SUBURBAN DEVELOPMENT AND ECONOMIC SEGREGATION IN THE 1990s

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ABSTRACT: *Suburbanization is changing the face of urban America. A common claim is that suburban sprawl has contributed to increasing levels of economic segregation, but few studies have directly tested this hypothesis. Using U.S. Census data for 1990 and 2000, this paper examines the trends in and the relationship between suburban development patterns and economic segregation in U.S. metropolitan areas. We find that economic segregation, as measured by the Neighborhood Sorting Index (NSI), declined during the 1990s, reversing the earlier trend. However, results from cross-sectional and fixed-effects regression models at the metropolitan level suggest that suburbanization, as measured by five different indicators, was a countervailing influence during the decade. Metropolitan areas that were suburbanizing more rapidly had smaller declines in economic segregation than comparable metropolitan areas.*

Although the suburbanization process is quite common in all wealthy western nations, metropolitan areas in the United States are remarkably spread out (Musterd & Ostendorf, 1998). Contrary to most of the world, American affluent and middle-class persons tend to live in suburbs, often quite distant from the urban core (Jackson, 1985). In U.S. metropolitan areas, there is a monotonic relationship between the median year of construction of a neighborhood's housing and the mean income of the households that occupy those neighborhoods (Jargowsky, 2002). Controlling for median year built, central city neighborhoods have consistently lower mean incomes than the comparable cohort of suburban neighborhoods. These considerations suggest that the suburban development process—"suburban sprawl" to its critics—may contribute to a greater geographic fragmentation of the population by social class, by creating homogeneous well-to-do neighborhoods on the periphery and leaving lower-income persons in central city and older suburban neighborhoods.

Segregation along economic lines increased in the United States from 1970 to 1990 (Abramson, Tobin, & VanderGoot, 1995). Both the inequality among regions and the economic segregation between cities and suburbs widened during this period (Dreier, Mollenkopf, & Swanstrom, 2001). Moreover, the economic segregation among suburbs also grew (Lucy & Phillips, 2000). As to neighborhood-level economic segregation, measured by the neighborhood sorting index, Jargowsky (1996) found a pronounced and "nearly ubiquitous trend toward increased economic

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segregation" within each racial or ethnic group between 1970 and 1990. Meanwhile, concentration of poverty in urban areas increased substantially during the same period. Using a 40% poverty standard, Jargowsky (1997) demonstrated that poverty areas grew in terms of the number of tracts, the total population, the percentage of the overall population, the percentage of poor persons living in them, and land area.

The 1990s, however, were a very different decade. Unemployment declined from 7.5% in 1992 to 4.0% in 2000, the lowest level since 1969 (Economic Report of the President, 2005, Table B-35). Poverty overall declined, as did concentration of poverty (Jargowsky, 2003; Kingsley & Pettit, 2003). At the same time, suburban development continued or even accelerated (Lang & Simmons, 2003).

We have two goals in this research: (1) to document the recent trends in economic segregation and suburbanization, and (2) to test whether suburban development patterns have played a significant role in the trend in economic segregation. Based upon 1990 and 2000 U.S. Census data, we examine these empirical questions by conducting a nationwide metropolitan-level analysis. Our study complements the existing literature by investigating the changes of economic segregation in U.S. metropolitan areas during the most recent decade. Furthermore, although suburban development is frequently discussed in the literature, there is no universally agreed measure of suburbanization or "sprawl." Instead of treating the concept of suburbanization as a monolith, our study identifies five plausible indicators of suburban development patterns, and tests their effects on changes in economic segregation.

We find that economic segregation, as measured by the Neighborhood Sorting Index (NSI), reversed course in the 1990s. Hispanics had the sharpest declines, from 0.501 in 1990 to 0.400 in 2000. For whites, the decline was from 0.454 to 0.377, and for blacks, from 0.479 to 0.408. However, we find that suburbanization was a countervailing influence during this period; in other words, the declines in economic segregation would probably have been even larger without the effects of suburban development.

Suburban development is clearly driven in large measure by the dynamics of demand and supply in the housing market; nevertheless, sprawl occurs in part because local governments in the United States encourage this form of development through highway construction, exclusionary zoning, and subdivision ordinances (Burchell et al., 1998). The findings of our study therefore have important implications for public policies regarding urban and metropolitan policy and community redevelopment. Moreover, economic segregation has important implications in all social policy arenas. For example, school attendance zones are geographic, and much school finance is local. Thus, our findings have clear implications in the education policy arena as well.

BACKGROUND

Measuring Economic Segregation

Economic segregation can be defined as "the spatial segregation of households by income or social class" (Jargowsky, 1996). The study of economic segregation in the United States emerges from and is influenced by the investigation of racial segregation. However, relative to segregation by race, residential segregation by income and other measures of socio-economic status has received much less attention by social scientists (Kain, 2000).

To measure economic segregation, many researchers have applied the index of dissimilarity (*D*) to variables that serve as proxies for social class (Duncan & Duncan, 1955; Erbe, 1975; Farley, 1977). For computation purposes, this measurement strategy requires that the social class variables be broken down into discrete categories. However, changes in the social meaning of these categories over time complicate longitudinal comparisons of economic segregation based on these variables (Jargowsky, 1996). For example, being a high school graduate means less in social and economic terms than it used to.

In contrast, household income offers greater comparability over time once inflation is taken into account. Massey and Eggers (1990) applied the index of dissimilarity to the income distribution. In an analysis of 60 large metropolitan areas, they defined four social classes based upon specific dollar thresholds: poverty income, lower-middle class, upper-middle class, and affluent. For each racial and ethnic group, they computed the six pair-wise indices of dissimilarity among the four social classes and averaged these indices to come up with a final aggregate measure. Based on this analysis, they found that interclass segregation among blacks increased over the 1970s in an often quite sharp manner, while it declined for whites, Hispanics, and Asians. These findings generally confirmed Wilson's hypothesis that black segregation by income increased between 1970 and 1980 (Wilson, 1987).

Income, however, is a continuous variable, and collapsing it into four categories discards information. More importantly, the introduction of dollar thresholds confounds changes in the income distribution with changes in spatial organization, because the thresholds are not independent of the mean and variance of the income distribution. Some other more advanced measures, such as the entropy index, could easily handle multiple categories, but are still affected by the shifts in the underlying distribution that change the meaning of the categories (Jargowsky, 1996).

