Gentrification and Neighborhood Housing Cycles: Will America's Future Downtowns Be Rich?

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Abstract

This paper identifies a new factor, the age of the housing stock, that affects where high- and low-income neighborhoods are located in U.S. cities. High-income households, driven by a high demand for housing services, will tend to locate in areas of the city where the housing stock is relatively young. Because cities develop and redevelop from the center outward over time, the location of these neighborhoods varies over the city's history. The model predicts a suburban location for the rich in an initial period, when young dwellings are found only in the suburbs, while predicting eventual gentrification once central redevelopment creates a young downtown housing stock. Controlling for other determinants of where the poor live (e.g. proximity to amenities and public transit), empirical work indicates that if the influence of spatial variation in dwelling ages were eliminated, central-city/suburban disparities in neighborhood economic status would be reduced by up to 10 percentage points. Model estimates further predict that between 2000 and 2020, central-city/suburban differences in economic status will narrow in cities of all sizes and especially so in the larger metropolitan areas as American cities become more gentrified.

Key words: Gentrification, Redevelopment, Income Stratification

JEL Codes: R14, R20, R31

1. Introduction

In the early 1980s, property values in parts of Harlem had fallen to new lows, with entire neighborhoods occupied by the poorest of the poor (e.g. Schaffer and Smith (1986)). Twenty years later, former President Bill Clinton established his office in Harlem amid a run-up in real estate values that saw many homes selling for over one million dollars.[1] Harlem's meteoric rise in economic status reflects the phenomenon of gentrification, under which previously poor central-city neighborhoods are revitalized by the in-migration of well-to-do households. While gentrification elicits different evaluations, being viewed as harmful by advocates of the displaced poor and desirable by its beneficiaries, the phenomenon has generated intense interest among policymakers, the press, and other observers.

Several questions about the gentrification process need answers. Is gentrification just a curiosity that affects a handful of neighborhoods and makes good material for newspaper feature stories? Or does the gentrification phenomenon reflect deep-seated forces that could substantially alter current location patterns in U.S. cities? If the latter view is correct, can the future course of gentrification be predicted by economic models that take account of the dynamic nature of the process? Such predictions are crucial for local policymakers and real-estate developers, who must plan for the future despite their limited ability to predict the city's evolution (see Rosenthal (2007), for example).

This paper develops and tests a new model of the gentrification process, providing fresh insights into how high- and low-income households locate within a city and how their locations evolve over time. Our starting point is the standard urban model developed by Alonso (1964),

[1] See, for example, *USA Today* (July 31, 2001, page 1), ``I'm home: Clinton opens Harlem Office," by Charisse Jones.

Mills (1967) and Muth (1969). That model has been largely successful in explaining the spatial patterns of land-use and real estate prices in cities. However, the model has met with less success in predicting the location patterns of different income groups. Under the U.S. pattern, higher-income households tend to live in the suburbs, with central-city incomes averaging only 40 percent suburban incomes, as will be shown later. However, gentrification can also generate rich neighborhoods close to the city center, complicating the location picture. In addition, in foreign cities, the U.S. pattern can be completely reversed, with the center occupied by the rich and the poor living in the suburbs (a notable example is Paris).

Within the standard model, the location of high-income households reflects the tension between two opposing forces: a housing-based pull toward the suburbs, where a high demand for housing can be satisfied at a low price per square foot, and a time-cost-based pull toward the center, where short commutes allow high-income households to save valuable time. Over the years, the U.S. location pattern has often been rationalized by viewing the housing-based force as dominant. However, Wheaton (1977) presented early empirical evidence showing that the two forces are approximately equal in size, implying an indeterminate pattern of location by income. Using newer evidence, Glaeser, Kahn, and Rappaport (2007) argue that the housing-based force is far weaker than the time-cost force, which would imply a central location for high-income households, contrary to the general U.S. pattern. Given such findings, a viable explanation for the strong (and often opposing) location patterns seen in U.S. and foreign cities must therefore be sought outside the basic urban model.

Two such explanations have been proposed in the literature: the attractive effect of amenities, both built and natural, and access to public transit. Brueckner, Thisse and Zenou (1999) argue that topographical and historical amenities in the city center (an attractive river or beachfront; beautiful buildings, monuments, and parks inherited from the past) may attract the rich more strongly than the poor, leading to a Paris-style location pattern (see also Kern (1981)).[2] Others have suggested that fiscal amenities matter, arguing that homogeneous suburban communities allow high-income households to escape redistributive central-city taxation while improving the quality of public goods (see, for example, Nechyba and Walsh (2004)).[3] LeRoy and Sonstelie (1983) instead argue that transport mode choice may provide the key to location patterns. They show that when a new, fast transport mode (the streetcar, or more recently, the automobile) is introduced and adopted only by the rich, the resulting weakening of the time-cost force for that group may cause them to abandon their initial central location and move to the suburbs, a reversal consistent with the current U.S. pattern. Glaeser et al. (2007) offer a related mode-based argument and evidence, pointing out that, because central cities have the high population densities required for convenient, frequent public-transit service, they naturally attract a poor population, which must rely on this transit mode.

Focusing on amenities and public-transit access helps to explain the relative locations of high- and low-income households. But, being mostly based on static models, neither argument allows for neighborhood evolution over time and its implications for household location. This paper fills that gap by developing a dynamic model where the age of dwellings plays a central role in determining location patterns. The model portrays a city that grows outward from the center over time, with dwellings in each location torn down and replaced at the end of a fixed lifespan. Under this growth process, new dwellings are always found at the city's edge, but they

[2]Vancouver, British Columbia shares some of the features of Paris, with high-income families clustered around a very scenic downtown and lower-income families occupying more distant suburbs. A similar pattern occurs in La Paz, Bolivia, in which the downtown is located at a lower elevation (11,000 feet) with a more moderate climate, in contrast to the suburbs, which extend up to a 14,000 foot plateau that suffers from a harsher climate.
[3]Glaeser, Kahn, and Rappport (2001) also argue that central cities offer more generous social services for the poor.

will also be present at interior locations when enough time has passed for redevelopment to occur.

Consistent with filtering models (e.g. Sweeney (1974), Weicher and Thibodeau (1988), Rosenthal (2007)), housing in the model is assumed to deteriorate with age, providing the impetus for redevelopment.[4] Because the demand for housing services increases with income, high-income households are then drawn to areas of the city where dwellings are relatively young. If such dwellings are found only in the suburbs, then this force pulls the rich outward while the time-cost-based force pulls them inward, setting up a tension analogous to that in the standard model. But when redevelopment brings new dwellings to central-city neighborhoods, both forces work in tandem, pulling the rich toward the center. In this case, high-income households are likely to be found at both central and suburban locations, yielding the gentrification pattern seen in some U.S. cities. A further implication of the model is that dwellings of a given age are more likely to be inhabited by a rich household closer to the CBD. As will become clear in section 2 below, this conclusion predicts a negative, rather than positive, association between income and distance, holding dwelling age fixed.

The empirical analysis, contained in sections 3 and 4 of the paper, is designed to test the simple hypothesis implied by this theoretical finding. The hypothesis states that *the usual positive association between income and distance may be weakened or reversed after controlling for the effect of dwelling ages*. The empirical work makes use of detailed census tract data for 331 individual MSAs. The initial regressions treat distance from the city center in linear fashion, implying a monotonic relationship between distance and a neighborhood's economic status. We

[4]Additional theoretical papers on filtering include Braid (1986) and Brueckner (1980). Other empirical studies include Galster (1996), Coulson and Bond (1990), and Aaronson (2001). For a micro-level empirical analysis of the economics of housing redevelopment, see Rosenthal and Helsley (1994). See Glaeser and Gyourko (2005) for a discussion of how falling housing demand in declining cities can preclude redevelopment.

then test for a positive correlation between income and distance like that seen in many previous studies. In all cases, the neighborhood's relative income (census tract mean income divided by the MSA mean) is our measure of local economic status. Based on this measure, the tendency for neighborhood economic status to rise with distance from the city center is confirmed in the 2000 data.

Next, we incrementally add to the model controls for local amenities and neighborhood-level access to public transit, factors that have previously been emphasized in the literature as affecting location patterns. Local amenities are taken into account by including school district fixed effects in the model. Access to public transit is measured using a dummy variable indicating whether more than 10 percent of the two-decade lagged (1980 versus 2000) census tract population relied on public transit when traveling to work. For cities of all sizes, each of these factors reduces the positive impact of distance on neighborhood economic status. In addition, in all instances, access to public transit reduces neighborhood economic status, consistent with the findings of Glaeser et al. (2007).

Our central hypothesis is tested by further adding information on the census tract's dwelling age distribution to the regression model. Summary measures in the data indicate that the average age of a dwelling falls as distance from the city center increases while the concentration of newly built homes rises. Both of these patterns indicate that younger dwellings are disproportionately located in the suburbs. As anticipated, controlling for this systematic spatial variation in dwelling ages further attenuates the effect of distance on neighborhood economic status for cities of all sizes, confirming our basic hypothesis. In the linear regressions, our point estimates actually imply a reversal of the "standard" relationship for most cities: a tract's economic status tends to fall rather than rise as distance increases, holding dwelling age fixed. These results suggest that high-income households would tend to live closer to the city centers were it not for old central housing stocks.

To allow for nonmonotonic distance effects, further regressions divide the city into narrow distance bands (1 or 2 miles in width), each represented by a dummy variable. This approach is useful because the rural areas of most MSAs tend to have lower economic status than the suburbs, yielding a humped-shaped relationship between distance and economic status that is obscured in the linear distance specification. Once again, the various control measures are added incrementally to the regression model, and the previous conclusions re-emerge. In particular, the spatial pattern of neighborhood economic status largely flattens over broad areas extending from the downtown through the suburbs of most cities, with the positive effect of distance no longer present. The controls for local amenities, access to public transit, and dwelling ages all contribute to this flattening effect. Dwelling ages in particular further reduce current central-city/suburban income differentials by up to 10 percentage points, holding constant other factors. These findings, based on a very flexible specification, further confirm our hypothesis that central-city/suburban income differences would be smaller were it not for systematic spatial variation in dwelling ages.

An important feature of our analysis is the use dwelling age and public transit measures from 1980, rather than 2000, to explain a tract's economic status in 2000. Using deeply lagged age measures helps to control for the possible endogenous character of dwelling ages. In addition, given this feature of the regression, the estimated model parameters can be used to forecast neighborhood economic status in 2020. This forecast, which is generated by substituting year-2000 measures in place of 1980 measures in the regression equation, yields an answer to the question posed in the title of this paper. The results show that, for metropolitan areas of all sizes, central-city/suburban differences in economic status are predicted to narrow over the next several decades, with the gap in the largest cities reduced by up to one-fourth. Moreover, these effects are driven primarily by shifts in the spatial distribution of dwelling ages, consistent with our model. Nevertheless, our estimates predict that the central cities of the largest cities in the U.S. are still expected to be poor on average relative to suburban communities.

2. Conceptual Model

2.1. City growth and the spatial pattern of dwelling ages

As explained in the introduction, the goal of the model is to analyze the spatial pattern of location by income in a city that grows over time while experiencing cyclical redevelopment as dwellings age and are periodically replaced. For simplicity, the city is monocentric, with all employment located in the center, and the dwelling lifespan is set exogenously at L years, where L is a fixed constant independent of time and location. The effect of variable dwelling lifespans is discussed below.

The fixed lifespan assumption generates a distinctive spatial pattern of dwelling ages in a growing city. To understand the pattern, suppose that the city comes into being as a dimensionless point at time $t = 0$ and expands outward as time progresses. Let $T(x)$ denote the "development-date function," which gives the date at which land x miles from the city center (CBD) is first developed, satisfying $T'(x) > 0$. Then consider a date t in the city's evolution before any redevelopment has occurred, a value satisfying $t < L$. At this date, dwelling ages follow the spatial pattern shown in Figure 1a, with new dwellings located at the city's edge and age rising as x decreases toward zero. Letting $a(x,t)$ denote age at distance x and time t, $a(x,t) = t - T(x)$ holds when $t < L$.

As time progresses beyond L, redevelopment occurs for dwellings near the CBD, which reach the end of their service lives and are replaced. These new dwellings then begin aging as time progresses, yielding the age pattern shown in Figure 1b. At this date, the central part of the city has second-generation dwellings, while first-generation dwellings occupy its outer area. As time progresses further, spatial expansion of the city continues, and these first and second-generation dwellings are replaced by second and third generations, yielding the age pattern shown in Figure 1c.[5] Thus, spatial growth along with the periodic redevelopment of dwellings generates a cyclical spatial pattern of dwelling ages, with $a(x,t)$ a periodic function of x when $t > L$.[6]

2.2. Additional elements of the model

To develop the implications of dwelling age patterns like those shown in Figure 1 for the pattern of location by income, the additional elements of the model must be specified, as follows. City residents consume a nonhousing tradable good, denoted c, and housing services, denoted h. The physical size q of each dwelling is fixed for simplicity, but housing services depend negatively on the dwelling's age a, with $h = h(a)$, where $h'(a) < 0$.[7] Preferences are given by the well-behaved utility function $v(c,h)$.[8]

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[5] The Figures are generated numerically using a parameterization presented in the appendix, and Figures 1a,b,c correspond to t values of 63.7, 114.3, and 187.5, respectively, in the underlying calculations (see below for details).

[6] For arbitrary t and x, $a(x,t) = \psi[(t - T(x))/L]L$, where the function $\psi(z)$ gives the fractional part of z (the value found by dropping z's integer part).

[7] The effect of allowing dwelling square footage to be a choice variable is considered in supplemental analysis that is available on request.

[8] This expression suppresses location specific non-tradable amenities (e.g. climate, school quality) for simplicity. Such attributes are taken into account in the empirical work by including school district fixed effects in the regressions.

