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On the Empirical Finding of a Higher Risk of Poverty in Rural Areas: Is Rural Residence Endogenous to Poverty?

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Research shows people are more likely to be poor in rural versus urban America. Does this phenomenon partly reflect that people who choose rural residence have unmeasured attributes related to human impoverishment? To address this question, two models are estimated using Panel Study of Income Dynamics data. A single-equation Probit model of individual poverty replicates the well-documented finding of higher poverty risk in rural places. However, an instrumental variables approach, accounting for correlation between rural residence and the poverty equation error term, finds no measured effect of rural location on poverty. Results suggest failure to correct for endogeneity or omitted variable bias may overestimate the "rural effect."

Key words: endogeneity, instrumental variables, omitted variable bias, poverty, rural

Introduction

The incidence of poverty is higher in nonmetropolitan (nonmetro) than metropolitan (metro) areas and, as shown in figure 1, this phenomenon is not new.[1] Regression analyses also document a rural welfare disadvantage.[2] In the literature, rural-urban differences are often identified by including a binary variable for nonmetro residence as a regressor in empirical models of individual or household poverty (e.g., Brown and Hirschl, 1995; Brown and Lichter, 2004; Cotter, 2002; Haynie and Gorman, 1999; Snyder and McLaughlin, 2004; Thompson and McDowell, 1994). Analysts then attempt to control for important individual/household characteristics (e.g., race and education of the household head, family structure) and contextual factors (e.g., county unemployment rate and region of residence) that influence well-being. Extant research shows that the

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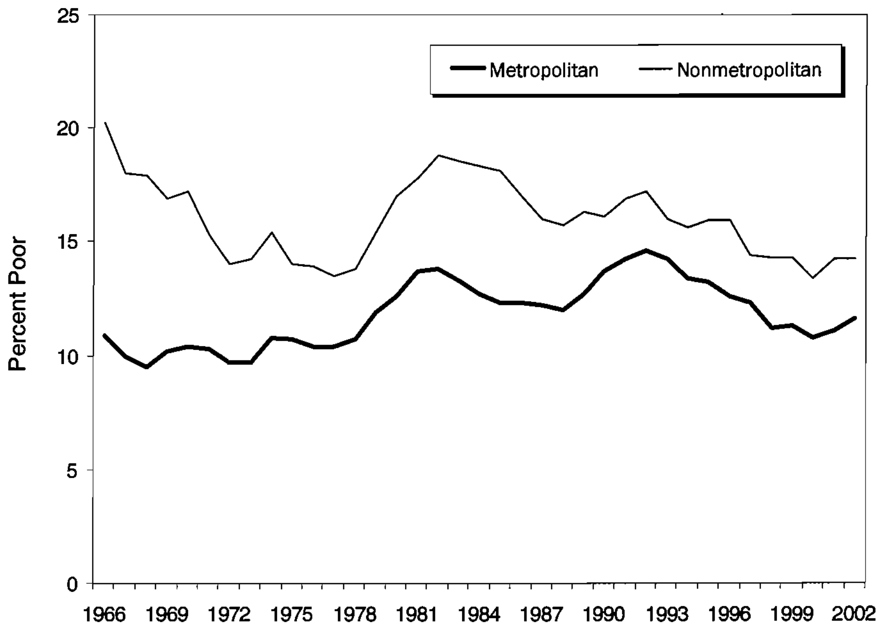
[1] The official poverty thresholds used to determine poverty incidence in the United States do not account for cost-of-living differences across space (e.g., region, metro/nonmetro county). Poverty analysts generally agree on the need to account for geographic cost-of-living differences, but data for such purposes are limited (Rural Poverty Research Center, 2003). It is expected that living costs are, on average, lower in rural versus urban locations, suggesting current measures of rural-urban differences in poverty prevalence could be biased. In fact, a recent paper by Jolliffe (2004) provides support for such a hypothesis. Jolliffe uses a spatial price index based on Fair Market Rents data. Accounting for cost-of-living differences across metro and nonmetro areas, he reports a complete reversal in the metro/nonmetro poverty rankings, with metropolitan poverty incidence being higher in every year from 1991–2002.

[2] The terms "nonmetro" and "rural" are used interchangeably in this paper to refer to counties outside of metropolitan areas.



Source: U.S. Bureau of the Census, *Current Population Survey* (annual March supplement).

Figure 1. People in poverty, by residence: 1966–2002

odds of being poor are between 1.2 to 2.3 times higher for individuals or households residing in rural compared with urban areas (Weber and Jensen, 2004, table 1).

While a strong positive correlation between rural residence and poverty is well documented in the literature, the direction of causation is indeterminate. Existing work assumes a causal link going from rural residence to poverty. Yet the positive association could partly reflect that poor people are attracted to rural areas, or are otherwise reluctant (or unable) to leave them. Furthermore, there may be unmeasured factors that determine both poverty and rural residence and, subsequently, the measured impact of rural residence may be partly spurious. In sum, the nature of the link between poverty and rural residence is an open question, and clearly merits further empirical exploration given its relevance to policy.

This study asks if the estimated rural effect on poverty partly reflects residential selection bias. Current models of rural poverty treat nonmetro residence as an exogenous variable. The validity of this assumption is questionable, because people have some degree of freedom to choose where they live. If people who decide to live in rural areas have unmeasured attributes which are related to human impoverishment, estimates of a rural effect can be biased. Consider, for example, that poverty models rarely control for whether an individual is geographically mobile. It is plausible that, compared to urban people, rural people are less mobile, having a preference for living close to their extended family and childhood friends. Geographic mobility may also be negatively correlated with poverty; those who are willing to move in search of employment may be less likely to be unemployed and poor. If mobility is negatively correlated with both poverty and rural residence, then the effect on poverty of living in a rural area could be overstated if one does not include a proxy variable for mobility in the empirical model.