The correlation ratio, also known as eta-squared or the segregation statistic, has been used to measure segregation (Bell, 1954; Farley, 1977; Schnare, 1980; Zoloth, 1976) and is particularly well suited to use with continuous variables (White, 1983). By applying the correlation ratio to income, Jargowsky (1996) developed the Neighborhood Sorting Index (NSI) as a measure of economic segregation. The NSI is defined as

$$\mathrm{NSI} = \frac{\sigma_N}{\sigma_H} = \frac{\sqrt{\dfrac{\sum_{n=1}^{N} h_n\,(\bar{y}_n - \bar{y})^2}{H}}}{\sqrt{\dfrac{\sum_{i=1}^{H} (y_i - \bar{y})^2}{H}}}$$

where y is household income, i indexes households, n indexes neighborhoods, h_n is the number of households in neighborhood n, and H and N are the total number of households and neighborhoods respectively. As in most segregation studies, census tracts were used as proxies of neighborhoods (White, 1987). The square of NSI is the between-tract variance over the total variance of household income. It "implicitly controls for the overall income level because it is based on deviations from mean household income and also controls for income inequality because it is expressed as a percentage of total income variance" (Jargowsky, 1996).

Jargowsky (1996) calculated NSI for many metropolitan areas in 1970, 1980 and 1990, and found the NSI increased steadily between 1970 and 1980 and again between 1980 and 1990 for whites, blacks, and Hispanics, though the values of the NSI were relatively modest. The increases are largest for blacks: the NSI increased more than 40% between 1970 and 1990, from 0.341 to 0.480. For whites, the NSI increased by about 10% in both 1970s and 1980s. For Hispanics, the NSI increased by about 9% in the 1970s and by 16% in the 1980s. In this study, we will adopt the NSI to measure economic segregation and examine the trends for the 1990s.

Characterizing Suburbanization Patterns

Suburban sprawl in United States has become a frequent topic for research and debate in urban studies. However, neither suburbanization nor sprawl is consistently defined in the current literature. "Sprawl" can be understood as a pejorative term describing a certain type of land development. For instance, Ewing (1997) defined sprawl as "the spread-out, skipped-over development that characterizes the non-central city metropolitan areas and non-metropolitan areas of the United States." Furthermore, he vividly portrayed the image of sprawl as "one- or two-story, single-family residential development on lots ranging in size from one-third to one acre, accompanied by strip commercial centers and industrial parks, also two stories or less in height and with a similar amount of land takings."

Sprawl is said to be unplanned and illogical development. It is criticized for being ugly, dehumanizing, and socially isolating. Sprawl is denounced for producing "jive-plastic commuter tract-home wastelands" that are a "wasteful, toxic, agoraphobic-inducing spectacle" (Kunstler, 1993). Sprawl is also defined by one or more examples of scattered or low-density patterns of urban development; Los Angeles often serves as a typical example in such case (Geddes, 1997). Moreover, scholars argue that unwanted externalities, such as traffic congestion, environmental contamination, and local fiscal disparities, are caused by sprawl (Downs, 1998, 1999; Popenoe, 1979).

In short, sprawl has become "the metaphor of choice for the shortcomings of the suburbs and the frustrations of central cities" (Galster et al., 2001). As Galster et al. (2001) argued, the term "sprawl" existing in current literature explains "everything and nothing." Some researchers have tried to produce a clearer conceptual definition and to construct a composite statistical measure of this concept. For example, Galster et al. (2001) defined sprawl as a pattern of land use in an urbanized area that exhibits low levels of some combination of eight distinct, objective dimensions: density, continuity, concentration, clustering, centrality, nuclearity, mixed uses, and proximity. Further, they selected 13 large areas from different regions of the country for a prototype test of their definition, for only housing sprawl. More specifically, they computed a *Z* score for each of the 13 urban areas (UAs) on each of six dimensions (density, concentration, clustering, centrality, nuclearity, and proximity), and then added the Z scores for each UA across all six dimensions to obtain a composite sprawl index. By this method, Galster et al. (2001) claimed some encouraging results that "comport with our firsthand knowledge of these areas, as well as the conventional wisdom." For example, they found that the UAs with the greatest degree of sprawl (or the lowest score on the composite index) were Atlanta, followed by Miami, Detroit, and Denver. The UAs with the lowest degree of sprawl were all older areas in the Northeast and Midwest: the New York area had the least sprawl, followed by Philadelphia, Chicago, and Boston. However, the composite score is somewhat opaque and composed of quite dissimilar elements.

Squires (2002) defined sprawl as "a pattern of urban and metropolitan growth that reflects low density, automobile-dependent, exclusionary new development on the fringe of settled areas often surrounding a deteriorating city." This definition does highlight several important aspects of suburbanization, such as low density, and the income characteristics of new construction. However, this concept is very broad and hard to measure operationally.

Instead of developing a comprehensive definition or a composite index, we identify five plausible indicators of different dimensions of suburbanization, based on the relevant literature. These indicators, discussed further below, are: the urban density gradient; population density; homogeneity of new growth; exclusionary zoning; and inaccessibility of jobs.

The *urban density gradient* has served as a measure of suburbanization for a long time. Clark (1951) estimated the exponential density functions for a variety of metropolitan areas and years, and observed a strong tendency for density functions to flatten through time. The exponential density function can be written as

$$d_x = d_0 e^{-bx} \tag{1}$$

where d_x is population density, x is distance from the metropolitan center, d_0 is density at or near the metropolitan center, and b (larger than 0 usually) is the density gradient. After taking the natural logarithm of both sides, the equation becomes

$$\ln d_x = \ln d_0 - bx. \tag{2}$$

The urban density gradient describes how the population per square mile of an area drops off with distance from the center of the metropolitan area (Berry et al., 1974; Berry & Gillard, 1977; Berry & Horton, 1970; Edmonston, 1975; Mills, 1972). It is the rate at which population density decreases as distance from an area's center increases (Schiller, 2004). A metropolitan area is said to become more suburbanized as the urban density gradient lessens (Mills, 1991; Schiller, 2004). If suburbanization contributes to economic segregation, then the expected effect of the urban density gradient on economic segregation is negative.

Density is one of the most widely used measures of suburbanization. Population density in the center areas of metropolises has continuously diminished as people relocate to outlying areas (Berry et al., 1974; Carruthers, 2003; Edmonston, Goldberg, & Mercer, 1984). However, the direct effects of density on income segregation remain unclear. On the one hand, increased density might lead to a fishbowl effect, where density crams people together, and creates more opportunities for mixing housing types and incomes within a small area. On the other hand, high densities also cause a residential sorting effect, where density leads to intense competition over urban space, and may end up contributing to income segregation rather than ameliorating it. Pendall and Carruthers (2003) found that density affected economic segregation in a nonlinear fashion: income segregation rose as density increased, but at a decreasing rate.

