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16. ABSTRACT

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The Decline in Mobility and its Impact on Passenger Travel

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Abstract

Residential mobility rates in the United States (US) have historically been among the highest in the world (Molloy et al, 2011); however, these rates have been in steady decline for the last two decades. While most planning models have ignored migration flows, and instead, focused on the characteristics of the population as drivers of travel demand, there are suggestions from other literatures that recent migrant may have different travel behavior than settled residents. This study is the first to explicitly model this relationship by merging data from the American Community Survey and the National Transit Database. The results also demonstrate that recent immigrant migrants to a metropolitan area are less likely to use transit if they arrived from another metropolitan area than those that have arrived directly from a foreign country.

1. INTRODUCTION

Residential mobility rates in the United States (US) have historically been among the highest in the world (Molloy et al, 2011). Annual one-year internal (i.e. move from one residential location to another within the country) migration rates were as high as 20%[1] between 1947 and 1983 (Frey, 2009), which have often been linked to high economic and social mobility among US residents. The rate, however, has steadily fallen over the past three decades, reaching a low of about 13% in recent years (Frey, 2009). Recent research (e.g. Cooke, 2011; Painter and Yu, 2014), including our own analysis of US CPS (Current Population Survey) and ACS (American Community Survey) data, suggests that this rate might have fallen even further since the 2008-09 recession.

Much research effort has gone into documenting the long-term decline in geographic mobility. We are, however, aware of only two papers – by Kaplan and Schulhofer-Wohl (2017), and by Molloy et al. (2014) – that have directly tested the determinants of the decline. Kaplan and Schulhofer-Wohl (2017) claim that economic restructuring, or the reduction in the geographic specificity of returns to skills, along with better information on the costs and benefits of migration partly explain the decline. Molloy et al. (2014) argue that the net benefit to changing employers has decreased, which in turn has made labor market and hence geographic transitions less desirable. Other possibilities suggested in the literature include: 1) decline in population within cohorts, characterized by certain combinations of sociodemographic characteristics such as age, that have traditionally been most mobile, 2) reduction in differences between consumption amenities across places, 3) stricter immigration policies that prohibit

[1] One-year internal migration rate of *m*% for a given year *y* means that *m*% of people surveyed in year *y* reported that their residence location was different in year *y* compared to year *(y-1)*.

frequent job-shifts among some non-citizens, and 4) fewer employment opportunities due to the recession (short-term).

Regardless of the causes of the decline, we observe that annual one-year inter-MSA (US Census-designated Metropolitan Statistical Area) migration rates have fallen across most demographic subgroups categorized by, for example, age, education, race, etc., over the 2005-2013 period (ACS data). Most notably, the rate for immigrants' (or the foreign-born, regardless of their immigration or citizenship status when surveyed) decreased by about 0.7 percentage points (from 3.1% to 2.4%)[2] compared to the relatively smaller decline of 0.2 percentage points (from 3.0% to 2.8%) for the US-born population. This steep decline in immigrants' mobility and its convergence to the native-born populations' level, particularly since the recession, is intriguing. A recent paper by Liu and Painter (2017) find that age and the reduction in new entrants can explain 40% of reduction in mobility for immigrants, but that demographic factors can explain little of the decline in mobility for native born populations during the last decade.

Geographic mobility (in- and out-migration) contributes to changes in the demographic, economic, and social makeups of cities. Monitoring and analysis of migration trends help planners and policy makers determine or forecast demand across urban sectors (e.g. housing, power, water, transportation, etc.) better, and devise strategies to adjust supply or manage demand. The effect of changes in geographic mobility on homeownership, for example, has been studied in the past (Painter and Yu, 2014). This work suggests that new entrants to regions are less likely to own. This is finding is much larger for those that are new entrants to the United States.

[2] Inter-MSA migration rate of *m*% for a given year *y* means that *m*% of people surveyed in year *y* reported that their MSA of residence was different in year *y* compared to year *(y-1)*.

The population mobility–transportation connection, however, remains unexplored, although the link is both straightforward and significant for urban policy. For example, in-migration of younger people, all else equal, could increase the demand for walking and bicycling in a city (McDonald, 2015 study is suggestive). Conversely, in-migration of high-income families, all else equal, could boost the demand for low-density suburban living and increase automobile travel. Both scenarios require policy interventions, such as building more sidewalks or bike-lanes and increasing gasoline taxes or parking fees, to promote broader transportation goals of increasing accessibility and reducing environmental impacts.