This study uses Panel Study of Income Dynamics (PSID) data to investigate the extent to which endogenous rural residence biases estimates of a rural effect on poverty. Two empirical models are estimated for comparative purposes. First, a single-equation Probit model of person poverty is estimated in order to replicate the well-documented finding of a higher risk of being poor in rural compared to urban places. An instrumental variables approach is then used to account for the possibility that rural residence is endogenous to poverty. The basic question addressed is: Does a rural effect persist when one accounts for rural residential choice? or does such an effect disappear, suggesting it may be an artifact of residential sorting? While this investigation focuses on links between rural residence and the risk of being poor, findings are relevant to a general body of research which measures place effects on individual behavior and well-being.

Data and Sampling Issues

This study uses data from the 1993 and 1994 waves of the Panel Study of Income Dynamics (PSID), a longitudinal survey that has followed a representative sample of approximately 5,000 families and their descendants since 1968 [see Brown, Duncan, and Stafford (1996) and Hill (1992) for detailed descriptions of the PSID]. The PSID family and individual files contain data on a wide range of topics including family structure and demographics, socioeconomic background, geographic mobility, employment, earnings, income, wealth, welfare participation, housework time, health, and food security. Due to the enormous value of nationally representative longitudinal data on economic and social issues, the PSID is one of the most widely used data sets in the world. The PSID data set is particularly useful for this analysis because it provides, for public use, information on nonmetro/metro residence for certain years.[3]

The focus of this study is on a subsample of the PSID data consisting of 18,869 individuals in 1993. The choice of 1993 as the analysis year was made because it is the most recent year for which all of the required data for the analyses are available. In selecting this analysis year, two main factors are relevant. First, a variable for nonmetro/metro residence is not available for all years; such a variable is provided only in 1968–1993, 1999, 2001, and 2003. Second, structural condition variables, such as the county unemployment rate, are provided only until 1993. The individual is the appropriate analysis unit for the current study, given the interest in studying the extent of poverty. However, an understanding of the correlates of poverty requires consideration of family-level factors, since poverty is measured at the family or household level in the United States. For this reason, the analyses here attribute to each person the characteristics of her or his family—that is, explanatory variables in the empirical models are primarily family level. One problem with using the individual as the analysis unit is the introduction of clustering in the data. To deal with non-independence of observations within families, calculated standard errors are clustered on the PSID family identifier variable.

[3] The main national surveys used for poverty research are the PSID, the Current Population Survey (CPS), the Survey of Income and Program Participation (SIPP), the National Longitudinal Survey of Youth (NLSY), and the National Survey of America's Families (NSAF). The CPS, similar to the PSID, provides public-use access to data on metro/nonmetro residence. However, in the CPS, a number of observations are suppressed for these area variables in order to protect the anonymity of respondents.

Two main criteria are used to arrive at the final sample of 18,869 individuals. First, only individuals who resided in the United States during the survey year are included in the sample. Second, those observations with incomplete data are dropped for the analysis variables.[4] A key question is whether imposing these sample selection rules introduces sample selection bias. Table 1 provides descriptive statistics for the variables used in the analysis for the full sample (this is possible only for variables for which complete data exist) and for the subsample. Note that sampling weights and variables identifying stratum and sampling error computation units are used to take account of the PSID sampling design and differential attrition, and to approximate nationally representative estimates. The test statistics shown in the last column of table 1 enable hypothesis testing for differences in means or differences in proportions. At a 0.10 significance level, we can reject the null hypothesis that means/proportions are the same for the full sample and the subsample in the case of five of the eight variables: (*a*) householder is disabled, (*b*) female headed family, (*c*) married couple family, (*d*) number of children in the family, and (*e*) presence of a young child in the family. The subsample appears to differ from the full sample, under-representing families with a disabled and female head and those with more children, and over-representing married couple families. This should be kept in mind in the interpretation of findings in later sections of the paper.

Empirical Analysis

Base Model Results for Person Poverty

The analysis begins by estimating an empirical model of person poverty in which nonmetro residence is assumed to be an exogenous variable, paralleling the common practice of existing work. The model is a single-equation Probit model of the form:

$$p = \alpha_0 + \alpha_1 \mathbf{x} + \alpha_2 u + \alpha_3 \mathbf{s} + \alpha_4 n + \varepsilon_1. \tag{1}$$

The dependent variable p is a binary variable indicating whether the individual is poor, defined as having before-tax family cash income less than or equal to 100% of the family-size conditioned official poverty thresholds. Explanatory variables are defined as follows. The vector $\mathbf{x}$ represents a set of family-level characteristics, including the age of the youngest child and number of family members; householder characteristics of age, race, gender, marital status, education, work experience, and disability status; and a background variable indicating whether the household head grew up in a city of any size, as opposed to growing up in a rural area, town, suburb, or combination of places.[5] Place-level variables are the county unemployment rate u, indicator variables for the state of residence $\mathbf{s}$, and a binary variable n indicating whether the county of residence is nonmetro.

[4] The initial sample was approximately 28,000 individuals. About 8,000 observations were dropped due to missing family income data. About 100 observations were eliminated because the family did not reside in the United States. An additional 200 individuals were excluded from the subsample due to missing data for place of residence (region or metro/nonmetro county). Finally, about 1,000 observations were dropped owing to missing data for family head's main race, education, work experience, or religious preference.

[5] This variable was initially meant for use as an instrument to identify the poverty model in later sections of the paper. However, the variable is highly correlated with poverty, and therefore it is important to include it as an explanatory variable in the poverty model.

