Institute for Research on Poverty
Discussion Paper no. 1106-96

The Effect of Neighborhood Characteristics on Young Adult Outcomes: Alternative Estimates

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September 1996

Abstract

We estimate a set of alternative models to examine the effect of neighborhood characteristics on outcomes among young adult women. The models are motivated by a concern that standard estimates of neighborhood effects may in part reflect the characteristics of families that reside in those neighborhoods. In addition to a “standard” model that includes controls for family background, we estimate fixed-effect models that also control for unobservable family characteristics that may affect young adult outcomes. To do this, we use a sample of sisters drawn from the Panel Study of Income Dynamics. In models that control for family background, we find evidence of neighborhood effects consistent with other recent work. In the fixed-effect models, however, there are no statistically significant effects that are consistent with standard hypotheses about neighborhood effects. The findings from this exploratory study suggest that one should be cautious about accepting findings of significant neighborhood effects derived from models that do not account for the possible selection of neighborhood.

The Effect of Neighborhood Characteristics on Young Adult Outcomes: Alternative Estimates

Spurred by William Julius Wilson's (1987) theory about the causes of ghetto poverty and underclass behavior, social scientists have renewed their attention to the question of whether neighborhood and peer characteristics affect the social and economic outcomes experienced by adolescent and young adults. Recent studies, such those of as Brewster (1994), Brooks-Gunn, Duncan, Klebanov and Sealand (1993), Case and Katz (1991), Crane (1991), Duncan (1994) and Sucoff (1996), report that a variety of such characteristics influence completed schooling and the likelihood that a teenager will drop out of school, not join the labor force, become sexually active or a parent, become involved in gang or criminal activity, or use drugs and alcohol.

As Aaronson (1995), Evans, Oates and Schwab (1992), Jencks and Mayer (1990), Tienda (1991), and several of the papers cited above observe, one may challenge these (and earlier) findings on a variety of methodological grounds. A central methodological concern is the difficulty of separating exogenous neighborhood effects from the effects of unobservable family characteristics that may be causally associated with neighborhood characteristics. Parents choose their children's neighborhoods; their choice is partly influenced by their observable and unobservable characteristics; their children's outcomes will be influenced by all of these factors. Hence, findings based upon methods that do not account for selection of the neighborhood may be biased. In a thoughtful and striking demonstration of this possibility, Evans et al. (1992) find significant peer-group effects on teenage pregnancy and the decision to drop out of school with a model that ignores this endogeneity, then show that these effects disappear when the endogeneity is taken into account.

In this paper, we use family and neighborhood data from the Panel Study of Income Dynamics (PSID) to estimate alternative models of the effect of neighborhood characteristics on outcomes among young adult women. In addition to a standard model that includes extensive controls for family background, we estimate fixed-effect models that also control for unobservable family characteristics that may affect young adult outcomes. We examine three important behaviors and outcomes among young adult women: whether a woman had a nonmarital birth, whether she obtained any postsecondary education, and her economic status, measured by her income-to-needs ratio. We estimate neighborhood effects in three specifications: with no controls, to obtain gross effects; with controls for observed family background variables, to obtain conventional net effects; and finally, with fixed-effects methods, to control for unobserved family effects. We focus on four measures of neighborhood quality from the 1970 and 1980 censuses, measured at the census tract level.

In the models that control for observed family background, we find evidence of neighborhood effects consistent with other recent work. The estimated neighborhood effects from fixed-effect models are much weaker, and we find no statistically significant effects that are consistent with standard hypotheses about neighborhood effects. These exploratory results suggest, therefore, that conventional estimates may overstate the effects of neighborhoods on young adult outcomes.

RECENT RESEARCH ON NEIGHBORHOOD EFFECTS

Jencks and Mayer's (1990) literature review identifies several mechanisms through which neighborhood characteristics might affect social behavior and outcomes. "Epidemic" or "contagion" models (Crane, 1991) indicate that a person's likelihood of engaging in a behavior, be it pro- or antisocial, is positively related to exposure to others engaged in the same behavior. "Collective socialization" models (Wilson, 1987) indicate that a neighborhood's adults provide role models for the neighborhood's children and monitor their behavior. "Institutional" models indicate that schools, police, social service agencies, and other local institutions affect children's behavior. In contrast, "competitive," "relative deprivation," and "cultural conflict" models conclude that more advantaged neighbors might be harmful to less advantaged children. For example, less advantaged children will tend to get relatively worse grades in schools largely populated with more advantaged children, and hence, may be more likely to drop out.