Declining inter-city mobility among immigrants affects urban passenger transportation. Travel surveys show that immigrants' travel behaviors and patterns are different from non-immigrants on average (e.g. see Tal and Handy, 2010). The difference, however, varies by immigrants' country of origin and length of stay in the US. The observed decline in mobility can significantly alter immigrant in- and out-flow patterns across cities. This can, in turn, alter the short- or long-term trends of the changes (net increase or decrease) in both the absolute numbers and proportions of immigrants across cities, thereby affecting aggregate travel demand trends in cities. Consequently, planning and public policies need to be reviewed and adjusted to accommodate or manage shifts in expected future demand.

Census/ACS and nationwide household travel survey (NHTS/NPTS)[3] data show that recent (<10 years in the US) immigrants constitute a significant proportion of transit users in the US. Therefore, changes (or decline in the present context) in geographic mobility among recent immigrants that alter their in- and out-migration rates[4] across cities can affect transit demand

[3] National Household Travel Surveys/Nationwide Personal Transportation Surveys
[4] e.g. rates in year *y* can be measured in terms of the numbers of recent immigrants who moved in and out of a given city between the years *(y-1)* and *y*, expressed as a fraction of the total population of the city in year *y*.

trends. Public transit agencies should therefore consider these mobility trends, among other factors, as they make forecasts and draft long range plans in anticipation of future demand.

In this paper, we present our research exploring how inter-urban migration of recent immigrants has affected aggregate public transit ridership change across US urban areas (Census-designated urbanized areas) over a five-year (2008-2013) period. By explicitly estimating the magnitude of change in ridership demand in an urban area in response to inflow of recent immigrants from other urban areas, all else remaining equal, our study helps determine the effect of declining geographic mobility of recent immigrants on urban transit. We confirm findings from other literature that among other factors, population density, transit supply, and the size of the recent immigrant population predict increased transit usage. The results also demonstrate that recent immigrant migrants to a metropolitan area are less likely to use transit if they arrived from another metropolitan area than those that have arrived directly from a foreign country.

2. LITERATURE REVIEW

Literature tells us that immigrants, particularly recent immigrants, significantly contribute to urban public transit demand in the US. All else equal, urban areas that have higher proportions of recent immigrants have higher transit patronage. Consequently, one can expect that urban areas experiencing net in-migration of recent immigrants, all else equal, will see year-over-year increases in transit demand (or ridership; latent demand if supply is inadequate). The influx of immigrants into California and the associated growth in transit ridership has been documented by Blumenberg and Evans (2010); Blumenberg and Norton (2010) have argued that almost all of the ridership growth in California since the 1980's can be attributed to in-migration of immigrants, and that falling immigration rates mean falling ridership for transit. Conversely, net out-migration will have the opposite effect on transit demand, and changes in migration patterns will change the demand vs. time curve, all else equal. Hence, changes in their geographic mobility that affect immigrant proportions relative to total populations (along with changes in the proportions over time) should independently impact aggregate transit demand (along with demand trends) across urban areas. In this section, we summarize recent studies that have underscored the immigrant-public transit connection.

Nationwide travel surveys can serve as the principal source of immigrants' travel behavior information. Tal and Handy (2010) analyzed 2001 National Household Travel Survey data to find that immigrants' travel behaviors and patterns are different, on average, from those born in the US. For example, the authors report that immigrants own fewer cars and drive fewer miles per year. They also find that immigrants from some parts of the world (e.g. Central and South America) are more likely to depend solely on public transit. Immigrants' length of stay in the US and country of origin, however, influence the nature and magnitude of difference in their travel patterns compared to the native born population. The authors find that immigrants slowly assimilate to US travel patterns after about five years. The country of origin effect is possibly derived from sociodemographic characteristics such as income and age, along with attitudes and cultures.

These results are analogous with those from studies analyzing Census commute mode choice data, such as Myers (1997) that used nationwide Census 1980 and 1990, Blumenberg and Shiki (2007) that used the 2000 Public Use Microdata Series from California, and Kim (2009) that used nationwide 2006 Public Use Microdata Series. These studies consistently reported recent immigrants to use public transit more and drive less compared to the US-born population as well as relatively longer-term immigrants. Myers (1997) found immigrant and US-born travel patterns to converge after immigrants spend 10 years in the US. Blumenberg and Shiki (2007) found the rate of assimilation to be a function of race or ethnicity; while Asians adopt the car relatively quickly, Hispanics use transit more than US-born commuters for as long as 20 years in the US. Kim (2009) found female immigrants to be less likely to use public transit than their male counterparts.

