Demography
DOI 10.1007/s13524-012-0137-5

Population Trends as a Counterweight to Central City Decline, 1950–2000

Leah Boustan · Allison Shertzer

Abstract The share of metropolitan residents living in central cities declined dramatically from 1950 to 2000. We argue that cities would have lost even further ground if not for demographic trends such as renewed immigration, delayed childbearing, and a decline in the share of households headed by veterans. We provide causal estimates of the effect of children on residential location using the birth of twins. The effect of veteran status is identified from a discontinuity in the probability of military service during and after the mass mobilization for World War II. Our results suggest that these changes in demographic composition were strong enough to bolster city population but not to fully counteract socioeconomic factors favoring suburban growth.

Keywords Central cities · Suburbanization · Population growth · Demographic factors

Introduction

The share of the metropolitan population in the United States living in a central city fell from 58 % in 1950 to 36 % in 2000. This suburbanization intensified residential segregation by race and income, hastened the contraction of the urban tax base, and augmented disparities in access to education and other locally provided public services (Baumol 1967; Benabou 1996; Fischer et al. 2004).

Many economic, political, and sociological trends contributed to the rapid growth of the suburbs, including rising real incomes among American households after World War II (Margo 1992); road-building programs that reduced the time cost of commuting from bedroom communities to the central city (Baum-Snow 2007);

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Published online: 11 September 2012

federal subsidies for the purchase of single-family homes through the underwriting of mortgages and the mortgage interest deduction (Jackson 1985); the relocation of employment opportunities to the suburban ring (Boustan and Margo 2009); and changes in the perceived benefits of urban residence owing to racial diversity, income disparities between cities and suburbs, and heightened crime rates (Boustan 2007, 2010; Cullen and Levitt 1999).

Alongside forces favoring suburbanization, a series of countervailing population trends bolstered the size of central cities. In this article, we identify four such demographic shifts: the growing share of the metropolitan population living in a household with a foreign-born household head; the growing share with an African American household head; the declining share in a household headed by a veteran of the Armed Forces; and the declining share of households containing a child under the age of 18. We also consider the life-cycle mobility of the large baby boom cohort from city to suburb (and back again) but find that it did not have a quantitatively meaningful effect on residential patterns.[1]

Central to our argument is the claim that demographic characteristics affect residential location.[2] However, residential location may also influence characteristics such as family size and veteran status. We employ instrumental variables to identify the causal effect of having an additional child or serving in the military on place of residence. In particular, we instrument for household size with the occurrence of twins on either the first or the second birth (Angrist and Evans 1998). We identify the effect of military service by comparing cohorts who came of age during and just after mass mobilization for World War II (Bound and Turner 2002; Fetter 2010; Page 2008).

We use our estimates to conduct a series of demographic counterfactuals. Overall, we find that, absent these population trends, the share of the metropolitan population living in the central city would have declined by an additional 10 % to 32 % from 1960 to 2000. Demographic changes were not strong enough to overcome the powerful economic and social forces favoring suburbanization.

Residential Mobility and the Decline of Central City Population

Trends in City and Suburban Population, 1940–2000

Panel a of Fig. 1 documents trends in city and suburban growth from 1940 to 2000 for the 103 metropolitan areas anchored by the largest central cities in 1940.[3] Over the

[1] A series of popular essays have pointed to the effect of demographic shifts on city growth (Ehrenhalt 2008; Leinberger 2008). To our knowledge, this is the first scholarly article to investigate this connection in detail.

[2] The relationship between demographic characteristics and residential location derive from a complex interaction between household preferences and institutional and social constraints. For example, veterans bought new homes in the suburbs not only because of their own preference for suburban residence but also because new housing construction after World War II was disproportionately located in the suburban ring. In documenting these relationships, we do not aim to distinguish between these demand-side and supply-side mechanisms.

[3] See Rappaport (2003) for a discussion of similar trends. Panel a of Fig. 1 includes all metropolitan areas anchored by a city that had at least 50,000 residents in either 1940 or 1970. The core analysis relies instead on metropolitan areas whose residents can be identified by location (city or suburb) in the census microdata in each year.

second half of the twentieth century, the share of metropolitan residents living in the central city fell from 58 % to 36 %.[4] With the exception of the 1970s, cities experienced positive population growth in each decade.[5] However, the suburban population grew at a substantially higher rate throughout the period, leading to a steady decline in the share of the metropolitan population living in central cities. The difference between city and suburban growth rates in each decade reflects net outflows from cities to suburbs as well as differences in the rates of in-migration and natural increase between cities and suburbs.

The metropolitan areas represented in Panel a of Fig. 1 are anchored by a mixture of growing and declining cities. Panels b and c display separate patterns of growth by city type. Despite differing *levels* of growth over this period, the *time pattern* of city and suburban growth is very similar across expanding and declining cities. The fastest rates of city growth were posted in the 1940s, when expanding cities grew by nearly 30 % and declining cities experienced their last decade of positive growth. The low point of city growth in both categories occurred in the 1970s. In the 1980s and 1990s, expanding cities again experienced positive growth, and the rate of population loss slowed in declining cities. Overall, the share of the metropolitan population living in the central city declined at a similar pace in both cities types.

Related Literature

This article contributes to two related literatures—one on urban decline and another on residential mobility. Studies in both areas emphasize the roles of race, nativity, and household structure in the determination of residential location.

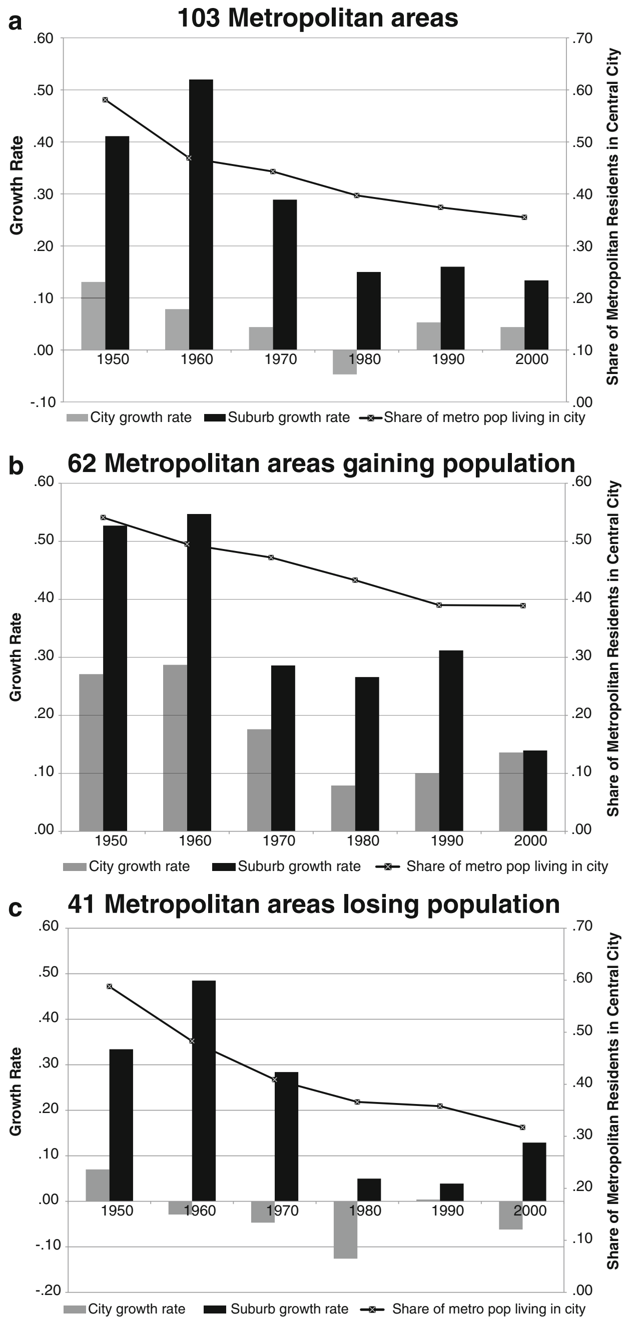
An extensive body of work documents that African Americans and the foreign-born are more likely than native-born, non-Hispanic whites to live in central cities (Massey and Denton 1993; Portes and Rumbaut 2001). Blacks moved in large numbers from the rural South to metropolitan areas from 1940 to 1970 (Boustan 2010; Gregory 2005). Over 80 % of these black migrants settled in central cities. Although black suburbanization began in earnest in the 1970s, sizable gaps in the residential locations of blacks and whites remain (Frey 1985; Schneider and Phelan 1993); 61.6 % of the black metropolitan population lived in central cities in 2000, compared with 26.1 % of whites.

