TRAFFIC AND SPRAWL:
EVIDENCE FROM U.S. COMMUTING, 1985 TO 1997*

by

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Abstract
The consequences of sprawl for travel behavior remain unclear. Theory suggests at least two possible commuting outcomes. As jobs decentralize and central employment areas congest, workers might shorten their commutes in time and distance by relocating to the suburbs. Or, the average commute could grow if residential choice is relatively inelastic with respect to job location, amenity explanations for residential and job location dominate, or as dual-worker households in polycentric labor markets become the norm.

Evidence on these questions is surprisingly rare and dated. For data on individual travel behavior, we use the American Housing Survey, a detailed individual-level panel survey for most major metropolitan areas of the US for several years between 1985 and 1997. Commute distance is regressed on a reduced form travel demand model, including U.S. Bureau of Economic Analysis measures of metro-wide employment deconcentration at the one-digit SIC industry level. The model specification conforms to urban form theory, the model estimation uses panel techniques, and the potential endogeneity of wages and land costs – as compensations for commute costs – is addressed statistically.

We find that the *more* suburbanized is employment – that is, the more sprawl – the *shorter* the average commute. There are strong differences by industry, however, that may reflect a combination of industry agglomeration effects, differential job location stability by industry, and historical transitions.

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1. Introduction

Perhaps the clearest structural evolution of metropolitan areas continues to be the decentralization of both employment and residents. During some periods, US central cities have grown at a lower rate than their suburbs; in others, the decline has been in absolute numbers. The overall trend is fairly strong, particularly over the past century:

> In 1940, only one of the ten largest cities in the US had a population density below 10,000 people per square mile (Los Angeles). In 1990, population density levels are below 7,500 people per square mile in seven of the ten largest cities… In 1940, the overwhelming number of urban jobs appears to have been close to the city center. In 1996, in the average metropolitan area only 16 percent of jobs are within 3 miles of the central business district. The medium density, driving city of today, has replaced the dense, walking city of the 19th century. (Glaeser and Kahn, 2001, p. 2)

Many questions remain regarding the consequences of this pattern of decentralization and deconcentration or, to use a broader term, sprawl. Much of the debate concerns environmental resource issues, such as the excess conversion of open land to urban uses and pollution problems associated with overdependence on cars. The potential social impacts are also widely discussed; for example, one of Putnam's (2000) "sprawl" explanations for social capital decline is that the suburban commute leaves less time and energy for social interaction.

Given continuing concerns with sprawl-related transportation issues, it is surprising how little is known about what sprawl means for driving behavior. On the other hand, it is less surprising given the complex nature of the relationship. For example, urban deconcentration consists of many different phenomena, including reduction in employment density, reduction in residential density, and a greater prevalence of discontinuous patterns of developed and undeveloped land. In this paper we focus on decentralization rather than deconcentration; in particular, the extent to which employment is decentralized, or dispersed, versus being clustered in the built-up portion of cities. More specifically, do people drive more or less when employment decentralizes?

We find that both the theoretical and empirical literatures reveal significant gaps. In general, theory provides several hypotheses but few if any hard conclusions, while empirical studies rarely formally test those hypotheses directly.

This paper addresses these gaps in several ways. It builds on established theories of urban structure and transportation behavior to develop the hypotheses of interest. These are tested on a rich data set for US metropolitan areas over the period 1985 to 1997. The model specification conforms to urban form theory, the model estimation uses panel techniques, and the potential endogeneity of wages and land costs is considered.

In brief, we find that the *more* suburbanized is employment – that is, the more sprawl – the *shorter* the average commute. There are strong differences by industry, however. The suburbanization of construction, wholesale, and service employment is associated with shorter commutes, while manufacturing and finance deconcentration (weakly) explain longer commutes. These results may reflect a combination of industry agglomeration effects, differential job location stability by industry, and historical transitions. While travel behavior is thus complex and nuanced, there is no evidence in these data that job decentralization lengthens the average journey to work. It appears to do the opposite.

Section 2 summarizes the theoretical and empirical literature. The following section discusses our data, while Section 4 describes those data in several key respects. Section 5 presents the empirical strategy. We analyze our results in Section 6, followed by conclusions.

2. Previous Research

Much of the theoretical literature assessing the effects of population and employment decentralization on commute length or travel behavior is based on the monocentric model (Alonso, 1964; Wheaton, 1979). In that model, the market for urban land is defined by the inelastic demand for a commute to the city center by one-worker households. Individual households choose where to live by trading off the value of access to their current jobs against the cost of housing, resulting in declining equilibrium housing rents and residential development densities as commute length increases (i.e., with increasing distance from the city center). Transportation costs thus determine rents, and the extent of decentralization is explained by the interaction of transportation costs with household preferences, given their financial and time budget constraints. In the end, suburban residents drive more but are compensated by lower rents per unit of land.