Exclusionary zoning could serve to promote economic segregation. The wide latitude and autonomy granted to local governments in the United States often encourage local jurisdictions to engage in a beggar-thy-neighbor competition with one another (Dreier, Mollenkopf, & Swanstrom, 2001). For instance, many suburbs set minimum lot sizes (such as one-half acre per home) that increase the cost of housing and rule out the construction of dense housing—not just apartment buildings, but also bungalow-style single-family homes (Mallach, 1984). Intentionally or unintentionally, local exclusionary zoning regulations, which are rooted in the fragmented arrangements of local jurisdictions, help isolate higher-income from lower-income families (Dreier, Mollenkopf, & Swanstrom, 2001). Thus, the expected influence of this indicator on economic segregation is positive.

Another indicator of suburbanization is the *homogeneity of new development*. White and affluent flight from city centers often entails relocation to newly developed, low-density areas at the urban fringe, where land use regulations enforce socioeconomic homogeneity (Carruthers, 2003; Squires, 2002). The stock aerial photograph, symbolic of suburban conformity, shows mile after mile of virtually identical households (Jargowsky, 2002). Usually protected by zoning regulations, these homogeneous new developments in suburban areas are "exclusionary" by their very nature, because households with lower incomes typically find that these housing units are difficult to afford (Dreier, Mollenkopf, & Swanstrom, 2001; Goldsmith & Blakely, 1992; Orfield, 1997). But homogeneous can mean homogeneously rich or homogenously poor, as in Bombay, where housing on the periphery is uniform and low income. The issue is whether the character of the newest housing has an impact on economic segregation.

The last indicator characterizing the contemporary suburbanization patterns is the *inaccessibility of jobs* to central-city residents. For instance, the spatial mismatch hypothesis argues that "sprawl" widened the gap between where people live and where jobs are located; the suburbanization process interacts with racial segregation in housing to create barriers to job access for low-income people (Kain, 1968; Ihlanfeldt & Sjoquist, 1998). As jobs follow wealthier people out to the suburbs, the remaining inner-city residents will find it harder to access new jobs on the periphery. Inaccessibility of jobs, as an indicator of suburbanization, is expected to be positively related to economic segregation.

EMPIRICAL TRENDS

Economic Segregation in the 1990s

By using the data from 1990 and 2000 Census Summary File 3 and 5% Public Use Microdata Sample (PUMS) data, our study investigates economic segregation over the last decade in all metropolitan areas in the United States. To best capture the dynamics of local housing and labor markets, stand-alone metropolitan statistical areas (MSAs) and primary metropolitan statistical areas (PMSAs) are chosen to define metropolitan areas in our study (Iceland, Weinberg, & Steinmetz, 2002). To compare income segregation between 1990 and 2000, we use constant 2000 metropolitan area geographic boundaries in our calculations. Furthermore, contemporaneous census tracts are used as proxies of neighborhoods in our study; use of contemporaneous tracts, rather than matched tracts, is important to ensure a consistent average neighborhood size in both periods. Census tracts are small statistical subdivisions of a county with generally stable boundaries and homogeneous population characteristics, economic status, and living conditions, at least at the time they were established (Ricketts & Sawhill, 1988; White, 1987).

There are 331 MSAs and PMSAs based on 2000 boundaries. Among these metropolitan areas, six are excluded from our analysis because five metropolitan areas contain no central city[1] and one has no suburbs.[2] Of the remaining 325 MSA/PMSAs, we further exclude those containing fewer than 10,000 households for a given racial/ethnic group. With fewer than that, there are too few neighborhoods or too sparsely populated neighborhoods to make a meaningful calculation of a neighborhood-based measure. We use household income information from the 5% PUMS data to estimate individual household income variance (the denominator of NSI) that is not given in Summary File 3.[3]

The findings are largely consistent with the dramatic decline of concentrated poverty during the same period. Jargowsky (2003) reported that the number of people living in high-poverty neighborhoods (where the poverty rate is 40% or higher) declined by 24%, or 2.5 million people, in the 1990s. Moreover, concentrated poverty, which is the share of the poor living in high-poverty neighborhoods, declined among all racial and ethnic groups, especially African Americans (Jargowsky, 2003; Kingsley & Pettit, 2003).

Table 1 shows the weighted mean NSI for non-Hispanic whites and other racial/ethnic groups, blacks, and Hispanics[4] for 1990 and 2000. The empirical results indicate a pronounced trend toward decreasing economic segregation, reversing the earlier finding. Figures in the upper panel are for all metropolitan areas in the United States for that group in that year; the lower panel demonstrates the weighted mean for a constant set of metropolitan areas. For instance, the 1990 NSI for blacks is 0.479, indicating the standard deviation of the distribution of neighborhood mean incomes is about half of the standard deviation of overall income distribution. Thus, between-neighborhood variance accounts for about 23% (0.479 squared) of the total variance in household income.

Between 1990 and 2000, economic segregation of all three groups declined substantially. White and other races declined 0.077, from 0.454 in 1990 to 0.377 in 2000, a decrease of nearly 17%.

TABLE 1

Neighborhood Sorting Index (NSI) by Racial and Ethnic Group: U.S. Metropolitan Areas (MSA/PMSAs), 1990 to 2000

	White and others		Black		Hispanic	
Sample and year	Mean	*N*	Mean	*N*	Mean NSI	*N*
All MSA/PMSAs						
1990	0.454	324	0.479	130	0.501	69
2000	0.377	324	0.408	143	0.400	99
Change	−0.077		−0.071		−0.101	
Constant set of MSA/PMSAs						
1990	0.454	324	0.479	130	0.501	69
2000	0.377	324	0.408	130	0.398	69
Change	−0.077		−0.071		−0.103	

Note. Includes metropolitan areas with 10,000 or more households for each racial or ethnic group indicated; means are weighted by number of households for each racial or ethnic group.

Economic segregation among blacks declined by a similar amount. The decline was greatest for Hispanics, whose NSI measure dropped 20% from 0.501 in 1990 to 0.400 in 2000. Figures in the lower panel hold the number of MSAs constant at the 1990 level; the results are essentially identical, implying that the findings are not driven by the inclusion of new metropolitan areas in the 2000 census.