Since the passage of the Immigration and Nationality Act of 1965, a large inflow of immigrants has settled in central cities (Frey 2005; Martin and Midgley 2003; Singer 2004). However, unlike European immigrants of the early twentieth century, new immigrant groups are increasingly "bypassing central cities and settling directly in suburbs" (Alba and Logan 1991:432). Despite this trend, immigrants from every sending country are still more likely than native-born whites to live in the central city (Alba et al. 1999).

A portion of these racial and ethnic differences can be explained by group disparities in socioeconomic status. In general, poor households are more likely to

[4] The share of the total population living in a central city declined only from 31 % in 1950 to 25 % in 2000 because the metropolitan shift to the suburbs was partially offset by rural-to-urban migration.

[5] The growth of central cities is partly driven by the expansion of city land area via annexation. In 1940, the average city in this sample was 48 square miles; by 2000, it had grown to 117 square miles.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103 Metropolitan areas
Growth Rate
Share of Metropolitan Residents in Central City
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City growth rate
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62 Metropolitan areas gaining population
Growth Rate
Share of Metropolitan Residents in Central City
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Growth Rate
Share of Metropolitan Residents in Central City
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City growth rate
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Share of metro pop living in city

◀**Fig. 1** City and suburban population growth by decade, 1940–2000, for 103 metropolitan areas (panel a) 62 metropolitan areas whose city gained population between 1940 and 2000 (panel b), and 41 metropolitan areas whose city lost population between 1940 and 2000 (panel c). Values refer to the decade ending in the census year on the *x*-axis. All metropolitan areas are anchored by a city that had at least 50,000 residents in either 1940 or 1970. City and county population are taken from the city and county data books. The 1970 county definitions of metropolitan areas are applied in all years. Suburban population is computed as the total metropolitan area population minus the city population

live in cities (Glaeser et al. 2008). However, notable differences in residential location by race and nativity remain even after controlling for income and education. This residual gap can be explained partly by the historical processes by which immigrant enclaves and majority black neighborhoods developed within central cities. Even today, some blacks and immigrants self-select into these areas to take advantage of familial or social networks or to enjoy community-specific amenities (Ihlanfeldt and Scafidi 2002; Thernstrom and Thernstrom 1997). In addition, African Americans and the foreign-born continue to face barriers that preclude suburban residence, including limited access to mortgage finance (Berkovec et al. 1996; Munnell et al. 1996; Ondrich et al. 1999).

Household structure is another important determinant of residential location. Married couples are more likely than other household types to live in the suburbs or to move to the suburbs in a given period conditional on living in a central city (Alba and Logan 1991; Frey and Kobrin 1982; South and Crowder 1997). The preference among married couples for suburban living is likely related to the association between marriage and childbearing. A large majority of married couples either currently live with children, have lived with children in the past, or are planning for children in the future. Therefore, married couples may place a higher premium on the larger lot sizes and the bundle of public goods, including higher quality public schools, available in the suburbs.[6]

At mid-century, veterans of the Second World War had access to housing benefits that encouraged homeownership and relocation to the suburbs. The Servicemen's Readjustment Act of 1944, commonly known as the G.I. Bill, included a mortgage program that allowed veterans to purchase a home with little or no down payment (Fetter 2010). The Veterans' Administration assisted 2.1 million veterans in purchasing homes between 1946 and 1950 alone, the majority of which were located in suburban areas (Bennett 1996:24). The civilian market for credit also expanded during this period, facilitated by the Federal Housing Administration (FHA).[7] Despite the expansion of credit in the civilian market, Vigdor (2006) found that eligible veterans were still seven percentage points more likely than nonveterans to own a home in 1970.

Journalists have speculated that the aging of the baby boom generation will lead retired couples to return to cities. Demographers, however, have been more skeptical (Frey 1993; Nelson 1988). Frey (2007), for example, argued that seniors are more

[6] The presence of higher quality public goods and more affluent neighborhoods in the suburbs is not an inherent feature of cities in developed economies. Indeed, many European cities are organized differently, with the most desirable neighborhoods located in the city center. For a U.S.-European comparison, see Brueckner et al. (1999).

[7] The FHA began insuring mortgages initiated by private lenders in the mid-1930s. As a result, mortgage rates fell from 6 % to 8 % in the 1920s to 2 % to 3 % in the 1940s (Jackson 1985:205).

likely to "age in place." Mobility rates among the elderly are very low; less than 5 % of Americans older than 65 move in a given year, compared with nearly 30 % of individuals in their early 20s. As a result, there is little increase in the probability of living in the central city after retirement. Thus, while we investigate the effect of the aging of the baby boom cohort on city population, the demographic literature leads us to believe that this force is unlikely to be particularly strong.

Demographic Correlates of Living in the Central City

Estimating Equation

Our goal is to examine whether demographic trends are quantitatively large enough to have bolstered the population of central cities despite the strong forces encouraging suburban growth. This section begins with the presentation of the demographic correlates of living in the central city. The analysis is based on individual records from the 1960–2000 censuses compiled by the Integrated Public Use Micro-data Series, or IPUMS (Ruggles et al. 2008).[8] Our sample includes all residents of metropolitan areas for whom place of residence (central city versus suburbs) is reported in the data. The fraction of the population that can be identified by place of residence shifts as the census privacy requirements change over time.[9] For robustness, we present results with a constant-geography sample.