Jencks and Mayer (1990) conclude that “the [pre-1990] literature ... does not ... warrant any strong generalizations about neighborhood effects” (p. 176). More recent studies by Brooks-Gunn et al. (1993), Case and Katz (1991), Crane (1991), Duncan (1994), and Osterman (1991) tend to support the view that neighborhood or peer effects characteristics influence some outcomes and some groups and operate in ways consistent with either the contagion or collective socialization models, though broad generalizations have yet to emerge from this line of research.

The common methodology in these studies is to regress an outcome on personal and family background variables and one or more neighborhood variables. The coefficients on the neighborhood variables are the neighborhood effects.

Such an approach, however, may not adequately capture the effect of unobservable characteristics that affect the way parents choose their communities. If, for example, parents who choose to live in disadvantaged neighborhoods tend to be parents who do not strongly encourage schooling or have few contacts in the labor market, an estimated neighborhood effect may in part reflect omitted family background variables. Estimation methods that do not take account of this will attribute the effect of the omitted background variables to the neighborhood. This is more likely when only a restricted set of background variables is available to the investigator. Parallel comments apply to the estimation and interpretation of observed peer-group variables.

Three recent studies suggest we take seriously the possible spuriousness of estimated neighborhood effects. Corcoran et al. (1992) use PSID data, with zip-code-level census data on four community characteristics, to analyze 25- to 32-year-old men’s earnings, wage rates, hours of work and family income. After controlling for an unusually wide array of personal and family background variables, they find that the percentage of community families receiving welfare is negatively associated with most outcomes, but that the other three community characteristics generally have negligible effects on all outcomes. The rich set of background variables may have washed out spurious neighborhood effects by largely eliminating the bias likely to arise when there is a limited set of background variables (i.e., owing to omitted variables).[1] In another study using the PSID, but with census tract data, Ginther, Haveman, and Wolfe (1993) find that neighborhood characteristics lose statistical significance as models include increasingly more family background variables.

Evans et al. (1992) argue that the choice of peer group is likely to be endogenous and, hence, simple regression models produce biased estimates. Their single-equation models of teenage pregnancy and school dropout show significant peer effects. With a simultaneous-equation model, the effects disappear.

Recent research, then, provides tantalizing evidence that the social and economic characteristics of neighborhoods have important impacts on youths' lives and the likelihood that they will become productive adults. Yet there are plausible methodological grounds to suspect the evidence.

EMPIRICAL PROCEDURES, DATA, AND MODEL SPECIFICATION

Methods. The model typically used to estimate neighborhood effects is of the general form

$$Y_i = \beta Z_i + \gamma N_i + \eta_i \, , \tag{1}$$

where Y is some outcome measure of interest, Z is a set of individual and family variables that affect Y, N is a measure of neighborhood quality, η is an error term, and β and γ are the corresponding parameter estimates.

There are a number of potential problems in this simple model. As Evans et al. (1992) note, N may be endogenous, since it reflects the location decisions of the parents. A related problem, on which we focus, is the presence of unobservable family effects that may affect the outcome and be correlated with neighborhood characteristics. Let F_j be the effect on outcome Y of growing up in family j, net of the observed family characteristics included in Z. For example, in a model of educational attainment, F might represent educational aspirations and encouragement of the parents. If F affects Y but is unobserved and thus omitted from the model, estimates will be biased as long as F and N are correlated. Suppose that more effective families (those with a higher value of F) and neighborhoods of better quality both affect an outcome in the same direction. Then, if more effective families tend to locate in better quality neighborhoods, the estimate of γ from a conventional model will be biased upward in absolute value, since it incorporates some of the impact of the unobserved family effect, F.

