(0.042)	(0.032)	(0.050)
Post-PRWORA Waiver	0.067	–0.062	0.186	0.038	–0.018
	(0.039)	(0.049)	(0.085)	(0.054)	(0.089)
Log Max AFDC/FSP Benefit	0.608	–0.479	0.946	0.832	–0.411
	(0.240)	(0.267)	(0.437)	(0.280)	(0.526)
Impact on Year Fixed Effects					
Aggregate unemployment	0.017	0.009	0.018	0.017	0.015
	(0.006)	(0.006)	(0.006)	(0.004)	(0.014)
Real GDP growth	–0.004	–0.004	0.009	–0.003	–0.012
	(0.003)	(0.004)	(0.005)	(0.004)	(0.009)
Trend	–0.022	–0.006	–0.050	–0.027	0.001
	(0.004)	(0.003)	(0.004)	(0.002)	(0.008)
Year trend × post 1990	0.015	–0.010	0.092	–0.011	–0.061
	(0.021)	(0.017)	(0.024)	(0.022)	(0.045)
Year trend × post 1992	–0.016	–0.084	0.090	–0.068	–0.113
	(0.035)	(0.038)	(0.031)	(0.024)	(0.078)
Year trend × post 1995	–0.075	–0.135	–0.019	–0.073	–0.238
	(0.057)	(0.044)	(0.051)	(0.032)	(0.125)

Notes: Robust standard errors are in parentheses. All regressions, based on three-year moving averages of CPS data from 1980–1999 for all 50 states and the District of Columbia, are weighted by the number of families in each state and control for year effects, state-specific fixed effects, and state-specific trends. For all families, the average cell size is 792.565, and the interquartile range is 289; for female-headed families, the mean cell size is 78.016, and the interquartile range is 30; for married couples, the mean cell size is 304.618, and the interquartile range is 139.5; for white families, the mean cell size is 690.228, and the interquartile range is 251.5; and for black families, the mean cell size is 139.28, and interquartile range is 111.

As we did in Table 2, in Table 4 we examine the determinants of squared poverty gaps with after-tax income, rather than with before-tax income. Overall, the results are similar to those presented in Table 3, although a few anomalous results appear. The unemployment rate has its expected countercyclical effect on the depth of after-tax poverty for female-headed families and black families, but it is acyclical overall because of the perverse procyclical effect for married-couple families. Growth in employment per capita, however, continues to exert strong antipoverty influences on after-tax poverty across families. Also, contrary to previous estimates, the state-funded EITC lowers after-tax squared poverty gaps for married couples and weakly lowers them for white families. The results in Table 4 indicate that PRWORA is strongly positively correlated with a higher after-tax squared poverty gap, although this is not true for female-headed and black families, for whom the effect is insignificant. The channel underlying the higher after-tax poverty, which is again driven by married-couple families, clearly merits further investigation.

Finally, in the aggregate, the 1990s witnessed substantial reductions in economic deprivation, as captured by the trend breaks in the time series. Summing the trend breaks indicates that the depths of after-tax poverty during the 1990s were nearly 50% lower for most groups under consideration. This finding should be compared with the before-tax situation in Table 3, where the declines were much more modest. This is partial evidence that the bulk of the improvement in economic status among America's poorest families occurred through the redistribution within the federal tax code, especially via the EITC.[25]

Simulations

Because median wages and wage inequality have a cyclical component and thus respond to changes in the unemployment rate and the employment growth rate, the interpretation of model estimates in Tables 1–4 containing the full set of macroeconomic indicators is muddied. To interpret our results further, we simulated the impact of the macroeconomy on poverty rates and poverty gaps before and after taxes and present the estimates in Table 5. We conducted the simulations for all families and for those groups with high concentrations of poverty: female-headed families and black families. In row 1 in each of the upper and lower panels, we list the predicted poverty rates and poverty gaps from our models with all the variables at their average values. In rows 2–5 of each panel, we present, using the coefficients from Tables 1–4, what would happen to the extent and depth of poverty if all states faced the unemployment rate, employment growth per capita, median wage, and 80–20 wage inequality at the levels observed in the trough of the 1980s recession (1982), the peak of the 1980s expansion (1989), the trough of the 1990s recession (1991), and the peak of the 1990s expansion (1999), holding the other variables at their mean levels. To aid in interpretation, we report the antilog of the dependent variables (i.e., the levels of the poverty rates and gaps).