Reasons why immigrants use public transit more than the US-born, at least initially, include: preference for relatively high-density low-cost residential neighborhoods and employment in urban centers that are primed for transit (Chatman, 2014); barriers to car use such as lack of drivers' licenses and high cost of car ownership and use (Donahue and Rodier, 2008); and habits or past experiences of transit use (Chatman and Klein, 2013). Regardless of the determinants of immigrants' travel behaviors, the implications that growing immigrant populations and changes in their preferences or choices (e.g. declining geographic mobility and resultant changes in migration patterns) have on urban traffic congestion and transit ridership are large and significant for policy (Blumenberg, 2009).

3. METHODS AND DATA

We use a linear ordinary least squares regression model to analyze changes in aggregate transit ridership across U.S. Census-designated "urbanized areas" (we use the term "urban areas" and the abbreviated form UA throughout the paper) over a 5-year period (2008-2013) as a function of the net in-migration rate of recent (10 years or less in the US) immigrants over the period across the urban areas. We include other variables that could also contribute to observed changes in ridership. In addition, we control for factors that need to be held constant in order to evaluate the effects of the various explanatory variables on the dependent variable. A cross-sectional study by Taylor et al. (2009) that explored the determinants of transit ridership across

US urban areas helped identify critical explanatory variables and influenced the general model form. Choice of urban area as the geographic unit of analysis is governed by the standard level of aggregation of annual transit supply and ridership data by the federal government. Choice of the study period is governed by availability of transit and sociodemographic data at the desired level of geographic detail. Variables and data sources are summarized in Table 1.

Table 1: Dependent and explanatory variables at the UA level, used in regression models

Variable name	Variable description	Data source
Dependent variable		
ΔRIDER	Estimated change in the number of annual unlinked transit passenger trips (all modes of transit, and both directly operated and contracted lines included, per the NTD). $\Delta RIDER = RIDER_{2013} - RIDER_{2008}$	US National Transit Database (NTD)
Mobility variable		
MIG	Average (over the 2008-2013 period) annual in-migration of recent (<10 years in the US) immigrants from other UAs. MIG_{iy} for a given UA *i* in a given year *y* is estimated (from ACS one-year microdata sample for year *y*) as the weighted total number of recent immigrants in *i* whose UA of residence in the previous year *(y-1)* was not *i*. $MIG_i = \sum_y MIG_{iy}/5$ (y=2008, 09,10,12, and 13)	ACS one-year microdata
Other sociodemographic variables		
ΔIMM10	Estimated change in the number of recent (<10 years in the US) immigrants (or foreign-born persons). $\Delta IMM10 = IMM10_{2013} - IMM10_{2008}$	ACS one-year microdata
ΔPOVERTY	Estimated change in the number of people with family income under the federal poverty line. $\Delta POVERTY = POVERTY_{2013} - POVERTY_{2008}$	ACS one-year microdata
ΔCOLLEGE	Estimated change in the number of college (or more) educated persons. $\Delta COLLEGE = COLLEGE_{2013} - COLLEGE_{2008}$	ACS one-year microdata
ΔUNEMP	Estimated change in the number of unemployed persons. $\Delta UNEMP = UNEMP_{2013} - UNEMP_{2008}$	ACS one-year microdata
ΔWHITE	Estimated change in the number of persons with race White only. $\Delta IMM10 = IMM10_{2013} - IMM10_{2008}$	ACS one-year microdata
Transportation supply variables		
ΔROADPERCAP	Change in road miles per capita. $\Delta ROADPERCAP = ROADPERCAP_{2013} - ROADPERCAP_{2008}$	FHWA highway statistics
ΔTRANSITMI	Change in annual transit revenue miles. $\Delta TRANSITMI = TRANSITMI_{2013} - TRANSITMI_{2008}$	US NTD
Control variables		
ΔPOP	Estimated change in population. $\Delta POP = POP_{2013} - POP_{2008}$	ACS one-year microdata
POP2008	Baseline population, capturing the effect of city size.	ACS one-year microdata
TRANSITMI2008	Baseline annual transit revenue miles, capturing the effect of the level of transit supply.	US NTD

Note: Individuals over 16 years of age are considered only; The NTD is available at transit.dot.gov/ntd; The ACS microdata is obtained from the IPUMS USA website, usa.ipums.org; FHWA highway statistics are available at www.fhwa.dot.gov/policyinformation/statistics.cfm; all websites last accessed in April 2017