Simple polycentric elaborations on the monocentric model permit workers to choose from among several job locations within a given metropolitan area, allowing greater average proximity between home and work (White 1976, 1988; Wieland 1987; Yinger 1992). In a world where employment exogenously disperses, this basic polycentric version of the monocentric model suggests that commutes should get shorter, *ceteris paribus*.[1] Recognizing that wages may play a role in compensating for the journey to work, in addition to land rents, increases the likelihood that firms face similar incentives (McMillen and Singell, 1992; Wheaton and Sivitanidou, 1992; Macek, et al., 2001; Senesky, 2001; Wheaton, 2002).

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[1] This is different from saying that commutes should or should not be reduced. Urban theory argues that traffic congestion by itself, as an externality, leads to overly decentralized monocentric cities, but that the consequences are less clear when employment also decentralizes (Anas, Arnott and Small, 1998). Thus, unpriced congestion by itself contributes to longer commutes, holding employment location fixed. However, the present study only explains commuting, not its welfare economics.

Gordon, Kumar and Richardson (1989b) provided one of the first empirical tests of the hypothesis that employment decentralization might reduce commuting, in this case commute duration. They used Census and satellite land use data for 82 US metropolitan areas to investigate the relationship between mean commute travel time by auto or transit and measures of urban form: total urban land area, proportion of workers in industrial and commercial employment respectively, and net land use density by type, controlling for area-wide average income.

In OLS regressions with commute duration as the independent variable, the authors found that spatially larger cities had longer commutes, while shorter commutes were associated with a higher proportion of industrial employment. Higher residential density was associated with commutes of longer duration. The proportion of employment in the central city was highly associated with commutes of longer duration. The authors concluded that both residential and employment dispersion lead to shorter commutes.

In a descriptive study appearing the same year, the authors again investigated some of the same relationships between congestion and urban form, this time focusing on city size (Gordon, Kumar, and Richardson 1989a). They used the Nationwide Personal Transportation Surveys of 1977 and 1983-84, along with 1970 and 1980 Census journey-to-work data, to argue that larger cities do not necessarily have longer commutes because much of that growth occurred in suburbs. They make two main points in support of this argument. First, in 1977, durations of commutes originating outside central cities were roughly equivalent regardless of city size, but those originating within central cities were uniformly lengthier with increasing city size. This suggests that only suburban dwellers avoided the effects of increasing city size. (The 1983 results were not as clearly defined but tended to show the same general pattern.) Second, commute speeds were higher in suburban areas.

Subsequent analytical work has mainly focused on disaggregate data.[2] Levine's 1990 dissertation and subsequent papers (Levine 1990, 1992, 1998) explain residential location choice, in search of shorter commutes to suburban employment, as a function of housing supply as well as demographic factors. He specifically finds that the commutes of low-income households *lengthen* as employment suburbanizes due to shortages of affordable household near these jobs. Another example is Dubin's use of 1977 Baltimore data to estimate a reduced-form supply and demand model of the *change* in commute time (Dubin 1991). She found that men, non-Blacks, higher income households, and those owning their home enjoyed shorter commutes. Notably, even while commute distances increased for most, travel times fell, especially in the less congested suburbs.

To summarize, commute length is a complex phenomenon that is especially challenging to investigate empirically. Investigating the connection between commute length and employment or residential decentralization is no less challenging. On the one hand, employers may benefit from shorter commutes by being able to reduce wages. However, this may be mitigated by the need for workers to take into account future job locations. Other important influences on residential location, such as local municipal and district services, and accessibility to other nonwork activities, may also come into play as decentralization occurs. The equilibrium outcome is an empirical matter, where the empirical strategy must account for the endogeneity of land rents, wages, employment status, and car access. We also need good measures of other demand/supply factors, such as socio-demographic characteristics at the individual level.

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[2] But see Chen, Ewing and Pendall (2002) for a recent effort aimed at explaining metropolitan level travel outcomes with aggregate sprawl measures.

3. Data

We rely on two main data sources: individual-level travel data and metropolitan-level employment data.