Table 1. Descriptive Statistics of Explanatory Variables, Full Sample and Subsample, 1993

Description of Variable		Full Sample (N = 28,324) Mean or Frequency / (Standard Error)[a]	Subsample (N = 18,869) Mean or Frequency / (Standard Error)[a]	Test Statistic[b]
Endogenous Variables:				
Poor (income ≤ official poverty threshold)[c]			0.1057	
Income-to-Needs Ratio (income/official poverty threshold)			4.3867	
County of residence is nonmetropolitan			0.2675	
Exogenous Variables:				
Age of householder (years)		44.8309 (0.2416)	45.1793 (0.2496)	-1.0031
Householder grew up in a city			0.3392	
Householder's education (years)			12.9544 (0.0778)	
Householder's work experience since age 18 (years)			17.7037 (0.2250)	
Householder is disabled		0.0307	0.0274	2.1121
Householder's main race is white			0.8516	
Family type:	Female headed	0.2210	0.2106	2.6682
	Male headed	0.0836	0.0846	-0.3654
	Married couple[d]	0.6954	0.7048	-2.1765
Number of adults in family		2.0165 (0.0166)	1.9981 (0.0172)	0.7676
Number of children in family		1.2059 (0.0213)	1.1524 (0.0235)	1.6891
Dummy for child < 6 years present in family		0.2656	0.2463	4.6860
County unemployment rate			7.2143 (0.1538)	
Identifying Instruments:				
Householder's first job is farmer			0.0230	
Householder's religion not well-represented in urban areas			0.0029	

[a] Means and standard errors are obtained using Stata's "svymean" command. The means are weighted by the PSID combined sample individual weight. To account for the stratified and clustered design of the PSID sampling procedure, standard errors are calculated using PSID stratum and sampling error computation units.

[b] The critical value ($\alpha = 0.05$) for the z-statistic (differences in proportions) and t-statistic (differences in means) is 1.96; the critical value ($\alpha = 0.10$) for the z-statistic (differences in proportions) and t-statistic (differences in means) is 1.65.

[c] The person poverty rate reported in the table is lower than the official poverty rate calculated using Current Population Survey (CPS) data, which is 15.1% for 1993. Stevens (1994) states that the consistently lower poverty rates calculated with PSID data appear to be the result of more thorough income reporting in the PSID compared with the CPS. For this reason, analysts employing PSID data sometimes use a more generous cut-off point, say 125% of the Census Bureau's official poverty threshold, in their poverty calculations (e.g., Iceland, 1997; Stevens, 1994). Following this approach, the measured person poverty rate with PSID data for 1993 is 14.4%, quite close to the official estimate.

[d] Household head is married or has a cohabitator with whom he/she has lived for at least one year.

The empirical model captures the main determinants of human impoverishment highlighted by poverty researchers (e.g., Rank, Yoon, and Hirschl, 2003; Schiller, 1995). One common view is that specific attributes of poor people, such as low levels of education or lack of competitive labor market skills, have brought about their poverty. From this individualist perspective, poverty is a consequence of individual decisions related to education, employment, and family structure; these decisions in turn have implications for economic well-being. Other observers argue that poverty is mainly the result of restricted educational, economic, and political opportunities. Restricted opportunities may be related to one's place of residence (e.g., neighborhood, county, region), or they may originate from discrimination on the basis of gender, race, or class. Thus, according to the restricted opportunity viewpoint, poverty is conditioned by forces beyond the control of individuals and families. These two explanations of poverty are here considered complementary, as reflected in equation (1).

Table 2 presents Probit results for the single-equation model of person poverty, reporting coefficients, robust standard errors, and marginal effects. Standard errors reported in the table use the Huber/White heteroskedasticity-consistent estimator of variance (Huber, 1967; White, 1980), and are adjusted for within-cluster (family) correlation with use of the PSID family identifier variable. Note that for binary variables, the marginal effects are interpreted as the percentage point change in the probability of poverty resulting from a discrete change in the explanatory variable. At standard test levels, all of the point estimates are individually significant. Parameter estimates for the variables *Age* and *Age Squared* indicate that age of the family head is negatively correlated with person poverty until the householder reaches the age of 44 years, at which point the correlation becomes positive. Results show a negative correlation between individual poverty and the householder's education and work experience. Findings suggest that people are more likely to be poor if they are part of a family headed by an individual who grew up in a city, has a temporary or permanent disability, is not white, and is not married. The risk of poverty is lower for individuals in families with more adult members and higher for individuals part of families with more children. Finally, place of residence appears to matter to poverty outcomes. Person poverty is positively correlated with the county unemployment rate and nonmetro residence. An important result of previous studies is confirmed—living in a rural area increases the risk of poverty, all else being equal. Nonmetro individuals have a 42% higher probability of being poor compared with people in metro areas.[6]

Testing for Endogeneity of Rural Residence

In the next section of the paper, a simultaneous equation model is estimated to account for rural residential choice. Prior to estimating this model, it is important to examine whether rural residence is in fact endogenous. Although economic theory suggests location of residence is a choice, it may still be possible to treat the nonmetro binary variable as (weakly) exogenous to poverty. To address this issue, an exogeneity test is conducted, using the approach proposed by Smith and Blundell (1986) for simultaneous limited dependent variable models. Performing the test requires finding at least one

[6]Marginal effects in the Probit model indicate percentage point rather than percentage change. To arrive at this percentage figure, the marginal effect was divided by the predicted probability of being poor (0.0929).