City residents commute to jobs at the CBD, and without loss of generality, time cost is assumed to be the only cost of commuting, with money cost suppressed. These time costs are generated in the simplest possible fashion by assuming that a minute of commute time reduces work time by a minute (leisure is thus fixed). Letting $y(t)$ denote "full" income at time t (the amount that would be earned with no commute), disposable income at location x is then $y(t)(1 - \tau x)$, where τ gives the fraction of the full work period lost per mile of commuting, assumed constant over time and across households for simplicity.[9] Time-t commuting cost from location x is thus $y(t)\tau x$. Letting R denote the rent per dwelling, the consumer's budget constraint is thus $c + R = y(t)(1 - \tau x)$.

Substituting for c in v, utility can be written $v[y(t)(1 - \tau x) - R, h(a)]$. The consumer's bid-rent function, denoted $R(x,t)$, gives willingness-to-pay at time t for the housing located at distance x, conditional on a fixed utility level $u(t)$. The bid-rent function is implicitly defined by the equation $v[y(t)(1 - \tau x) - R, h(a(x,t))] = u(t)$, and differentiating this equality to find its slope function yields[10]

$$\frac{\partial R}{\partial x} = -\tau y(t) + \frac{v_h}{v_c} h'(a) \frac{\partial a}{\partial x}. \qquad (1)$$

Recalling that housing services decrease with age ($h'(a) < 0$) and that age decreases with distance between the jumps in Figure 1 ($\partial a/\partial x < 0$), the second expression in (1) is positive between these jumps. Since the first expression in (1) is negative, the bid-rent function can either slope up or down over the distance ranges between the jumps. The intuitive explanation is that, while higher commuting costs make more distant locations less desirable, lower dwelling

[9] In the empirical work, we allow for spatial variation in proximity to public transit, given that such access likely affects commute times differently for high- and low-income households.
[10] Note that since $a(x,t)$ is not differentiable at the jump points in Figure 1, (1) applies between these points.

ages have the reverse effect. Since the implications of the model do not depend whether (1) is positive or negative, the slope is assumed to be positive.

2.3. Patterns of location by income: preliminaries

To investigate the pattern of location by income, let the city contain two income groups, denoted rich and poor, with incomes $y_r(t)$ and $y_p(t) < y_r(t)$. Each group has its own bid-rent function, with the slopes given by

$$\frac{\partial R_r}{\partial x} = -\tau y_r(t) + \left(\frac{v_h}{v_c}\right)_r h'(a)\frac{\partial a}{\partial x} \tag{2}$$

$$\frac{\partial R_p}{\partial x} = -\tau y_p(t) + \left(\frac{v_h}{v_c}\right)_p h'(a)\frac{\partial a}{\partial x}, \tag{3}$$

where the subscripts on the MRS expressions indicate evaluation at different consumption bundles.

To proceed further, suppose that Figure 2a applies, so that the city has just a single generation of dwellings, with the age function exhibiting no jumps (this Figure is an enlarged version of 1a). Then, the location pattern of the two groups depends on the relative magnitudes of the slopes in (2) and (3), following well-known principles. If the rich slope is algebraically larger, then the rich outbid the poor for housing far from the CBD, with the poor outbidding the rich close to the CBD. Given that the rich have a higher time cost, this outcome requires a strong income effect on housing demand, with the MRS expression in (2) suitably larger than the expression in (3). When the bid-rent curves are upward sloping, as assumed, the resulting slope relationship means that the rich curve is steeper than the poor curve. Examples of the resulting bid-rent curves are shown in Figure 2b, with the dotted curve belonging to the rich group.[11]

When Figure 2b is relevant, the model generates a traditional US-style location pattern, with the rich living in the suburbs, occupying newer dwellings, and the poor living close to the CBD, occupying older dwellings. This pattern is shown in Figure 2c, which graphs an index identifying the group residing at each location, with the value 1 corresponding to the rich and the value -1 corresponding to the poor. For convenience, the pattern shown in Figures 2b and 2c is referred to subsequently as the "rich-decentralization" case.

2.4. Gentrification

The key insight from the model comes from investigating how redevelopment of centrally-located dwellings, as captured in the transition from Figure 1a to 1b, affects the pattern of location by income. The analysis shows that, in the rich-decentralization case, central redevelopment leads to gentrification, with some rich consumers moving into the central city.

To establish this property of the model, focus on a time $\hat{t}$ that satisfies $L < \hat{t} < 2L$, so that the city has two generations of dwellings, as in Figure 1b. Let $\hat{b} = b(\hat{t})$ denote the distance to the edge of the city, let $\hat{y}_r = y_r(\hat{t})$ and $\hat{y}_p = y_p(\hat{t})$ denote incomes, and let $\hat{x}$ denote the location of the boundary between the second and first-generation dwellings. Furthermore, let $\hat{x}^-$ and $\hat{x}^+$ denote locations just to the left and right of $\hat{x}$.

Consider first the behavior of bid-rents as they cross the $\hat{x}$ boundary, focusing on the rich curve. These bid-rents must adjust to equate rich utilities at locations $\hat{x}^-$ and $\hat{x}^+$,

[11] If the bid-rent curves were instead downward sloping, Figure 2 would show the rich curve as less steep. Again, the sign of the slopes is not crucial for the analysis.

responding to the jump in age (and hence the drop in housing services) across these two locations. Thus, the equality

$$v\left[\hat{y}_r(1-\tau\hat{x})-R_r(\hat{x}^-),h(0)\right]=v\left[\hat{y}_r(1-\tau\hat{x})-R_r(\hat{x}^+),h(L)\right] \tag{4}$$

must hold, where the $\hat{t}$ argument of R_r is suppressed for notational simplicity. But since $h(0)$, housing services from a new dwelling, is much larger than $h(L)$, services from a dwelling that is about to be replaced, (4) implies that $R_r(\hat{x}^-)$ must be much larger than $R_r(\hat{x}^+)$ to equate utilities between $\hat{x}^-$ and $\hat{x}^+$. Since the same conclusion holds for the poor bid-rent, it follows that both curves show downward discontinuities at $\hat{x}$, as seen in Figure 3b.

The key second conclusion is that, if the rich outbid the poor for housing at $\hat{b}$, then they must also offer a higher bid for housing at $\hat{x}^-$. Thus, when the city has second generation dwellings, the rich must outbid the poor for the outermost of these dwellings, implying the existence of an area of gentrification in the central city. To establish this conclusion, note that since dwelling age equals zero at both $\hat{x}^-$ and $\hat{b}$, consumption of c must be the same at both locations in order to equate utilities, a condition that holds for both rich and poor. From the budget constraints, the bid-rents must then satisfy the following equalities:

$$\hat{y}_r(1-\tau\hat{x})-R_r(\hat{x}^-)=\hat{y}_r(1-\tau\hat{x})-R_r(\hat{b}) \tag{5}$$

$$\hat{y}_p(1-\tau\hat{x})-R_p(\hat{x}^-)=\hat{y}_p(1-\tau\hat{x})-R_p(\hat{b}) \tag{6}$$

But subtracting (5) from (6) yields

$$R_r(\hat{x}^-)-R_p(\hat{x}^-)=R_r(\hat{b})-R_p(\hat{b})+(\hat{y}_r-\hat{y}_p)\tau(\hat{b}-\hat{x})>R_r(\hat{b})-R_p(\hat{b}). \tag{7}$$

Eq. (7) says that the rich-poor bid-rent differential is larger at $\hat{x}^-$ than at $\hat{b}$, which implies that if the rich outbid the poor at $\hat{b}$, as assumed, they also do so at $\hat{x}^-$. This conclusion follows because the willingness-to-pay for housing of a given age rises faster for the rich than for the poor as distance falls, a consequence of their higher time costs. Summarizing yields

> **Proposition 1.** *When a growing urban area characterized by the rich-decentralization case reaches the point where second-generation dwellings appear in the central city, gentrification occurs, with rich households occupying the outermost of these dwellings.*

Proposition 1 is illustrated in Figure 3.[12] Note that, while the rich-poor bid-rent differential is larger at $\hat{x}$ than at $\hat{b}$ from (8), this difference is small enough to be indistinguishable in the Figure (recall that the dotted curve belongs to the rich).

Figure 3 shows a gentrification pattern in which rich households occupy all of the city's second-generation dwellings, outbidding the poor over the entire distance range between the CBD and $\hat{x}$. This occupancy pattern is shown in Figure 3c. However, as the second-generation area expands outward, with its oldest dwellings approaching the end of their lifespans, poor households will reappear in the city center, being the highest bidders for dwellings in the innermost area. This outcome in shown in Figure 4. Thus, moving away from the CBD, the location pattern is poor-rich-poor-rich, as seen in Figure 4c.[13] This pattern easily generalizes to a city with an arbitrary number of dwelling generations.[14]

For a location pattern like that in Figure 4c, an additional result on occupancy patterns across dwelling generations can be established via the argument leading to Proposition 1:

> **Proposition 2.** *In the rich-decentralization case, the critical age beyond which occupancy switches from rich to poor is older in higher generation dwellings (those closer to the city center) at a given date in the city's evolution.*

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[12] Figure 3 pertains to a slightly earlier time than Figure 1b ($t = 88.3$).

[13] Cyclical location patterns like this one also emerge in the model of Bartolomé and Ross (2004), where the existence of separate political jurisdictions can disrupt an otherwise monotonic relation between income and distance. Note that Figure 4 pertains to $t = 120.8$.

[14] It should be noted that the model also admits an extreme version of gentrification, where the rich entirely disappear from the suburbs once the second-generation area in the central city becomes large enough. In this case, the initial location pattern by income is completely reversed, at least temporarily, as the city grows.

(see the appendix for a proof). Thus, the oldest dwelling containing a rich household will be older in the central area of a city, where higher-generation dwellings are located, than in the suburbs. The intuitive reason is that, because willingness-to-pay (holding age constant) rises relatively rapidly for the rich as distance falls, they are able to bid away housing farther up into the age distribution in later generation dwellings, which are located closer to the CBD.

Note that the Proposition 2 can be restated in a converse form, which will prove to be useful in motivating the empirical work. In particular, the Proposition implies that, looking across dwellings of a common age at different locations within the city, dwellings closer to the CBD are more likely to have a rich occupant. This conclusion follows because the critical age below which dwellings remain in rich hands is higher near the CBD

As can be seen from the above analysis, all the qualitative properties of the model can be derived from a given $T(x)$ function, which yields the age function $a(x,t)$. The $T(x)$ function is itself generated endogenously by considering the development decision of the landowner, and the appendix presents this calculation for a special case. Parameterization of this special case then yields the diagrams shown in the previous Figures.

2.5 Empirical implications

What guidance does the model provide for empirical work on spatial income patterns in cities? Consider first the model's implications for a standard regression, which estimates the simple relationship between neighborhood income and distance. If the gentrification area in Figure 3 is small relative to the suburban high-income area, then a linear regression of income on distance could well produce the usual positive coefficient, although a more flexible specification would show a U-shaped relationship. But now consider the effect of including dwelling age along with distance in the regression model. First, given normality of housing and the model's assumption that housing services fall with age, the coefficient on age should be negative, indicating that neighborhoods with old housing have lower incomes, holding distance fixed. Second, the model implies that the coefficient on distance should be negative, not positive, indicating the neighborhood income *falls with distance*, holding dwelling age fixed. This conclusion follows from Proposition 2, which implies that dwellings of a given age are more likely to have a rich occupant closer to the CBD, or equivalently, more likely to have a poor occupant farther from the CBD. The conclusion generates the main empirical hypothesis tested in the paper: *in a regression that controls for variation in dwelling ages, the usual positive association between neighborhood economic status and distance should be weakened or reversed.*

3. Econometric Model

3.1. Basics

In contrast to the model, real cities have a distribution of incomes rather than just two income groups, and dwellings in any given neighborhood do not share a common age but instead exhibit a range of ages. Recognizing these facts, the dependent variable for the regression analysis is a census tract's "relative income" in 2000, equal to average household income in the tract divided by the average household income for the MSA. Income values are taken from the Geolytics Inc. Neighborhood Change (NCBD) database. In addition, dwelling ages are captured by the age distribution of the housing stock, as measured by the shares of dwellings in a tract falling in different age ranges (see below). A tract's distance from the CBD is computed using geographic information software (MapInfo and MapBasic), and the CBD is identified using a population-density criterion.[15]

To test the hypothesis stated above, we estimate regressions of the following form:

$$y_{2000,j} = D(x_j, z_j; \delta_x, \eta) + \omega_{2000,j} \qquad (9)$$

and

$$y_{2000,j} = D(x_j, z_j; \delta_x, \eta) + A(a_{2000,j}; \theta) + \omega_{2000,j} \ . \qquad (10)$$

In these regressions, $y_{2000,j}$ is the relative income of neighborhood *j* (census tract *j*) in 2000, x_j is the tract's distance from the CBD, and z_j represents local amenities and access to public transit, as previously emphasized in the literature. The variable $a_{2000,j}$, represents the age distribution of the housing stock in 2000 and is new to this paper, while $\omega_{2000,j}$ is an error term. In addition, D is a function that describes the relationship between neighborhood economic status, distance and z_j, with δ_x representing the distance coefficient(s) and η the coefficients on *z*. Similarly, A is a function describing the relationship between economic status and age, with θ being a coefficient vector.

Our strategy is to estimate both (9) and (10) and then to compare the effects of distance on neighborhood economic status. If the dwelling age distribution has no effect on the pattern of location by income, holding the other covariates constant, then the estimated influence of distance should be alike in the two regressions. But when dwelling age matters, the coefficients from (10) reveal how neighborhood economic status would vary with distance if dwelling age and the other covariates were held constant. Our hypothesis predicts that the usual positive effect of distance on economic status should be weakened or reversed.

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[15]The CBD is defined as the centroid of the Census tract with the highest population density, and distance from the centroid of a given tract to this point is computed. Population densities are generated using GIS software.