Our dependent variable is an indicator equal to 1 if the respondent lives in a central city. We pool individual records from 1960 to 2000 and estimate Eq. 1 using a probit specification:

$$\mathbf{I}(= 1 \text{ if } \mathit{central} \text{ city})_{iact} = \alpha + \mathbf{X}_i\,'\Gamma + \nu_t + \mu_a + \theta_c + \varepsilon_{iact}. \tag{1}$$

The subscript i indexes individuals who are a years old in census year t and belong to birth cohort c. The regression includes fixed effects for census years (ν_t); individual years of age (μ_a); and five cohort intervals, each representing roughly 20 years of birth cohorts (θ_c).[10] $\mathbf{X}_i$ is a vector of characteristics for the household in which individual i resides.

In the baseline equation, $\mathbf{X}_i$ contains indicators for the race, nativity, and veteran status of the household head and a dummy variable for the presence of children in the household. We define a child as anyone who is 18 years of age or younger regardless of his or her relationship to the household head. In alternative specifications, we allow

[8] We do not include data from 1950 because, in that year, veteran status was asked only of individuals on the sample line.

[9] IPUMS does not report central city status if doing so would allow users to identify geographic areas with fewer than 100,000–250,000 residents. Depending on the year, we are able to identify observations from between 91 and 143 metropolitan areas. We analyze a consistent sample of 109 metropolitan areas in Table 2.

[10] The five cohort groups in the main specification were born in 1869–1910, 1911–1930, 1931–1950, 1951–1970, and after 1970 (the omitted category). We identify age, period, and cohort effects by constraining that cohort effect to be identical within these 20-year intervals. Results are robust to using finer cohort groups.

residential location to vary with the number of children in the household and add indicators for being married or being an "empty nester." Households are considered to be "empty nesters" if one member reports having had children but there are no children currently present.

To the best of our knowledge, our study is the first to estimate the age profile of city residence within birth cohorts over time. Vigdor (2006) and others reported age profiles of city residence constructed from single cross sections. These profiles likely overstate the probability that the elderly will "return" to the central city by conflating age and cohort effects. For example, individuals who were 70 years old in 1970 were born in 1900 and came of age before the diffusion of the automobile and the large-scale suburban growth of the post–World War II period. Therefore, the elderly in 1970 may have been more likely to live in central cities for both life-cycle and cohort-specific reasons.

Probit Results

Our estimating equation produces two sets of results: the age profile of city residence over the life cycle, and the relationship between the other demographic characteristics in the vector X and the probability of living in the central city. We report the age profile of city residence in Fig. 2, which graphs the average marginal effects by single years of age from Eq. 1 (plus the constant). The probability that a metropolitan resident lives in the central city peaks between the ages of 20 and 22. Many individuals then leave the central city in their late 20s and 30s. The lowest probability of city residence occurs at the age of 55, an age at which households are likely to have

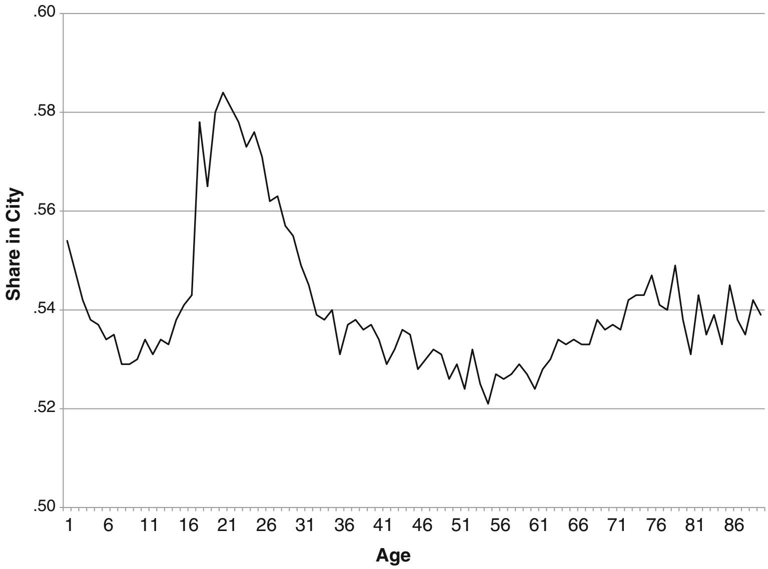

Fig. 2 Probability of living in city conditional on being in metropolitan area by age. We plot the constant plus the average marginal effects of the single years of age indicators in Eq. (1). The underlying regression equation also contains indicators for four birth cohorts, four census years, and controls for the presence of children in the household and the race, nativity, and veteran status of the household head. *Source:* IPUMS, 1960–2000

children and to be able to afford the larger homes available in the suburbs. After that point, individuals slowly return to the city.

Table 1 presents the average marginal effects relating the other demographic characteristics to the probability of living in the central city. Members of immigrant-headed households are 17 percentage points more likely than native-born whites to live in the central city. The excess probability of living in a city is even higher for individuals living with a black household head (37 percentage points). In contrast, members of households headed by a veteran of the U.S. Armed Forces are 4.6 percentage points *less* likely to live in a central city.[11] Consistent with recent work on immigrant locations, we find that the relationship between immigration status and residential location changes over time. In 1960, immigrant households, most of whose household heads were born in Europe, were only 14 percentage points more likely than the native-born to live in a central city. By 1990, the immigrant-native gap increased to 20 points, before declining again to 16 points in 2000 as immigrants began to suburbanize or to bypass the central city altogether.

In the first column, we estimate the effect of children on residential location with a single dummy variable for the presence of any child in the household. Households with at least one child are 7.6 percentage points less likely to live in the central city. The second column replaces the indicator variable with a linear measure of the number of children in the household. Each child appears to depress the likelihood of living in the central city by 1.2 percentage points. Together, these estimates suggest that the relationship between the presence of children and residential location is nonlinear and, in particular, that the first few children are most strongly associated with leaving the central city. To further explore this nonlinearity, the third column adds dummy variables for having exactly one child and for having two or more children in the household. Relative to households with no children, households with one child are 5.9 percentage points less likely to live in the central city, and households with two or more children are 8.7 percentage points less likely to live in the central city.[12]

Households may not instantaneously adjust their residential location decisions on the basis of current composition. Rather, some households may move to the suburbs in anticipation of having children, and some households that once contained children may remain in the suburbs even after the children leave home. We provide evidence consistent with this life-cycle perspective in columns 4 and 5. Column 4 uses two indicator variables to summarize household composition: one for the current presence of children and another for empty nesters. The omitted category is individuals living in households that have never (to date) included children. Compared with this omitted category, households with children present are 7.6 points and empty nesters are 1.2 points more likely to be suburban residents.[13]

[11] A portion of the relationship between veteran status and residential location is driven by an association between having served in the military and being married (see column 5).

[12] We experimented with adding a richer set of dummy variables and found no statistical difference between having two versus three children, three versus four children, and so on.

[13] The "empty nest" indicator is available from 1960 to 1990 only. In these years, the census added a question about children ever born for all women who were at least 15 years old.