Our simulations from the 1980s are as one would expect: the values for the head count and squared poverty gap are higher than average for the trough in 1982 and lower than average for the peak in 1989. Across all measures, the recession of the early 1990s resulted in lower poverty than did the trough of 1982, highlighting the severity of the 1980s recession. For example, the predicted poverty rate for female-headed families and

25. In results not presented, we examined Haveman and Schwabish's (2000) claim that the link between macroeconomic performance and poverty was tighter in the 1990s than in the 1980s. We found no change in the relationship between unemployment and poverty, but a stronger link between employment growth and poverty, especially among female heads of families. Likewise, we examined how our results differ from a pure time-series analysis. We found that in the aggregate time series, poverty dependence is stronger, that is, the coefficient on the lagged dependent variable is larger, which is consistent with the typical greater autocorrelation found in time-series data compared with panel or cross-sectional data. In the time-series results, though, both employment growth and wage inequality were poorly determined, and state welfare reforms were obviously unidentified.

Table 4. Estimates of the Impact of Macroeconomic Performance and Social Policies on After-Tax Squared Poverty Gaps

Variable	All Families	Female-Headed Families	Married-Couple Families	White Families	Black Families
Poverty ($t-2$)	0.170	0.017	0.146	0.119	0.131
	(0.058)	(0.051)	(0.053)	(0.063)	(0.084)
Unemployment Rate	0.009	0.035	–0.060	–0.010	0.052
	(0.014)	(0.008)	(0.027)	(0.017)	(0.017)
Growth in Employment per Capita	–0.037	–0.009	–0.067	–0.049	–0.039
	(0.012)	(0.007)	(0.025)	(0.015)	(0.016)
Median Wage	–0.150	–0.137	0.294	–0.018	–0.275
	(0.166)	(0.070)	(0.207)	(0.143)	(0.076)
Median Wage, Squared	0.006	0.006	–0.008	–0.001	0.009
	(0.006)	(0.004)	(0.007)	(0.004)	(0.003)
Ratio of 80th to 20th Wages	0.086	0.079	0.170	0.302	0.024
	(0.103)	(0.021)	(0.110)	(0.129)	(0.023)
Log of State-Federal EITC	–0.068	0.058	–0.407	–0.111	0.037
	(0.063)	(0.044)	(0.138)	(0.087)	(0.053)
Log of State-Federal Minimum Wage	–0.006	–0.032	0.049	0.010	0.001
	(0.027)	(0.021)	(0.059)	(0.030)	(0.024)
Pre-PRWORA Waiver	0.204	0.009	0.305	0.237	–0.073
	(0.073)	(0.044)	(0.143)	(0.095)	(0.091)
Post-PRWORA Waiver	0.382	–0.044	0.876	0.387	0.136
	(0.118)	(0.063)	(0.208)	(0.147)	(0.137)
Log Max AFDC/FSP Benefit	0.773	–0.459	1.031	1.176	–0.423
	(0.565)	(0.297)	(0.991)	(0.677)	(0.598)
Impact on Year Fixed Effects					
Aggregate unemployment	0.070	0.014	0.101	0.092	0.032
	(0.025)	(0.008)	(0.045)	(0.025)	(0.023)
Real GDP growth	–0.045	–0.007	–0.043	–0.056	–0.025
	(0.020)	(0.006)	(0.032)	(0.020)	(0.015)
Trend	–0.028	–0.009	–0.084	–0.044	0.000
	(0.013)	(0.005)	(0.025)	(0.013)	(0.014)
Year trend × post 1990	–0.139	–0.016	–0.044	–0.248	–0.101
	(0.068)	(0.024)	(0.137)	(0.067)	(0.073)
Year trend × post 1992	0.081	–0.100	0.340	0.039	–0.129
	(0.124)	(0.056)	(0.225)	(0.128)	(0.128)
Year trend × post 1995	–0.473	–0.225	–0.752	–0.585	–0.381
	(0.179)	(0.063)	(0.379)	(0.176)	(0.218)

Notes: Robust standard errors are in parentheses. All regressions, based on three-year moving averages of CPS data from 1980–1999 for all 50 states and the District of Columbia, are weighted by the number of families in each state and control for year effects, state-specific fixed effects, and state-specific trends. For all families, the average cell size is 792.565, and the interquartile range is 289; for female-headed families, the mean cell size is 78.016, and the interquartile range is 30; for married couples, the mean cell size is 304.618, and the interquartile range is 139.5; for white families, the mean cell size is 690.228, and the interquartile range is 251.5; and for black families, the mean cell size is 139.28, and interquartile range is 111.