3.2. Endogenous dwelling ages

While the model above treats the timing of housing redevelopment and thus dwelling ages as exogenous, redevelopment would occur endogenously in a more general model. In such a model, the housing developer would compare the gain in profit from continuing to operate an aging dwelling to the lost profit (in present value terms) from a slight delay in redevelopment. The dwelling is replaced when these two quantities are equal.[16]

Although endogenous redevelopment by itself would not materially change the spatial implications of the analysis, adding a stochastic element to such a model is a potential source of econometric trouble. For example, suppose unobservable neighborhood amenities exist and that they are valued more highly by the rich than by the poor. Then, since the bid-rent offer of the rich will be relatively high in a high-amenity neighborhood, developers will be eager to redevelop its aging dwellings, switching occupancy from the poor to the high-paying rich. Under this scenario, the error term $\omega_{j,2000}$ will tend to be large in high-amenity neighborhoods, indicating that they have higher than average incomes. But since quick redevelopment means that dwelling ages will also tend to be low, the result is a negative correlation between age and the error term, leading to inconsistent parameter estimates in (10).

Our solution to this problem is to instrument for the potentially endogenous age distribution of the housing stock. We do so by replacing the age variables $a_{j,2000}$ in (10) with the census tract dwelling age distribution from 1980, $a_{j,1980}$.[17] Lagged age measures are good instruments because, for dwellings not redeveloped, the year-2000 age is simply the year-1980 age plus 20. This link ensures correlation between the 1980 and year-2000 dwelling age

[16]Brueckner (1981) models this behavior in a steady-state environment without urban growth, while Braid (2000) treats the more-difficult non-steady-state case. While both models assume a single income group, their approach could be adapted to the present two-group case, with some additional complications.

distributions. In addition, the 1980 dwelling age distribution will be exogenous to unanticipated amenity shocks in the subsequent twenty years, reducing correlation between the 1980 age measures and $\omega_{j,2000}$ in (10).[18]

3.3 Additional covariates

As explained above, the regressions also include controls for local amenities and access to public transit. To control for local amenities, we include fixed effects for each of the several thousand unified school districts represented in the sample. While these fixed effects control for school quality, they also capture innumerable local natural and produced amenities that are correlated with the quality of a school district (e.g. local crime rates, public services, parks, scenic settings, etc.). In total there are 50,511 census tracts in the sample, belonging to 1,979 unified school districts and 331 MSAs. On average, therefore, there are roughly 25 census tracts per school district, representing a population of approximately 100,000 people.

To control for tract-level access to public transit, we specify a variable that equals one if 10 percent or more of the census tract's commuters used public transit and zero otherwise. The discrete nature of this measure captures the idea that, if a very small share of the local population uses public transit for commuting, then public transit is not very accessible from that location. Because we set the cutoff for this variable at a low level, we treat public-transit access measured in this fashion as exogenous. To further ensure exogeneity, we compute the public transit variable using 1980 data, analogous to our treatment of the census tract's dwelling age distribution. An added bonus to using lagged measures for both these variables is that it enables

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[17] In an earlier version of the paper we also estimated models using 1990 controls for dwelling ages. Results were similar to those presented here based on 1980 values.

[18] Employing the lagged age measures is possible because our data have been uniformly coded to year-2000 census tract boundaries, circumventing intercensus boundary changes. See the website for Geolytics Inc. for detail: www.geolytics.com.

us to forecast change in neighborhood economic status between 2000 and 2020, as will be clarified below.

4. Estimation Results

4.1 The linear specification

Because of its familiarity from a host of past studies, we first present results based on a linear specification of the D function in (9), initially using MSA fixed effects as the single other covariate. While this specification reveals average patterns in the data, it is restrictive in imposing a monotonic relationship between distance and economic status. This deficiency is remedied in a more flexible specification used in the following subsection.

Table 1 presents results for the linear distance regressions, in which census tracts from all MSAs are pooled together. We begin in column 1 with the restricted version of the model described in expression (10). This specification implicitly assumes that within-MSA amenities, access to public transit, and the local distribution of dwelling ages do not affect spatial variation in neighborhood economic status. Consistent with conventional wisdom, the coefficient on the distance variable is positive and significant. The point estimate implies that a one-mile increase in distance from the CBD increases the relative income of a neighborhood by 0.0018, indicating that income rises relative to the MSA mean by about 0.2 percent.[19]

Column 2 begins to assess the determinants of spatial variation in neighborhood economic status in earnest. In this column, we replace MSA fixed effects with school district fixed effects to better control for local amenities. This change reduces the distance coefficient to essentially zero. A zero effect reflects the presence of attractive local amenities (especially school quality) in the suburbs, which tend to draw income families away from the downtown. Controlling for that effect leads to a more uniform spatial distribution of income.

Column 3 adds public-transit access to the model. This variable's coefficient is negative and strongly significant, confirming that neighborhoods with good transit access tend to attract poor households, as found by Glaeser et al. (2007). Observe also that the coefficient on distance remains positive and close to zero.

Column 4 extends the model further by adding the dwelling age measures. These measures consist of the percentages of census-tract dwellings that were 0 to 4 years, 5 to 9, 10 to 19, 20 to 29, and 40 or more years in age in 1980. The age 30-to-39 category is omitted to avoid perfect collinearity.

The age coefficients confirm the basic premise of our model by showing that neighborhood economic status indeed depends on the age of the housing stock. Column 4 shows that an increase in the dwelling share in any age category outside the 30-39 year range leads to higher economic status for the census tract, with the effect strongest for newly produced housing in 1980 (the age 1-4 year category). Recognizing that dwellings in the 30-39 age range in 1980 will advance into the oldest age category by 2000, this coefficient pattern makes sense. Being the oldest in 2000, these dwellings will disproportionately attract poor households in that year, depressing neighborhood status. Therefore, an increase in the 1980 share of dwellings *outside* the 30-39 age category should, by reducing the tract's share of the oldest dwellings in 2000, raise

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[19] While that magnitude is small, we show shortly that treating distance in a linear manner greatly understates central-city/suburban income differentials. That bias is corrected shortly when we adopt a more flexible function form.

economic status in that year, a pattern that is confirmed by the positive coefficients for all the other age categories.[20]

The most important finding from column 4 is that, after controlling for dwelling ages, the distance coefficient switches sign, becoming negative. The emergence of a negative coefficient, which is equal to -0.0010 and highly significant, confirms our basic hypothesis, establishing that the suburban locational tendency of higher-income households would be weakened or reversed if the effect of spatial variation in dwelling ages were eliminated. Note that this sign reversal implies that dwelling age is negatively correlated with distance, a fact that is confirmed below. Thus, failing to control for dwelling ages leads to an upward bias in the effect of distance on neighborhood economic status, even after controlling for local amenities and access to public transit.

To explore these patterns further, we next stratify the model by city size. This stratification is appropriate given that Table 1 pools cities with vastly different spatial sizes, and because larger cities tend to be older, thus having different spatial distributions of dwelling ages. Accordingly, Tables 2a and 2b present separate regressions for four different size categories of MSAs. Table 2a shows results for cities with fewer than 100 census tracts and cities with 100 to 500 tracts, with Table 2b showing results for cities with 500 to 1,000 tracts and more than 1,000 tracts. The structure of these tables is otherwise the same as for Table 1.

Consistent with the previous results, the regressions with only MSA fixed effects show economic status rising with distance in all but the third size class of MSAs. In addition, for all

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[20] In contrast to the above argument, dwellings of unusually high quality may often be destined for extensive care and long lives. As a result, for some very old dwellings, housing quality may actually be quite high (as with historic buildings). However, historic dwellings of unusually high quality are likely a small subset of all dwellings over 40 years in age, justifying the view that the oldest age category disproportionately attracts poor households. Further support for this view comes from observing that the newest housing in 1980 has by far the most positive impact on neighborhood economic status in 2000. See Rosenthal (2007a and 2007b) for further discussion on these points.

four city size categories, successively controlling for local amenities (through school district fixed effects) and public-transit access weakens the positive association between distance and neighborhood economic status. This weakening continues with the addition of dwelling ages to the model. Once dwelling ages are added, distance has a negative and usually significant impact on neighborhood economic status. The only exception is in the largest city size category, where the distance coefficient is marginally significant (its t-ratio is -1.31). These results once again affirm our basic hypothesis.[21]

4.2. The distance-dummy specification

In this section, we adopt a more flexible specification of the D function by breaking distance into discrete bands, each represented by a dummy variable. The D function is now given by

$$\begin{aligned} \mathrm{D}(x_{\mathrm{j}}, z_{\mathrm{j}}; \delta_x, \eta) = \mathrm{D}(z_{\mathrm{j}}; \eta) + \delta_{0to1} x_{0\mathrm{to}1} + \ldots + \delta_{9to10} x_{9\mathrm{to}10} \\ + \delta_{10to12} x_{10\mathrm{to}12} + \ldots + \delta_{38to40} x_{38\mathrm{to}40} + \delta_{40\mathrm{orMore}} x_{40\mathrm{orMore}} \end{aligned} \tag{11}$$

where the distance bands are one mile in width up to 10 miles and two miles in width from 10 to 40 miles, with 40-miles-or-more as the final category (the x's denote the dummy variables). Note that the 0-to-1 mile distance range is omitted, so that the remaining dummy coefficients indicate tract economic status relative to the city center. The term $\mathrm{D}(z_{\mathrm{j}}; \eta)$ represents controls for local amenities and public-transit access, as before.

To begin exploring the results under this new specification, consider first Figure 5a, which plots average dwelling age within each distance band for the four city-size categories. As asserted above, average dwelling age generally declines with distance from the CBD for each

[21] Observe also that for each city size class in Tables 2a and 2b, public-transit access has a significantly negative impact on local economic status and the age coefficients show the same pattern as in Table 1.

MSA size category, with some notable variation across categories. The interval of lower ages near the CBDs of the largest cities testifies to the kind of central redevelopment envisioned in the model of section 2. The flat range of average age in the largest cities is also noteworthy, as is the rapid age decline with distance in the smallest category. Figure 5b displays analogous plots of the percentage of the housing stock constructed in the previous ten years (between 1990 and 2000). By this measure, it is also clear that newer housing is disproportionately concentrated in the suburbs. Failing to control for these patterns leads to an overstatement of the effect of distance on neighborhood economic status, as seen in Tables 1 and 2.

Figures 6a-6d display the spatial pattern of relative incomes in the four MSA size categories, providing the analog to the basic regressions in Table 2 for the distance-dummy specification. In each figure, the heavy black curve shows the pattern of distance-dummy coefficients from a regression that includes MSA fixed effects as the only control apart from the distance dummies. The thin dotted black curve shows the effect of replacing MSA with school district fixed effects, our control for local amenities. The thin dotted light colored line shows the impact of further controlling for public-transit access. Finally, the thick light colored curve shows the pattern of distance dummy coefficients when the dwelling age measures are added to the regression.

We focus first on the heavy black lines as these show the actual spatial distribution of neighborhood economic status, allowing that distribution to be shifted up or down by the MSA fixed effects. Observe that, among MSAs of all sizes, neighborhood economic status varies in a humped shaped pattern with distance from the center: relative income is low close to the CBD, rises sharply with distance, achieves a global peak at some intermediate distance, and then declines in outlying portions of the metropolitan area that are likely rural in nature. Observe also that the peak in neighborhood economic status lies farther away from the CBD in the largest cities. Also noteworthy is that, in the largest MSA class, economic status tends to fall with distance near the CBD before rising, apparently indicating ongoing gentrification close to the center of the largest cities. A final important observation is that the difference in economic status between central cities and suburbs is large. Among the largest cities, for example, from trough to peak, the central-city/suburban difference in economic status is nearly 50 percent of MSA mean income.[22]

Controlling for local amenities through school district fixed effects generally flattens the spatial pattern of neighborhood economic status, but by varying degrees depending on city size. The effect of controlling for local amenities is slight among the smallest MSAs (Figure 6a), but the impact is greater in the larger cities (Figures 6c and 6d). This larger impact may be indicative of more extensive Tiebout sorting among larger cities, which have many school districts.

When public-transit access is added to the regressions, the distance-dummy coefficient patterns show little change in the smallest MSAs (Figure 6a). In contrast, public-transit access notably weakens the tendency of neighborhood economic status to rise with distance in the two middle city size categories, as seen in Figures 6b and 6c. Among the largest MSAs, controlling for public-transit access has a relatively small additional impact on economic status, except in neighborhoods close to the city center. For these locations, removing the influence of public-transit access increases economic status, consistent again with the findings of Glaeser et al (2007).

[22] This number is derived by subtracting the minimum relative income value on the black curve in Figure 6d from the maximum value.

The effect of adding dwelling age controls is our principal focus, and Figures 6a-6d show that the presence of the dwelling age measures further attenuates the effect of distance on economic status. The incremental effect is largest among the smallest MSAs. As seen in Figure 6a, for these cities, the thick light colored curve is roughly ten percentage points below the other curves over most distance bands, and is considerably flatter. Effects are similar qualitatively but of lesser magnitude among the larger cities, as seen in Figures 6b-6d. For these cities, controlling for dwelling ages reduces neighborhood economic status most in the higher income portions of the MSAs, and typically by roughly 5 percentage points. Once again, these patterns imply that central-city/suburban economic income differentials would be reduced if the effect of spatial variation in dwelling ages were eliminated.[23]

The magnitude of the combined effects of local amenities, public-transit access, and dwelling ages is also important. For cities of all four size categories, eliminating spatial variation in dwelling ages, as well as making public-transit access and county amenities uniform, would reduce central-city/suburban differences in economic status by roughly half of the actual disparity.[24] Although only a portion of this reduction in income disparities can be attributed to the dwelling-age effect, the patterns seen in Figures 6a-6d nevertheless confirm the paper's basic hypothesis.