Table 1 Demographic characteristics and the probability of living in a central city, 1960–2000 (dependent variable = 1 if live in central city)

RHS Variable	(1)	(2)	(3)	(4)	(5)	(6)
Any Children in Household	−0.076 (0.001)			−0.079 (0.001)	−0.042 (0.001)	−0.074 (0.001)
Number of Children		−0.012 (0.000)				
One Child			−0.059 (0.001)			
2+ Children			−0.087 (0.001)			
Empty Nester				−0.012 (0.001)	0.009 (0.001)	
Married					−0.114 (0.001)	
Head Is Foreign-born	0.170 (0.001)	0.169 (0.001)	0.171 (0.001)	0.172 (0.001)	0.179 (0.001)	0.234 (0.001)
Head Is Black	0.367 (0.001)	0.369 (0.001)	0.367 (0.001)	0.372 (0.001)	0.351 (0.001)	0.281 (0.001)
Head Is Veteran	−0.046 (0.001)	−0.048 (0.001)	−0.045 (0.001)	−0.045 (0.001)	−0.018 (0.001)	−0.008 (0.000)
Constant	0.564 (0.003)	0.519 (0.003)	0.506 (0.003)	0.567 (0.004)	0.629 (0.004)	0.391 (0.003)
N	5,742,225	5,742,225	5,742,225	4,536,450	4,536,450	8,406,585
Pseudo-R^2	.99	.99	.99	.99	.99	.99

Notes: Coefficients from probit estimation of Eq. (1), with standard errors in parentheses. Regressions also contains four birth cohort dummy variables, four census year dummy variables, dummy variables for single years of age between 1 and 90, and an indicator for being older than 90. Household members are assigned the race, nativity, and veteran status of the household head. Cox-Snell pseudo-R^2 statistics are reported in the last row. In columns 1–5, the sample contains all residents of metropolitan areas for whom place of residence (central city or suburb) is known. Columns 4 and 5 do not include data from the year 2000 because the variable "child ever born" used to construct the "empty nester" indicator is not available in that year. Column 6 adds all nonmetropolitan residents to the sample.

Column 5 proxies for the full life-cycle effect of having children by adding an indicator for being married, relative either to never having been married or to being divorced or widowed.[14] Married individuals are 11.4 percentage points more likely than singles of the same age to live in the suburbs.[15] Because marital status is highly correlated with the presence of children, the independent effects of both the current and prior presence of children in the household decline substantially.

In the final column, we expand the sample to include nonmetropolitan households. In this case, the coefficients can be interpreted as the effect of each demographic factor on the probability of living in the central city, relative to living either in a suburb or in a nonmetropolitan area. The largest change in this expanded sample is the effect of veteran status on residential location. Although veterans are 4 percentage points more likely to live in a *suburb* relative to a city, they are only 1 percentage point more likely to live in a *noncity* relative to a city. In other words, veterans are least likely to live in cities or nonmetropolitan areas and most likely to live in the suburbs.

Thus far, we have relied on the set of metropolitan areas for which residential locations (city versus suburb) is known in the census microdata. This geography presents two concerns. First, information on place of residence is available for a varying set of metropolitan areas in the microdata in each year. Second, the boundaries of each metropolitan area, which is composed of one or more contiguous counties, can expand over time as the Census Bureau adds peripheral counties to existing area definitions.[16] Table 2 presents results using two samples that impose consistent geography in the 1980 and 2000 census years.[17]

For comparison, the first column of Table 2 uses all metropolitan observations in 1980 and 2000. Columns 2 and 3 restrict the sample to the 109 metropolitan areas for which at least *one* underlying county is large enough to be separately identified in the microdata in both years. The second column includes all observations from these 109 metropolitan areas; in the third column, we restrict our attention to observations from counties that are identified in both years. The coefficients are qualitatively unchanged in both columns, suggesting that the results are not sensitive to the set of included metropolitan areas and that residence patterns in peripheral counties are similar to those in core suburban counties. If anything, demographic characteristics are slightly stronger predictors of city residence in these (larger) metropolitan areas, perhaps because they have a sharper distinction between city and suburb.

[14] We include divorced and widowed household heads in the control group because these marital transitions often lead to residential mobility. Results are similar if we instead compare the ever- and never-married.

[15] Married men have higher labor-market earnings than their single counterparts, which may allow married couples to afford a suburban residence (Ginther and Zavodny 2001; Korenman and Neumark 1991). When we control for household income, the effect on marital status declines by 20 % (results not shown). We interpret the results in column 5 as the total effect of marriage on residential location, including a potential earnings channel.

[16] For example, between 1980 and 1990, the number of counties included in the average metropolitan area increased from 4.6 to 7.5.

[17] We exclude 1960 and 1970 because of additional data restrictions in these years. In 1960, the microdata do not report metropolitan area of residence; and in 1970, either place of residence (central city versus suburb) or metropolitan area is known, but not both.

Table 2 Demographic correlates of living in the central city using constant city and metropolitan area samples, 1980 and 2000 (dependent variable = 1 if live in central city)

RHS Variable	All Metros (1)	Consistent Set of Metros (2)	Consistent Set of Metros With Consistent Boundaries (3)
Any Children in Household	−0.079	−0.084	−0.088
	(0.000)	(0.000)	(0.000)
Head Is Foreign-born	0.163	0.176	0.178
	(0.000)	(0.000)	(0.000)
Head Is Black	0.340	0.373	0.363
	(0.000)	(0.000)	(0.000)
Head Is Veteran	−0.055	−0.062	−0.063
	(0.000)	(0.000)	(0.000)
Constant	0.450	0.613	0.561
	(0.003)	(0.003)	(0.004)
N	11,938,826	8,830,474	8,393,237
Pseudo-R^2	.87	.92	.90

Notes: Coefficients from probit estimation of Eq. (1), with standard errors in parentheses. Regressions also contains four birth cohort dummy variables, one census year dummy variable, dummy variables for single years of age between 1 and 90, and an indicator for being older than 90. Household members are assigned the race, nativity, and veteran status of the household head. Cox-Snell pseudo-R^2 statistics are reported in the last row. In column 1, the sample contains all residents of metropolitan areas for whom place of residence (central city or suburb) is known. In column 2, the sample contains residents of the 109 metropolitan areas for which at least one underlying county is large enough to be separately identified in the microdata in both years. In the third column, the sample contains the same 109 metropolitan areas and imposes the 1980 metropolitan area county definitions in both 1980 and 2000.

Instrumental Variables: The Effect of Children and Veteran Status on City Residence

In the previous section, we estimated the effects of five demographic characteristics—age, race, nativity, veteran status, and the presence of children—on residential location. Ultimately, our goal is to use these estimates to infer how changes in the demographic composition of the metropolitan population have affected the size of central cities over the past half-century. However, before doing so, we must determine that the estimates indeed reveal the effect of characteristics on residential location and not the other way around. In so doing, we distinguish between predetermined characteristics, such as race and nativity, and mutable characteristics like veteran status and childbearing. For example, suburbanites may be encouraged by the child-friendly environment to have an additional child. Furthermore, both childbearing and veteran status may be correlated with other characteristics that are associated with living in the suburbs. As a result, we employ instrumental variables (IV) to estimate the causal effect of having an additional child or serving in the military on place of residence.