Table 2. Probit Results for Single-Equation Model of Person Poverty

Variable	Coefficient	Robust Std. Error	Marginal Effect
Constant	0.1887	0.4635	
Age of householder (years)	-0.0285*	0.0099	-0.0047
Age squared	0.0003*	0.0001	0.0001
Householder grew up in a city	0.1652*	0.0547	0.0282
Education of householder (years)	-0.1422*	0.0112	-0.0236
Work experience of householder (years)	-0.0095*	0.0033	-0.0016
Householder is disabled	0.8817*	0.1131	0.2303
Householder's main race is white	-0.4997*	0.0578	-0.0899
Family type (married couple excluded):			
Female headed	0.8539*	0.0632	0.1873
Male headed	0.5534*	0.0820	0.1222
Number of adults	-0.1718*	0.0400	-0.0286
Number of children	0.1793*	0.0234	0.0298
Child < 6 years present in household	0.1747*	0.0634	0.0303
County unemployment rate	0.0333*	0.0132	0.0055
Observed Nonmetro Residence	0.2201*	0.0745	0.0394

No. of observations	= 18,869
Wald $\chi^2_{[56]}$	= 1,155.97[a]
Wald $\chi^2_{[42]}$	= 94.01[b]
Pseudo-R^2	= 0.32

Notes: An asterisk (*) denotes significance at the 0.05 probability level or better. Standard errors are Huber/White robust standard errors adjusted for clustering on family.

[a]The Wald test for joint significance of the explanatory variables is distributed as a χ^2 with a critical value of 67.50 for 56 degrees of freedom at the 0.05 probability level.

[b]Not shown in the table are parameter estimates for state fixed effects. Eight of the state binary variables were dropped from the analysis due to collinearity. The Wald test for joint significance of the state binary variables is distributed as a χ^2 with a critical value of 55.76 for 42 degrees of freedom at the 0.05 probability level.

instrument to identify the poverty equation. Two variables are used for this purpose. The first is a binary variable for whether the family head has a religious preference uncommon in urban locations; this includes Amish, Mennonite, Church of God, Disciples of Christ, and Church of Christ. A second identifying instrument is an indicator variable for whether the householder's first occupation was farmer.

A key concern is the quality of the chosen instruments. To be valid, they should be highly correlated with rural residence and uncorrelated with the error term of the poverty model. A Probit model in which the dependent variable is nonmetro residence and explanatory variables are the two identifying instruments finds that the variables are both strong predictors of rural residence. The marginal effect on nonmetro residence of rural religious preference and of farming profession is 0.20 and 0.37, respectively. These variables are both statistically significant at the 0.05 probability level.

The second condition for instrument validity is more difficult to assess. To examine the direct effects of identifying instruments on poverty, a Probit model is estimated in which the dependent variable is a binary variable for person poverty and explanatory variables are the two identifying instruments. The rural religious preference variable has a marginal effect of 0.07 and a p-value of 0.39. The farmer as first occupation variable has a marginal effect of 0.01 and a p-value of 0.88. In sum, identifying instruments do not have direct effects on person poverty. This is reassuring, but does not necessarily imply that the instruments are uncorrelated with the poverty model error term. In a later section of the paper, a Sargan test of overidentifying restrictions provides further support for instrument validity (see table 5).

Having selected identifying instruments, we now turn to a Smith-Blundell test to examine whether rural residence is exogenous to poverty. The test is implemented with the "probexog" program written by Baum (1999) using the Stata language. The conducted test is one for exclusion of residuals from an auxiliary regression of rural residence on all exogenous variables and the two instruments described above. Under the null hypothesis, the auxiliary regression residuals are not predictors of poverty. Results for the test indicate that one can reject statistical exogeneity of nonmetro residence at the 0.10 probability level; the p-value for the test statistic is 0.086. This may suggest instrumental variables methods are warranted for estimating the relationship between nonmetro residence and poverty.

Accounting for Endogeneity of Rural Residence

A maintained hypothesis is that current estimates of a rural effect on poverty are biased because residence in a nonmetro area is a choice influenced by unobserved individual characteristics. There are two main ways for dealing with the residential selection problem. The first, instrumental variables, or two-stage least squares (2SLS), identifies and exploits an exogenous source of variation in residential choice; the second, fixed-effects strategies, involves introducing controls for individual heterogeneity (Weinberg, Reagan, and Yankow, 2004). Each approach has advantages and drawbacks but, for the purposes here, instrumental variables (2SLS) is more appropriate.[7] It is important to point out that the instrumental variables approach can also correct for omitted variable bias of general form. Specifically, it can reduce any bias related to omission of variables that determine poverty and are correlated with (though not determinants of) nonmetro residence.

A two-stage instrumental variables approach is used to account for potential endogeneity of rural residence. This is the method proposed by Newey (1987), and results in consistent point estimates. In the first stage, a Probit model of nonmetro residence is estimated. The model assumes the probability an individual resides in a nonmetro location n is a function of family-level variables that determine poverty $\mathbf{x}$, the county unemployment rate u, state binary variables $\mathbf{s}$, and a set of identifying instruments $\mathbf{z}$ assumed to affect residential choice but not whether an individual is poor. The residential choice model is expressed as:

$$n = \beta_0 + \beta_1 \mathbf{x} + \beta_2 u + \beta_3 \mathbf{s} + \beta_4 \mathbf{z} + \varepsilon_2. \tag{2}$$

[7] One fixed-effects approach involves the use of data from multiple siblings of families to difference out fixed family effects (e.g., Aaronson, 1998). This helps reduce the bias associated with unobserved family factors that influence both neighborhood choice and other individual behaviors; but the method is data intensive and is not particularly useful for studies measuring contemporaneous neighborhood effects on adults. Panel data regression with individual fixed effects has also been used in the attempt to distinguish causal neighborhood effects from neighborhood choice (e.g., Weinberg, Reagan, and Yankow, 2004). This approach can only account for neighborhood selection related to time-invariant individual factors (Dietz, 2002).