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[23] A further feature of these figures warrants a brief comment. Notice especially that for the two larger MSA size categories, the distance-income relationship is nearly flat over an extended range from roughly 8 to 28 miles beyond the city center. This pattern could be reflective of Wheaton's (1977) original conclusion that, as household income rises, the housing-based pull toward the cheaper suburbs (where high housing quantities are more easily purchased) approximately cancels the time-cost-based pull toward the center. An alternative explanation, however, is that suburban employment subcenters have become common, and especially among larger cities (e.g. Garreau (1991)). The presence of these subcenters allows workers to choose whether to commute inward to the downtown or outward toward more-outlying centers, weakening the CBD/suburban distinction in the household's location problem.

[24] Again, this number comes from subtracting the minimum value from the maximum relative income value on the thick light colored curve.

4.4. Will future downtowns be rich?

We return now to the question posed in the title of this paper: will the downtowns of U.S. cities achieve higher economic status in the future? Our first step is to compute within-sample fitted values for year-2000 census tract relative income using the estimated model coefficients. These values are denoted as $\hat{y}(z_{1980}, a_{1980}; \hat{\eta}, \hat{\delta}, \hat{\theta})$, where the tract and location subscripts are suppressed for convenience. Next, we assume that the model parameters do not change between 2000 and 2020. We then simulate 2020 census tract relative incomes twice, in each case by updating data used in the model from 1980 values to year 2000 values. In the first instance, we update both the dwelling age and public transit variables to obtain $\hat{y}(z_{2000}, a_{2000}; \hat{\eta}, \hat{\delta}, \hat{\theta})$. This is our best prediction of year-2020 census tract relative income. In the second instance, we update only the dwelling ages, leaving the public transit variables at their 1980 values to obtain $\hat{y}(z_{1980}, a_{2000}; \hat{\eta}, \hat{\delta}, \hat{\theta})$. This is our best estimate of year-2020 relative income if only dwelling ages changed. All of these predicted relative incomes are then scaled to ensure that each measure averages to 1 across census tracts, consistent with the manner in which the dependent variable in the model is created.[25]

Differencing $\hat{y}(z_{1980}, a_{2000}; \hat{\eta}, \hat{\delta}, \hat{\theta})$ and $\hat{y}(z_{1980}, a_{1980}; \hat{\eta}, \hat{\delta}, \hat{\theta})$ isolates the impact of changes in dwelling ages between 1980 and 2000 on the predicted change in relative income between 2000 and 2020. Differencing $\hat{y}(z_{2000}, a_{2000}; \hat{\eta}, \hat{\delta}, \hat{\theta})$ and $\hat{y}(z_{1980}, a_{2000}; \hat{\eta}, \hat{\delta}, \hat{\theta})$ isolates the further

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[25] Because the data used in creating the simulated relative incomes differs from that used to estimate the model parameters, the simulated relative incomes typically average across census tracts to values that differ from 1. Implicitly, this outcome has the effect of changing the MSA average level of income. But it should be emphasized that our model only addresses spatial variation in income *relative* to the downtown. It says nothing about income levels. Accordingly, all of our simulated relative income measures for 2020 are scaled to ensure that they average across tracts to 1, consistent with the manner in which the dependent variable is created.

impact of changes in public-transit access between 1980 and 2000. In both instances, notice that local amenities are held constant at their year-2000 values, as measured by the school district fixed effects.

To illustrate the predicted change in neighborhood economic status, we next run ordinary least squares regressions of the predicted relative incomes on the distance dummies. The corresponding distance dummies reflect the average value of the predicted relative income in a given distance band. Those values are plotted in Figures 7a-7d for the different MSA size categories. Several patterns are apparent.

For each MSA size class, forecast central-city/suburban income differentials are expected to narrow between 2000 and 2020. This outcome can be seen by comparing the black line with dots to the light colored line with dots. Observe also that the magnitude and geographic scope of the narrowing increases with the size category of the MSA. Among the smallest MSAs (Figure 7a), the narrowing is roughly 5 percentage points relative to MSA mean income, roughly one-sixth of the year-2000 central-city/suburban differential. Moreover, that narrowing is concentrated in the 2 to 6 mile range from the city center. Among the largest MSAs (Figure 7d), the narrowing is roughly 11 percentage points of MSA mean income, a decline of roughly one-fourth relative to the year-2000 central-city/suburban income differential. In this instance, most of the shift in relative income occurs from 10 to 38 miles from the center.

The plots in Figures 7a-7d also identify the independent contributions of public-transit access and dwelling ages to these forecast changes in relative income. This decomposition is accomplished by differencing the predicted relative incomes as outlined above. The contribution of public-transit access is shown by the light colored dashed line, while the contribution from dwelling ages is shown by the black dashed line. For each MSA size class, public-transit access contributes little to the forecast change in central-city/suburban income differentials. The largest effect is for the third MSA size category (500 to 1,000 census tracts), but even there the contribution from transit access is quite small. Instead, virtually all of the forecast change in 2000 to 2020 central-city/suburban income differentials is driven by changes in the spatial distribution of dwelling ages.

These results reaffirm important ideas discussed at the outset of this paper, as follows. Several mechanisms are believed to contribute to central-city/suburban differences in neighborhood economic status. These include local amenities, public-transit access, and the spatial distribution of dwelling ages. Results presented earlier (e.g. Tables 2a-2d and Figues 6a-6d) confirm the importance of each of these mechanisms. However, of these factors, only dwelling ages have an inherently dynamic component that is expected to drive systematic change in the spatial distribution of income within an MSA. This dynamic component is reflected in changes in the dwelling age distribution over the 1980-2000 period which account for the relatively large age contribution to the predicted changes in neighborhood economic status.

A final point evident from the patterns in Figures 7a-7d concerns whether America's future downtowns will be rich. Even though our model forecasts a reduction in the central-city/suburban disparity in economic status, central cities will nevertheless continue to be poorer than suburban communities.

5. Conclusion

This paper has provided new insights into the process of gentrification by identifying a new factor, the age of the housing stock, that affects patterns of location by income. Since their high demand for housing services draws high-income households toward newer dwellings, such households will tend to be found in areas of the city where the housing stock is relatively young.

While these areas will lie in the suburbs over much of a city's history, eventual redevelopment of aging dwellings in the center creates a young downtown housing stock that attracts high-income households, leading to gentrification. Moreover, gentrification is ultimately driven by the passage of time and the associated aging and obsolescence of the housing stock. This explanation is markedly different from the recent amenity and transit-based explanations for the location of poor neighborhoods, which are predominantly static in nature. Neither of these mechanisms implies systematic change in the location of poor neighborhoods over time.

The empirical work in the paper provides substantial support for our new perspective. Results show that local amenities, access to public transit, and dwelling ages all contribute to central-city/suburban differences in economic status. If spatial variation in these factors were eliminated, our estimates indicate that central-city/suburban disparities in neighborhood economic status would be reduced by up to half within American cities. Dwelling ages contribute significantly to that effect in cities of all sizes, with effects ranging from 5 to 10 percent of MSA mean income. Thus, if the housing age distribution were made uniform across space, reducing average dwelling ages in the central city and raising them in the suburbs, then neighborhood economic status would shift in response, rising in the center and falling in the suburbs.

Drawing on two-decade lagged measures of dwelling ages further allows us to forecast future changes in economic status in response to evolution of the housing stock through 2020. For MSAs of all sizes, the model forecasts a narrowing of central-city/suburban income differentials that increases with the size of the MSAs. Among the largest cities, suburban income relative to the city center is expected to decline by roughly 10 percent of MSA mean income between 2000 and 2020, roughly one-fourth of the year-2000 disparity. While this shift implies ongoing gentrification, central-city neighborhoods are nevertheless expected to remain poor, on average, relative to the suburbs.

Explaining patterns of location by income is a task that has absorbed substantial research effort by urban economists over the years. Despite this effort, consensus on a robust explanation for observed location patterns has been elusive. We hope that the perspective offered in this paper advances the state of this inquiry while providing a glimpse into the future of American cities.

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Table 1: Neighborhood Economic Status For All MSAs
(Dependent variable is Year 2000 census tract average income relative to MSA average income; t-ratios in parentheses are based on robust standard errors)

	Control for Distance	Control for Local Amenities	Control for Public Transit[a]	Control for Dwelling Age[b]
Distance (Miles from city center)	0.0018	0.00005	0.00011	-0.0010
	(9.08)	(0.23)	(0.45)	(-4.28)
Access to public transit in 1980			-0.1525	-0.0945
			(-18.06)	(-10.76)
% Dwellings 0 to 4 years in 1980				0.9967
				(34.01)
% Dwellings 5 to 9 years in 1980				0.3219
				(9.98)
% Dwellings 10 to 19 years in 1980				0.6197
				(21.49)
% Dwellings 20 to 29 years in 1980				0.6501
				(18.22)
% Dwellings 40 or more years in 1980				0.5365
				(16.66)
Constant	0.9772	0.9994	1.0329	0.4765
	(300.09)	(289.96)	(235.93)	(18.53)
Observations	50,511	50,511	48,735	48,437
MSA Fixed Effects	331	-	-	-
Unified School District Fixed Effects	-	1,979	1,927	1,925
Adj. R-Square	0.0016	0.1228	0.1372	0.1747
RMSE	0.4646	0.4345	0.4268	0.4168

[a]Public-transit access equals 1 if 10 percent or more of the tract's population in 1980 used public transit, and 0 otherwise.
[b]Dwelling ages are measured as of 1980. The omitted category is age 30 to 39 years.

Table 2a: Neighborhood Economic Status By MSA Size
(Dependent variable is Year 2000 census tract average income relative to MSA average income;
t-ratios in parentheses are based on robust standard errors)

	MSAs With Fewer Than 100 Tracts				MSAs With 100 to 500 Tracts			
	Control for Distance	Control for Local Amenities	Control for Public Transit[a]	Control for Dwelling Age[b]	Control for Distance	Control for Local Amenities	Control for Public Transit[a]	Control for Dwelling Age[b]
Distance (Miles from city center)	0.0014	0.0002	0.0006	-0.0014	0.0017	0.00001	-0.0003	-0.0010
	(3.41)	(0.33)	(1.09)	(-2.66)	(5.74)	(0.04)	(-0.63)	(-2.09)
Access to public transit in 1980			-0.2406	-0.1632			-0.2222	-0.1789
			(-9.97)	(-6.81)			(-18.81)	(-14.04)
% Dwellings 0 to 4 years in 1980				1.1951				0.9367
				(16.99)				(20.58)
% Dwellings 5 to 9 years in 1980				0.4470				0.1340
				(5.23)				(2.64)
% Dwellings 10 to 19 years in 1980				0.8572				0.5004
				(12.57)				(11.29)
% Dwellings 20 to 29 years in 1980				0.7386				0.5020
				(7.77)				(8.89)
% Dwellings 40 or more years in 1980				0.5132				0.4300
				(6.36)				(8.23)
Constant	0.9877	0.9987	1.0083	0.3526	0.9789	0.9999	1.0389	0.5871
	(199.88)	(179.72)	(162.91)	(5.53)	(203.10)	(173.14)	(144.96)	(14.34)
Observations	10,223	10,223	9,827	9,813	20,004	20,004	19,357	19,265
MSA Fixed Effects	217	-	-	-	91	-	-	-
Unified School District Fixed Effects	-	980	953	953	-	1,172	1,126	1,122
Adj. R-Square	0.0012	0.0461	0.0625	0.1711	0.0016	0.1394	0.1643	0.2038
RMSE	0.3422	0.3308	0.3304	0.3106	0.4376	0.4055	0.4003	0.3899

[a]Public-transit access equals 1 if 10 percent or more of the tract's population in 1980 used public transit, and 0 otherwise.
[b]Dwelling ages are measured as of 1980. The omitted category is age 30 to 39 years.

Table 2b: Neighborhood Economic Status By MSA Size
(Dependent variable is Year 2000 census tract average income relative to MSA average income;
t-ratios in parentheses are based on robust standard errors)

	MSAs With 500 to 1000 Tracts				MSAs With More Than 1000 Tracts			
	Control for Distance	Control for Local Amenities	Control for Public Transit[a]	Control for Dwelling Age[b]	Control for Distance	Control for Local Amenities	Control for Public Transit[a]	Control for Dwelling Age[b]
Distance (Miles from city center)	-0.0001	-0.0002	-0.0015	-0.0031	0.0053	0.0034	0.0020	-0.0012
	(-0.39)	(-0.33)	(-1.95)	(-4.17)	(9.13)	(4.29)	(2.20)	(-1.31)
Access to public transit in 1980			-0.1653	-0.1455			-0.1186	-0.1085
			(-8.51)	(-6.82)			(-5.74)	(-5.05)
% Dwellings 0 to 4 years in 1980				0.7163				1.0423
				(10.06)				(14.73)
% Dwellings 5 to 9 years in 1980				0.2042				0.4207
				(2.70)				(5.88)
% Dwellings 10 to 19 years in 1980				0.2611				0.6480
				(3.55)				(10.78)
% Dwellings 20 to 29 years in 1980				0.3533				0.6794
				(4.03)				(10.26)
% Dwellings 40 or more years in 1980				0.4719				0.6210
				(6.10)				(9.74)
Constant	1.0023	1.0036	1.0591	0.7076	0.9226	0.9494	1.0347	0.5125
	(134.22)	(82.03)	(71.18)	(10.41)	(88.73)	(66.51)	(49.67)	(10.48)
Observations	10,669	10,669	10,357	10,276	9,615	9,615	9,194	9,083
MSA Fixed Effects	17	-	-	-	6	-	-	-
Unified School District Fixed Effects	-	672	649	643	-	352	340	336
Adj. R-Square	0.0000	0.2214	0.2299	0.2561	0.0085	0.2339	0.2672	0.2860
RMSE	0.4836	0.4264	0.4245	0.4169	0.5893	0.5179	0.4898	0.4836

[a]Public-transit accesss equal 1 if 10 percent or more of the tract's population in 1980 used public transit, and 0 otherwise.
[b]Dwelling ages are measured as of 1980. The omitted category is age 30 to 39 years.