Veteran Status

According to our probit estimates, veterans are less likely than nonveterans to live in the central city. One explanation for this pattern is that veterans were offered generous housing benefits that provided the resources necessary to buy single-family housing in the suburbs. However, this relationship could also be generated by omitted variables that are correlated with veteran status. Veteran status is determined by a combination of individual initiative (whether or not to enlist) and military selection. Men who suffer from health ailments or who are cognitively impaired are less likely to serve in the military. At various points, the military also offered deferments to men who were enrolled in college, employed in a war industry, or working in the agricultural sector. Any of these factors may be correlated with later residential location.

Our goal is to estimate the direct effect of veteran status on residential location while minimizing these other confounding factors. To do so, we focus on the era of mass mobilization for World War II between 1940 and 1945, a period in which the probability of military service was strongly influenced by external events. Figure 3 reports the share of white, native-born men from the 1915–1934 birth cohorts who served in the Armed Forces. The probability of military service increased from 50 % for men born in 1915 to over 80 % for the men born between 1919 and the third quarter of 1927, who were 18 to 26 years old at the end of World War II. The probability of military service then declined from 83 % to 70 % because men born after the third quarter of 1927 were too young to participate in World War II (although many of them served in Korea).

With this history in mind, we compare the veteran status and residential location of adjacent birth cohorts who faced different probabilities of military service because of the timing of World War II. In this setting, we treat quarter of birth as an instrument for the probability of military service.[18] In particular, in our first-stage equation, the probability of military service is a function of a linear trend in quarter of birth and a dummy variable for being born before the fourth quarter of 1927. We estimate

$$\mathbf{I}(= 1 \text{ if veteran})_{it} = \alpha + \beta\mathbf{I}(= 1 \text{ if born before 4th Q 1927})_{it} + \gamma(\text{quarter of birth})_{it} + \nu_t + \varepsilon_{it}. \quad (2)$$

The validity of our instrument depends on the assumption that being born after the third quarter of 1927 affects residential location only via its influence on veteran status. The linear trend in quarter of birth accounts for other factors that may have increased residential location over time, such as rising real incomes. The sample is restricted to white, male, native-born heads of household from the birth cohorts of 1919 through 1932. As a robustness exercise, we consider different birth years as starting and ending points of the comparison window.

Table 3 presents the coefficients from our first- and second-stage equations, which are estimated using linear probability models.[19] As is clear in Fig. 3, men born before the fourth quarter of 1927 were 13 percentage points more likely than men born after that period to have served in the Armed Forces. In this restricted sample, our OLS estimates suggest that being a veteran increases the probability of living in the suburbs by 3 percentage points (a slightly smaller effect than in the full sample in Table 1). When we instrument for veteran status, the relationship between military

[18] Bound and Turner (2002) and Page (2008) used a similar approach to study the effect of the G.I. Bill on educational attainment. Because quarter of birth is available in the census only from 1960 to 1980, we focus on these years in our IV estimation.

[19] We also estimated this system using a seemingly unrelated bivariate probit model. The IV probit results are quite similar to the two-stage least squares estimates (coefficient = −0.076, SE = 0.030). Results are also similar when we shorten or lengthen the treatment window.

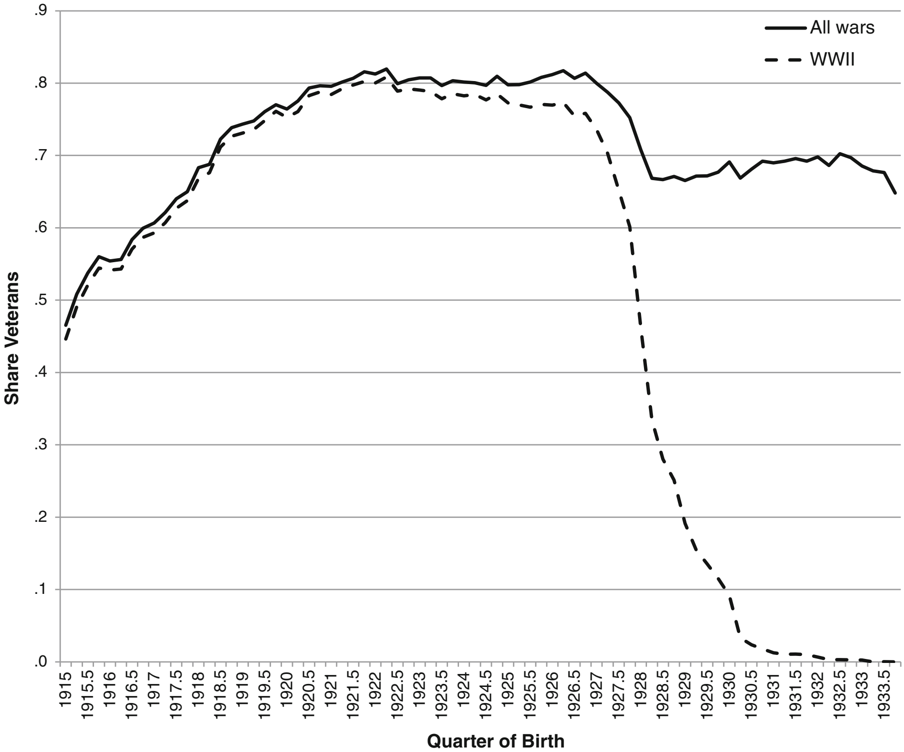

Fig. 3 Share of white men serving in armed forces by year and quarter of birth. Sample includes all white, native-born men from the 1960–1980 1 % IPUMS samples

service and suburban residence increases to 6.7 percentage points. Given the size of the standard errors, however, we cannot reject that the OLS and IV coefficients are the same.

Table 3 IV estimates of the effect of veteran status on place of residence, 1960–1980

	Birth Cohorts: 1919–1929		Birth Cohorts: 1919–1932	
RHS Variable	OLS	IV	OLS	IV
A. First Stage: Dependent Variable = 1 if Veteran				
Born between 1919 and 1927		0.132 (0.003)		0.130 (0.003)
B. Second Stage: Dependent Variable = 1 if Live in City				
Veteran	−0.029 (0.003)	−0.067 (0.029)	−0.029 (0.002)	−0.069 (0.028)
N	188,734	188,734	237,968	237,968
R^2	.02	.02	.03	.03

Notes: The sample is restricted to white, native-born male heads of household for whom place of residence (central city or suburb) is known. Regressions include a linear trend in quarters of birth, an indicator for children present in household, and dummy variables for 1970 and 1980 census years.