Besides the notable economic boom, the United States also experienced significant changes in housing and community development policy during the last decade that could have contributed to reduction in economic segregation. There was a great deal of gentrification and other forms of neighborhood revitalization (Sohmer & Lang, 2003), much of it encouraged by public incentives and public/private partnerships. By reinvesting in city centers and deteriorating neighborhoods, this development changes the essential character and flavor of inner-city neighborhoods by encouraging higher income households to displace lower income residents. Furthermore, numerous local governments and jurisdictions have adopted the goals of creating mixed-income, diverse, integrated communities, providing housing for a diverse labor force, and connecting residents in high-poverty neighborhoods to opportunities. All these programs and policies, along with the economic prosperity that broadly distributed economic gains across neighborhoods, are possible contributors to reducing economic segregation from 1990 to 2000.

Trends in Suburbanization Indicators

Urban Density Gradient

The first suburbanization indicator is the urban density gradient, which describes how population per square mile of an area drops off with distance from the center of the city (Berry et al., 1974). The basic descriptive statistics of this indicator are shown in Table 2. Most importantly, the national mean of urban density gradients of all MSA/PMSAs falls from 0.141 in 1990 to 0.132 in 2000; a drop about 6.38%. This decline is also quite widespread. Among the 325 metropolitan areas, 250 of them (about 76.85%) experienced declines in urban density gradient.

The observed decline in urban density gradient over the last decade is consistent with the historical trend documented by numerous studies. Berry and Horton (1970) recognized the phenomenon that density gradient falls over time as the second axiom in intra-urban structure and growth, and demonstrated the diminishing density gradients in London from 1801 to 1941 and in Chicago from 1860 to 1950. Mills (1972) reported the average population density gradients of Baltimore, Milwaukee, Philadelphia, and Rochester dropped from 1.22 in 1880 all the way

TABLE 2

Descriptive Statistics of Suburbanization Indicators: 1990 to 2000

		Mean	SD	Minimum	Maximum
Urban density gradient	1990	0.141	0.073	0.002	0.613
	2000	0.132	0.070	−0.040	0.657
Gross population density	1990	1001.606	1644.454	4.500	11855.27
	2000	1064.831	1752.25	5.412	13043.62
Number of governments per 100,000 households	1990	14.716	9.639	2.555	79.051
	2000	13.714	9.261	2.596	75.195
Homogeneity of new development	1990	1.635	0.479	1.351	3.420
	2000	1.658	0.157	1.353	3.232
Average commuting time	1990	21.622	5.099	11.359	36.383
	2000	24.483	5.560	13.254	39.915

Note. Means are weighted by the number of total households in MSA/PMSAs.

down to 0.31 in 1963. Berry et al. (1974) showed the density gradients of the eight U.S. urban regions they investigated all declining from 1950 to 1970. As Berry and Gillard (1977, 1) indicated, "counter-urbanization has replaced urbanization as the dominant force shaping the nation's settlement patterns."

Population Density

Population density serves as the second suburbanization indicator. As shown in Table 2, the gross population densities among U.S. metropolitan areas vary in a wide range. In 1990, Jersey City, New Jersey contained the highest density in this country of approximately 11,855 people per square mile, while Flagstaff, Arizona/Utah had only about four persons per square mile.

Regarding the change of population density in U.S. metropolitan areas over the last decade, Table 2 seems to indicate an increasing trend: the national mean rises from about 1,002 people per square mile in 1990 to roughly 1,065 persons per square mile in 2000. However, this increase is only an artifact caused by the consistent metropolitan boundary approach used in the calculations. By assigning the fixed 2000 metropolitan boundaries to 1990 data, the land area in every metropolitan area actually remains constant. Thus, the increase in national mean of population density is more a reflection of the urban population growth during the last ten years, rather than a sign of greater metropolitan crowdedness considering urban boundary expansions that actually happened.

Nevertheless, the gross population density based upon the consistent metropolitan boundary approach is still appropriate for cross-sectional analyses, since it efficiently captures the existing differences of population concentration between metropolitan areas at a static point in time. For instance, Dallas with 569 people per square mile in 2000 is much less dense than New York where 8,159 persons live in every square mile. However, for the first-difference fixed-effects estimations that focus on changes of variables over time, this computation method may lead to measurement bias. The estimated central city density is a better measure of density change (d_0 in Equation (1)) and thus is used as a substitute for gross population density in the first-difference fixed-effects analyses. The estimated central city density is estimated by the urban density gradient calculation discussed above, as the intercept in the estimation equation.

The average estimated central city densities declines from 3,433 in 1990 to 3,381 in 2000. Among the 325 metropolitan areas, 183 of them experienced declines in estimated central city density, despite the fact that they were experiencing population increases that would tend to increase central densities, all else equal. This trend is generally consistent with the findings about diminishing central city density in the course of time from many earlier studies (Berry et al., 1974; Edmonston, Goldberg, & Mercer, 1984).

Exclusionary Zoning

Another suburbanization indicator is the exclusivity of local zoning regimes. Unfortunately, there is no easy way to summarize the zoning status of an entire metropolitan area, which normally spans many cities, counties, and even states. Due to the lack of suitable national data for exclusionary zoning at the metropolitan level, the number of local governments per 100,000 households within a given metropolitan area serves as a proxy variable. A greater number of governments may have more incentives and more opportunities to use zoning competitively (Tiebout, 1956).

Table 2 shows this proxy variable varies in a wide range. For example, the minimum value of 2.55 was found in New York in 1990, and Joplin, Missouri had the highest value at 79.05. Moreover, the national mean of this variable declined moderately from 14.7 in 1990 to 13.7 in 2000.

Homogeneity of New Development

The fourth suburbanization indicator is the homogeneity of new development. To measure homogeneity, we first rank neighborhoods within metropolitan areas by the age of their housing stock. Within the newest 10% of the neighborhoods in each metropolitan area, we form the ratio of the third to first quartile of housing value, and compute a weighted average of these for each metropolitan area. The expected sign is negative, because a lower ratio means more homogeneity, and an expectation of higher economic segregation. As illustrated in Table 2, this indicator obtains very similar national mean values in 1990 and in 2000. Also, the 2000 calculation yields smaller numbers in both range and standard deviation, representing the most newly developed suburban neighborhoods between different metropolitan areas in 2000 are more similar to each other in terms of the homogeneity of their housing units.

Inaccessibility of Jobs

The inaccessibility of jobs serves as the fifth indicator of suburbanization, which is measured by the length of average daily commuting time of the central city residents[5] in metropolitan areas in this study. Suburban residents might also have longer commutes to get downtown, but they may also benefit from the increasing suburbanization of jobs. Restricting the measure to central city residents avoids this complication. More suburbanized metropolitan areas will have longer commute times for their central city residents and are anticipated to have higher economic segregation.