A standard approach to a problem like this is to use fixed-effects methods to eliminate the effect of F_j, and estimate the model by using multiple observations on Y, Z, and N for a given F_j. Fixed-effect models have been used, for example, to measure the effect of a teen birth on socioeconomic outcomes (Geronimus and Korenman, 1992; Hoffman, Foster, and Furstenberg, 1993) as well as in other contexts (Ashenfelter and Krueger, 1994, Griliches, 1979). There are, however, no published applications of this method to the estimation of neighborhood effects.[2] In this application, fixed-effect estimation requires information on two or more siblings per family. Identification of γ is then achieved by differences in N between or among siblings who share a common value of F_j but different values of N.

<u>Data and Model Specification</u>

Our data come from the Panel Study of Income Dynamics (PSID). The PSID is a nationally representative survey of approximately 7,000 families, who have been interviewed annually since 1968. The PSID is well-suited for sibling analyses because all individuals who resided in PSID households in 1968 are sample members and remain part of the sample when they form independent households. As a result, the sample includes a large number of young adults with siblings who were originally sampled as children in PSID households. Although the data set has some limitations that we note below, it does include the key elements necessary for our analyses—information on neighborhood characteristics and on socioeconomic outcomes for a sizable sample of sisters.

The measures of neighborhood quality come from the 1970 and 1980 censuses, which have been matched with addresses in the PSID at the census-tract level.[3] We use four measures of neighborhood characteristics similar to those considered in previous research: the percentage of families with children headed by a single female; the percentage of families receiving public assistance; the percentage of low-income families (< $5,000 in 1970, $10,000 in 1980); and the percentage of middle/upper income families (> $15,000 in 1970, $30,000 in 1980).[4] These income levels identify approximately the bottom 20 percent and the top 25 percent of all families.

To analyze the effect of neighborhood attributes on outcomes, we focus on a sample of young women with sisters who are also part of the PSID sample.[5] The outcome measures we use are standard measures of attainment for young women: whether a woman had a nonmarital birth, whether she obtained any postsecondary schooling, and her income-needs ratio, which measures her own earnings and those of her spouse, if she is married, relative to the poverty threshold. The contagion and socialization models suggest that higher values on the measures regarding female family head, public assistance, and low income will be associated with a higher likelihood of a nonmarital birth, a lower likelihood of obtaining postsecondary schooling and a lower income-needs ratio, while higher values on the income measure will be associated with each outcome in the opposite way.

Ideally, we would like to have information on neighborhood characteristics over a substantial portion of each woman's childhood, so that we could construct a measure of average childhood neighborhood quality or even examine at what ages neighborhood quality mattered most. Unfortunately, we are constrained by the data available in the PSID and also by the need to measure outcomes at a meaningful young adult age. Because the PSID does not contain any locational information prior to 1967, it is not possible to construct neighborhood information at young ages for many individuals. Even if that information were available, the relevant years would be sufficiently distant from 1970 to make the 1970 census tract data misleading. Restricting the sample to persons for whom earlier ages fall after 1967 is possible, but this would reduce the sample size substantially and many of them would still be quite young at the point outcomes are measured.

We have, therefore, chosen to measure neighborhood quality in the later teen years—as an average across the census tracts in which each woman lived when she was aged 16 to 18. Further, we limit the sample to women who were between ages 25 and 36 in 1987 (the year in which we measure outcomes)[6] and who had one or more sisters in the sample. This age range means that ages 16–18 fall between 1967–69 for the oldest women and 1978–80 for the youngest women. With these restrictions, we are left with a sample of 614 women from 265 families. Within-family variation in neighborhood characteristics of sisters can exist either because the family moved or because the neighborhood changed over time, and as a result a younger sister experienced a different neighborhood from that of an older sister.[7]

To estimate the effect of neighborhood characteristics on adult outcomes, we use three specifications. First, we estimate a model using only a constant and one of the neighborhood measures to obtain a gross effect. Then we add a set of major personal and family background variables, including whether the mother and father were high school graduates, whether the respondent grew up primarily in a two-parent family, her race (black or not), her age in 1987, and the parental family's average income-to-needs ratio, measured over the teenage years. This yields an estimated neighborhood effect, net of measured background. Finally, we specify a family fixed-effects model that estimates the neighborhood effect based on the within-family differences in outcomes, neighborhood characteristics, and personal characteristics that can vary among sisters (i.e., age, mean income-to-needs during her teen years, and being raised primarily in a two-parent family).