The fact that the IV estimates are larger (in absolute value) than their OLS counterparts implies that, at least during World War II, veterans were selected on attributes that were *positively correlated* with living in the central city. This pattern is consistent with draft exemptions for farmers and agricultural workers. Acemoglu et al. (2004) showed that mobilization rates of prime-age men during World War II were lowest in the agricultural states of the Great Plains and the cotton South. Furthermore, in the 1950 census, young veterans of World War II were far less likely than nonveterans to live on a farm (6.7 % versus 15.7 %).

The Presence of Children

In our probit estimation, we find that households with children are less likely to live in the central city. One explanation for this result is that having children increases the demand for the larger housing units, presence of open space, and higher quality public schools available in the suburbs. However, this finding could be contaminated by either omitted variables bias or reverse causality. During this period, rich households had fewer children and were more likely to live in the suburbs, which may bias downward the relationship between having children and living in the suburbs. On the other hand, suburban residence could directly influence a household's preferences for optimal family size. The attitudes of friends and neighbors in the suburbs may encourage households to have an additional child. In this case, our estimate would overstate the effect of having children on moving to the suburbs.

We use the birth of twins as an instrument for the number of children in a household. Angrist and Evans (1998) argued that, conditional on the age and race of the mother, twinning is an exogenous event. We focus on the period 1960–1980 because of the availability of quarter-of-birth data used to identify twin pairs in these years. In addition, these years pre-date the development of infertility treatments that have increased the probability of twinning for (the nonrandom set of) mothers who seek medical intervention. A large literature has used twinning to study the effect of family size on women's and children's outcomes (Angrist and Evans 1998; Black et al. 2005; Bronars and Grogger 1994; Rosenzweig and Wolpin 1980a, b).

We define two children in the same household with the same quarter and year of birth as a pair of twins. In households with at least one birth, 0.5 % have twins on the first birth. Our first-stage equations relate the presence of twins to various measures of the number of children in the household. For example, for households with at least one birth, we estimate

$$\text{Number of children}_{it} = \alpha + \beta \mathbf{I}(= 1 \text{ if twin on first birth})_{it} + \nu_t + \varepsilon_{it}. \qquad (3)$$

Alongside the standard controls included in Eq. 1, we also control for the race and age of the mother. One limitation of this approach is that households must have at least one birth event in order to be at risk for having twins. Table 1 demonstrates that the shift from no children to one child is an important determinant of residential location; however, twinning cannot be used as an instrument for the presence of the first child in a household.

Table 4 presents the coefficients from our first- and second-stage equations, estimated using linear probability models.[20] The raw data indicate that among households with at least one birth, those with a singleton on the first birth have an average of 2.58 children, whereas those with a twin on the first birth have 3.34 children. Accordingly, we estimate that having a set of twins on the first birth event increases completed household size by 0.7 children (column 2). Much of the difference in completed family size arises from the (obvious) fact that the vast majority of households with twins on the first birth have at least two children, whereas only 73.4 % of households with a singleton first birth have an additional child. Consistent with this figure, we estimate that having twins increases the probability of having two children by 25.6 percentage points (column 4).

In the restricted sample of households with at least one child, the OLS estimate implies that each additional child reduces the likelihood of living in the central city by 0.5 percentage points. The effect of family size on residential location more than doubles when we instrument for the number of children with the occurrence of twins on first birth. The larger IV estimates suggest that households with many children have unobserved characteristics that are otherwise *positively associated* with living in the central city; for example, large households may have a lower socioeconomic status. In this case, the OLS and IV estimates lie outside of each other's confidence intervals. The IV results, however, are derived from the subset of households with at least one child and may not be generalizable to the full population. We therefore conduct our counterfactual simulations using both the probit and IV estimates.

Counterfactual Effects of Demographic Composition on Central City Population

Demographic Characteristics of the Metropolitan Population, 1940–2000

Thus far, we have documented a relationship between demographic characteristics and residential location, which suggests that the size of cities is closely tied to population trends. Figure 4 displays trends in this set of demographic characteristics over time. With the immigration restrictions of the mid-twentieth century, the share of the metropolitan population living in an immigrant-headed household fell from 30 % in 1940 to 10 % in 1970. After the expansion of immigration quotas in 1965, this share returned to nearly 30 % by 2000. Given that the foreign-born are more likely to live in central cities, we expect this pattern to contribute to population growth in central cities from 1970 onward. As blacks from the rural South migrated to industrial cities, the share of the metropolitan population living with a black household head increased from 8 % in 1940 to 17 % in 2000. Again, this pattern would likely bolster city population.

After servicemen returned from World War II, the share of the metropolitan population living in a veteran-headed household spiked from less than 5 % in 1940 to nearly 50 % in 1960 and 1970. Since 1970, the veteran share has declined to just over 10 % by 2000. The share of households with a child present has also declined

[20] We also experimented with appropriate probit methods (bivariate probit or probit IV). However, the resulting estimates were either implausibly large or implausibly small. Therefore, we follow Angrist and Pischke (2009:204–05) in reporting the more robust two-stage least squares coefficients.

Table 4 IV estimates of the presence of children in the household on place of residence, 1960–1980

	Households With at Least 1 Birth				Households With at Least 2 Births			
	OLS	IV	OLS	IV	OLS	IV	OLS	IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. First stage: Dependent Variable = Number Children (or indicator)								
Twins on 1st (2nd) birth		0.705		0.256		0.924		0.408
		(0.012)		(0.004)		(0.010)		(0.004)
B. Second Stage: Dependent Variable = 1 if Live in City								
Number of children	−0.005	−0.019			0.000	0.004		
	(0.000)	(0.006)			0.000	(0.004)		
2+ children			−0.032	−0.052				
			(0.001)	(0.015)				
3+ children							−0.006	0.008
							(0.001)	(0.009)
R^2	.10	.10	.11	.10	.10	.10	.10	.10

Notes: N = 2,372,595 for one birth and 1,746,963 for two births. Twins are defined as two children in the same household with the same year and quarter of birth. Sample selection and specifications follow the notes to Table 1.