The last panel in Table 2 indicates that the national mean of daily commuting time for central city residents was 21.6 minutes in 1990. After ten years, the national mean value becomes 24.5 minutes, increasing more than 13%. This rising trend is quite widespread as well; among the 325 metropolitan areas, 322 of them (99%!) experienced increases in daily commuting length. The central city residents of Bismarck, North Dakota and Grand Forks, North Dakota/Minnesota had the least daily travel time at 11.4 minutes in 1990 and 13.3 minutes in 2000 respectively. For both years, New York central city residents suffered the longest average commutes of 36.4 and 39.9 minutes daily.

Overall, based upon the discussions above, the suburbanization indicators portray a picture of suburban development in the metropolitan areas in this country during the last decade. The urban density gradients and the estimated central city densities for most metropolitan areas declined.

The length of average daily commuting time of the central city residents increased slightly. Our empirical work also reveals that these suburbanization indicators have diverse regional variations (Yang, 2005). More important, all these indicators have strong corrrelations with the size of metropolitan areas. These correlations between suburbanization indicators and metropolitan size portend multicollinearity problems in the cross-sectional regression analyses discussed below.

ANALYSIS

Modeling Strategy

To investigate the effects of suburbanization indicators on economic segregation, we employ two series of regression models based upon different estimation techniques. The cross-sectional regressions emphasize differences between metropolitan areas, with explicit controls for many possibly confounding variables. However, not all differences among metropolitan areas that might cause bias (those that are both correlated with suburbanization and have an independent effect on economic segregation) can be easily controlled. The fixed-effects estimations focus on the changes within metropolitan areas over the 1990s, and implicitly control for all factors that are constant over time within metropolitan areas during that period.

For the cross-sectional regressions, we calculate the Neighborhood Sorting Index (NSI) in both 1990 and 2000 for non-Hispanic whites and other racial/ethnic groups, blacks, and Hispanics respectively, for all metropolitan areas in United States that had at least 10,000 households for that group in that year. Considering the limited degrees of freedom, the six resulting sets of NSI (three race groups, two time periods) are pooled together as dependent variables. In the first-difference, fixed-effects models, we calculate the metropolitan-level changes in NSI for non-Hispanic Whites and other racial/ethnic groups, Blacks, and Hispanics in the 1990s. The three resulting sets of changes in the NSI are pooled. To deal with the substantial heteroskedasticity associated with size of the metropolitan area, all the cross-sectional regressions are weighted by the square root of the number of households entering into the denominator of the dependent variable (Maddala, 1977, 268–69). Similarly, all of the fixed-effects models are weighted by the square root of the average number of 1990 and 2000 total households in metropolitan areas. Additionally, because of the pooling method, observations from the same metropolitan area are not independent; the standard errors are adjusted to allow for covariance of these observations.

Control Variables

In addition to the suburbanization indicators, we employ control variables of a number of different types, as discussed below.

Metropolitan Context

Metropolitan areas with more elderly residents are likely to have less economic segregation because people in this group are less mobile, even though they have accumulated wealth (Pendall & Carruthers, 2003). Economic segregation may be lower for metropolitan areas with more old housing stock, since there are usually more neighborhoods with mixed housing types available (Pendall & Carruthers, 2003). Compared to smaller households, larger households are less mobile and not able to quickly respond to and reflect rapid shifts in metropolitan economic patterns, rents, and housing prices. Therefore, metropolitan areas with more large-size households should have less economic segregation (Pendall & Carruthers, 2003).

A rapid influx of new households puts pressure on the housing market and will reduce economic segregation at least in short run. A variable for in-migration, measured as the proportion of metropolitan area residents aged five years or older who lived outside the metropolitan area five years before the base year of the decade, is included to address this impact. The coefficient of this variable is expected to be negative. In contrast, a high rate of internal turnover in the housing market may advance the ecological process and increase economic segregation. This effect is captured by the proportion of metropolitan area residents (except for the in-migrants) who moved within the metropolitan area in the previous five years.

There are many other idiosyncratic features of metropolitan areas, such as geophysical configuration, room for expansion, accumulated housing stock, historical ownership patterns, institutional arrangements, information networks, and so on, that may influence economic segregation. We cannot hope to control for all of them, but many such variables would vary by region of the country. In the cross-sectional models (Table 3), a set of eight dummy variables for U.S. census divisions are included to capture such features if they are correlated with regional geography. In

TABLE 3

Neighborhood Sorting Index: Pooled Cross-Sectional Weighted Least Squares Regression Results, U.S. Metropolitan Areas, 1990 and 2000

Variables	Model 1	Model 2	Model 3	Model 4	Std. coef.
Constant	0.454***	0.388***	−0.682***	−0.516***	- -
Black	0.025**	−0.001	0.002	−0.007	−0.024
Hispanic	0.047***	0.009	0.011	0.005	0.014
Year=2000	−0.077***	−0.100***	−0.148***	−0.143***	−0.760
Black ∗ (Year = 2000)	0.006	0.008	0.008	0.010	0.026
Hispanic ∗ (Year = 2000)	−0.024***	−0.018*	−0.018	−0.016*	−0.038
Suburbanization indicators					
Density gradient		−0.101**		−0.040	−0.030
Density/100		0.002*		0.001	0.100
$(\text{Density}/100)^2$		−0.003***		−0.001*	−0.184
No. of governments		−0.001***		−0.001*	−0.063
Homogenity new. dev.		−0.043***		−0.030**	−0.053
Avg. travel time		0.008***		0.001	0.063
Metropolitan context[a]					
Elderly			0.002**	0.002	0.056
Old housing stock			−0.001	−0.0003	−0.038
Households > 4			0.000	−0.00002	−0.001
In-migration			−0.003***	−0.003***	−0.186
Internal turnover			0.006***	0.005***	0.268
Structural economic characteristics					
Mean HH Inc. ($1,000s)			0.011***	0.009***	0.946
Mean HH $\text{Inc.}^2/100$			−0.005***	−0.004**	−0.510
Manufacturing share			0.000	−0.0003	−0.021
Managerial/professional			0.001*	0.001	0.064
Social distance					
Poverty rate			0.000	0.0001	0.008
Gini index			1.078***	1.037***	0.342
R^2	0.1830	0.5299	0.6795	0.6883	
N	1,089	1,089	1,089	1,089	

Notes. $^{*}p < 0.10$; $^{**}p < 0.05$; $^{***}p < 0.01$ (two-tailed tests).
[a]Models 3 and 4 include eight census division dummies, not shown separately.