To keep the exposition manageable, we present results for a model in which the neighborhood effects enter in a linear fashion. This constrains the effect of a one-unit change in a neighborhood characteristic to be independent of the level of the neighborhood characteristic. We have also estimated models with nonlinear effects, using a spline specification in which the effects are permitted to vary above and below the mean. Because the basic findings are so similar across models, we do not present those estimates here.[8]

We estimate the models using OLS regression and fixed-effects regression methods for the income-needs ratio and logit and fixed-effects logit for the two dichotomous variables. This fixed-effects logit model is Chamberlain's conditional logit model. As Chamberlain (1980) shows, only families in which there is within-family difference in outcomes contribute to the likelihood function. This reduces the sample size, because observations from families in which all sisters have the same outcome must be dropped. The computational formulas for the elements of the likelihood function are shown in Maddala (1987), and differ according to the number of sisters in the family and the number of sisters with the particular outcome. In the simplest case of a family with two sisters, the probability that the second sister has a particular outcome and the first does not is

$$\text{Prob}\ (0,1) = \exp[\beta(Z_1 - Z_2) + \gamma(N_1 - N_2)] \ / \ [1+\exp(\beta(Z_1 - Z_2) + \gamma(N_1 - N_2))], \qquad (2)$$

where the subscripts denote the two sisters. The corresponding terms in the likelihood function for families with three or more sisters are constructed in a similar, but more complex, way, with a distinct term for each possible combination of outcomes.[9] Note that this is a standard logit probability in which the explanatory variables are in difference form. The effect of any variable that does not differ between sisters (for example, race or mother's education) cannot be estimated, nor are there estimates of the fixed effects.

In most previous fixed-effect logit estimations (see, e.g., Geronimus and Korenman, 1992; Hoffman, Foster, and Furstenberg, 1993), researchers have selected two sisters per family, regardless of family size, constructed the terms in (2) manually, and then estimated the model via standard logit. Using two sisters in larger families causes two potential problems: first, the choice among sisters is often arbitrary; second, restricting the number of sisters tends to reduce within-family variation in the dependent variable, which is important when the dependent variable is dichotomous. We take advantage of new software to estimate fixed-effect logit models using all sisters in each family.[10]

FINDINGS

Table 1 provides basic descriptive information about the sample. The figures shown in the table are weighted by the inverse of the PSID selection probabilities, and thus are population estimates. More than 40 percent of the women obtained some postsecondary education, 27 percent had a nonmarital birth, and the mean income-needs ratio as young adults was 3.78. The middle panel shows the family background variables we use, and the bottom panel shows the neighborhood measures. The mean woman lived in a neighborhood where one-seventh of the families with children had a female head, 7 percent of families received public assistance income, one-fifth of families had low incomes, and nearly one-fourth were in the middle/upper-income category.

Table 2 shows that substantial variation in neighborhood characteristics occurs within families. As indicated by the first row, within-family variation exists in about two-thirds of the families according to each neighborhood measure. The next two rows show the average within-family difference for all families and for those where the difference is non-zero. For families with two sisters, the within-family difference is simply the arithmetic difference between them in their neighborhood measures; for larger families, it is computed as the simple average across all pairs of sisters.[11] When they exist, mean within-family differences are reasonably large, ranging from 22 to 69 percent of the sample means shown in Table 1.

TABLE 1

Sample Characteristics, Women Age 25–37 in 1987 with Sisters, Panel Study of Income Dynamics

Dependent and Independent Variables	Mean[a]
Outcomes	
Postsecondary education	.43
Nonmarital birth	.27
Income-needs ratio	3.78
Personal and Family Characteristics	
Mother's education < high school	.37
Father's education < high school	.42
Grew up in two-parent family	.78
Age	31.1
Black	.17
Family income-needs ratio[b]	3.04
Neighborhood Characteristics	
Female-headed families with children(%)	13.85
Families receiving Public Assistance (%)	7.27
Low-income families[c] (%)	20.56
Middle/upper-income families[d] (%)	23.36
Sample size: individuals	614
Sample size: families	265

[a]All means computed using PSID sample weights.
[b]Measured for parental family, averaged over teen years.
[c]Families with income < $5,000 in 1970 or < $10,000 in 1980.
[d]Families with income > $15,000 in 1970 or > $30,000 in 1980.