Data on worker commute distance and demographics

For data on workers and their commute lengths, we employ the American Housing Survey (AHS), a panel of housing units (mainly) surveyed every two years by the Census Bureau for the US Department of Housing and Urban Development. We use the seven waves for 1985, 1987, 1989, 1991, 1993, 1995 and 1997. Between 11,000 and 15,000 respondents are included per year. Rich detail is provided on the residents of the housing units, including the reported distance to work (duration is not available for all but one of these years), income, education, marital status, ethnicity, age, gender, and other demographic and economic characteristics. Even greater detail on the condition and characteristics of the housing unit is recorded, though not utilized at this stage of the research.

Eliminating group quarters, mobile homes, units from non-metropolitan areas (as defined by the Census Bureau), households who reported spending less than one percent or more than 150 percent of their income on housing, observations where the number of occupants was not reported, vacant structures, nonurban structures, and observations for which the SMSA is not provided leaves 185,085 total observations in the seven waves, representing 42,380 distinct housing units, or about two-thirds of the full sample.

Metropolitan employment and population data

The BEA Regional Economic Information System (REIS) provides county-level estimates of employment by industrial sector (one-digit Standard Industrial Classification code level) for the years 1969 to 2000. The REIS includes both regular wage and salary employment and other kinds of employment, such as proprietors' employment and self-employed people.

The source for wage and salary employment is the federal Covered Employment and Wages Program, a joint program with the Bureau of Labor Statistics (BLS) and the State Employment Security Agencies, which administer unemployment insurance programs. The states get the information from reports issued monthly by employers subject to unemployment insurance regulations. The employment figures include both full-time and part-time jobs.

The BEA augments the BLS wage and salary employment data with estimates for self-employment and proprietors' employment that are based partly on IRS income tax reporting. Self-employment and proprietors' employment can be a substantial fraction of total employment.[3]

In addition to employment by county by one-digit SIC code, the REIS tables include mid-year Census estimates of population by county.

Metropolitan area (MA) definitions

This study uses the official US metropolitan area definitions promulgated by the Office of Management and Budget in 1999 and revised January 28, 2002, which are the same

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[3] See http://www.bea.doc.gov/bea/regional/articles/lapi6992.htm.

definitions used for Census 2000. The definitions are provided by the Bureau of the Census, and include a classification of counties within metropolitan areas as "central" or "outlying."[4]

Metropolitan areas must include a city of at least 50,000 inhabitants; alternately, they can include a Census-defined "urbanized area" of 50,000 or more if the county containing the urbanized area has at least 100,000 inhabitants.

The classification of intra-metropolitan-area counties is of particular interest to this study. The Census definitions of central and outlying counties are extended. In brief, central counties are those with at least 50 percent of the population of a designated central city or 50 percent of whose population lives in a Census-defined urbanized area.[5] Counties bordering the central county or counties are designated as outlying counties, and included within the metropolitan area, based on commuting patterns and density. For example, if at least half of the residents of the nearby county commute to the central counties to work, and certain threshold density and population requirements are met, the county is designated as outlying and included within the metropolitan area. Peripheral counties whose commuting is at a lower level still qualify as outlying if their density is at a sufficiently high level, if a substantial enough fraction of their population is within the Census-designated urbanized area, or if recent population growth in the area has been high between the previous two decennial censuses.

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[4] See http://www.census.gov/population/estimates/metro-city/99mfips.txt.

[5] Central cities include the largest city in an MSA; any additional cities with more than 250,000 population or 100,000 workers; any additional cities of at least 25,000 in size with a large employment base (greater than 75 percent of population), 40 percent of whose employed residents work there; and secondary non-contiguous urbanized areas that meet certain criteria. See http://www.census.gov/population/www/estimates/mastand.html, Part III, Section 4.

Metropolitan employment and population variables

Metropolitan employment and population were calculated by summing the BEA county data using the definitions of metropolitan areas described above. The metropolitan employment and population decentralization variables were constructed by dividing the employment in designated central counties by that in all of the counties within the metropolitan area, and representing this number as a percentage in whole numbers. These variables were in turn assigned to respondents based on the match between the AHS designation of metropolitan areas, based on the 1983 OMB/Census definition, and the current (1999) OMB/Census designation. In cases of Consolidated Metropolitan Statistical Areas, which are metropolitan areas made up of more than one Primary Metropolitan Statistical Area (PMSA), the total employment in the CMSA was used and respondents from more than one PMSA were assigned the same decentralization variable values.

Unlike typical measures of decentralization based on distance from a city center, this measure allows non-contiguous or spatially extensive urbanization of sufficient density to be designated as core urbanization. Only employment in counties without large enough agglomerations or sufficiently high density is designated as decentralized or dispersed. The measure is conservative, although somewhat crude due to its reliance on county geography. It can be thought of as a measure of the share of employment in the most dispersed form—outside of urbanized counties altogether, though adjacent to them.