Table 3. First-Stage Probit Results for Rural/Urban Residential Choice

Variable	Coefficient	Robust Std. Error	Marginal Effect
Constant	-3.6696*	0.4412	
Age of householder (years)	-0.0255*	0.0100	-0.0056
Age squared	0.0002*	0.0001	0.0001
Householder grew up in a city	-0.8338*	0.0610	-0.1672
Education of householder (years)	-0.0516*	0.0097	-0.0114
Work experience of householder (years)	-0.0002	0.0032	-0.00003
Householder is disabled	0.1155	0.1357	0.0270
Householder's main race is white	0.7723*	0.0696	0.1558
Family type (married couple excluded):			
Female headed	-0.3416*	0.0758	-0.0682
Male headed	-0.2011*	0.0825	-0.0404
Number of adults	-0.0850*	0.0413	-0.0188
Number of children	0.0201	0.0241	0.0044
Child < 6 years present in household	0.0066	0.0652	0.0015
County unemployment rate	0.3349*	0.0153	0.0739
Identifying instruments:			
Householder's first job is farmer	0.4976*	0.1561	0.1380
Householder rural religion	-0.0823	0.2672	-0.0174

No. of observations	= 18,869
Wald $\chi^2_{[62]}$	= 1,441.32[a]
Wald $\chi^2_{[47]}$	= 1,088.34[b]
Wald $\chi^2_{[2]}$	= 10.28[c]
Pseudo-R^2	= 0.42

Notes: An asterisk (*) denotes significance at the 0.05 probability level or better. Standard errors are Huber/White robust standard errors adjusted for clustering on family.

[a]The Wald test for joint significance of the explanatory variables is distributed as a χ^2 with a critical value of 67.50 for 62 degrees of freedom at the 0.05 probability level.

[b]Not shown in the table are parameter estimates for state fixed effects. Three of the state binary variables were dropped from the analysis due to collinearity. The Wald test for joint significance of the state binary variables is distributed as a χ^2 with a critical value of 61.66 for 47 degrees of freedom at the 0.05 probability level.

[c]The Wald test for joint significance of the instruments is distributed as a χ^2 with a critical value of 5.99 for two degrees of freedom at the 0.05 probability level.

In the second stage, the Probit model of person poverty [equation (1)] is estimated, where observed nonmetro residence is replaced with the predicted probability of nonmetro residence. The latter variable should be purged of its potentially spurious correlation with omitted variables.

Table 3 presents results from the first-stage regression. At standard test levels, most of the parameter estimates are individually significant, and a Wald test indicates joint significance of explanatory variables. The point estimates for householder age suggest that age of the family head is negatively correlated with rural residence until the householder reaches the age of 65 years, at which point the correlation becomes positive. Rural residence is less likely if the family head grew up in a city, as opposed to a rural area, suburb, or combination of places. Findings display a negative association between nonmetro residence and education of the householder, consistent with Census Bureau

Table 4. Second-Stage Probit Results for Person Poverty

Variable	Coefficient	Robust Std. Error	Marginal Effect
Constant	0.2033	0.4587	
Age of householder (years)	-0.0291*	0.0100	-0.0049
Age squared	0.0003*	0.0001	0.0001
Householder grew up in a city	0.1522*	0.0659	0.0260
Education of householder (years)	-0.1422*	0.0114	-0.0238
Work experience of householder (years)	-0.0097*	0.0033	-0.0016
Householder is disabled	0.8821*	0.1135	0.2313
Householder's main race is white	-0.4862*	0.0661	-0.0878
Family type (married couple excluded):			
Female headed	0.8460*	0.0655	0.1860
Male headed	0.5469*	0.0825	0.1211
Number of adults	-0.1717*	0.0402	-0.0287
Number of children	0.1801*	0.0232	0.0301
Child < 6 years present in household	0.1743*	0.0632	0.0304
County unemployment rate	0.0379	0.0202	0.0063
Predicted Nonmetro Residence	0.1482	0.2453	0.0248

No. of observations = 18,869
Wald $\chi^2_{[56]}$ = 1,151.15[a]
Wald $\chi^2_{[42]}$ = 95.91[b]
Pseudo-R^2 = 0.32

Notes: An asterisk (*) denotes significance at the 0.05 probability level or better. Standard errors are Huber/White robust standard errors adjusted for clustering on family.

[a]The Wald test for joint significance of the explanatory variables is distributed as a χ^2 with a critical value of 67.50 for 56 degrees of freedom at the 0.05 probability level.

[b]Not shown in the table are parameter estimates for state fixed effects. Eight of the state binary variables were dropped from the analysis due to collinearity. The Wald test for joint significance of the state binary variables is distributed as a χ^2 with a critical value of 55.76 for 42 degrees of freedom at the 0.05 probability level.

data documenting lower educational attainment in nonmetro compared to metro areas. Results reported in table 3 suggest that people living in nonmetro counties are more likely to be part of a family headed by an individual who is white and married. The number of adult family members is negatively correlated with nonmetro residence. Findings show a positive association between rural residence and the county unemployment rate. Turning to results for the identifying instruments, the probability of rural residence is higher if the householder's first occupation was farming. The second identifying instrument—religious preference—has an unexpected negative sign, and is not statistically significant. As shown at the bottom of table 3, the two instruments are jointly significant at the 0.05 probability level.

Table 4 presents second-stage Probit results for person poverty. As above, standard errors reported use the Huber/White estimator of variance and adjust for within-cluster (family) correlation. Focusing first on all explanatory variables other than the nonmetro residence binary variable, we observe that coefficient estimates are very similar for the two models. The sign on each of these variables is the same across equations, and differences in magnitude are quite small. In addition, the set of statistically significant explanatory variables is essentially the same for the base model and the two-stage Probit model (the exception is that the county unemployment variable is statistically significant at the 0.05 probability level in the base model only).

