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Spatial inequality and poverty among American children

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Abstract. National-level statistics often mask extreme spatial differentiation in child poverty. Using county-level data from the 1990 US decennial census summary tape file, we show that child poverty is distributed unevenly over geographic space. Child poverty is concentrated in counties in Appalachia, the Mississippi Delta, and the southern 'black belt'. Child poverty rates are strongly influenced by the local industrial composition (e.g., agriculture and manufacturing), but the effects are largely indirect, operating primarily through reduced employment opportunities among adult workers. High county unemployment and underemployment rates contribute directly to children's economic deprivation, as well as indirectly by undermining the formation and stability of two-parent families. Our results highlight existing spatial differentiation and inequality in children's economic well-being, and provide a point of departure for additional research on the geography of child poverty.

Key words: Children, Poverty, Spatial inequality, USA

Introduction

The apparent intractability of poverty in America is responsible for the continuing and often heated debates about welfare reform and anti-poverty legislation (Corbett 1993). The poverty rate in the USA of nearly 15 percent in 1992 was only slightly lower than that observed during the early 1980s recession, while the number of poor people (37 million) exceeded that of any year since the early 1960s (US Bureau of the Census 1993). The problem is particularly acute among America's children (Bane & Ellwood 1989; Bianchi 1990; Eggebeen & Lichter 1991). In 1990, children constituted about 36 percent of the poor population. Over one-in-five children today are poor, a poverty rate higher than that of any other age-segment of the population. Since 1990, the child poverty rate has resumed its upward trajectory, while the economic circumstances of the elderly population continued to improve (US Bureau of the Census 1993).

Much of the increase in child poverty has been attributed to changing family structure, especially to the rise in the proportion of children living in 'at risk' female-headed families (Bianchi & McArthur 1991; Eggebeen & Lichter 1991). Demographic explanations that emphasize family structure alone, however, are incomplete (Children's Defense Fund 1991). National-level statistics often mask extreme spatial differentiation in child poverty, a pattern that cannot be explained by the substantially lower geographic variability in family structure.[1] Indeed, local labor market conditions (e.g., industrial composition) affect family and children's economic well-being directly through the wage structure and parental employment opportunities, as well as indirectly by discouraging marriage and destablizing existing marital unions (Conger et al. 1990; Lichter & McLaughlin 1995; Wilson 1987). The main purpose of this paper is to evaluate this hypothesis using county-level data from the 1990 US decennial census.

Ours is an important task for several reasons. First, we provide needed balance to current perspectives that emphasize growing family disorganization as the primary cause of children's current economic deprivation. Second, poverty is typically evaluated as a characteristic of individuals and families (Eggebeen & Lichter 1991), rather than as a reflection of the apparently growing ascriptive constraints of place (i.e., the local opportunity structure). Third, our focus on all US counties, including nonmetropolitan (nonmetro) counties, provides a necessary corrective to the current preoccupation with poverty in metropolitan (metro) areas and central cities (e.g., Eggers & Massey 1991; Frey 1995).

Economic opportunity and child poverty

Uneven regional development increasingly characterizes the US labor market (Bluestone & Harrison 1982; Colclough 1988). Some labor markets have emerged as centers for capital investment and economic expansion (i.e., the so-called 'winners'), while others – especially those in rural and declining urban industrial areas – have remained economically stagnant in the 1990s. The historical trend favoring spatial economic homogeneity (e.g., declines in inter-state income inequality) has been replaced by a new pattern of increasing economic inequality over counties, states, and regions (Lichter 1993; Falk & Lyson 1993). Frey (1995), for example, has documented the sharper economic divisions over geographic space, i.e., a new 'balkanization' of social and demographic groups.

Uneven economic development is perhaps best reflected in place-to-place differentials in industrial structure that contribute to spatial inequality in wages, unemployment, and underemployment. Indeed, local-area differences in industrial composition account for a substantial share of the difference in unemployment and underemployment rates across geographic space and demographic groups (Eggers & Massey 1991; Lichter & Costanzo 1987; Tigges & Tootle 1993). County employment in manufacturing, for example, is associated with higher wages and lower unemployment and underemployment (Tomaskovic-Devey 1987; Tickamyer & Tickamyer 1988). Agriculture- and mining-counties, on the other hand, are characterized by disproportionately high rates of unemployment.

The effects of industrial structure on poverty are thus arguably indirect, i.e., they are mediated through adult or parental unemployment and underemployment. But the evidence to date is mixed or inconclusive. For example, Tomaskovic-Devey (1987) found no significant direct effect of industrial structure (i.e., defined in terms of county employment rates and average earnings) on poverty rates in counties in South Carolina. Tickamyer & Tickamyer (1988), on the other hand, reported that the percentage employed in agriculture and mining was associated with high unemployment and underemployment rates, which in turn exacerbated poverty among families in Appalachian counties. The percentage employed in manufacturing also affected poverty indirectly (through unemployment), but the relationship was weak.