Matching metropolitan employment data to AHS respondents

The AHS data do not include information about geographical area when the information might enable identification due to small numbers. Therefore part of the total sample has metropolitan area identification codes suppressed. All such respondents are necessarily dropped from the analysis. This likely tends to bias the sample in favor of larger metropolitan areas.

The AHS definitions for New England metropolitan areas are based on aggregations of cities, towns, and some county portions, making them unsuitable for use with county-level BEA REIS employment estimates. In the absence of an alternative method, this required that respondents in 12 New England metropolitan areas were dropped from the analysis.[6]

Particular hypothesis-related criteria required that other metropolitan areas be dropped from the analysis. First, only individuals living in metropolitan areas with more than one county could be considered under the decentralization hypothesis. Second, only those living in metropolitan areas with clearly identified core/periphery areas corresponding to county lines could be used. Some metropolitan areas including more than one county had no counties defined as outlying.

4. Descriptive Results

As shown in Figure 1 below, the share of both total population and total employment in outlying counties of metropolitan areas has increased over the 12-year period. During this period, the percentage of the population residing in outlying counties increased from less than 9

[6] The metropolitan areas dropped were the Bangor (ME) MSA; the Barnstable-Yarmouth (MA) MSA; the Boston-Worcester-Lawrence (MA, NH, ME, & CT) CMSA; the Burlington (VT) MSA; the Hartford (CT) MSA; the Lewiston-Auburn (ME) MSA; the New London-Norwich (CT-RI) MSA; the New York-Northern New Jersey-Long Island (NY, NJ, CT, & PA) CMSA; the Pittsfield (MA) MSA; the Portland (ME) MSA; the Providence-Fall River-Warwick (RI & MA) MSA; and the Springfield (MA) MSA .

percent to about ten percent. The percentage of dispersal outside central counties is relatively low but also climbed during this period—increasing from about 6.2 percent in 1985 to about 7.3 percent overall by 1997.