TABLE 2

Overall and Within-Family Variation in Neighborhood Characteristics

	Neighborhood Characteristic			
	% Female-Headed Families	% Families Receiving Public Assistance	% Low-Income Families	% Middle/Upper-Income Families
Families with within-family difference	69.5%	64.5%	69.9%	67.7%
Mean within-family difference	5.6	3.2	4.9	3.5
Mean within-family difference (if > 0)	8.1	5.0	7.0	5.2

Table 3 provides the key regression findings for the neighborhood variables. To conserve space, we present only the estimated coefficients for the neighborhood variables.[12] Before discussing them, we note that estimates of the effects of the personal and family background variables in the regressions were consistent with the previous literature. More parental education and family income increased the income-needs ratio and the probability of postsecondary schooling, and decreased the probability of a nonmarital birth. Race (black) had a statistically significant negative effect on the income-needs ratio and a positive effect on the other two outcomes. Growing up primarily in a two-parent family had a significant negative effect on the schooling and nonmarital birth outcomes and a positive one on income-to-needs.

Consider first the very simple models in row 1 of each panel. Every neighborhood characteristic is statistically significant with the anticipated sign in all 12 models. Being raised in a neighborhood with a higher proportion of female-headed families is associated with a lower income-needs ratio in adulthood, a lower probability of obtaining postsecondary schooling, and a higher probability of a nonmarital birth. Neighborhoods with relatively more families on public assistance or with low incomes show the identical set of associations with the outcomes. And all the neighborhood effects are significant, with reversed signs, for the middle/upper income neighborhood characteristic.

The various neighborhood effects have an interesting pattern of gross effects. The proportion of families receiving public assistance has the biggest estimated effect for two of the outcomes and is just .001 below the highest effect for the other outcome. Except for postsecondary schooling, the effects of the two income measures are equal and opposite; for schooling, the effect of living with middle/upper income neighbors is 67 percent larger in absolute value than the effect of low-income neighbors.

Further, the coefficients yield reasonably large quantitative effects. Suppose, for example, that we compare predicted outcomes for women from neighborhoods that differ by ten percentage points.[13]

TABLE 3

**Effect of Neighborhood Characteristics on Adult Outcomes
(Standard Errors in Parentheses)**

	Neighborhood Characteristic			
Dependent Variable and Model Specification	% Female-Headed Families	% Families Receiving Public Assistance	% Low-Income Families	% Middle/Upper-Income Families
Income-Needs Ratio				
No controls	-.066**	-.108**	-.062**	.060**
(N=614 persons)	(.010)	(.013)	(.028)	(.006)
Family background	-.021*	-.027*	-.017**	.028**
(N=614 persons)	(.011)	(.015)	(.008)	(.007)
Fixed-effect	.031	-.002	-.005	.005
(N=265 families)	(.028)	(.047)	(.022)	(.020)
Postsecondary Schooling				
No controls	-.024**	-.031**	-.019**	.032**
(N=614 persons)	(.009)	(.012)	(.006)	(.005)
Family background	-.021 **	-.003	.001	.016**
(N=614 persons)	(.011)	(.015)	(.008)	(.006)
Fixed-effect	-.005	.008	.010	.004
(N=107 families)	(.020)	(.032)	(.021)	(.035)
Nonmarital Birth				
No controls	.081**	.100**	.040**	-.036**
(N=614 persons)	(.011)	(.014)	(.006)	(.007)
Family background	.049**	.040**	.001	-.003
(N=614 persons)	(.012)	(.016)	(.009)	(.084)
Fixed-effect	-.013	.006	-.058**	.102**
(N=87 families)	(.022)	(.036)	(.028)	(.040)

* Statistically significant at 10% level.
** Statistically significant at 5% level.