Evidence of direct relationships between the industrial structure and poverty is substantial (Eggers & Massey 1991; Lichter & McLaughlin 1995), although differences in measurement and study designs have often led to inconsistencies in substantive conclusions. For example, Tickamyer & Duncan (1990) found a positive association between resource-based economies and local-area poverty. Weinberg's national study (1987) of inter-county variation in poverty revealed a strong negative relationship between percent employed in manufacturing and poverty. Tigges (1987) found a negative effect for core transformative manufacturing and low income, but a positive effect of periphery manufacturing. A similar inconsistency has been reported for service employment. For example, county poverty rates were positively associated with the percentage employed in service industries in South Carolina (Tomaskovic-Devey 1987), while Reif (1987) showed no relationship with periphery service employment. Moreover, Lichter & McLaughlin (1995) reported a significant negative association in 1980 but a positive relationship in 1990 for all US counties.

Our question is how industrial structure affects spatial variation in county-level poverty rates for children, and, unlike previous research, how these effects are mediated by employment opportunities and family structure. Unfortunately, studies of the determinants of spatial inequality are in surprisingly short supply, despite large geographic variations in children's poverty rates (US Bureau of the Census 1993). An exception is the study by Rexroat (1989), who found that poverty rates among black children living in northern SMSAs were on average larger than those in southern SMSAs. She attributed these differences largely to spatial differentials in employment opportunities rather than to regional differences in family structure. Her study underscores the importance of labor markets in evaluating spatial variation in child poverty.

Other studies emphasize that the effects of industrial structure and employment on child poverty are also mediated through family formation and marital instability (Easterlin 1987; Conger et al. 1990). The stability of the family is presumably undermined in depressed labor markets. In urban areas, for example, Wilson (1987) argued that depressed labor markets were associated with low marriage rates and high nonmarital fertility; i.e., high unemployment rates reduced the pool of 'marriageable' men and discouraged marriage. Indeed, Hernandez (1994) found a strong relationship between recessionary periods and the growth of female-headed families. In economically depressed areas, high percentages of female-headed families are likely to aggravate the child poverty problem.

In sum, the main objective of this paper is to document spatial variation in child poverty rates and evaluate how the economic circumstances of American children are affected by local-area economic opportunities available to their parents. Specifically, we estimate multivariate models that assess the direct and indirect effects (through family structure) of local labor market conditions and employment opportunities on county child poverty rates. To our knowledge, this is the first study to evaluate the extent and etiology of child poverty over geographic space.

Methods

Data and measurement. The data for this study are from the 1990 United States Census Summary Tape File 3C (US Bureau of the Census 1991). The county – rather than the neighborhood or community – is a useful level of analysis for examining aggregate-level phenomena (Bloomquist & Summers 1982; Stafford & Fossett 1989). Counties are the smallest geographic units having comparatively fewer boundary changes over time and, unlike communities, counties provide full geographic coverage of the USA.

Our dependent variable is the proportion of children aged 17 or younger living in the county with family incomes below the poverty line in 1989. The official poverty rate is based on family money income as it relates to an absolute income threshold that varies by family size and composition (Ruggles 1990). Noncash or in-kind benefits, such as food stamps, medical insurance, and public housing, are excluded.

We consider the effects of several local economic indicators: county industrial composition, unemployment, and underemployment. County industrial composition is measured by the proportion of employed persons in the following categories: extractive, nondurable manufacturing, miscellaneous services (e.g. personal services, entertainment, and recreation services), and professional services (e.g. finance, insurance, real estate, health services, educational services, public administration).[2] Our review of previous studies implies that county-level child poverty rates will be positively associated with the percentage employed in resource-based industries and in low-wage manufacturing (i.e. nondurable manufacturing) and service industries.

Much of the effect of industrial composition is expected to be mediated through local employment opportunities and family structure. Unemployment is measured by the proportion of all workers (aged 16 and over) in the county's civilian labor force with any episode of unemployment during 1989. Underemployment is measured as the proportion of all men (aged 16 and over) in a county who worked less than 35 hours a week and/or less than 27 weeks during 1989.[3] Since many women voluntarily choose to work part-time, our measure of underemployment is restricted to men's experiences. Family structure is defined as the county proportion of families with children which are headed by females (with no husband present).

Our multivariate analysis also includes several control variables. The minority composition is measured by the county proportion non-Hispanic black and the proportion Hispanic; each is expected to be positively associated with county child poverty rates (Lichter & McLaughlin 1995). Educational composition is measured by the proportion of adults (aged 18 or older) in the county with a high school education or less. Regional location (South = 1) and metro status (metro = 1) are hypothesized to be negatively associated with child poverty (Lichter & Eggebeen 1992). Finally, because workers often live in one county but work in another, we control for the proportion of the workers (aged 16 and over) employed within the county. We expect that counties with a greater proportion of commuters will have lower child poverty rates.