Figure 1

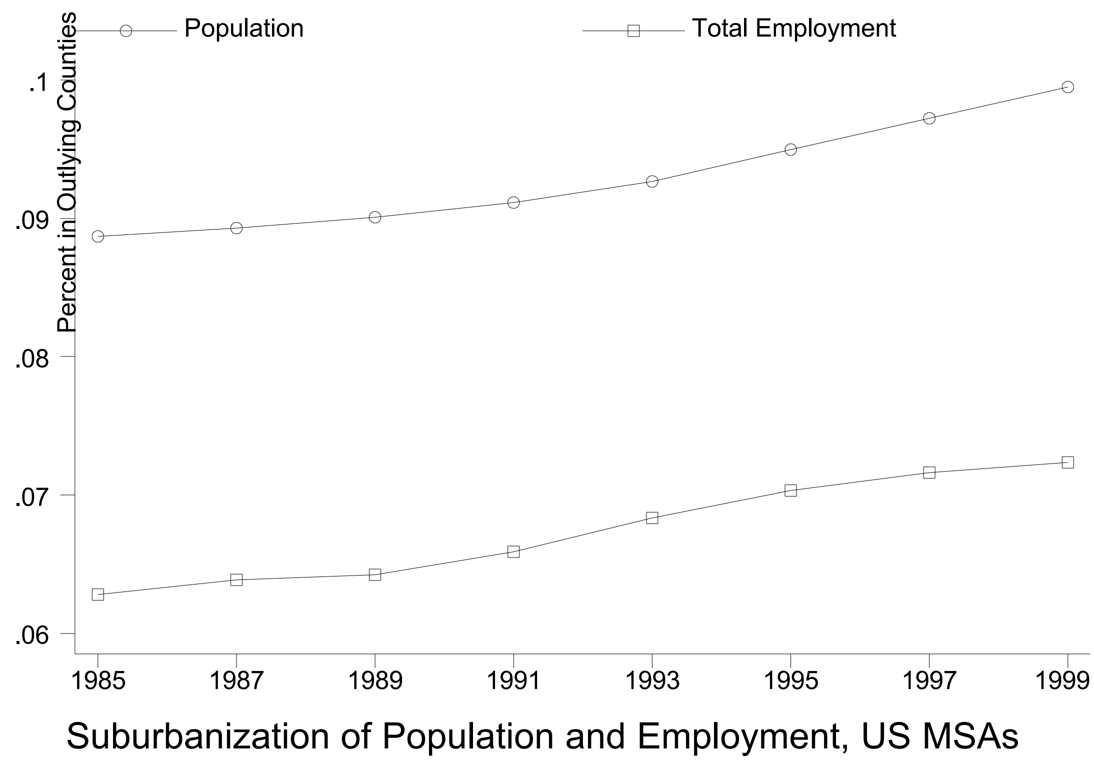
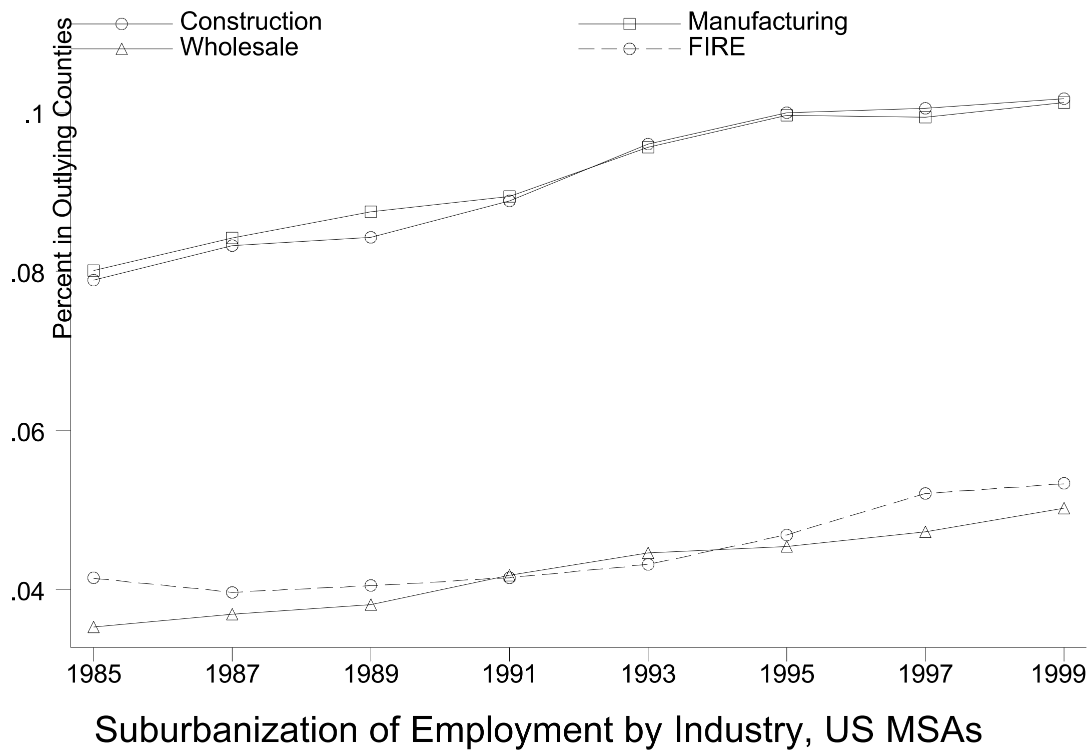


Figure 2 shows the share of employment in outlying counties for four industries defined using one-digit Standard Industrial Classification codes. Nationwide, manufacturing and FIRE (finance, insurance, and real estate) are substantially more dispersed than construction and wholesale employment.

Figure 2



Concurrent with this trend towards decentralization, the average commute distance showed a trend towards *increasing* distance over time in our sample of AHS respondents (Table 1). The average commute lengthened slightly through this period, though it declined somewhat from 1995 to 1997.

Table 1

Commute by Year
(AHS sample)

1985	10.7 miles
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1991	11.1
1993	11.0
1995	11.9
1997	11.3

There is some variance by Census region, as shown in Table 2. AHS respondents in selected metropolitan areas located in the northeast, northwest, and southern states have a similar average commute length, ranging from 11.2 to 11.5 miles. Those in the western Census region have a shorter commute length.[7]

Table 2

Commute by Region, 1985 to 1997
(AHS sample)

NE	11.2
MW	11.5
South	11.3
West	10.6

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[7] However, this sample excludes the two largest California CMSAs, for the data availability reasons given above, which likely reduces the average commute distance for the western region.

It is clear that demographic characteristics play a role in commute length. For example, as shown in Table 3, men have longer commutes than women, and residents of owner-occupied units have longer commutes than renters.

Table 3

Commute by Sex and Tenure, 1985 to 1997 (AHS sample)

Male	12.0
Female	9.7
Owner-Occupied	12.4
Renter	9.8

5. Empirical Strategy and Variables

To examine the issue of how sprawl influences the commute, we estimate commute length as a function of employment dispersal, individual occupation, and life-cycle factors. Other independent variables include demographic and economic characteristics of the individuals and households, such as income, the presence of dual earners, sex, race/ethnicity (African Americans, Asian Americans, and Latinos are represented with indicator variables), and educational attainment.[8] In addition to the respondent's age (and a squared age term to reflect nonlinearity in the relationship between age and travel), life cycle characteristics are included (whether the respondent is married and the number of children in the household).