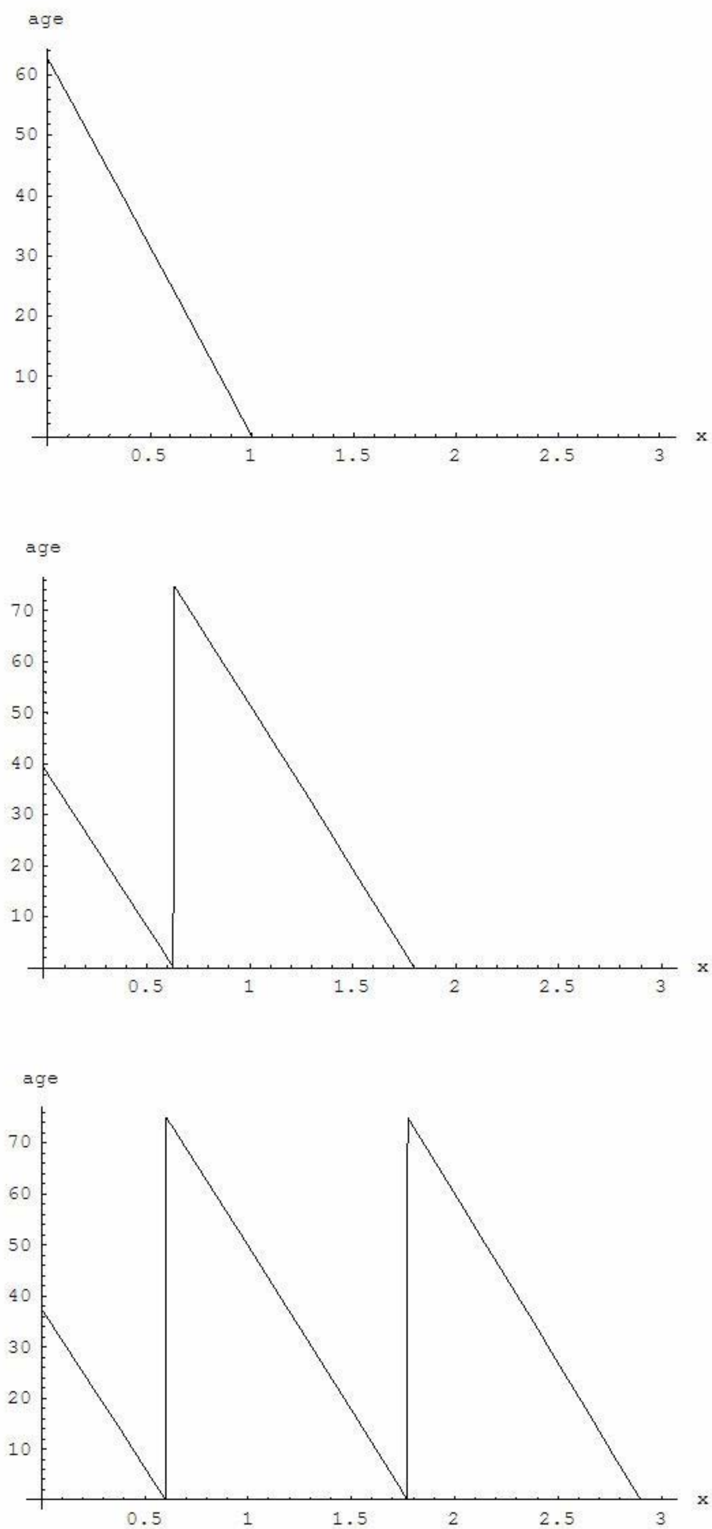

Fig. 1a

Fig. 1b

Fig. 1c

Figure 1: Dwelling Age Patterns

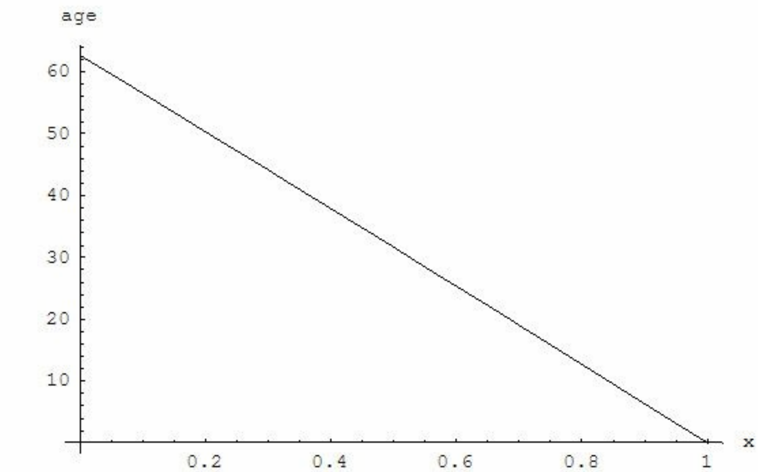

Fig. 2a

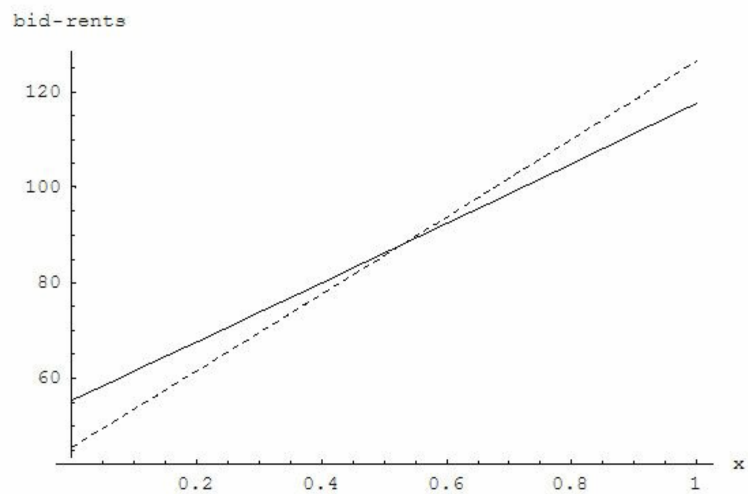

Fig. 2b

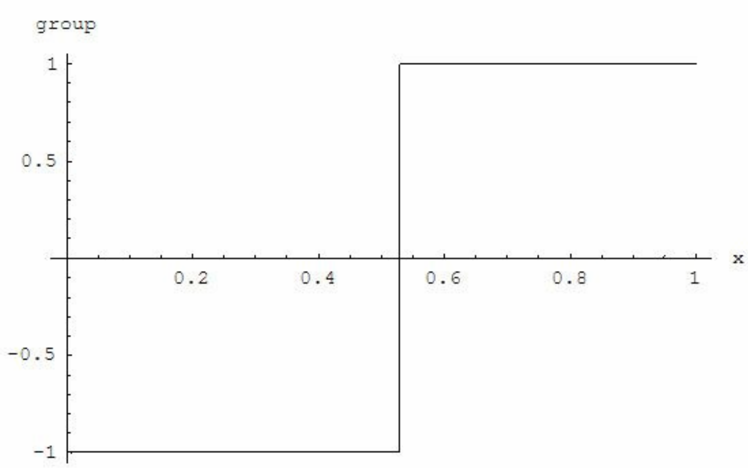

Fig. 2c

Figure 2: Rich in the Suburbs

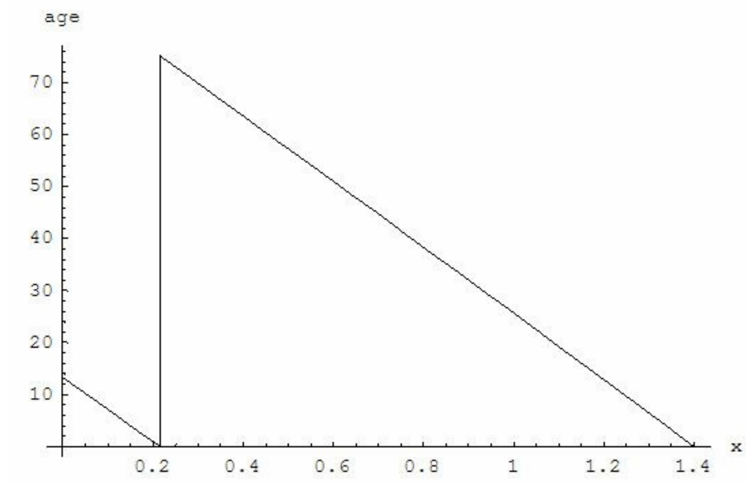

Fig. 3a

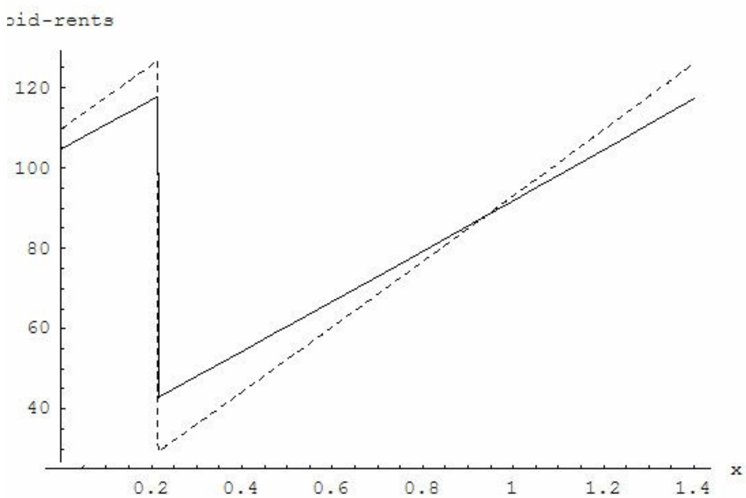

Fig. 3b

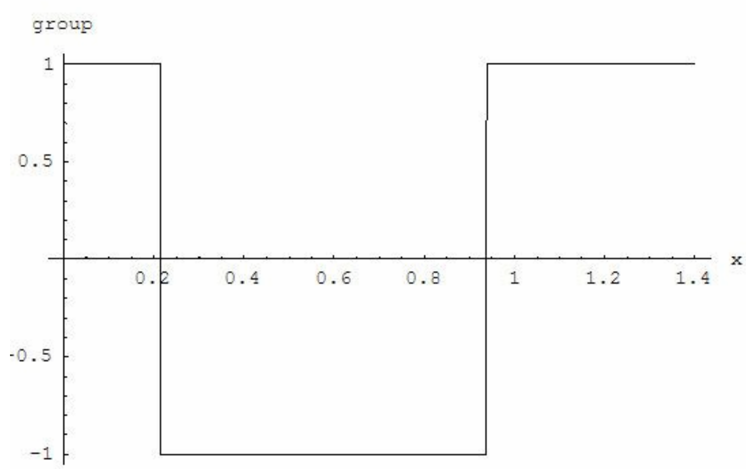

Fig. 3c

Figure 3: Gentrification

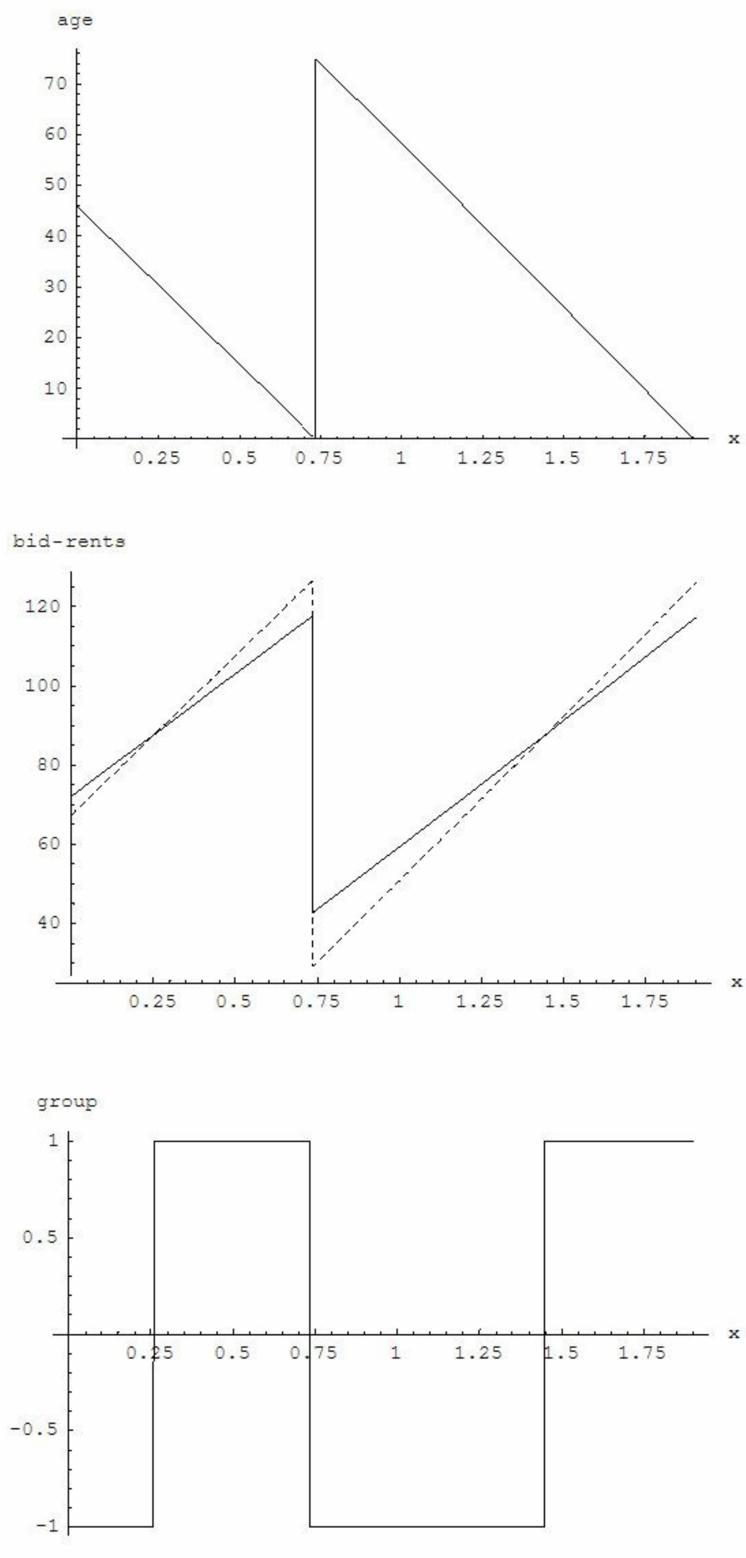

Fig. 4a

Fig. 4b

Fig. 4c

Figure 4: Poor Regain the Center

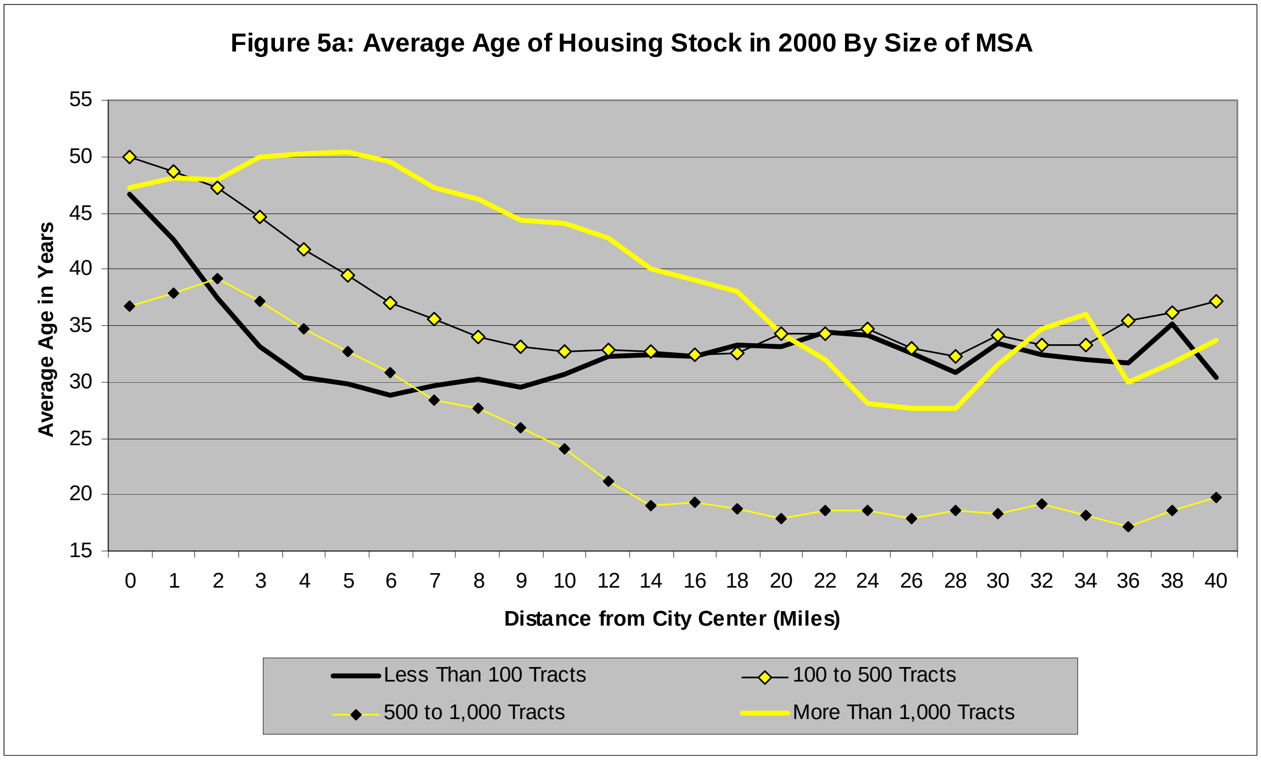
Figure 5a: Average Age of Housing Stock in 2000 By Size of MSA
55
50
45
40
35
30
25
20
15
Average Age in Years
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
Less Than 100 Tracts
100 to 500 Tracts