Analytic approach. Our analysis proceeds as follows. First, a county-level US map of child poverty rates visually illustrates the striking economic inequality over geographic space. Second, means (for each independent variable) are calculated for counties with low and high rates of child poverty; this exercise provides a useful socio-demographic profile of economic 'winners' and 'losers'. Third, logit regression models provide a basis for evaluating the comparative effects of labor market conditions and family strcuture on child poverty.[4]

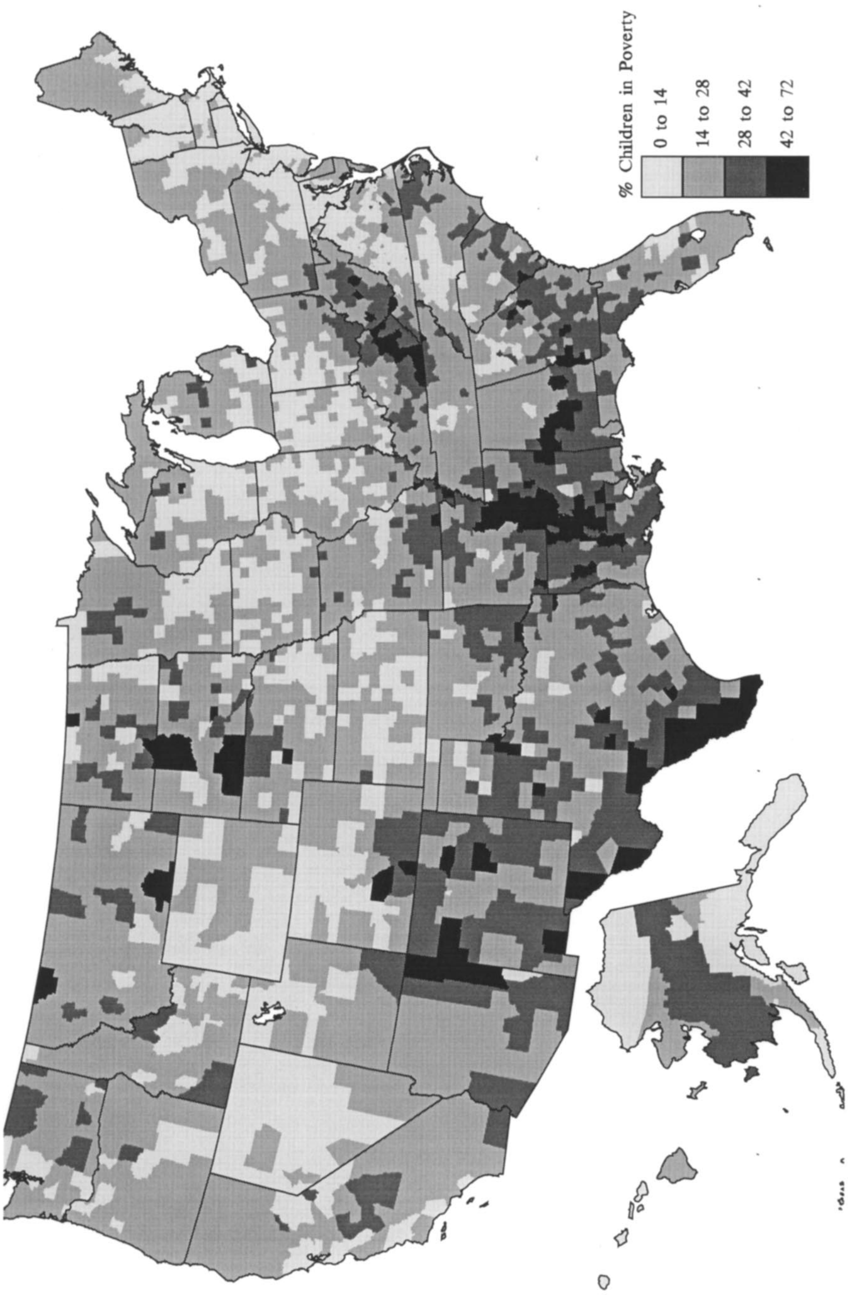

Figure 1. Percentage of children in poverty.

Findings

Geographic variation in child poverty rates

The county-level map in Figure 1 shows that child poverty rates are highest in Appalachia, the Mississippi Delta, the southern 'black belt', and the lower Rio Grande Valley. Over 40 percent of all children in counties of these regions typically live in poverty. Parts of South Dakota and the Four Corners area of the Southwest also exhibit high child poverty rates, a pattern consistent with high concentrations of Native Americans. Child poverty is lowest in counties in the Midwest, New England, and parts of the mid-Atlantic region. These counties have child poverty rates well below the national rate of 21 percent.