TABLE 4

Changes in the Neighborhood Sorting Index: Pooled First-Difference Fixed-Effects Weighted Least Squares Regression Results, U.S. Metropolitan Areas, 1990–2000

Variables (changes)	Model 1	Model 2	Model 3	Model 4	Std. coef.
Constant	−0.075***	−0.085***	−0.089***	−0.090***	- -
Black	0.005	0.006	0.020**	0.021**	0.146
Hispanic	−0.032***	−0.035***	−0.024**	−0.024**	−0.142
Suburbanization indicators					
Density gradient		−0.446**		−0.224	−0.099
ln(Cen. city density)		0.527***		0.306*	1.019
(ln(Cen. city density))^2		−0.028***		−0.016*	−0.924
No. of governments		0.001**		0.001***	0.085
Homogenity new. dev.		−0.026**		−0.009	−0.031
Avg. travel time		0.003		0.002	0.050
Metropolitan context					
Elderly			−0.010***	−0.008***	−0.137
Households > 4			−0.010***	−0.009***	−0.229
Ln(population)			0.114***	0.079**	0.168
In-migration			−0.004***	−0.003***	−0.219
Internal turnover			−0.003**	−0.002	−0.081
Structural economic characteristics					
Mean HH Inc. ($1,000s)			0.001	0.0006	0.077
Manufacturing share			−0.004**	−0.003**	−0.106
Managerial/professional			0.0006	−0.0003	−0.010
Social distance					
Poverty rate			0.003*	0.004*	0.182
Gini index			−0.753***	−0.721***	−0.201
R^2	0.0370	0.1044	0.1485	0.1650	
N	523	523	523	523	

Note. $^*p < 0.10$; $^{**}p < 0.05$; $^{***}p < 0.01$ (two-tailed tests).

the fixed-effects models (Table 4), such factors are implicitly controlled if they are constant over time.

The size of metropolitan area, measured by the log of population, is expected to have a positive effect on economic segregation since large metropolitan areas may have greater internal differentiation of neighborhoods than do small areas (Hoch, 1987). However, high correlations between this variable and the suburban indicators present a serious multicollinearity problem with this variable. Controlling for metropolitan area size is more easily accomplished in the fixed effects models.

Structural Economic Characteristics

A second set of factors targets on the local opportunity structure—the structural economic characteristics of the metropolitan area. Overall, mean household income of a metropolitan area may have a nonlinear impact on economic segregation (Jargowsky, 1996). An increase in mean household income may come about in a variety of ways; for example, the wealthy households may become wealthier, or the poor residents may be catching up. The expected effect of mean household income is ambiguous.

Economic restructuring has affected various features of urban spatial structure (Frey & Speare, 1988; Kasarda, 1985; Kleinberg, 1995; Squires, 2002), and it may also affect economic segregation within racial/ethnic groups. A smaller share of jobs in manufacturing and a growing share of jobs in the management and professional related occupations may increase income inequality, as well as economic segregation by drawing skilled minority individuals away from traditional minority enclaves to jobs in dispersed locations (Jargowsky, 1996). Studies show that deindustrialization has increased inner-city distress (Galster, Mincy & Tobin, 1997), even as the expansion of knowledge-based industries has helped ameliorate it (Glaeser, 1999). Therefore, declines in the manufacturing employment proportion should increase economic segregation as new firms locate in a more dispersed pattern with concurrent adjustments in the residence patterns of employees (Jargowsky, 1996). Since rising skill requirements can accentuate social class differences within racial/ethnic groups, decrease group cohesion, and increase economic segregation within racial/ethnic communities (Wilson, 1980), higher management and professional related occupation ratios may lead to increased economic segregation.

Social Distance and Inequality

The final set of control variables focus on social distance and inequality. The social distance within each racial/ethnic group is measured by poverty rate. According to middle-class flight hypothesis (Wilson, 1987), a higher poverty rate within a racial/ethnic group should produce an increase in economic segregation, because it may encourage more privileged group members to isolate themselves spatially (Massey & Eggers, 1993; Jargowsky, 1996). Furthermore, rapid increases in income inequality since the late 1960s may increase both overall residential segregation among income groups and economic segregation within racial and ethnic groups (Massey & Eggers, 1993). Thus, metropolitan areas with a larger Gini coefficient, an index of income inequality, should have higher economic segregation.

Descriptive statistics for the cross-sectional and fixed-effects models are shown in Appendix Tables A1 and A2, respectively.

Cross-Sectional Models

Table 3 presents the results for the cross-sectional weighted least squares regression models. In Model 1, the pooled NSI is regressed against a set of dummy variables for race, year, and their interaction terms. This regression, in effect, recreates the upper panel in Table 1, with the coefficients representing unconditional differences between the cells of that panel. For example, the constant term is 0.454, the average NSI value for white and other races in 1990. The black coefficient of 0.025 is the difference in the NSI between blacks and the white and other groups in 1990. The Year 2000 coefficient of -0.077 shows the sharp decline in the NSI shown in Table 1.

Model 2 adds five indicators characterizing suburbanization development patterns. Except for the number of local governments per 100,000 households, all of the coefficients of these indicators demonstrate expected directions, and all are statistically significant at $p < 0.10$ or better. The urban density gradient is negatively associated with economic segregation levels; more suburbanized metropolitan areas with less steep urban density gradients have higher levels of economic segregation. This negative effect is statistically significant. The gross population density affects economic segregation in a nonlinear fashion as the literature suggested. Specifically, the coefficient of density is positive while the coefficient of its square term appears negative, indicating that economic segregation tends to rise when density increases, but at a decreasing rate (Gujarati, 2003, 226–29). Both of these coefficients are significant. The homogeneity of new development is measured by the weighted mean of housing value quartile ratios in the most recently developed suburban neighborhoods; smaller values of this indicator represent more homogeneous development. The highly significant negative coefficient of this variable indicates metropolitan areas with more homogeneous newly built suburban neighborhoods have a higher level of economic segregation as expected. Furthermore, the coefficient of the inaccessibility of jobs is positive and highly significant. The average daily commuting time of central city residents increases by 1 more minute in a metropolitan area; its economic segregation level rises by 0.008.

By controlling for these suburbanization indicators, the coefficient of the year dummy variable became more negative, changing from −0.077 to −0.100 and was still highly significant. This change suggests that the decrease in economic segregation would have been stronger if the suburbanization indicators had been constant. Additionally, the coefficients for the dummy variables for blacks and Hispanics decrease in magnitude and become insignificant after the introduction of variables controlling for suburbanization patterns.

The black and Hispanic coefficients decline sharply and lose their statistical significance when the suburbanization indicators are entered into the model. The control variables in Model 3 have the same effect. This suggests that the baseline differences in economic segregation among the racial groups are explained by the structural features of the metropolitan areas in which they reside, including suburban development patterns.