500 to 1,000 Tracts
More Than 1,000 Tracts

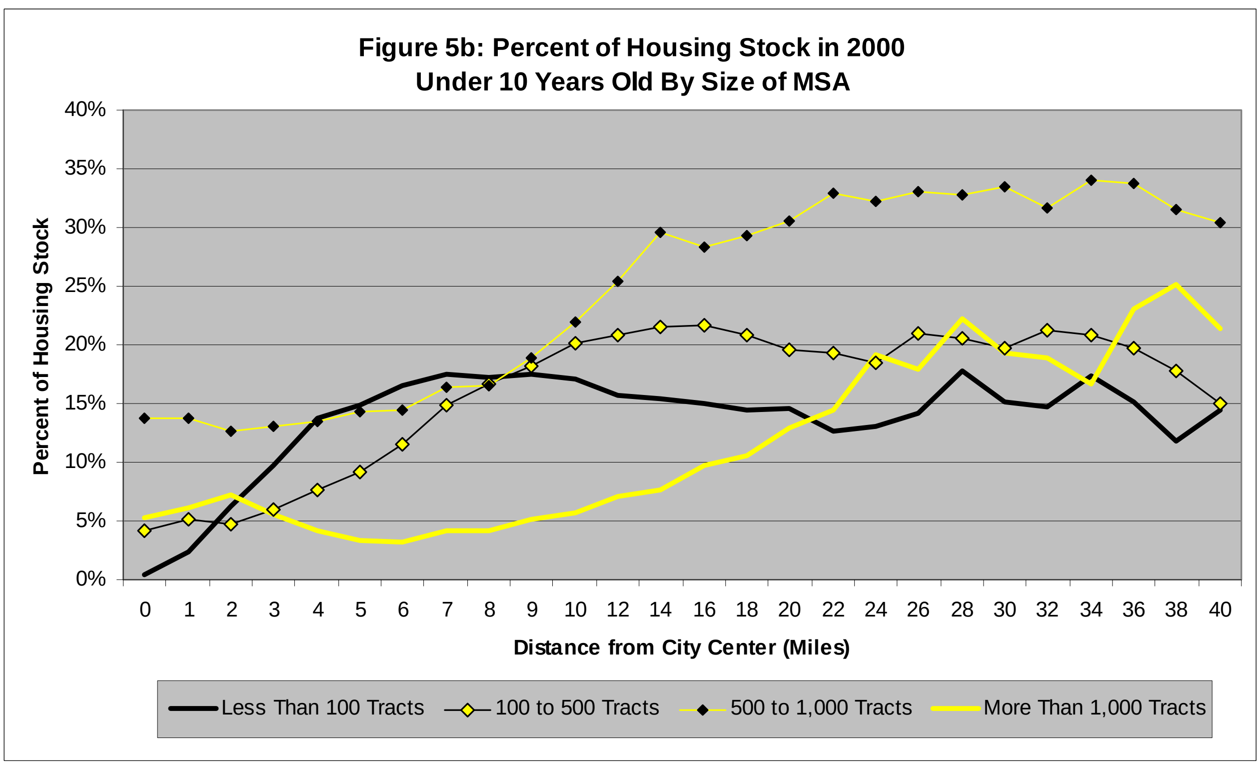
Figure 5b: Percent of Housing Stock in 2000
Under 10 Years Old By Size of MSA
40%
35%
30%
25%
20%
15%
10%
5%
0%
Percent of Housing Stock
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
Less Than 100 Tracts
100 to 500 Tracts
500 to 1,000 Tracts
More Than 1,000 Tracts

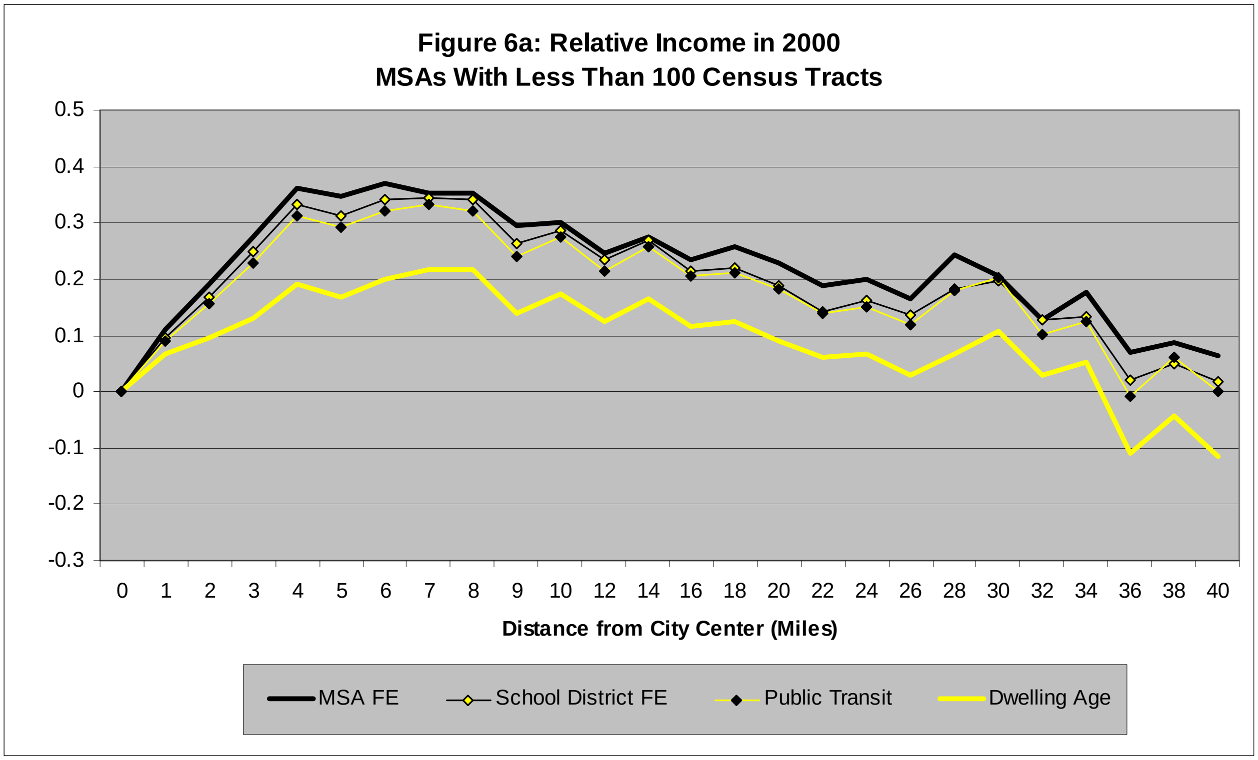

Figure 6a: Relative Income in 2000
MSAs With Less Than 100 Census Tracts
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
MSA FE
School District FE
Public Transit
Dwelling Age

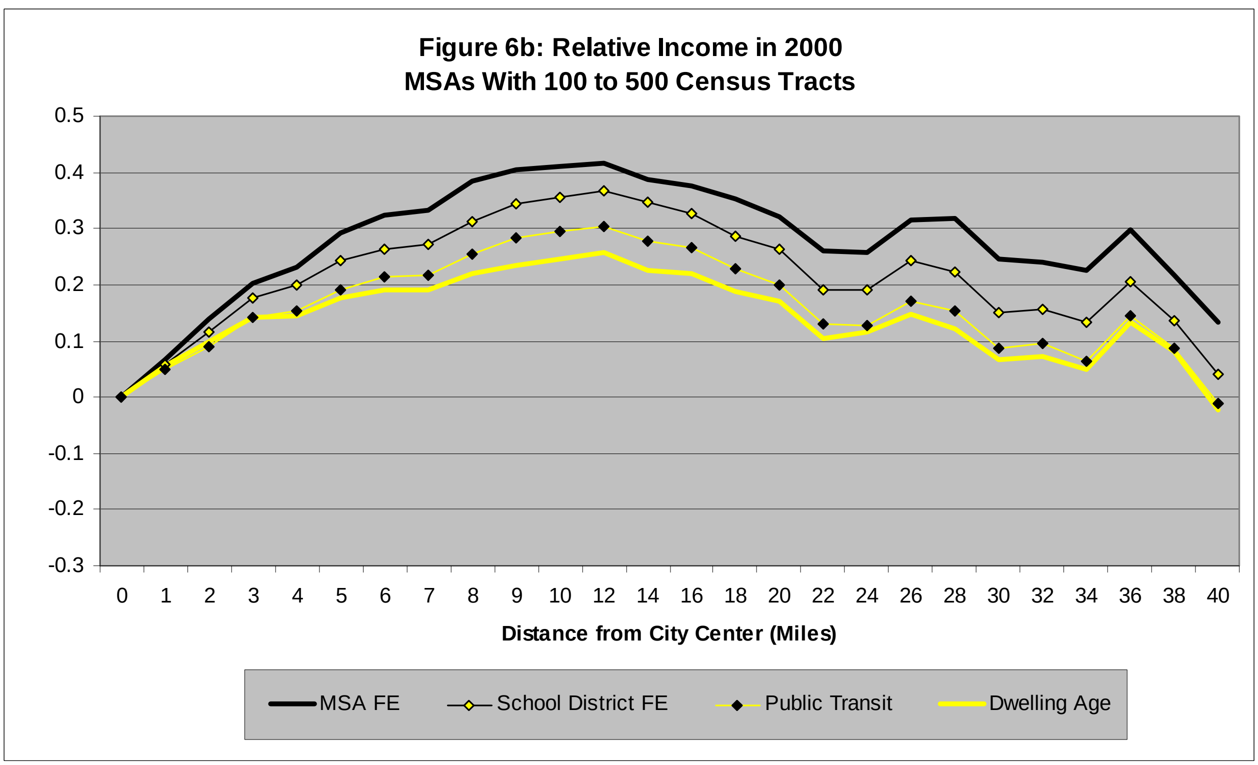

Figure 6b: Relative Income in 2000
MSAs With 100 to 500 Census Tracts
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
MSA FE
School District FE
Public Transit
Dwelling Age