Table 1 provides profiles of counties with very low child poverty rates and very high poverty rates. As expected, the economic 'losers' exhibit the most disadvantaged socio-demographic profiles. For example, the proportion black is only 0.03 in very low poverty counties, but is 0.27 in very high poverty counties. The proportion unemployed also displays a positive relationship with county child poverty rates, increasing about threefold from the very low category to the very high category. Not surprisingly, the proportion of female headed families with children more than doubles between the lowest and the highest poverty categories. Clearly, counties with very high child poverty rates are much more disadvantaged socially and economically than counties with very low poverty rates.

Explaining geographic variation in child poverty

Table 2 presents the results of the logit regression analysis of the proportion of children living in poverty. The baseline model (model 1) evaluates the effects of the industrial composition variables (and other control variables). Each is significantly related to the proportion of children living in poverty in counties. For example, employment in extractive industries and professional services is positively associated with child poverty, while nondurable manufacturing and miscellaneous service employment are negatively related to county rates of child poverty.

The inclusion of unemployment and underemployment variables in model 2 eliminates the significant effects of the industrial variables (as shown in model 1), except in the case of the proportion employed in extractive industries. The implication is straightforward. Most of the effect of industrial structure on child poverty is mediated through unemployment and underemployment.

Model 3 in Table 2 evaluates the mediating effects of family structure on child poverty rates. The results indicate that a one percentage point increase in the percentage of female headed families is associated with a

Table 1. Means of poverty related factors for very low to very high child poverty counties

Variable	Child poverty rates in counties: Very low (0–14%)	Low (14–28%)	High (28–42%)	Very high (42–72%)
Number of counties (N)	798	1645	551	146
Proportion black (non-Hispanic)	0.032	0.067	0.166	0.273
Proportion Hispanic	0.024	0.034	0.075	0.145
Proportion with a high school education or less	0.571	0.646	0.715	0.748
Employed in the county	0.675	0.743	0.726	0.740
South[a]	0.216	0.441	0.740	0.822
Met[a]	0.430	0.216	0.091	0.034
Industry				
Extractive	0.150	0.174	0.191	0.209
Nondurable manufg.	0.073	0.082	0.091	0.079
Miscellaneous services	0.079	0.073	0.070	0.069
Professional services	0.317	0.303	0.302	0.320
Proportion of labor force unemployed	0.047	0.064	0.087	0.129
Proportion of male working population underemployed	0.219	0.244	0.266	0.302
Proportion of female headed families with children	0.122	0.148	0.190	0.253

[a] Proportion of counties exhibiting each locational characteristic within each category poverty.

0.85 percentage-point increase in the county child poverty rate – nearly a one-to-one correspondence.[5] The direction and magnitude of the effects of the industrial variables are similar to that in column 1 except for the effect of extractive industries, which triples in this model (0.568 vs. 1.614).

Model 4 – the full model – reveals several noteworthy results. First, the direct effect of extractive industries on child poverty is stronger than in models 1 and 2 (which exclude family structure). This implies that the more traditional family structures observed in counties with high concentrations in extractive industries may mask the high poverty in these counties. Simply, traditional families (i.e., married couples) buffer local economic dislocations associated with counties dependent on extractive industries.

Second, the coefficients for unemployment and underemployment are smaller with the inclusion of family structure (compare model 4 to model 2). This

Table 2. Logit regression models of the proportion of children in poverty in counties (N = 3136)

Variable	(1)	(2)	(3)	(4)
Proportion black	1.830	1.475	−0.091[ns]	0.212
(non-Hispanic)	(0.050)[a]	(0.040)	(0.066)	(0.054)
Proportion Hispanic	1.380	0.831	1.151	0.778
	(0.042)	(0.035)	(0.036)	(0.031)
Proportion with a high	3.268	1.954	2.190	1.460
school education or less	(0.076)	(0.069)	(0.069)	(0.063)
Employed in the county	1.028	0.702	0.782	0.596
	(0.036)	(0.028)	(0.030)	(0.025)
South	0.049	0.089	0.192	0.180
	(0.013)	(0.010)	(0.011)	(0.009)
Met	−0.139	−0.006[ns]	−0.179	−0.061
	(0.018)	(0.014)	(0.015)	(0.013)
Industry				
Extractive	0.568	0.557	1.614	1.287
	(0.146)	(0.112)	(0.124)	(0.101)
Nondurable manufg.	−1.593	−0.268[ns]	−1.154	−0.216[ns]
	(0.170)	(0.136)	(0.142)	(0.119)
Miscellaneous services	−0.701	0.282[ns]	−1.831	−0.691
	(0.267)	(0.204)	(0.224)	(0.182)
Professional services	1.562	0.077[ns]	0.150[ns]	−0.604
	(0.123)	(0.101)	(0.108)	(0.091)
Proportion of labor		7.378		5.984
force unemployed		(0.261)		(0.234)
Proportion of male working		3.211		2.555
population underemployed		(0.132)		(0.118)
Proportion of female headed			5.085	3.517
families with children			(0.135)	(0.115)
Intercept	−4.811	−4.770	−4.320	−4.435
Adjusted R-square	0.725	0.842	0.811	0.878