For the income-needs ratio, this translates into a difference of about 0.6–1.0 points (or between $7,800 and $13,000, using the 1996 federal poverty guideline for a family of three). For postsecondary schooling, a ten-point difference yields a 5 to 7 percentage point difference in the predicted probability. For a nonmarital birth, the effects are about 20 percentage points for the female-headed and public assistance variables and less than ten points for the two income variables.

The estimates in row 2 of each panel are from models containing family background variables along with the neighborhood characteristic. As one would expect if neighborhood characteristics are correlated with family background, every estimated neighborhood effect falls in absolute value. The average decline is 67 percent. The estimated effects of the proportion of families receiving public assistance, which was the largest in the models without controls, fall quite substantially, as do the estimates for the proportion of low-income families. Still, of the 12 neighborhood variable coefficients, eight remain significant with the expected signs and with substantively important quantitative effects. Statistically significant neighborhood effects remain for all of the neighborhood measures for income-needs, for female-headed families and middle/upper-income families in the postsecondary schooling regression, and for female-headed families and public assistance families in the nonmarital birth regression. From such evidence one would tend to conclude that neighborhood characteristics do affect young women's outcomes, independently of family background.

Row 3 of each panel contains results from fixed-effect models. Note the sample sizes shown in the stub for each variable. For the income-needs ratio outcome, we can use all 265 families, but for the dichotomous schooling and nonmarital birth outcomes, only families with outcome differences among sisters can be included in the estimation. This reduces sample sizes substantially, to 107 families for the postsecondary model and 87 for the nonmarital birth model.

If unobservable characteristics affect the way parents select their communities and affect their children's outcomes, then both the magnitude and significance level of a neighborhood effect derived from a cross-section model that does not account for this unobserved heterogeneity would decline under a fixed-effect specification. Our findings strongly support this expectation. The estimated neighborhood effects differ markedly from those derived from models that control just for observed family background. *All eight of the statistically significant effects in the cross-section models become statistically insignificant.* There are also two anomalous effects for the nonmarital birth outcome—the two income class variables become significant, but with signs contrary to expectation.

These results suggest that neighborhood effects tend to disappear once family effects are taken into account. The pattern in the way the neighborhood effects change as we move from simple to more sophisticated models is consistent with the view that neighborhood characteristics may largely be proxies for unobserved family background characteristics that affect choice of neighborhood. The evidence from the family fixed-effect models suggests one should view with caution findings that neighborhood effects matter which are derived from standard cross-section models.[14]

We nevertheless think the findings in Table 3 should be viewed as tentative, for several reasons. The models estimated in this study are parsimonious, and, because of data limitations, each woman's neighborhood characteristics were measured only when she was between ages 16 and 18. If neighborhood characteristics have strong effects at younger ages or accumulate over many years, the neighborhood variables may be poorly measured. Note, however, that these two concerns apply to both the standard and fixed-effect models. The point to be emphasized is that even with these limitations, there are consistent neighborhood effects with the usual cross-section approach, but when the models are replicated using a fixed-effect approach, these effects evaporate.

We also note two limitations to the fixed-effect approach. The small sample sizes in two of the fixed-effect models make it difficult to estimate effects reliably. This is an inevitable problem of fixed-effect models of dichotomous variables where there is limited within-family variation. Yet the same findings appear in the income-needs ratio model, which has a much larger sample size. Also, the fixed-effect approach assumes that a parental decision to move to a different neighborhood is exogenous to the ways in which the characteristics of the new neighborhood affect their daughters. To the extent this assumption is inaccurate, the fixed-effect estimates may be biased.

SUMMARY AND CONCLUSION

We used sisters from the PSID to examine the extent to which estimates of neighborhood effects are robust to control for observable and unobservable family background characteristics. We examined three outcomes for young adult women and four neighborhood variables commonly used in the literature.

In models that do not control for unobserved family effects, we find evidence of neighborhood effects consistent with other recent work. For example, the percentage of families with children that are headed by a single female has significant effects on all three outcomes.