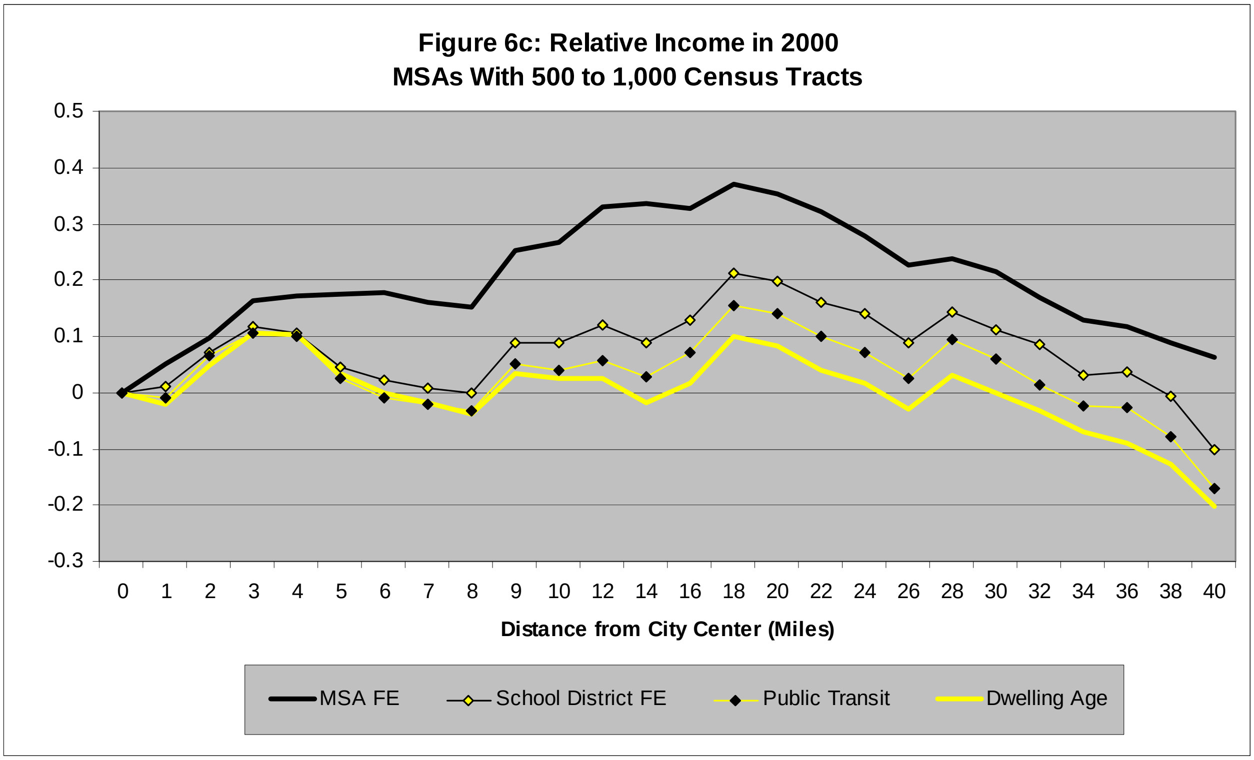
Figure 6c: Relative Income in 2000
MSAs With 500 to 1,000 Census Tracts
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
MSA FE
School District FE
Public Transit
Dwelling Age

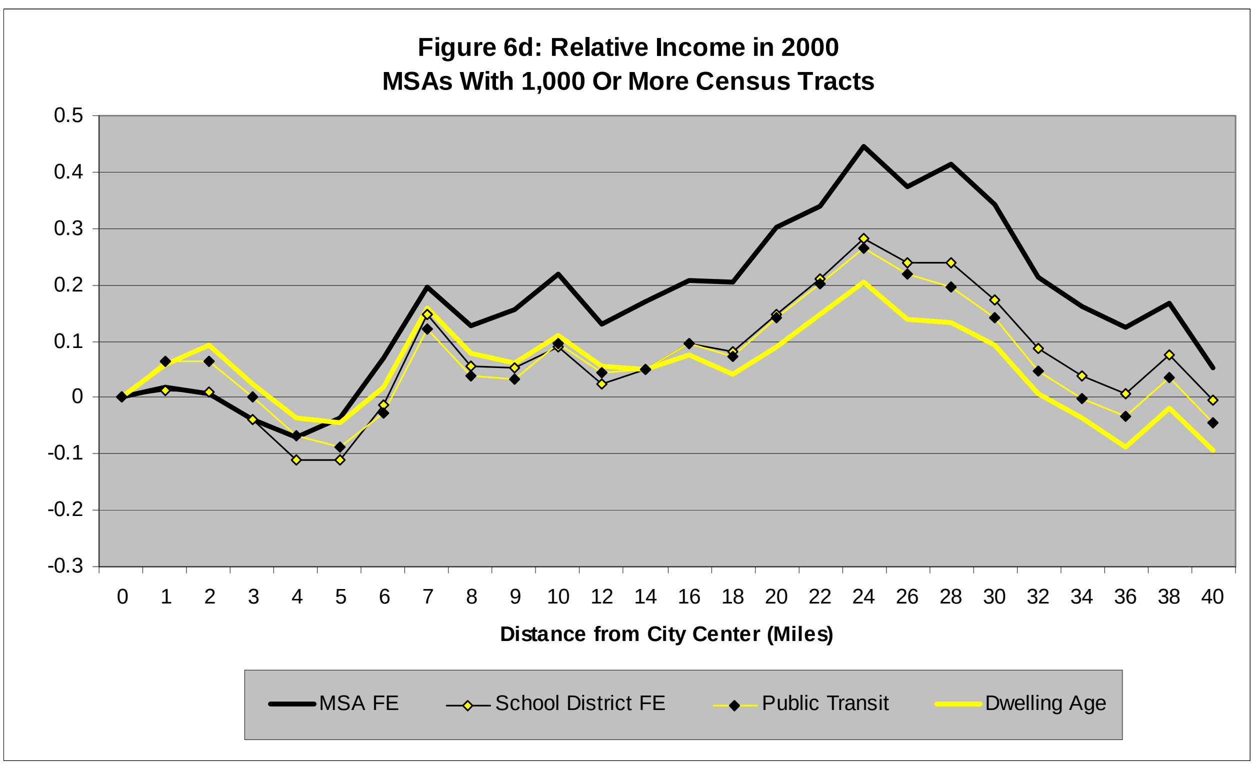
Figure 6d: Relative Income in 2000
MSAs With 1,000 Or More Census Tracts
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
-0.3
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
MSA FE
School District FE
Public Transit
Dwelling Age

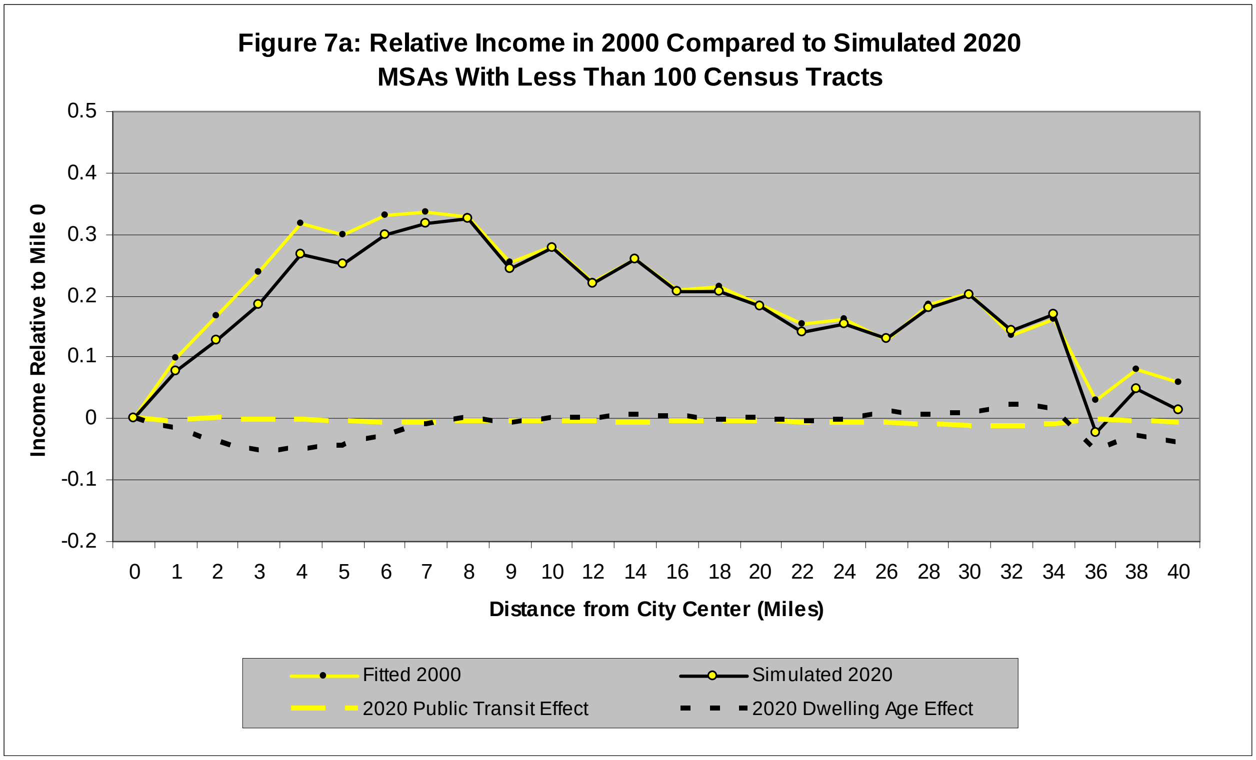
Figure 7a: Relative Income in 2000 Compared to Simulated 2020
MSAs With Less Than 100 Census Tracts
Income Relative to Mile 0
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
Fitted 2000
Simulated 2020
2020 Public Transit Effect
2020 Dwelling Age Effect

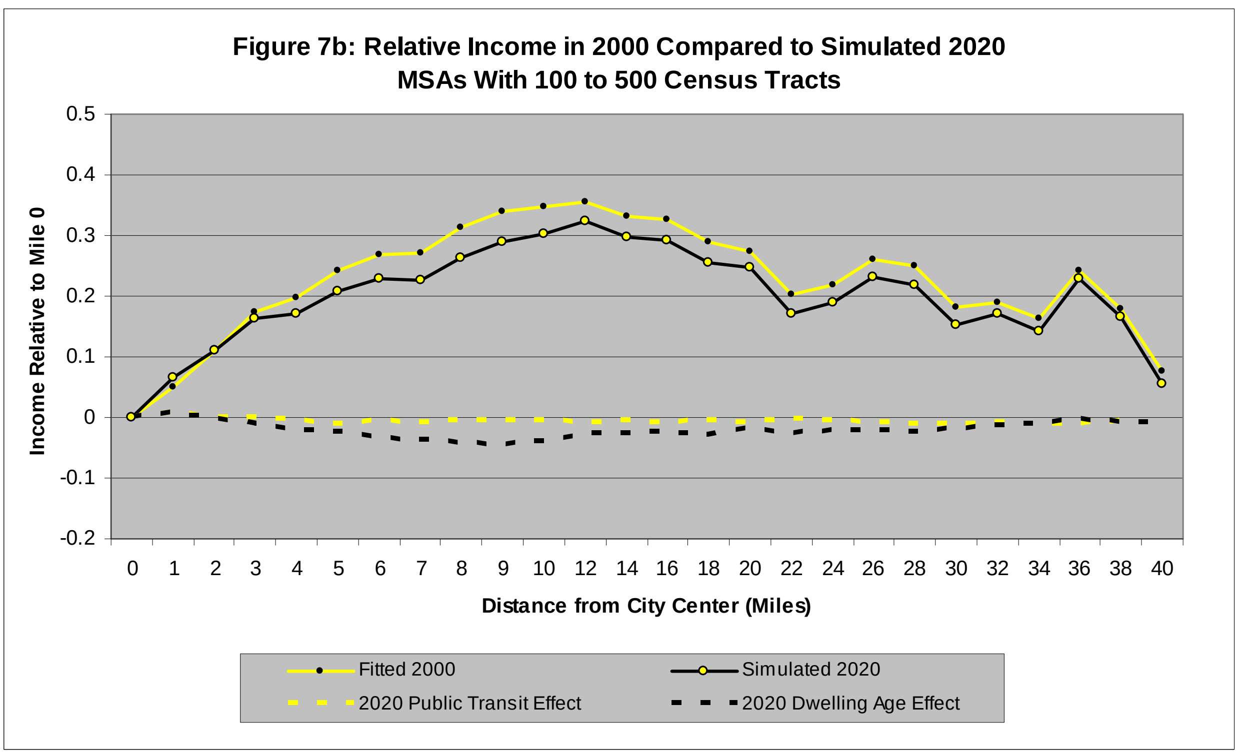
Figure 7b: Relative Income in 2000 Compared to Simulated 2020
MSAs With 100 to 500 Census Tracts
Income Relative to Mile 0
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
Fitted 2000
Simulated 2020
2020 Public Transit Effect
2020 Dwelling Age Effect