[a] The numbers in parentheses are standard deviations.
[ns] Not significant at the 0.01 level.

result buttresses Wilson's (1987) argument that the effect of local unemployment on poverty is indirect, at least in part. That is, employment dislocations contribute to the child poverty problem by undermining the formation and stability of 'low risk' two-parent families. This is not the entire story, however. A one percentage point increase in the percentage unemployed is associated with a one percentage increase (i.e., 1.003) in the poverty rate, *even after* *controlling for family structure and other socioeconomic characteristics of the county*.

Third, the size of the family structure coefficient is reduced in the full model (compare model 4 with model 3). A one percentage point increase in the percentage female headed families with children is associated with a 0.59 percentage point increase in child poverty. This compares with 0.85 percentage point increase implied in model 3. Clearly, the effect of family structure on child poverty is important, but nevertheless is smaller than the effect of the county unemployment rate. Moreover, the relationship between family structure and child poverty cannot be evaluated in isolation of local labor market conditions.

Spatial variation in child poverty within metro and nonmetro areas

The preceding analysis indicates that labor market structure and family composition account for much of the geographic variability in the economic status of American children. It is also clear, however, that the extent and etiology of child poverty is potentially different in different types of counties. For example, the negative effect of metro residence shown in Table 2 (model 4) indicates that even after controlling for labor market conditions, family composition, and other control variables,[6] metro children's poverty rate is only 0.94 of the rate of nonmetro children. This finding is consistent with previous research that reveals the relative economic disadvantage of nonmetro adults and children compared to their metro counterparts (Jensen & Tienda 1989; Lichter & Eggebeen 1992).

The fact that nonmetro child poverty rates are higher than metro child poverty rates is paradoxical in light of current perceptions about the strength of family and kinship ties in rural America and a strong work ethic. This implies that the etiology or causes of child poverty are different in rural and urban counties.[7] Economic returns to manufacturing employment – in the form of lower child poverty – may be lower in nonmetro areas; i.e., manufacturing employment pays less well in nonmetro than metro areas (Lichter & McLaughlin 1995). Indeed, McLaughlin & Perman (1991) showed that economic returns to employment and education are lower for nonmetro workers than for metro workers. Likewise, poverty among workers is higher in nonmetro than metro areas, a fact reflecting lower economic returns to employment and lower labor demand in these areas (Lichter et al. 1994). The implication is (1) that the uneven geographic distribution of economic and demographic characteristics contributes to spatial differences in child poverty (i.e., a composition effect), and/or (2) that the effects of these characteristics on county child poverty may differ over geographic location (i.e., a rate effect).

Our focus on nonmetro and metro areas provides an interesting test case for evaluating this hypothesis. Model 1 (Table 3) of the proportion of children living in poverty in nonmetro areas gives the results of the baseline model. All of the coefficients, except for the coefficient of miscellaneous services, are significant at the 0.01 level. The addition of unemployment and underemployment in Model 2 renders the industrial structure coefficients insignificant (except in the case of extractive industries). This suggests that much of the effect of nondurable manufacturing industries, miscellaneous service industries, and professional services industries on nonmetro child poverty are mediated through unemployment and underemployment. This is consistent with the results found for all counties (see Table 2).

Model 3 in Table 3 shows that family composition also is an important determinant of child poverty in nonmetro areas. A one-percentage point change in the proportion of female headed families increases the nonmetro child poverty rate by 1.02, an effect larger than present in Model 3 of Table 2 (0.85). Clearly, female-headed families play a much stronger role influencing child poverty rates in nonmetro areas than is the case for child poverty rates within the USA as a whole.

The full and final model (Model 4 in Table 3) also supports conclusions that are similar to those drawn from the results in Table 2. The coefficients for proportion unemployed, proportion underemployed, and proportion of female headed families are smaller in the full model, although the relative effect of the proportion unemployed to the proportion of female headed households is smaller than in the full model in Table 2. Nevertheless, nonmetro child poverty is mostly affected by labor market factors. The combined effects of the proportion unemployed, the proportion of men underemployed, and the proportion employed in extractive industries outweigh the effect of family structure.