When examining the parameter estimates for the nonmetro residence variable, findings are quite striking. Whereas the rural effect is positive and statistically significant in the base model, in the two-stage Probit model this effect disappears. The point estimate for nonmetro residence is 38% smaller in the two-stage model compared with the base model estimate. Furthermore, predicted nonmetro residence is not statistically significant at standard test levels in the two-stage Probit model. In short, accounting for endogenous rural residence and omitted variable bias of general form, there is no measured effect on person poverty of living in a nonmetro location. This finding parallels results reported by Evans, Oates, and Schwab (1992) who found that when one accounts for peer group choice in empirical models of teenage pregnancy and school dropout behavior, estimated peer group effects vanish. The finding here of no measured rural effect is at odds, however, with a large literature that documents a rural disadvantage in poverty outcomes [see Weber and Jensen (2004) for a review of studies].

To check the robustness of results to alternative specifications, a 2SLS model is estimated where the first-stage dependent variable is the dichotomous nonmetro residence variable used earlier, and the second-stage outcome variable is the ratio of family income to the Census annual needs standard. The income-to-needs variable is a family-size adjusted measure of family income and is continuous. While 2SLS models are usually implemented in settings where all endogenous variables are continuous, Angrist (2001) shows that 2SLS yields consistent point estimates even for the case of binary endogenous variables.

Table 5 presents 2SLS estimates for the income-to-needs ratio in which, as above, identifying instruments are binary variables indicating: (*a*) the householder's first occupation was farming, and (*b*) he/she has a religious preference that is not well represented in urban settings. Results for an ordinary least squares (OLS) model are also provided in table 5 for comparative purposes. Findings for a rural effect are qualitatively similar to those of the Probit models presented earlier. The OLS model in which nonmetro residence is assumed exogenous finds that people in nonmetro counties have a lower income-to-needs ratio compared with people in metro counties, and this result is statistically significant at the 0.05 probability level. In other words, findings indicate a higher risk of poverty in rural versus urban areas.

The 2SLS model is intended to account for any bias related to endogenous nonmetro residence or, more generally, for omission of variables that determine the income-to-needs ratio and are correlated with rurality. Focusing on results for the nonmetro residence variable, the absolute value of the point estimate is considerably smaller in size in the 2SLS model versus the OLS model. In addition, the standard error of the estimate is very large relative to the point estimate for the 2SLS model. In short, findings for the 2SLS income-to-needs model support those of the two-stage Probit model of person poverty. Results of the 2SLS model appear to indicate that living in a nonmetro area has no measured effect on the probability a person is poor.

Table 5 also reports the Sargan statistic for overidentifying restrictions which can be used to test for validity of the two identifying instruments. Under the null hypothesis that the instruments are uncorrelated with the error term of the poverty equation, the Sargan statistic is distributed as a χ^2 with degrees of freedom equal to the number of

Table 5. Ordinary Least Squares and Two-Stage Least Squares Results for Family Income-to-Needs Ratio, Individuals, 1993

	Ordinary Least Squares		Two-Stage Least Squares	
Variable	Coefficient	Robust Std. Error	Coefficient	Robust Std. Error
Constant	-2.9464	1.7714	-4.7487*	1.0589
Age of householder (years)	0.1476*	0.0211	0.1510*	0.0235
Age squared	-0.0013*	0.0002	-0.0014*	0.0002
Householder grew up in a city	0.2101	0.1240	0.3059	0.2964
Education of householder (years)	0.3884*	0.0262	0.3952*	0.0311
Work experience of householder (years)	0.0058	0.0072	0.0060	0.0071
Householder is disabled	-0.9514*	0.2307	-0.9693*	0.2405
Householder's main race is white	1.1088*	0.1262	1.0179*	0.2911
Family type (married couple excluded):				
Female headed	-1.5423*	0.1006	-1.5014*	0.1505
Male headed	-0.8504*	0.2286	-0.8246*	0.2297
Number of adults	-0.0360	0.0587	-0.0272	0.0640
Number of children	-0.4159*	0.0433	-0.4178*	0.0441
Child < 6 years present in household	-0.0641	0.1513	-0.0644	0.1512
County unemployment rate	-0.0845*	0.0230	-0.1278	0.1263
Nonmetro Residence	-0.6440*	0.1261	-0.0003	1.8179
No. of observations	18,869		18,869	
F-statistic (64, 6,832)[a]	30.74		11.05	
Sargan statistic[b]			0.87	
R^2	0.20		0.20	

Notes: An asterisk (*) denotes significance at the 0.05 probability level or better. Standard errors are Huber/White robust standard errors adjusted for clustering on family. Not shown in the table are parameter estimates for state fixed effects.

[a] Test for joint significance of the explanatory variables, distributed as an F-statistic with a critical value of 1.31 for 64 (numerator) and 6,832 (denominator) degrees of freedom at the 0.05 probability level.

[b] Sargan test of overidentifying restrictions, distributed as a χ^2 with a critical value of 3.84 for one degree of freedom at the 0.05 probability level.

overidentifying restrictions (the number of instruments less the number of regressors) (Davidson and MacKinnon, 1993).[8] The calculated Sargan statistic is 0.87, while the critical $\chi^2_{[1]}$ at the 0.05 probability level is 3.84. Thus, the null hypothesis cannot be rejected. In sum, the instruments appear to be valid.