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[8] Although only one individual from each eligible housing unit is used in the analysis to avoid the need to correct the standard errors for clustering within households.

Household characteristics that are expected to affect chosen commute length include housing tenure (renter or owner) and the number of automobiles owned by the household. Housing costs and income are included as explanatory variables. These are instrumented in later regressions.

Finally, in addition to variables representing both total employment decentralization and decentralization of employment by industry (percentages represented as whole numbers), we include several other measures of urban form: population decentralization (using the same county-based Census definition as for employment decentralization), total population and total employment, and land area of the counties in the metropolitan area. A single intra-metropolitan variable is included: whether the housing unit of the respondent is located in the central business district. Indicator variables for the Census regions are included to reflect differences in weather and transportation infrastructure that might vary by US region. A variable representing an independent time trend (YEAR) is included. All variables are listed with their labels in Table 3.

Several estimation strategies are used to investigate the hypothesized relationships between commute distance and employment decentralization. The initial analysis is carried out using a single equation and ordinary least squares (OLS), and all observations are treated as independent despite the repeated nature of the panel. Because income and housing cost are used as explanatory variables in this model, subsequent models correct for the potential endogeneity of these variables by using a two-stage instrumental variables technique. In later models we use panel regression techniques that separately account for cross-sectional variation (between the workers in different housing units) and variation over time (for each housing unit, for all workers living in the unit over the twelve-year period).

Table 4: Variable Labels

Label	Definition
Commute	One-way commute length in miles (log)
lr_hcost	Monthly housing costs (constant $, log)
ly_per	Household income per person (log)
ly_per2	Household income squared (log)
married	Married = 1
educ	Years of education
latin	Hispanic = 1
black	Black = 1
asian	Asian, PI = 1
AGE	Age
age2	Age squared
male	Male = 1
pchild	Proportion of household members that are children
apart	Apartment dweller = 1
CARS	Number of cars
MOVED	Year last moved
ZADULT	Number of adults in household
renter	Renter = 1
pop	Population of metropolitan area
employ	Employment in metropolitan area
landarea	Land area of metropolitan area
cmsaflag	CMSA = 1
cbd	CBD = 1
ne	Northeast = 1
mw	Midwest = 1
west	West = 1
YEAR	Year
sub_emp	Total suburban employment share in metro
mfsub_emp	Manufacturing employment share in suburbs
srsubub_emp	Service employment share
cons_emp	Construction employment share
whole_emp	Wholesale employment share
retail_emp	Retail employment share
fire_emp	Finance, Real Estate, Insurance employment share
gov_emp	Government employment share

6. Results

The results of the first simple regression are shown in Table 5. The dispersal of employment metro-wide is represented as the percentage share of employment in outlying counties within the metropolitan area, a figure ranging between 1.2 and 38.4 percent in the pooled panel sample of individuals.

Commute length is shorter the more suburbanized all employment. The coefficient on SUB_EMP is -0.299, implying that a five percent increase in the amount of employment in a metropolitan area's outlying counties will lead to a 1.5 percent reduction in the average commute distance, equivalent to a reduction of less than a tenth of a mile.

A separate time trend for increased commute length, as represented by the coefficient on YEAR, implies that despite the dampening effect of employment dispersal over this 12-year period, unexplained causes are lengthening commutes by about one percent per year.

Table 5
OLS Regression of Logged Commute Distance Using Pooled Panel

RHS variables	Coefficient	Std. Error	t
lr_hcost	0.023	0.011	2.13
ly_per	0.861	0.103	8.38
ly_per2	-0.041	0.006	-7.32
married	0.038	0.014	2.70
educ	-0.003	0.003	-0.84
latin	0.025	0.021	1.17
black	0.130	0.016	8.34
asian	-0.002	0.031	-0.06
AGE	-0.005	0.002	-2.25
age2	0.000	0.000	-1.39
male	0.097	0.013	7.58
pchild	0.246	0.029	8.37
apart	-0.155	0.016	-9.67
CARS	0.082	0.007	11.94
MOVED	0.000	0.000	-2.25
ZADULT	0.030	0.008	3.56
renter	-0.077	0.016	-4.90
pop	0.000	0.000	-2.05
employ	0.000	0.000	2.09
landarea	0.000	0.000	5.47
cmsaflag	0.052	0.016	3.32
cbd	-0.236	0.011	-20.64
ne	-0.160	0.027	-5.91
mw	-0.020	0.022	-0.93
west	0.066	0.022	2.99
YEAR	0.011	0.002	7.45
sub_emp	-0.300	0.110	-2.73
Constant	-3.534	0.499	-7.09