Table 4 presents the results of the logit regression models of metro child poverty. Differences from the nonmetro results are largely a matter of degree rather than kind. For example, the effect of the proportion with a high school education or less on child poverty in nonmetro areas is about 1.7 times larger than that observed for metro areas. This difference is statistically significant (Z = 5.235; $p < 0.1$).[8] The effect of the proportion employed in extractive industries is about 1.5 times larger than that observed for metro children (Z = 1.962; $p < 0.050$). The effect of the proportion of the labor force unemployed on child poverty in metro areas, however, is about 2 times larger than that observed for nonmetro children. This difference also is statistically significant ($Z = -9.27; p < 0.01$). Clearly, these differences contribute to the inequality of children between metro and nonmetro areas.

Table 3. Logit regression models of the proportion of children in poverty in nonmetro counties (N = 2383)

Variable	(1)	(2)	(3)	(4)
Proportion black	1.540	1.333	−0.220	0.218
(non-Hispanic)	(0.057)[a]	(0.045)	(0.076)	(0.065)
Proportion Hispanic	0.935	0.614	0.763	0.572
	(0.062)	(0.050)	(0.054)	(0.045)
Proportion with a high	3.202	1.953	2.550	1.842
school education or less	(0.105)	(0.097)	(0.092)	(0.088)
Employed in the county	0.726	0.556	0.269	0.291
	(0.058)	(0.046)	(0.052)	(0.043)
South	0.173	0.161	0.246	0.209
	(0.019)	(0.015)	(0.017)	(0.014)
Industry				
Extractive	0.934	1.082	1.747	1.560
	(0.135)	(0.107)	(0.119)	(0.010)
Nondurable manufg.	−1.119	−0.329[ns]	−1.118	−0.522
	(0.173)	(0.138)	(0.148)	(0.126)
Miscellaneous services	0.461[ns]	0.619[ns]	−0.719[ns]	−0.154[ns]
	(0.386)	(0.302)	(0.331)	(0.277)
Professional services	2.695	0.356[ns]	1.614	0.055[ns]
	(0.164)	(0.143)	(0.144)	(0.130)
Proportion of labor		6.230		4.644
force unemployed		(0.248)		(0.237)
Proportion of male working		2.514		2.264
population underemployed		(0.140)		(0.128)
Proportion of female headed			5.774	3.842
families with children			(0.194)	(0.173)
Intercept	−5.076	−4.589	−4.832	−4.594
Adjusted R-square	0.616	0.767	0.721	0.807

[a] The numbers in parentheses are standard deviations.
[ns] Not significant at the 0.01 level.

This also means that the diffusion of urban economic activities to the hinterland will significantly reduce but not eliminate rural child poverty. Indeed, as a final objective, we examine how child poverty rates would hypothetically change if nonmetro areas were more economically developed (i.e., industrial and employment patterns paralleled metro areas). Table 5 reports the results of a regression standardization, one in which metro social and economic characteristics are substituted for nonmetro characteristics (Clogg and Eliason 1986). The observed mean log-odds of poverty for nonmetro children is

Table 4. Logit regression models of the proportion of children in poverty in metro counties (N = 753)

Variable	(1)	(2)	(3)	(4)
Proportion black	2.027	1.191	−0.144[ns]	−0.049[ns]
(non-Hispanic)	(0.109)[a]	(0.096)	(0.133)	(0.114)
Proportion Hispanic	1.404	0.612	1.125	0.603
	(0.077)	(0.069)	(0.062)	(0.060)
Proportion with a high	3.103	1.517	1.865	1.063
school education or less	(0.153)	(0.134)	(0.133)	(0.120)
Employed in the county	1.051	0.670	0.834	0.628
	(0.065)	(0.050)	(0.052)	(0.043)
South	−0.017[ns]	0.104	0.140	0.180
	(0.025)	(0.019)	(0.021)	(0.017)
Industry				
Extractive	0.898[ns]	−0.027[ns]	2.219	1.075
	(0.373)	(0.274)	(0.300)	(0.247)
Nondurable manufg.	−1.625	0.169[ns]	−0.630[ns]	0.435[ns]
	(0.416)	(0.318)	(0.331)	(0.275)
Miscellaneous services	−1.347	0.443[ns]	−2.173	−0.502[ns]
	(0.506)	(0.374)	(0.400)	(0.329)
Professional services	1.145	0.097[ns]	−0.252[ns]	−0.504
	(0.243)	(0.195)	(0.202)	(0.73)
Proportion of labor		10.832		8.614
force unemployed		(0.688)		(0.610)
Proportion of male working		3.560		2.515
population underemployed		(0.291)		(0.260)
Proportion of female headed			5.368	3.485
families with children			(0.251)	(0.219)
Intercept	−4.706	−4.777	−4.310	−4.496
Adjusted R-square	0.764	0.867	0.843	0.901

[a] The numbers in parentheses are standard deviations.
[ns] Not significant at the 0.01 level.