Models 3 and 4 include the control variables discussed above. Model 3 includes the control variables alone, and Model 4 includes both the suburbanization indicators and the control variables. Standardized coefficients for Model 4 are also shown to facilitate comparisons across variables. The inclusion of control variables reduces but does not eliminate the effect of suburbanization on economic segregation. The density gradient, the linear density term, and average travel time lose statistical significance; the other suburbanization variables retain significance.

Some of the findings for the control variables are of interest as well. The proportion of the elderly, aged housing stock, and large households was not significant in the full model. Metropolitan areas that had more in-migration had lower economic segregation as predicted. More internal turnover in the metropolitan housing market seems to facilitate higher economic segregation, also as predicted. The metropolitan area's average household income affects economic segregation in a nonlinear fashion. The coefficient for mean household income is positive whereas the coefficient for its square term is negative, indicating that economic segregation rises when average household income increases, but at a decreasing rate. Metropolitan areas with a large share of jobs in the manufacturing sector had smaller values in NSI. The industrial and occupation structure variables were not significant in the full model. Both the social distance measures had the expected sign (positive), but only the coefficient for Gini coefficient was statistically significant, supporting the hypothesis that increasing income inequality leads to higher economic segregation within racial/ethnic groups.

By comparing Models 3 and 4, we see that adding the suburbanization variables to a model including the control variables adds little to the R^2 of the model. Recall that these models include dummy variables for Census divisions (New England, Mid-Atlantic, etc.). Thus, only differences in the suburbanization variables within divisions can contribute to the coefficients in those models. Even so, an F-test indicates that the suburbanization measures are jointly significant ($F = 4.03$, $p < 0.001$).

One control variable not included in the regressions is the log of metropolitan area population. All of the suburbanization variables are significantly correlated with size, at high levels ranging from 0.3347 (density) to 0.7624 (average travel time). When the log of population is added to the regression (not shown), the suburban variables as a group remain significant ($F = 5.2, p < 0.001$), but the signs and pattern of significant coefficients change erratically. While the results in Models 1 through 4 are suggestive, the cross-sectional results are plagued by possible omitted variable bias based on the exclusion of metropolitan area size. The fixed-effects models, presented below, implicitly control for metropolitan area size as well as other variables that were constant within metropolitan areas between 1990 and 2000.

Fixed-Effects Models

Table 4 presents the results for the set of first-difference, fixed-effects weighted least squares regression models. Such models implicitly control for many idiosyncratic features of metropolitan areas that may influence economic segregation and that are unchanging over the period in question (such as geophysical configuration, room for expansion, accumulated housing stock, and historical ownership patterns), even if they are not quantifiable (such as institutional structure, information networks, and cultural aspects). The fixed-effects estimation structure thus helps to minimize the potential left-out variable bias. Variables that are constant over time must be dropped from these models, such as the housing stock built before 1939 and the census division dummy variables. The general size of the metropolitan areas is also implicitly controlled: the fact that New York is huge and Yuba, California is relatively small did not change between 1990 and 2000. The change in the log of population (roughly, the percentage change in population) can be included in these models, since the rate of population *growth* is not so highly correlated with the other variables.

As shown in Table 4, Model 1 includes only the dummy variables for race and ethnic groups. The constant in this regression represents the component of the change in economic segregation common to all these four racial/ethnic groups in the 1990s—a decrease about 0.075. There was no significant difference between blacks and the base case, whites and others, in the magnitude of the decrease. However, economic segregation for Hispanics declined by 0.032 more than the common decrease.

Model 2 adds the changes in indicators characterizing suburbanization development patterns. All of the coefficients of these indicators demonstrate expected directions and most of them are statistically significant. The changes in the urban density gradient are negatively associated with the changes in economic segregation; and the coefficient is significant. As the urban density gradient lessens, a metropolitan area is considered to be more suburbanized, and a smaller decrease (or larger increase) in economic segregation is observed over the last decade. As expected, density affects the changes in economic segregation in a nonlinear fashion: the coefficient for the changes in estimated central city density is positive whereas the coefficient for its square term appears negative; both of these coefficients are highly significant.

The coefficient for exclusionary zoning—the number of governments per 100,000 persons—is positive as anticipated and significant. The significant negative coefficient for the changes in the homogeneity of new development indicates that economic segregation will increase more (or decrease less) if the most recently developed neighborhoods in suburbs are more homogeneous in terms of their housing values. Finally, the coefficient for the change in the inaccessibility of jobs is positive as expected, but not significant. The introduction of changes in the suburbanization indicators made the constant term increase in magnitude from –0.075 to –0.085, again suggesting that the decrease in economic segregation would have been larger if the suburbanization indicators had been constant over the 1990s.

Parallel to the cross-sectional models shown above, Models 3 and 4 in Table 4 include the control variables, with and without the suburbanization variables. In this full model, the coefficients for all suburbanization indicators remain in the expected directions. However, the coefficients for the estimated central city density and its square term are only significant at the borderline ($p < 0.10$), and the coefficient for changes in the exclusivity of the local zoning regimes remains highly significant. For the full model, standardized coefficients are also shown.

The control variables perform largely as expected. Cities with larger increases in the proportions of the aged population over the decade had larger decreases in the NSI. In addition, metropolitan areas with increases in the proportions of large-size households over the 1990s had larger decreases in economic segregation. Both the coefficients for these variables are significant. Population increases had a significant positive coefficient, indicating that metropolitan areas with larger population expansions over the decade had smaller decreases in the NSI. As anticipated, cities with higher proportions of residents moving into the area had larger decreases in economic segregation over the 1990s. The change in the proportion of residents moving within the metropolitan area, however, was not statistically significant in the full model.

The change in mean household income over the decade did not have a significant effect on the change in economic segregation. As expected, declines in the share of jobs in the manufacturing sector, which were the norm, are associated with increases in income segregation, even after controlling for the percentage of change in mean income over the decade. The effect of increases in the share of jobs in professional and management-related occupations, however, was not statistically significant in any of the models.

The coefficient for changes in the group's poverty rate has the expected positive coefficient and is statistically significant. Increased poverty within the racial and ethnic groups thus produced smaller decreases (or larger increases) in economic segregation, supporting the hypothesis of middle-class flight (Wilson, 1987). The Gini coefficient is negative and statistically significant. This finding is contrary to the findings in the cross-sectional estimations and our expectation that increased overall income inequality leads to increases in economic segregation within racial/ethnic groups. One possible explanation is that the unexpected sign on this coefficient reflects a short-run disequilibrium condition. Annual income probably changes faster than do residential patterns. If the overall variance in the distribution of household income changes faster than persons can change neighborhoods to reflect their new economic status, the between-neighborhood proportion of that variance may temporarily dip.