N = 45,082
F(27, 45054) = 123.06
Adj R-squared = 0.0681

We can decompose this result by breaking out employment by industry (see Table 6). The effect of suburban employment share varies by industry. Construction and wholesale employment dispersal are associated with shorter commutes, while manufacturing and government employment dispersal with longer commutes. Retail and service employment does not appear to be strongly associated with commute length.

These differences may be due to the pattern of clustering characteristic of these industries. Specifically, construction and wholesale employment may not be as clustered within a given county-level pattern of dispersion, while manufacturing and government employment may be. There is evidence that certain kinds of manufacturing firms (particularly, small manufacturers in the more technologically advanced industries) tend to cluster to realize Marshallian agglomeration economies. Meanwhile, retail and service firms do cluster to some extent, but because they are population-serving they tend to also be pulled out to follow the more dispersed pattern of the residential development that they serve.

This clustering explanation for the regression results shown in this initial research makes sense only to the extent that industry sector clustering patterns drive overall commute patterns, because we have not controlled for occupational characteristics of residents in this analysis. Exploring this relationship is a subject for future research.

Table 6
OLS on Log Commute Distance Using Pooled Panel, With Industry By Sector

RHS variables	Coefficient	Std. Error	t
lr_hcost	0.025	0.011	2.28
ly_per	0.858	0.103	8.35
ly_per2	-0.041	0.006	-7.28
married	0.037	0.014	2.64
educ	-0.002	0.003	-0.67
latin	0.025	0.021	1.17
black	0.131	0.016	8.39
asian	-0.001	0.031	-0.03
AGE	-0.005	0.002	-2.24
age2	0.000	0.000	-1.38
male	0.097	0.013	7.54
pchild	0.244	0.029	8.29
apart	-0.156	0.016	-9.75
CARS	0.081	0.007	11.81
MOVED	0.000	0.000	-2.32
ZADULT	0.030	0.008	3.61
renter	-0.077	0.016	-4.88
pop	0.000	0.000	-2.63
employ	0.000	0.000	2.74
landarea	0.000	0.000	5.32
cmsaflag	0.067	0.016	3.95
cbd	-0.236	0.011	-20.44
ne	-0.164	0.027	-5.27
mw	-0.022	0.022	-0.98
west	0.052	0.022	2.29
YEAR	0.011	0.002	7.02
mfsub_emp	0.381	0.153	2.48
srsubub_emp	-0.407	0.534	-0.76
cons_emp	0.025	0.011	-2.17
whole_emp	0.858	0.103	-4.44
retail_emp	-0.041	0.006	1.61
fire_emp	0.037	0.014	-0.60
gov_emp	-0.002	0.003	2.36
Constant	0.025	0.021	-7.06

N = 45,082
F(33, 45048) = 101.92
Adj R-squared = 0.0688

We also estimated models controlling for the endogeneity of housing costs, using an instrumental variables procedure in which predicted housing costs are used in place of actual housing costs. Predicted housing costs are obtained by using coefficients on the independent variables from a simple hedonic price model. Table 7 shows the results of this model. While the model is clearly an improvement on the previous one, the coefficient on SUB_EMP is more or less the same as in the simple ordinary least squares model.

Table 7
Two-Stage Least Squares on Log Commute Distance Using Pooled Panel

RHS variables	Coefficient	Std. Error	t
lr_hcost	-0.055	0.037	-1.47
ly_per	0.862	0.103	8.38
ly_per2	-0.040	0.006	-7.09
married	0.042	0.014	2.96
educ	0.001	0.003	0.17
latin	0.020	0.021	0.94
black	0.122	0.016	7.64
asian	0.001	0.031	0.03
AGE	-0.005	0.002	-2.03
age2	0.000	0.000	-1.83
male	0.097	0.013	7.59
pchild	0.293	0.036	8.06
apart	-0.163	0.016	-9.91
CARS	0.085	0.007	12.12
MOVED	0.000	0.000	-2.56
ZADULT	0.039	0.009	4.17
renter	-0.079	0.016	-4.98
pop	0.000	0.000	-2.57
employ	0.000	0.000	2.67
landarea	0.000	0.000	5.44
cmsaflag	0.051	0.016	3.27
cbd	-0.245	0.012	-20.20
ne	-0.169	0.027	-6.18
mw	-0.031	0.022	-1.38
west	0.057	0.022	2.55
YEAR	0.011	0.002	7.51
sub_emp	-0.303	0.110	-2.76
Constant	-3.214	0.520	-6.18

N = 45,082
F(33, 45054) = 122.84
Adj R-squared = 0.0671

Finally, controlling for the clustering of individuals within housing units does not change the results importantly in Table 8.