−1.185. This suggests that the probability of poverty among children living in nonmetro areas is about 31 percent.[9] The observed mean log-odds of poverty for metro children is −1.475. Children living in metro areas, on average, have about a 23 percent probability of living in poverty.

Our analysis gives the nonmetro means of the log-odds of poverty, standardized on various metro characteristics. These results indicate the extent to which these new or ‘standardized’ means differ from the observed metro mean. For example, the replacement of observed means in nonmetro counties

Table 5. Standardized means (log-odds of poverty) under varying assumptions

Assumptions for standardization	Standardized means for nonmetro children
Observed mean[a]	−1.185
(nonmetro minus metro)	(0.290)
Metro baseline means	
(rest of means are nonmetro)	−1.495[b]
nonmetro constant	(−0.020)
Metro employment means	
(rest of means are nonmetro)	−1.273
nonmetro constant	(0.202)
Metro family means	
(rest of means are nonmetro)	−1.015
nonmetro constant	(0.460)
Metro baseline and employment	
(rest of means are nonmetro)	−1.584[b]
nonmetro constant	(−0.109)
Metro baseline and family	
(rest of means are nonmetro)	−1.326
nonmetro constant	(0.149)
Metro employment and family	
(rest of means are nonmetro)	−1.104
nonmetro constant	(0.371)
All metro means	−1.415
nonmetro constant	(0.060)

[a] The mean log-odds of poverty for metro children is −1.475. Differences between the standardized means and the metro means appear in parentheses.
[b] Adjusted means that place nonmetro below metro.

with the baseline means of metro characteristics (model 1 results) suggests that metro-nonmetro differences would be virtually nonexistent. The substantive implication is that nonmetro child poverty would be much like metro child poverty if nonmetro areas had the same industrial structure (and demographic composition) as metro areas. Moreover, adjustment for employment differences (model 2) yields an adjusted nonmetro mean log-odds of child poverty (i.e., −1.584) that is actually lower than the −1.475 observed in metro counties. Clearly, the current industrial mix and the lack of employment opportunities place nonmetro children at a comparative economic disadvantage and exacerbate the child poverty problem.

Discussion and conclusions

A disproportionate and growing share of American children are poor (Hogan & Lichter 1995). The changing economic circumstances of children are linked in obvious ways to parental employment patterns and family living arrangements (e.g., single-parent families). At the same time, employment opportunities and family patterns are typically discussed or studied in isolation from the geographic context or local labor market in which they are embedded. Surprisingly little attention has been given to how industrial restructuring is linked to spatial economic inequalities among America's children. Our study addresses this void.

We showed that child poverty is distributed unevenly over geographic space. Child poverty is concentrated in economically depressed areas, such as Appalachia, the Mississippi Delta, and the southern 'black belt'. More important, our analysis of spatial inequality clearly reasserted the primacy of local labor market conditions in influencing county child poverty rates. Local industrial structure has direct effects on child poverty (presumably through parental wages), as well as indirect effects through parents' ability to find work and to maintain a stable family life for their children. Children's lives – and their economic circumstances – are tied to the fortunes of the communities in which they reside. Policies that promote community economic development are anti-poverty policies (Madden 1996).

This fact is too often ignored in current welfare reform debates that focuses on promoting work among single mothers and the two-parent family ideal (Corbett 1993). Indeed, our results raise fundamental issues about whether public policy and anti-poverty legislation should be directed at impoverished families and individuals or at impoverished communities. Over the past decade or so, a shift in perspective away from community (e.g., community block grants, community action, etc.) toward the individual (e.g., employment and training programs) has occurred. Our results imply that a balanced approach is now needed. Although community economic development efforts sometimes fail to benefit the indigenous population most in need (i.e., a rising tide may no longer lift all boats), our results clearly suggest that economic development (i.e., as reflected in industrial structure) affects child poverty rates by promoting better wages, more employment opportunities, and 'stronger' families.

Our analysis also shows that the mix of policies that target poor communities and poor families may have different benefits for children in different parts of the USA. This conclusion was supported in our disaggregated analysis of child poverty in nonmetro and metro counties. Overall, the results supported previous research showing that children in nonmetro areas are more economically disadvantaged than their metro counterparts, largely because job opportunities are quantitatively and qualitatively different in rural than urban labor markets (Lichter & Eggebeen 1992). In rural areas, welfare reform aimed at promoting work among rural single mothers is unlikely to be successful in the absence of jobs that pay a family wage.

The substantive implications of our results are straightforward. If the types of well-paying, stable jobs found in metro areas also existed in nonmetro counties, nonmetro child poverty would be reduced substantially. Differences in the etiology of child poverty across geographic space implies different policy prescriptions for different areas. Policies that promote industrial diversification may be a positive step that ultimately lowers nonmetro child poverty rates. At the same time, policies that 'strengthened' the family may be more efficacious in urban than in rural areas. One policy implication is that recently proposed welfare and economic development programs, funded through the mechanism of state block grants, may allow new anti-poverty initiatives to be tailored to the specific needs and circumstances of poor children. They may also create greater geographic balkanization among children as resource-rich states further distance themselves from resource-poor states in welfare generosity and the effectiveness of economic development strategies.