Table 5. Simulations of the Impact of Macroeconomic Performance on Before and After-Tax Poverty Rates and Squared Poverty Gaps For Selected Families

	Before Taxes			After Taxes		
Variable	All Families	Female-Headed Families	Black Families	All Families	Female-Headed Families	Black Families
Head Count						
Average levels	9.752	42.105	25.393	9.233	39.031	24.092
Values from the trough of the 1980s recession	10.673	46.044	32.184	10.166	42.852	30.546
Values from the peak of the 1980s expansion	9.293	41.059	24.705	8.870	38.273	23.613
Values from the trough of the 1990s recession	10.072	42.548	27.260	9.613	39.650	26.028
Values from the peak of the 1990s expansion	9.579	39.175	24.361	9.140	36.469	23.269
Squared Poverty Gaps						
Average levels	2.241	11.672	6.582	2.481	10.887	6.367
Values from the trough of the 1980s recession	2.519	12.899	8.120	2.841	12.162	7.571
Values from the peak of the 1980s expansion	2.203	11.789	6.600	2.772	11.120	6.320
Values from the trough of the 1990s recession	2.354	12.134	7.412	2.785	11.441	7.138
Values from the peak of the 1990s expansion	2.249	11.297	6.582	2.752	10.675	6.257

Notes: All results are multiplied by 100. Simulations are based on the results from the models in Tables 1–4. The troughs of the 1980s and 1990s recessions are 1982 and 1991, respectively, while the peaks of the expansions are 1989 and 1999, respectively.

black families before and after taxes was nearly 5 percentage points higher in the early 1980s than in the 1990s. However, despite the record expansion observed in the 1990s, our simulations of the peak of the 1990s expansion indicate that poverty rates and gaps among all families were higher than in a simulation of the peak of the 1980s expansion. This finding may reflect the continuing tempering influence of rising inequality on reductions in poverty. Again, though, focusing on all families alone and ignoring subgroups paints a distorted picture. While families as a whole were not predicted to be as well off at the end of the 1990s, female-headed families and families that were headed by black persons were predicted to have made substantial reductions in both the extent and depth of poverty. In large part, this positive development was driven by the gains in median wages among these two groups.

CONCLUSION

To summarize our findings, we return to the issues raised in the introduction—namely, is long-run economic growth an effective antipoverty tool? Are transfer policies that are targeted specifically at the low-income population key to reducing poverty? Are the anti-poverty impacts of economic growth and social policies uniform across subsets of the population; that is, does a "rising tide lift all boats"? Our results indicate that the answer to the first two questions is a resounding yes, while the answer to the third is a qualified no.

A strong macroeconomy at both the state and national levels reduces the number of families with incomes below the poverty line and the severity of poverty. The magnitude and source of these antipoverty effects, however, are not uniform across family structures and racial groups or necessarily over time. During the 1980s and 1990s, poverty among married-couple families and families that were headed by white persons tended to respond to changes in both the unemployment rate and the employment growth rate, but not to changes in the median wage. On the other hand, it was not until the high-growth economy of the 1990s that employment growth for female heads of families surged and was reflected in lower poverty rates and gaps. A further distinction is that the business cycle has a countercyclical effect on after-tax poverty rates and before- and after-tax poverty gaps among families that are headed by black persons, but has no effect on before-tax poverty rates, although there is limited evidence that this link became stronger in the 1990s. What is clear is that the increase in the median wage over the past decade substantially reduced the extent and depth of poverty among female-headed and black families. Rising inequality, however, tempered the gains made against poverty overall and especially among married-couple families and white families.

While the macroeconomy continues to have a major impact on poverty, there is an array of other programs with antipoverty goals. The discussions surrounding the optimal mix of these policies are predicated on the implicit trade-off between the efficiency and the equity of the policies. The EITC is a perfect case in point. The EITC, which redistributes income to low-income workers and thus reduces after-tax income inequality, is believed to improve labor-market participation by drawing nonworkers into the labor force. At the same time, over part of the EITC range, the subsidy and associated implicit tax rate lower the incentive to work and thus exacerbate labor-market inefficiencies. There is some evidence that, in the aggregate, the EITC reduces the supply of labor (Hoffman and Seidman 1988). Thus, if it does not play a positive antipoverty role, then support for the program will likely diminish. Our results, particularly those based on the auxiliary model, suggest that expansions in the federal EITC were a significant contributor to the declines in poverty over the past decade, especially in reducing the after-tax squared poverty gap. Likewise, while the potential disemployment effects of the minimum wage are well known, we found that states with more generous minimum wages relative to the national level had lower poverty rates during the period under study, although the effect was economically small.

Assessments of the contribution of welfare reform to the declines in poverty face similar equity and efficiency considerations. Many of the waivers adopted by states prior to PRWORA were intended to "make work (and welfare) pay," while others were simply designed to make welfare unattractive, which may or may not have induced a positive labor-supply response. We found limited and mixed evidence that (after-tax) poverty was lower among female-headed families and black families after welfare waivers and PRWORA were implemented. An area that is in need of further research is the increase in poverty among married-couple families after welfare reform.

Collectively, our results indicate the importance of policies that foster economic growth and policies that redistribute income. Economic growth clearly reduces the depth of poverty in the United States. But since this growth seems to reduce the number of black families who are poor only after income taxes and credits are accounted for, it suggests the importance of social policy. Policies that foster further gains in labor productivity, coupled with a redistributive tax policy like the EITC, are likely to lead to the further amelioration of poverty in the United States.

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