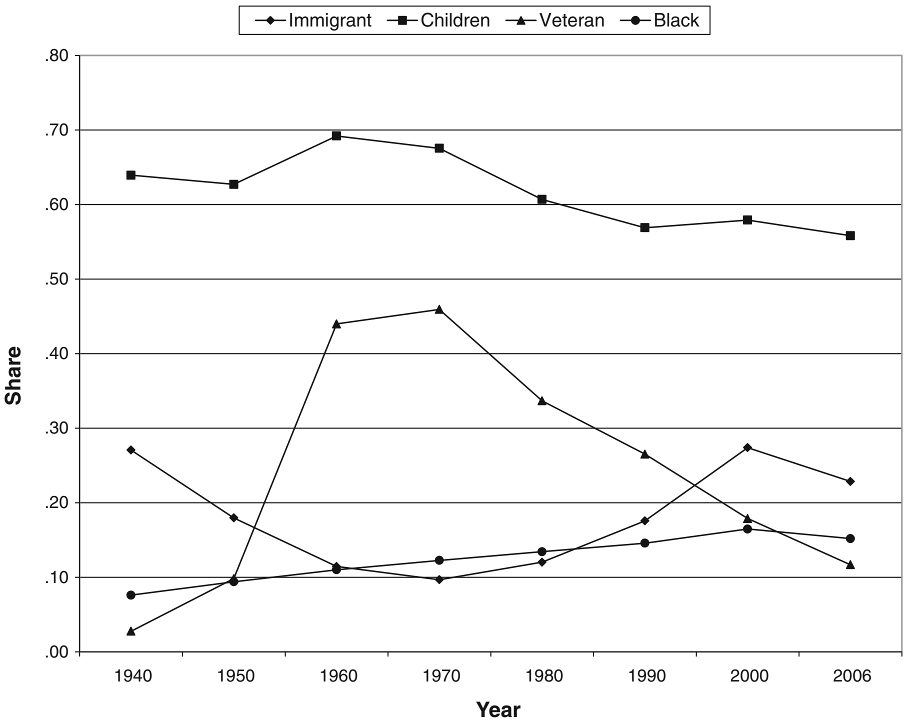

Fig. 4 Characteristics of the metropolitan population, 1940–2000: Race, nativity, veteran status and presence of children. Sample contains all metropolitan area residents for whom place of residence (central city versus suburb) is known in the given census year. Household members are assigned the characteristics of the household head. *Source:* IPUMS, 1940–2000

substantially since mid-century. While 70% of households had at least one child in 1960, this share declined to 58 % in 2000. Both trends favor the city.

The age structure of the population also shifted notably from 1940 to 2000 with the birth and aging of the large baby boom cohort (born 1946–1964). Yet we find that, because of the rapid swings in the age profile of city residence (Fig. 2), the aging of this cohort had little effect on city growth. In essence, there is no decade in which the baby boom generation has been clustered in either a peak or a valley of the city residence profile. In 1980, for example, many of the baby boomers were in their early 20s and lived in central cities. However, at the same time, others in the cohort were still in their teenage years or had entered their early 30s and therefore tended to live in the suburbs.

We explored the possibility of age effects on city growth by imposing a counterfactual flat age profile and predicting the resulting city share of the metropolitan population in each decade. In particular, we allowed each age between 0 and 70 to contain 1.3 % of population and constrained older ages to each hold 0.3 % of population. Using the estimated age effects in Fig. 2, we then predicted the share of the population that would be living in the central city under this counterfactual age profile. We found no meaningful difference between the predicted and the actual city shares, leading us to turn our attention to the other set of demographic factors (results available upon request).

Counterfactual Simulations

Table 5 uses the coefficient estimates relating demographic trends to city residence to provide counterfactual statements about how much further the share of metropolitan residents living in central cities *would have declined* between 1960 and 2000 if not for these demographic moderators. In the simplest exercise, we use the probit or IV coefficients to consider the extent to which central city population increased through each of the four demographic channels. We then allow for the fact that new arrivals may result in the departures of some existing residents, either directly through white or native flight or indirectly via an increase in city housing prices.

The first row of Table 5 presents the actual change in the share of metropolitan residents living in central cities from 1960 to 2000. The share of the metropolitan population living in a central city declined by 17.8 percentage points over this period. The magnitude of this decline reflects a combination of (a) socioeconomic forces favoring suburban residence and (b) the demographic trends favoring city residence.

The second row of Table 5 isolates the role of changes in demographic composition. In particular, we ask how the share of the metropolitan population living in a central city would have changed if changes in demographic composition had been the only relevant factor over this period (and if new arrivals did not generate corresponding departures). In this case, the city share of metropolitan population would have increased by 6.4–8.3 percentage points from 1960 to 2000. The low and high points of this range measure the relationship between demography and residential location using the probit or IV coefficients, respectively.[21]

[21] Because our IV regressions are estimated on selected samples, we use the ratio between the OLS and IV estimates in Tables 3 and 4 to scale the coefficients for the whole population from Table 1. Specifically, we augment the veterans coefficient by 2.3 (−0.067 / –0.029 from Table 3, columns 1 and 2), and we augment the "any child" coefficient by 1.6 (−0.052 / –0.032 from Table 4, columns 3 and 4).

Table 5 Demographic contributions to city population growth: Counterfactual scenarios

	Share of Metropolitan Population Living in Central Cities			
	Level in 1960	Level in 2000	Change (probit)	Change (IV)
1. Actual City Share	51.3	33.5	−17.8	−17.8
2. Counterfactual Share, Gross Population Flows	51.3	57.7	6.4	8.3 [5.4, 11.5]
3. Contributions to Gross Counterfactual				
a. Foreign born	Increased 16.3 points		2.8	2.8
b. Black	Increased 4.3 points		1.6	1.6
c. Veteran	Declined by 25.8 points		1.2	2.5 [0.4, 5.0]
d. Children in household	Declined by 11.1 points		0.8	1.4 [0.6, 2.1]
4. Counterfactual Share, Net Population Flows	51.3	53.4	2.1	3.4
5. Contributions to Net Counterfactual				
a. Foreign-born			0.6	0.6
b. Black			0.0	0.0
c. Veteran			0.8	1.8 [0.3,3.5]
d. Children in household			0.5	1.0 [0.4, 1.5]

Notes: Row 1 reports the actual share of metropolitan area residents who report living the central city from IPUMS samples. Row 2 presents the counterfactual share of the metropolitan population living in central cities under a scenario in which demographic composition is the only factor allowed to change between 1960 and 2000. The counterfactual in column 3 is based on the probit regression in Table 1; the counterfactual in column 4 is derived from the IV coefficients in Tables 3 and 4. See footnote 23 for the details on translating the IV coefficients for use in the counterfactual. For the IV-based counterfactuals, we report the 95 % confidence intervals are shown in brackets. Rows 3a–3d indicate the contribution of each demographic factor to the overall counterfactual in row 2. Row 4 reports the results of a modified counterfactual simulation that allows for the fact that new arrivals may lead to the departures of some existing residents, either through white/native flight or through an increase in housing prices. Rows 5a–5d indicate the contribution of each demographic factor to the net counterfactual in row 4.

Rows 3a through 3d illustrate how each demographic characteristic contributes to the total counterfactual change in city population. That is, the sum of the entries in rows 3a through 3d is equal to the total counterfactual change in the city population share in row 2. To generate these values, we multiply the total change in the variable in question from 1960 to 2000 by the estimated effect of that variable on the probability of living in the central city. For example, the share of households headed by an immigrant increased by 16.3 percentage points from 1960 to 2000 and, as a result, the share of metropolitan residents who lived in the central city rose by 2.8 percentage points (16.3 × 0.170, coefficient estimate from Table 1, column 1).

The most important demographic trend contributing to the growth of city population is renewed immigration, followed by the increase in black population. The IV estimates suggest a greater role for the decline in veteran status and in the share of households with children present. However, we caution that the IV estimates are less precisely estimated. Therefore, we report 95 % confidence intervals around each of the counterfactual effects calculated from the IV estimates. Note that if we had instead selected 1940 or 1950 as a starting point for this exercise, the role of racial composition and the presence of children would have been larger.