In the final analysis, our study provides a point of departure for additional studies of child poverty and growing spatial economic differentiation. The uneven geographic distribution of child poverty clearly implies that children's economic circumstances are tied directly to the economic fortunes of the communities in which they and their parents live. This is a seemly self-evident assertion, but one rarely acknowledged in previous empirical studies of child poverty. This is particularly unfortunate at a time when the apparently growing economic stratification among states, counties, communities, and neighborhoods threatens to transform America's political landscape (Frey 1995; Lichter 1993).

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Notes

1. For example, the poverty rate in Alabama was 31.7% (for 1983–1987), but only 6.4% in New Hampshire. At the same time, the percentage of children living in a single-parent

family was 27.0% (in 1980) in Alabama, but 17.3% in New Hampshire. The child poverty in Alabama is almost 5 times that of New Hampshire, yet the proportion of children living in a single parent families was less than 0.5 times that of New Hampshire.

2. Originally, the following grouping of industries was considered to represent the industrial structure: extractive, nondurable manufacturing, durable manufacturing, trade, transportation, miscellaneous services, and professional services. Because the sum of the proportion employed in all these industries adds to one, one of these categories had to be eliminated from the model. Durable manufacturing was eliminated because it was responsible for more collinearity problems than the other industrial categories. After entering the other variables into the model, the variance inflation factors and the condition indices returned to unacceptably high levels. Several models were run eliminating one variable at a time and then groups of variables. Multicollinearity was significantly reduced by eliminating the industrial categories of trade and transportation.
3. The drawback of this underemployment measure is that it does not take into account the underemployment due to low-wages (Lichter & Landry 1991). The direct effect of industries on the proportion of children in poverty, however, is likely to compensate for the crudeness of this measure. A significant, positive effect of an industry category, net of unemployment and underemployment effects, suggests that the low-wages characterized by the particular industry are probably affecting poverty rates through this type of underemployment.
4. The dependent variable is the proportion of children living in the county who were considered to be poor in 1989 (call this p). Because this measure is a proportion, all the responses fall between zero and one. In order to linearize the relationship between our dependent and independent variables, the logarithm of $(p/1-p)$ is taken. This logit transformation does not address the issue of unequal error variances associated with a bounded dependent variable. In order to correct for this heteroskedasticity, the dependent variable is weighted by the number of children in the county $(n) * p * (1-p)$. Weighted least squares regression is used for the modeling in this study because it is appropriate for bounded dependent variables of this sort.
5. In a regression model of $\ln(p/1-p)$, each independent variable has nonlinear effects over the range of the dependent variable, i.e., the county child poverty rate in this case. A common practice is to evaluate the effects of a unit change in the independent variable at the mean of the dependent variable (in this case 0.213). The effect is calculated as: $(p)(1-p)(b_i)$, where p is the mean proportion of children who are poor and b is the logit coefficient.
6. $e^{-0.061} = 0.9418$; since the metro coefficient is coded one for metro status and zero for nonmetro status, the chances of children living in poverty in nonmetro counties is 5.8 percent, i.e., $[(1-0.9418) * 100 = 5.8]$ greater than for children living in metro areas.
7. It is commonly assumed that the cost of living differs between metro and nonmetro areas. Fuguitt, Brown & Beale (1989) suggest that the income differential between metro and nonmetro families likely to have children under 17 is about 10 percent. Therefore, even adjusting for cost of living differences, we do not expect the metro-nonmetro gap in child poverty rates to be eliminated.
8. To identify statistically significant differences between the coefficients in the nonmetro and metro models, z scores are calculated as:

$$z = (B_n - B_m)/\text{sqrt}[s^2(B_n) + s^2(B_m)]$$

where $z \sim N(0,1)$ under H_0: $B_n = B_m$; B_n is the estimated coefficient (say for the proportion of the population with a high school education or less) in the nonmetro model; B_m is the estimated coefficient in the metro model; and the s^2s are the respective squared standard errors of the coefficients in the nonmetro and metro models (Lichter & Landry 1991).

9. The chances of nonmetro children living in poverty are different from the prevalence rate (0.230) because the regression standardizations are based upon coefficients and means obtained from weighted least squares regression models. The difference between the chances of nonmetro children living in poverty and the chances of metro children living in poverty, here, is about 8 percent. This corresponds to the difference between the nonmetro and metro prevalence rates of child poverty (0.230 vs. 0.158). Therefore, these results do not appear to be problematic.

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