Conclusion

This study used Panel Study of Income Dynamics data (N = 18,869 individuals) to examine whether living in a rural area is associated with a higher risk of poverty in the United States. Two main estimation strategies were employed. First, a single-equation Probit model was estimated of the probability an individual is poor as a function of family characteristics, county unemployment rate, state fixed effects, and nonmetro

[8] The Sargan statistic is equal to the number of observations times the R^2 from a regression in which the residuals of the instrumental variables estimate of the poverty equation are regressed on the instruments (Davidson and MacKinnon, 1993).

residence. This model assumes that nonmetro residence is an exogenous variable, paralleling common practice in the rural poverty literature. Empirical results confirm the well-documented finding of a higher risk of poverty in rural places. Findings indicate nonmetro people have a 42% higher probability of being poor compared with people in metro areas.

The second modeling approach involved estimating a two-stage Probit model which accounts for bias related to omitted variables that determine poverty and are correlated with nonmetro residence. In the first stage, nonmetro residence was modeled as a function of family characteristics, county unemployment rate, state of residence, and a set of identifying instruments. In the second stage, predicted nonmetro residence from the first-stage regression replaced observed rural residence in the model of poverty probability. Findings of the two-stage Probit model suggest nonmetro residence has no measured effect on person poverty, all else being equal. Tests for the validity of instruments used to identify the poverty model provide some support for the choice of instruments. A Smith-Blundell test for exogeneity of rural residence appears to indicate that living in a nonmetro area is a choice, suggesting the two-stage Probit model provides a more reliable means to measure the links between rural residence and poverty compared with the single-equation Probit model. In tandem, empirical findings show that failure to account for residential endogeneity and omitted variable bias of general form leads to overestimation of the effect of rural residence on person poverty.

The findings of this study may indicate that people with certain attributes related to human impoverishment choose to live in rural places or are reluctant (or unable) to leave them. Such a finding is consistent with results reported by Nord (1998) who uses 1990 Census data to examine the effect on the geographic distribution of poverty of county-to-county migration of the poor and the nonpoor. His study shows more poor people moving into than out of persistent poverty nonmetro counties during the analysis period (1985–1990), a pattern that reinforced the preexisting spatial concentration of poverty.

Clearly, empirical findings of the present study should not be taken as definitive. Results here do not rule out the possibility that living in a rural area is a factor which causes poverty in the United States. Study findings highlight the need to test and, if necessary, correct for endogeneity in the econometric measurement of the effects of rural residence on poverty outcomes. Future work using other nationally representative data sets, covering a range of analysis years, and employing alternative estimation strategies to correct for residential endogeneity will enable an improved assessment of the extent to which there exists a rural disadvantage in welfare outcomes in the United States. The answer to this question has implications for future research on rural poverty. If empirical studies suggest that rural residence is an important determinant of poverty, then a key area for research might be to improve our understanding of the specific structural conditions that foster rural poverty. If, however, accounting for residential selection, there exists no measured effect of nonmetro residence on poverty, then at least two research questions seem particularly important. One, why do people with certain attributes related to human impoverishment choose to live in rural places? And two, what combination of human-capital and community-strengthening policies is most likely to reduce rural poverty and its unfavorable consequences?

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On the Empirical Finding of a Higher Risk of Poverty in Rural Areas: Is Rural Residence Endogenous to Poverty? COMMENT

Thomas A. Hirschl

Dr. Monica Fisher's paper theorizes that the observed rate of rural poverty is higher than urban poverty for one of two possible reasons: (*a*) because of unfavorable social conditions in rural communities, or (*b*) because individuals who choose to reside in rural communities have unmeasured characteristics that make them more likely to be poor. A series of econometric tests using a cross-section of the Panel Study of Income Dynamics suggest that the underlying reason is (*b*), that rural poverty is higher because individuals predisposed to being poor choose to reside in rural communities. Dr. Fisher states that "the findings of the present study may indicate that people with certain attributes related to human impoverishment choose to live in rural places or are reluctant (or unable) to leave them" (p. 197, this journal issue). Presumably if these "at-risk" folks chose to live elsewhere, rural poverty rates would decline to urban poverty rates.

While I appreciate the clarity and statistical rigor of this paper, I take issue with the overall approach, and therefore also with the conclusion. First, it would seem important to identify what it is about rural communities that makes their members more likely to be poor. Is it labor market disadvantage, distance from the metropolitan center, or something else? If one wishes to argue that A causes Y as opposed to B causing Y, then the competing causal processes need to be spelled out, and appropriate indicators of each process should be measured, and then competitively tested. In the absence of concrete hypotheses about the causal process, one is "shooting in the dark." Surely it is not the census designation "nonmetropolitan" that causes poverty.

In her monograph on rural poverty Cynthia Duncan (1999) finds that persistent poverty is intrinsic to the social structure of some rural communities, but not to others. To simplify Duncan's analysis, some rural communities (such as those found in Appalachia and the Mississippi Delta) have a history of social inequality and chronically underutilized labor, whereas other rural places (such as those found in the upper Midwest) have a history of social equality and labor market success. Her study further suggests that individual behavior within these divergently structured communities tends to reinforce pre-existing structures. To the extent Duncan's generalizations are correct, it is inappropriate to posit a covering law for all nonmetropolitan places that does not key into specific place characteristics related to poverty causation. Hence, the spatial category "nonmetropolitan" is an imperfect proxy for rural community poverty.

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From this point of view it is unsurprising that different measures of "nonmetropolitan" (fitted versus observed) have differing effects upon poverty.

To the extent one agrees with Duncan (1999) that some (but not all) rural labor markets embody structural disadvantage, then it becomes possible to interpret the present study as supporting the structural disadvantage thesis. Inspecting the coefficients for county unemployment rate in tables 2 and 4, I note they are both positive and statistically significant. This suggests that community labor market status affects the odds that an individual will be poor. Thus, social context as well as individual characteristics predict the likelihood of individual impoverishment. If this statement is correct, we might then ask whether any of these predictive relationships reflect a "fundamental cause" (Lieberson, 1985), roughly defined as a causation that persists for long periods of time.