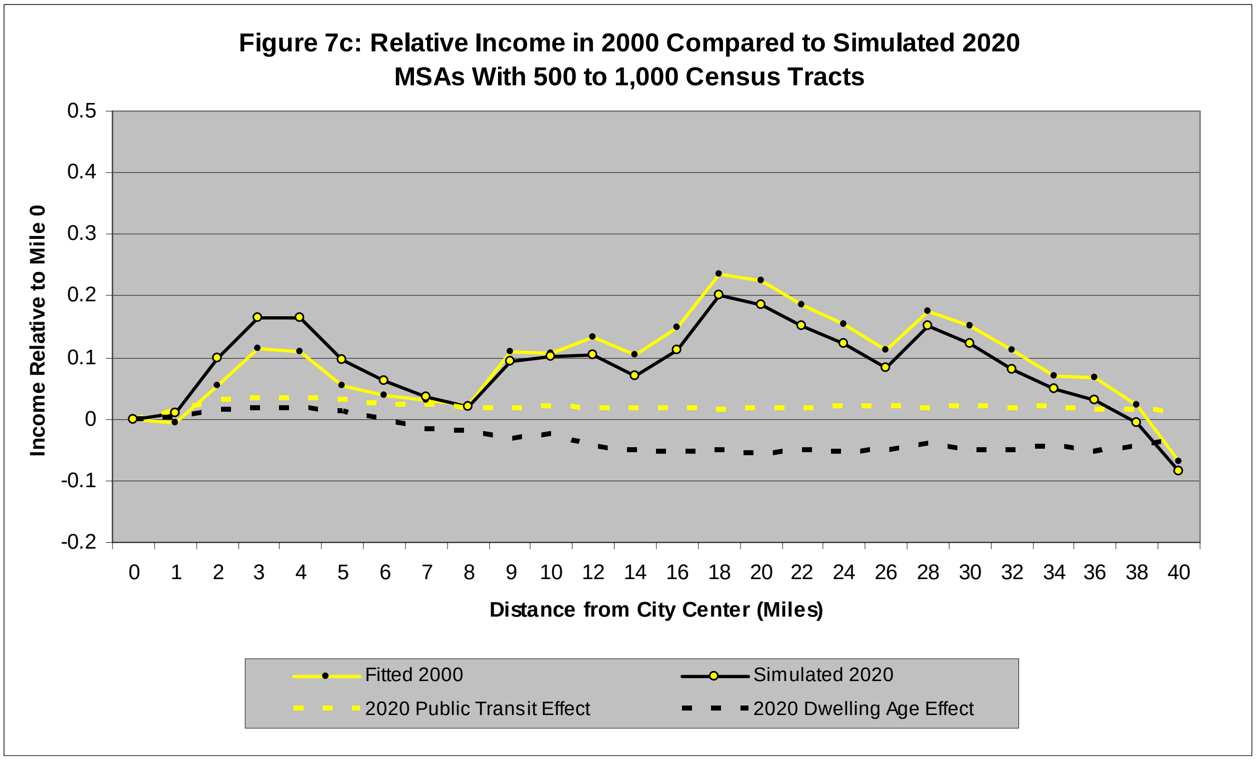
Figure 7c: Relative Income in 2000 Compared to Simulated 2020
MSAs With 500 to 1,000 Census Tracts
Income Relative to Mile 0
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
Fitted 2000
Simulated 2020
2020 Public Transit Effect
2020 Dwelling Age Effect

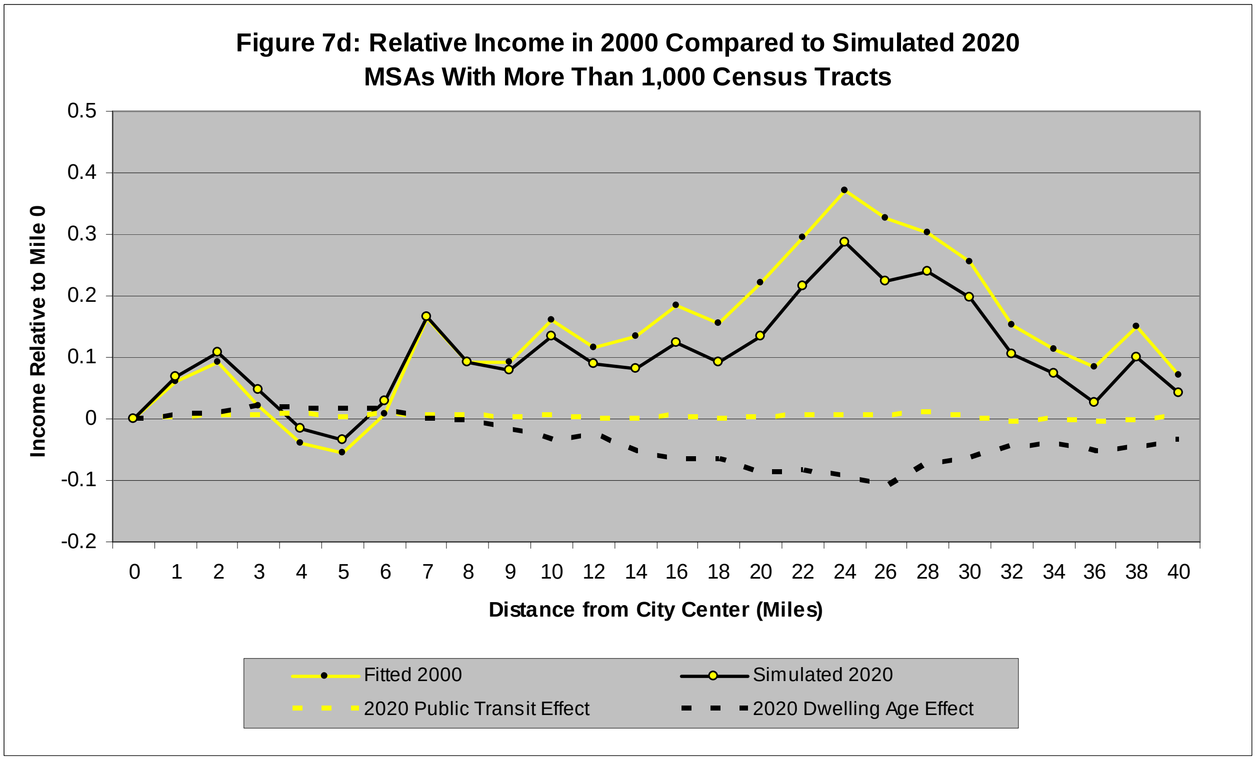
Figure 7d: Relative Income in 2000 Compared to Simulated 2020
MSAs With More Than 1,000 Census Tracts
Income Relative to Mile 0
0.5
0.4
0.3
0.2
0.1
0
-0.1
-0.2
0 1 2 3 4 5 6 7 8 9 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40
Distance from City Center (Miles)
Fitted 2000
Simulated 2020
2020 Public Transit Effect
2020 Dwelling Age Effect

Appendix A: Proofs and Additional Derivations

Proof of Proposition 2

Suppose the city has two dwelling generations at time $\hat{t}$, and consider the location $\tilde{x}_1$ in the first-generation area where rich and poor bid rents are equal. The dwelling age at this location, denoted $\tilde{a}_1$, is the first-generation critical age. Next, find the location, denoted $\breve{x} < \tilde{x}_1$, in the second-generation area where age equals $\tilde{a}_1$. Repeating the steps leading to (7) for the present case yields

$$R_r(\breve{x}) - R_p(\breve{x}) = R_r(\tilde{x}_1) - R_p(\tilde{x}_1) + (\hat{y}_r - \hat{y}_p)\tau(\tilde{x}_1 - \breve{x}) = (\hat{y}_r - \hat{y}_p)\tau(\tilde{x}_1 - \breve{x}) > 0, \quad \text{(A1)}$$

which shows that the rich outbid the poor for dwellings at $\breve{x}$. As a result, the location where second-generation bids are equal, denoted $\tilde{x}_2$, satisfies $\tilde{x}_2 < \breve{x}$, and the dwelling age $\tilde{a}_2$ at that location (the second-generation critical age), satisfies $\tilde{a}_2 > \tilde{a}_1$. The argument clearly generalizes to an arbitrary number of generations, establishing the Proposition.

Derivation of $T(x)$ and simulation parameter values

To make the analysis tractable, preferences over c and age a are assumed to be linear, with utility given by $c - \alpha_r a$ for the rich and $c - \alpha_p a$ for the poor.[26] Let $u_r(t)$ and $u_p(t)$ denote the time paths of utility for the two groups. These time paths are assumed to be exogenous, reflecting the assumption of an open-city model. Then, age-conditional bid-rent functions for the two groups are given by $S_r(x,t,a) = y_r(t)(1 - \tau x) - \alpha_r a - u_r(t)$ and $S_p(x,t,a) = y_p(t)(1 - \tau x) - \alpha_p a - u_p(t)$. The unconditional bid-rent function $R(x,t)$ comes from substituting $a(x,t)$, a function that is yet to be determined, into these age-conditional functions.

[26]Two aspects of this formulation deserve note. First, preferences are written directly in terms of age rather than housing services h. Second, preferences differ between the groups, with the age parameters $\alpha_p > 0$ and $\alpha_r > \alpha_p$

The next assumption is that income and utility increase linearly over time at common rates, with t = 0 intercepts given by $\bar{y}_r, \bar{y}_p, \bar{u}_r$, and $\bar{u}_p$. Substituting into S_p and S_r above and setting the two bid-rents equal yields the following expression for the critical age at location x:

$$\tilde{a}(x) = \frac{(\bar{y}_r - \bar{y}_p)(1-\tau x) - (\bar{u}_r - \bar{u}_p)}{\alpha_r - \alpha_p}. \qquad \text{(A2)}$$

Note that since the critical-age solution does not depend on t, the age at which a dwelling switches from rich to poor occupancy is independent of when it was built, depending only on location.[27] The critical age rises as x falls, validating Proposition 2 for this special case.

Using the above information, the present value of revenue as of the construction date from a single dwelling built at time t at location x is given by

$$\int_0^{\tilde{a}(x)} S_r(x, a+t, a) e^{-\rho a} da + \int_{\tilde{a}(x)}^{L} S_p(x, a+t, a) e^{-\rho a} da, \qquad \text{(A3)}$$

where ρ is the discount rate. Note in (A3) that calendar time is represented by the construction date t plus age a. Evaluating the integrals in (A3) using (A2) and the bid-rent formulas yields an expression equal to $g(x) + f(x)t$, where g and f are complicated functions, with $f > 0$.[28]

For land first developed at time T, subsequent redevelopment dates are given by $t = T + iL$, for i = 1,2,3,… Using (A3), the present value of revenue as of date T from the entire sequence of dwellings is then given by

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different for the rich and poor. This setup thus overturns the common-preference assumption in the general model above, but it does generate the key income-related difference in MRS's that appears in (3) and (4).

[27]The reason is that the linear income and utility time trends cancel in the calculation.

[28]The function $f(x)$ is given by $\left[\beta(1-\tau x) - \gamma\right]\left(1 - e^{-\rho L}\right)/\rho$, where β and γ are the income and utility time trends respectively. It is assumed that $\beta(1-\tau x) - \gamma > 0$ holds over the relevant range of x values so that f is positive. In addition,

$$g(x) = \left[1-\tau x)\bar{y}_r - \bar{u}_r\right]\left(1 - e^{-\rho\tilde{a}(x)}\right)/\rho + \left[1-\tau x)\bar{y}_p - \bar{u}_p\right]\left(e^{-\rho\tilde{a}(x)} - e^{-\rho L}\right)/\rho - \alpha_r\left[1 - e^{-\rho\tilde{a}(x)} - \rho\tilde{a}(x)e^{-\rho\tilde{a}(x)}\right]/\rho^2$$
$$- \alpha_p\left[\rho\tilde{a}(x)e^{-\rho\tilde{a}(x)} - \rho L e^{-\rho L} + e^{-\rho\tilde{a}(x)} - e^{-\rho L}\right]/\rho^2 + \left[\beta(1-\tau x)\right]\left(1 - e^{-\rho L} - \rho L e^{-\rho L}\right)/\rho^2$$

$$H(T) \equiv \sum_{i=0}^{\infty} \left[g(x) + f(x)(T + iL) \right] e^{-\rho iL}. \qquad \text{(A4)}$$

To avoid inessential complexity, it is assumed that a unit of land accommodates exactly one dwelling regardless of location (yielding constant density) and that construction costs are zero, so that (A4) directly gives the present value of profit per unit of developed land. Then, the development date $T(x)$ for land at location x is found by choosing T to maximize $\int_0^T r_a e^{-\rho t} dt + e^{-\rho T} H(T)$ where r_a is agricultural land rent. It is easily seen that the first-order condition for this problem requires r_a/ρ to equal (A4) with $1/\rho + T + iL$ in place of $T + iL$. Since the resulting expression is increasing in T and decreasing in x, it follows that $T'(x) > 0$.[29]

The simulation results in the Figures use the following parameter values: $L = 75$, $\rho = .05$, $\tau = .004$, $\bar{y}_r = 100$, $\bar{y}_p = 70$, $\bar{u}_r = 50$, $\bar{u}_p = 30$, $\alpha_r = 1.3$, $\alpha_p = 1.0$, $\beta = .04$, $\gamma = .03$ (the latter parameters are the income and utility time trends). In addition, r_a is set at a value that yields $T(0) = 0$, so that the city starts as a point at time zero.

As usual in an open-city model, the exogenous utility paths generate endogenous time paths for the populations of both groups, which are not presented here. Note, however, that the simulation can also be viewed as portraying a closed city whose population evolves exogenously along the given paths, with the now-endogenous utility paths equal, by construction, to the ones specified in the simulation.

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[29]The fact that the modified (A4) is decreasing in x can be seen by differentiating (A3) and recalling that the functions S_r and S_p are decreasing in x.