None of these results is robust, however, under fixed-effect methods. In the fixed-effect models, there are no statistically significant results that are consistent with standard hypotheses about neighborhood effects. Every statistically significant neighborhood variable in the cross-section models becomes insignificant in the fixed-effect models. Some become significant, but in directions contrary to expectation. The findings from this exploratory study suggest, therefore, that one should be cautious about accepting findings of significant neighborhood effects derived from cross-section models which do not account for the possible selection of neighborhood. Put differently, the effects of families may be substantially more important than the effects of neighborhoods. Evans et al. (1992) reach a similar conclusion using a different methodology for dealing with selection based on unobserved family characteristics.

This study examined only three of a large set of social behaviors and outcomes for which neighborhood characteristics may be important determinants. Findings for other dependent variables and further methodological exploration are needed before we will have an adequate understanding of the relationships among parental characteristics, neighborhood choice, and the influence of neighborhoods on social behavior.

Endnotes

[1]The failure to find neighborhood effects as strong as those in studies which used tract-level data might reflect the crudeness of zip-code-level data for measuring neighborhood characteristics. Or perhaps neighborhoods have less influence on men who are 25 to 32 years old than on adolescents.

[2]Aaronson (1995) takes an approach similar to ours in a study of neighborhood effects on educational outcomes.

[3]For addresses in geographic areas that were tracted in the 1980 census, tracts are identified, where possible, using the GBF-DIME address listing file. Nearly all of the land in metropolitan areas is tracted, and a few states are fully tracted in nonmetropolitan areas as well. In most states, some nontracted areas (mainly cities in nonmetropolitan areas) were blocked. Aggregations of blocks, called Block Numbering Areas (BNAs), provide a reasonable analog to tracts. In untracted and unblocked areas, the enumeration district in which the address fell was identified. Enumeration districts are the basic work units of individual census enumerators and are probably somewhat less like a neighborhood than tracts or BNAs. They tend to be larger geographically than tracts, although they usually contain fewer people (an average of 600 in 1980). Addresses listed as post office boxes, rural route, or general delivery, or addresses in tracted areas that were too new to show up in the address listing file were assigned to minor civil divisions (MCDs), a much larger geographic unit. MCDs are typically townships, but state governments have, in widely varying degrees, also designated some cities and villages as MCDs.

[4]We also examined the effects of the poverty rate for nonelderly families and the proportion of residents who were black, but we do not present those results here, since they had insignificant effects in all multivariate specifications.

[5]Including brothers would be possible for some of the outcomes we considered, but it would complicate some of the comparisons.

[6]Using data for 1987 has several advantages for the sample and the variables we use. First, it permits us to use the fertility histories that were collected in 1985. Second, sample size is preserved. Using more recent years results in a substantially smaller sample of siblings because of sample attrition since, whenever one sister in a two-sister family becomes a non-response unit, the entire family is lost for the fixed-effect model.

[7]For all years between 1967 and 1974, we assigned the 1970 Census value for the appropriate Census Tract; for 1975 and thereafter, we used the corresponding 1980 Census values.

[8]Those estimates are available from the authors. We also experimented with three-segment splines (variously defined), logarithmic, and quadratic specifications to check the sensitivity of the estimates to functional form and to explore the importance of nonlinear relationships. None of these variations changed the basic tenor of the findings reported here.

[9]For the three-sister case, there are six terms in the likelihood function: three for the cases in which each of the sisters alone has the particular outcome and three more for each combination of two sisters with the outcome.

[10]Our 265 families include 197 with two sisters, 54 with three sisters, 12 with four, and 2 with six. The fixed-effect logit model is estimated using LIMDEP 7.0.

[11]For example, for the families with four sisters, there are six comparisons; for the two families with six sisters, there are ten.

[12]The full estimates for all 36 regressions are available upon request.

[13]Ten points is approximately equal to the standard deviation for the proportion of female-headed families, a bit greater than the standard deviation for the proportion receiving public assistance, and less than the standard deviation for the two income measures.

[14]Estimates from the spline function followed exactly the pattern as the linear effects, with, however, somewhat more indication of signs contrary to expectations.

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