The simple counterfactuals discussed thus far do not account for the possibility that some existing residents might have left central cities as new households arrived, either because of a direct distaste for living near black or immigrant neighbors or because of the indirect effect of arrivals on city housing prices. The fourth row of Table 5 presents a modified counterfactual using estimates of white/native flight and housing price responses from the literature to calculate the *net effect* of these demographic shifts on central city population.

Recent studies of white and native flight have found nearly one-for-one displacement rates from cities or urban neighborhoods. Boustan (2010) showed that, over a single decade, one southern black arrival into a central city led to 2.5 nonblack departures. Over the long run, new arrivals partially compensate for initial white flight, and one black arrival is associated with one nonblack departure. Saiz and Wachter (2011) found that one immigrant arrival into a census tract led to the departure of 0.68 native, non-Hispanic whites. Borjas (2006) presented a similar estimate for the native workforce at the metropolitan level (0.61 departures).[22] Adjusting for native white departures completely offsets the effect of black arrivals and reduces the net effect of immigrant arrivals by 70 %. In this case, the 4.4-percentage-point gross increase in the share of the metropolitan population located in the central city owing to black and immigrant arrivals would have led to a *net* increase of only 0.9 percentage points in the city share [(= (2.8 points × 0.32) + (1.6 points × 0.0)].

The other two demographic groups—veterans and households without children—are not likely to prompt specific outflows. However, their presence could affect the probability that other households remain in (or move to) a city indirectly via the housing market. A larger population in the city can increase housing prices, and higher prices may in turn encourage existing residents to leave or deter other new residents from moving in.[23] The best empirical evidence on the effect of population growth on housing prices is based on variation in immigrant arrival rates across metropolitan areas.[24] Saiz (2007) found that a 1 % *net* increase in metropolitan population increases housing prices by 0.77 %.

[22] Similar work by Crowder et al. (2011) has shown that, for the native-born, the log odds of moving increase with the foreign-born share of the neighborhood.

[23] The extent of price increases depends on the elasticity of housing supply; at the extreme, prices will not respond to in-migration if each new arrival is met with the construction of a new housing unit.

[24] An earlier literature used total cohort size to assess the relationship between population growth and housing prices. Mankiw and Weil (1989) documented large effects of the entry of baby boomers into the housing market in the United States, while Engelhardt and Poterba (1991) found no effect in Canada. This evidence is hard to interpret because it is based on national time series.

Assessing the effect of higher housing prices on out-migration is an empirical challenge. In the raw data, there is no correlation between housing rents and net migration at the state level (Coen-Pirani 2010). However, this null effect is likely due to the fact that states with higher housing prices also have a more productive set of industries (and hence higher wages) or a more valuable set of amenities (Roback 1982). Saks (2008) wrote one of the few papers to (indirectly) estimate the effect of an exogenous change in housing prices on migration rates, in this case driven by variation in zoning regulations across cities. She found that, in places with strict zoning rules, in-migration in response to a given labor-demand shock is dampened by a corresponding increase in housing prices; her estimates imply that a 1 % increase in housing prices reduces any given in-migration flow by 0.4 %.

According to our probit estimates, the three demographic forces under consideration increased gross city population by 3.0 percentage points, or by 7.5 %.[25] Using estimates from the literature, we propose that housing prices would have increased by 6 %, leading to a subsequent 2 % population decline. In other words, a 7.5 % increase in *gross* city population would have resulted in only a 5.5 % *net* increase in the number of city residents owing to the effect of population growth on housing prices. Converting these percentages back into shares, these calculations imply that, net of white flight and price-induced out-migration, demographic factors would have increased the city share of the metropolitan population by 2.1 percentage points (Table 5, row 4). Rows 5a through 5d use a similar logic to illustrate how each demographic trend contributes to the net counterfactual.

Overall, we conclude that, absent these demographic shifts, central city population would have declined by 10 % to 32 % more than it did between 1960 and 2000. The range of these estimates depends on the estimation method used (probit or IV) and assumptions about how existing residents would have reacted to these new arrivals. Our preferred estimate uses the IV coefficients to calculate the effect of each demographic factor on city population, while also accounting for responsive out-migration. By this estimate, the share of metropolitan residents living in the city would have declined by an additional 3.4 percentage points, or 16 %, from 1960 to 2000 if not for the demographic counterweight (16 % = 3.4 additional points / [17.8 actual points + 3.4 additional points]). In other words, shifts in demographic composition helped to maintain city population but were not strong enough to compensate for the powerful forces favoring population growth in the suburbs.

Conclusion

The share of the metropolitan population living in central cities has declined sharply over the past 60 years. In this article, we show that, absent changes in the demographic composition of the metropolitan population, this share would have fallen even further. In particular, city population has been bolstered by the in-migration of

[25] Over this period, around 40 % of metropolitan residents lived in a central city. Therefore, city population would have needed to increase by 7.5 % in order for the share of the metropolitan area living in a central city to increase by 3 percentage points.

southern blacks from 1940 to 1980, the expansion of international immigration after 1965, and a decline in the share of households with children or headed by a veteran.

We provide new estimates of the relationship between each of these demographic characteristics and the likelihood of living in the central city. Our analysis distinguishes between predetermined characteristics, such as race and nativity, and endogenous characteristics like veteran status and childbearing. We instrument for veteran status by comparing birth cohorts of men coming of age during and just after the mass mobilization for World War II. We use the arrival of twins to instrument for the number of children in a household. In both cases, the IV estimates are larger in absolute value than the corresponding probit estimates.

Two counterfactual simulations are used to assess the effect of these demographic factors on city population from 1960 to 2000. Our simplest approach predicts the gross number of new residents in central cities using our coefficient estimates and the trends in each characteristic over time. We then allow for the fact that new arrivals may lead to the departures of some existing residents, either directly through white or native flight or indirectly via an increase in city housing prices. The counterfactuals indicate that changes in demographic composition increased the share of the metropolitan population living in the central city between 10 % and 32 %. Our preferred estimate, which relies on the IV coefficients and allows for responsive out-migration, suggests that the share of the metropolitan population living in the central city would have fallen by 16 % more than it did if not for these demographic trends. Changes in demographic composition were strong enough only to attenuate, not to reverse, a relative decline in city population driven by economic and social factors favoring suburbanization.

Our national focus may miss localized instances of gentrification in certain neighborhoods or within particular cities that have been fueled in part by demographic trends (Vigdor 2002). For example, the interaction between demographic forces—particularly delayed childbearing, longer life expectancies, and rising incomes in top income brackets—may have contributed to the renaissance of "superstar" cities like New York City and San Francisco (Gyourko et al. 2006). Understanding variation in the role of demography across different types of cities remains an important avenue for future research.

Acknowledgments We thank Pascaline Dupas and Fran Kobrin Goldscheider for motivating conversations. Adriana Lleras-Muney, Stewart Tolnay, and participants at the 2010 annual meetings of the Population Association of America and the UCLA KALER group provided helpful suggestions.

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