Table 8
Two-Stage Least Squares on Log Commute Distance Controlling for Panel Effects

RHS variables	Coefficient	Std. Error	t
lr_hcost	-0.037	0.044	-0.84
ly_per	0.729	0.101	7.19
ly_per2	-0.033	0.006	-6.04
married	0.027	0.014	1.88
educ	0.005	0.003	1.37
latin	0.018	0.023	0.78
black	0.106	0.018	5.88
asian	-0.041	0.034	-1.21
AGE	-0.001	0.003	-0.35
age2	0.000	0.000	-2.72
male	0.099	0.013	7.47
pchild	0.256	0.036	7.10
apart	-0.160	0.019	-8.63
CARS	0.070	0.007	10.21
MOVED	0.000	0.000	-2.63
ZADULT	0.034	0.009	3.59
renter	-0.079	0.017	-4.60
pop	0.000	0.000	-1.39
employ	0.000	0.000	1.48
landarea	0.000	0.000	3.90
cmsaflag	0.057	0.020	2.91
cbd	-0.250	0.015	-16.40
ne	-0.172	0.034	-5.05
mw	-0.017	0.028	-0.59
west	0.072	0.028	2.54
YEAR	0.010	0.001	7.50
sub_emp	-0.335	0.139	-2.40
Constant	-2.650	0.523	-5.06

N = 45,082
R-sq: within = 0.0081
Obs per group: min = 1
between = 0.0908 avg = 2.3

7. Concluding comments

Our evidence supports the argument that suburbanized employment – that is, the more sprawl – is associated with *shorter* commutes on average. This is not to say that commutes are shortening as cities expand their footprint; indeed, they seem to be slowly lengthening (Kahn, 2000). It does suggest that, in the aggregate, the marginal effect of job suburbanization is to mitigate this effect. Other factors, such as rising incomes, are generating more and longer trips to work. However, the direct effect of job suburbanization appears to be to bring jobs and workers closer. Put another way, the average commute would be longer still if jobs were not suburbanizing.

A closer look reveals that this is not a transparent outcome. There are strong differences by industry, for example. The suburbanization of construction, wholesale, and service employment is associated with shorter commutes, while manufacturing and finance deconcentration (weakly) explain longer commutes. These results may reflect a combination of industry agglomeration effects, differential job location stability by industry, and historical transitions.

What does this all mean? First, our study is exploratory and thus preliminary. Many loose ends remain, suggesting numerous ways to refine and extend our understanding of these relationships. On the one hand, the AHS data do not allow us to test how commute *duration* has changed. This ignores the substantial role of congestion in urban form and behavior debates (Solow, 1973; Wheaton, 1998). Other evidence suggests that congestion may be lower in outlying areas, so the dispersal of employment to outlying counties within metropolitan areas may actually reduce commute duration more than it does commute distance. Alternatively, if job dispersal to outlying counties is associated with a higher amount of non-work travel by affluent households living in those areas, commute duration could increase despite a slight decline in commute distance. This is because peak period travel duration is only partially accounted for by work trips; non-work trips make up a growing majority of trips made during peak periods (Giuliano, 1991).

In addition, while these results partially support the Crane (1996) hypothesis that commutes are longer for individuals with greater uncertainty about their future job locations and higher moving costs, our data do not permit us to explore this hypothesis in detail.

Finally, even if further study provides consistent support for the "sprawl shortens commutes" story, the public policy implications require more discussion. On the one hand, there is the issue of whether, on net, jobs follow workers to the suburbs or vice versa. The evidence is mixed on either count, suggesting the simultaneity of this process is challenging to nail down empirically. If jobs returned to the city, is it possible that workers would follow suit? On the other hand, neither the influence of urban form on travel behavior nor the merits of concentrated versus dispersed urban growth are well understood. The former is yet another complex set of nuanced behaviors awaiting better data and empirical strategies (Boarnet and Crane, 2001). Regarding the latter, we do not know how the net social and economic costs of sprawl, however measured, compare with their benefits (Crane and Greenstein, 2002). A key purpose of our study is to clarify the underlying behaviors needed for both kinds of evaluations.

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