CONCLUSION

Rapid suburbanization is changing the face of urban America. Suburban development is not only transforming the geospatial landscape of U.S. metropolitan areas, but is also triggering a variety of changes in the social, economic, and political arenas. Many commentators have bemoaned the single-income character of much suburban development, but as yet few studies directly linked suburban development with a causal impact on economic segregation.

To investigate this question, we examined the trend in economic segregation and the patterns of suburbanization during the last decade, and then modeled the relationship between these two using U.S. Census and PUMS data.

Economic segregation, as measured by the NSI, decreased significantly for all racial and ethnic groups during the last decade, reversing the earlier increasing trend from 1970 to 1990. This is consistent with the trends for other indicators of socioeconomic segregation (Jargowsky, 2003; Jargowsky & Yang, 2006; Kingsley & Pettit, 2003).

Since there is little agreement on how suburbanization ought to be measured, we used five specific metropolitan-level indicators suggested by the relevant literature. These suburbanization indicators are urban density gradient, population density, homogeneity of new growth, exclusivity of local zoning (proxied by the number of governments), and inaccessibility of jobs (proxied by average travel time of central city residents). Together, they painted a picture of suburban development: both the urban density gradients and the estimated central city densities for most metropolitan areas declined; the length of average daily commuting time of the central city residents increased.

Empirical results from cross-sectional regressions and fixed-effects estimations suggest that contemporary suburbanization patterns do contribute to economic segregation, although the particular indicators that are significant is sensitive to model specification. In baseline cross-sectional and fixed-effects models, the coefficients for the suburbanization indicators have significant effects on economic segregation. With a large set of controls, including dummies for census divisions, the suburbanization indicators were reduced in magnitude and either became insignificant or achieved lower levels of statistical significance. In addition, the estimates from the cross-sectional models may be biased by the omission of metropolitan area size due to the high correlations between size and the suburbanization variables. The first-difference fixed-effect estimations, which implicitly control for metropolitan area size and features related to it, generally supported the results from the cross-sectional models.

Several suburbanization indicators appeared to be individually significant even in the fully controlled models. For example, the coefficient for the homogeneity of new development obtained the anticipated sign (negative) in the pooled cross-sectional regressions and remained individually significant in the full model. Moreover, the zoning regimes proxy variable had the expected sign and was significant, even after the inclusion of controls, in the fixed-effect model.

The empirical results presented here suggest that the contemporary suburbanization process does play a role in the general trend toward higher levels of economic segregation. True, there were decreases in economic segregation in the 1990s, but the results presented here suggest that those declines would have been even larger without the countervailing effect of American-style suburban development. It will be hard to replicate the economy of the 1990s, while at the same time suburban development has continued on the same course. The result is likely to be a resumption in the longer-term trend of growing economic segregation.

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ENDNOTES

1 The five metropolitan areas having no central city are Bergen, NJ; Brazoria, TX; Middlesex, NJ; Monmouth, NJ; and Nassau, NY.

2 The metropolitan area having no real suburbs is Anchorage, AK, where the central city is coextensive with the metropolitan area.

3 Individuals in PUMs areas that span MSA boundaries were included in all the spanned MSAs, with sample weights adjusted proportional to the population proportions in the spanned MSAs, based on estimates obtained from the Master Area Block Level Equivalency (MABLE) Geographic Correspondence Engine. See http://www.oseda.missouri.edu/plue/geocorr/.

4 This configuration of race groups provides the best match of inconsistent categories in the 1990 and 2000 Census data on income by race at the census tract level.

5 "Central city residents" particularly refer to employed population that are 16 years old or older, work outside home, and reside in central cities within a metropolitan area.

APPENDIX

TABLE A1

Descriptive Statistics for the Cross-Sectional Models

Variable	Mean	SD	Minimum	Maximum
Neighborhood Sorting Index	0.384	0.117	0.129	0.986
Indicator variables				
Black	0.251	0.434	0	1
Hispanic	0.154	0.361	0	1
Year = 2000	0.52	0.5	0	1
Suburbanization				
Density gradient	0.177	0.1	−0.04	0.657
Density	5.182	11.35	0.045	130.436
No. of governments	18.297	11.934	2.555	79.051
Homogenity new. dev.	1.706	0.236	1.351	3.42
Avg. travel time	19.881	4.204	11.359	39.915
Metropolitan context				
Elderly	12.28	3.293	4.375	34.712
Old housing stock	14.234	10.658	0.436	50.759
Households > 4	10.796	3.53	5.044	34.08
In-migration	20.053	7.246	6.825	54.438
Internal turnover	35.29	4.576	21.883	49.159
Structural economic characteristics				
Mean HH Inc. ($1,000s)	51.935	10.259	32.402	134.981
Manufacturing share	14.963	6.484	2.844	44.04
Managerial/professional	31.259	5.096	18.916	50.183
Social distance				
Poverty rate	16.227	9.872	2.833	52.768
Gini index	0.433	0.028	0.346	0.535

TABLE A2

Descriptive Statistics for the Fixed-Effects Models

Variable (changes)	Mean	SD	Minimum	Maximum
Neighborhood Sorting Index	−0.072	0.071	−0.47	0.233
Black	0.249	0.433	0	1
Hispanic	0.132	0.339	0	1
Suburbanization indicators				
Density gradient	−0.011	0.028	−0.132	0.192
ln(Cen. city density)	−0.006	0.192	−0.915	0.998
No. of governments	−0.85	4.585	−12.858	37.698
Homogenity new. dev.	0.007	0.219	−1.463	1.066
Avg. travel time	2.692	1.257	−2.311	8.43
Metropolitan context				
Elderly	0.188	0.902	−4.672	3.261
Households > 4	−0.251	1.252	−3.307	3.569
Ln(population)	0.128	0.099	−0.077	0.606
In-migration	−1.784	3.091	−16.654	4.603
Internal turnover	−0.413	2.161	−6.469	7.216
Structural economic characteristics				
Mean HH Inc. ($1,000s)	10.499	5.516	−3.254	32.262
Manufacturing share	−2.755	1.863	−10.463	7.086
Managerial/professional	3.259	1.768	−2.196	9.51
Social distance				
Poverty rate	−2.136	3.005	−12.515	5.041
Gini index	0.027	0.014	−0.036	0.079

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