From my point of view, rural/urban poverty differentials are important because they may provide clues about the fundamental causes of poverty, and the way in which people respond to these causes. This viewpoint is premised upon the notion that poverty is not an individual characteristic, but rather a social state that individuals move in and out of over their life course (cf. Rank and Hirschl, 2001). It is from this perspective that I ask why would families with "certain attributes" choose to live in, or be reluctant to leave, rural places? Having re-read Dr. Fisher's paper several times, I am at a loss to answer this question.

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On the Empirical Finding of a Higher Risk of Poverty in Rural Areas: Is Rural Residence Endogenous to Poverty? REPLY

Monica Fisher

In his commentary on my paper in the current issue, Hirschl poses the sorts of questions I anticipated my article would stimulate. His main points of concern are (*a*) the aim of the paper, (*b*) the research hypothesis, and (*c*) one aspect of the empirical analysis. I respond to each issue in turn.

First, Hirschl argues that my paper sets out to determine the singular explanation for the higher observed rate of poverty in rural than in urban areas of the United States. Is this phenomenon a result of less favorable economic and social conditions in rural communities (the "structural condition hypothesis"), or is it a function of poor people "choosing" to reside in rural places (the "residential sorting hypothesis")?[1]

My paper indeed focuses on a single explanation of high rural poverty—the residential sorting hypothesis. Never, however, is the claim made that this is the sole explanation for such a complex phenomenon. I ask if the disproportionate poverty in rural areas *partly* reflects that those with personal attributes related to human impoverishment are attracted to rural places, or are otherwise reluctant (or unable) to leave them. In posing this question, I do not aim to disprove the structural condition argument. Instead, by investigating a largely overlooked yet plausible explanation for rural poverty, I complement a large literature documenting the role of social and economic context in persistent rural poverty. Certainly a problem as enduring as poverty has numerous causes.

While the rural poverty literature has long emphasized the structural condition hypothesis [see Weber and Jensen (2004) for a review], only two studies have explored the residential sorting hypothesis. Nord (1998) used 1990 Census data to examine the effect on the geographic distribution of poverty of the county-to-county migration of poor and nonpoor. He found that more poor people moved into than out of persistent poverty nonmetro counties between 1985 and 1990, a pattern that reinforced poverty's pre-existing spatial concentration. Fitchen's (1995) in-depth interviews with low-income families in upstate New York tell a similar story. Her case-study community, a rural area facing economic decline, was found to be a migration destination for poor urban families. The lack of jobs in the community did not appear to deter low-income migrants.

Hirschl's second main concern centers on the study's hypothesis that poor people "choose" rural living. Hirschl asks, "why would families with 'certain attributes' choose

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[1]The structural condition hypothesis ascribes a causal role to place of residence. According to this view, otherwise identical individuals will have lower economic well-being in rural compared to urban settings due to the spatial distribution of economic and social opportunities (Tomaskovic-Devey, 1987). The residential sorting perspective, by contrast, posits that, holding constant human capital attainment, individuals' prospects for economic prosperity are independent of where they live.

to live in, or be reluctant to leave, rural places?" I did not address this question, although my findings certainly warrant future research on it. Here I will speculate. It is conceivable, as argued by Nord (1998), that individuals with low education and limited work experience are drawn to places that offer opportunities matching their skills and needs—for example, communities with a high share of entry-level positions and where living costs are low. Low-skill occupations continue to make up a higher percentage of total jobs in rural areas (42%) than in the nation as a whole (35.5%) (Gibbs, Kusmin, and Cromartie, 2004). And studies show that living costs are substantially lower in nonmetro than in metro areas (e.g., Kurre, 2003; Nord, 2000). Fitchen's (1995) interviews with poor urban migrants, described above, reveal that the main attraction of her case-study community was its inexpensive rental housing. Finally, rural places may appeal to those with low earning capacity because of the possibilities for informal work. Studies document a range of informal employment activities in rural communities that help the poor weather income shortfalls (e.g., Jensen, Cornwell, and Findeis, 1995). In some regions, such work features more prominently in the livelihood strategies of rural than of urban residents (Tickamyer and Wood, 1998).

Hirschl's third comment is "it is inappropriate to posit a covering law for all nonmetropolitan places that does not key into specific place characteristics related to poverty causation." I fully agree that a key drawback of my analysis is the implicit assumption that rural places are homogeneous. This assumption is a common one in the empirical rural poverty literature (e.g., Brown and Hirschl, 1995; Cotter, 2002), a reflection of data limitations, and is unlikely to be valid. As articulated by Miller, Farmer, and Clarke (1994, p. 3), "If you've seen one rural community, you've seen one rural community.... Thus, to speak of a singular rural America is folly." Future research should examine whether the finding that people with low income capacity choose rural residence is robust across regions and for rural areas with varying characteristics (e.g., high versus low amenity counties, and remote-rural places versus rural areas adjacent to metro areas). Conducting this type of investigation will require access to confidential data sets with identification codes for respondents' place of residence.

The residential sorting hypothesis may be construed by some as blaming low-income individuals for their condition, but its investigation is warranted from both research and policy standpoints. If poor people do have a tendency to sort themselves into rural areas, a key area for future research is to understand the factors that drive such residential choices (e.g., spatial living cost differences or the geographic distribution of entry-level work). Empirical studies can also inform anti-poverty policy, suggesting what combinations of human-capital and community-strengthening policies are most likely to reduce rural poverty and its unfavorable consequences.

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