Computation of the *MIG* measure at the urban area level is not straightforward. The geographic identifier of the current place of residence of an individual in ACS microdata samples is PUMA (Public Use Microdata Area), not urban area. The geographic identifier of the place of residence in the previous year (for those who moved into the current PUMA) is the MIGPUMA (Migration PUMA), not PUMA or urban area. Moreover, Census-designated PUMA/MIGPUMA and urban area boundaries do not coincide, and their relative sizes vary. The boundaries are also periodically changed. In order to estimate the mobility measure at the aggregate urban area level, our spatial unit of analysis, we performed additional geoprocessing. Since we also used multi-year ACS microdata to estimate other urban area level sociodemographic variables used in the models, establishing the PUMA-UA spatial association was essential.

We used Census TIGER/Line Shapefiles to establish the relation between urban areas, PUMAs and MIGPUMAs for each year within the study period.[5] PUMAs and MIGPUMAs are easily relatable – a MIGPUMA is comprised of one more PUMAs, and their boundaries coincide with each other and with their constituent census tracts or counties. Urban area boundaries, however, are somewhat arbitrarily defined by urban development patterns and not by standard administrative boundaries.

Urban area boundaries were updated by the Census Bureau in 2010, but the NTD used 2000 urban area designations for their 2010 dataset. PUMA/MIGPUMA boundaries were updated by the Census Bureau in 2012. As a result, there are two groups of years (2008-2010, and 2012-2013) in which NTD and ACS microdata share version-consistent, albeit incongruent, underlying geographies. We use these years only, and do not include 2011 ACS microdata or

[5] The PUMA-MIGPUMA-UA spatial relationship for each year is important for estimating the average annual in-migration measure over the 2008-2013 period for each urban area.

NTD data in our analysis. The urban area level annual in-migration data is averaged across the five usable years.

To map residence PUMAs of ACS microdata individuals to urban area of residence in a given year, we use the following logic: 1) if residence PUMA falls completely inside a UA → assign the UA as residence; 2) if residence PUMA partially intersects with only one UA, or if only one UA falls completely inside the PUMA → assign the UA as residence if the UA covers >50% of the PUMA's area, else drop the observation; 3) if residence PUMA intersects with multiple UAs → assign the UA with greatest overlap as residence provided that the overlap area is >50% of the PUMA's area, else drop the observation.

Next, in order to ascertain whether a mover with an assigned urban area of residence migrated from outside the urban area over the past year, we overlaid MIGPUMA boundaries over urban area boundaries. We found that some urban areas intersect multiple MIGPUMAs and vice versa. We classify an individual as an in-migrant if their origin MIGPUMA (residence location in the previous year) does not intersect the assigned urban area of current residence.

The above data processing steps help identify, for individual (person-level) observations in each of the five selected ACS microdata samples, both the urban area of current residence and urban area of residence in the previous year (if applicable), and hence estimate urban area level aggregate sociodemographic as well as migration variables.

Note that it is impossible to accurately ascertain outmigration from one urban area to another because: 1) the ACS is a sample survey where person-weights are required to estimate population characteristics, and 2) the ACS is not a repeated panel survey. This data limitation, however, does not affect the interpretation of our analysis because if both *ΔIMM10* and *MIG* are known, the average annual recent immigrant out-migration value can be estimated fairly accurately.[6]

4. ANALYSIS AND DISCUSSION

Table 2 summarizes descriptive statistics of the variables used in regression analysis. Data are available for 338 out of 486 Census (2010) designated urban areas for most variables. For each of the remaining urban areas, either there was at least one year within the study period for which transit supply and ridership data was not available from the NTD, or there was at least one ACS survey year in which no surveyed individual could be assigned to the urban area based on our PUMA-UA match criteria (see section on data and methods for details). Moreover, since road supply data was available for 294 out of the 338 selected urban areas, our regression models include 294 urban area observations.

Table 2: Descriptive statistics

Variable	N	Mean	Std. Dev.	Min	Max
ΔRIDER	338	366557.10	6400049.00	-24200000	102000000
MIG	338	3723.34	10049.67	0	121157
ΔIMM10	338	-4883.44	23557.83	-277953	27116
ΔPOVERTY	338	6074.30	30065.00	-75870	284647
ΔCOLLEGE	338	75201.81	140531.60	-103413	1172103
ΔUNEMP	338	9103.55	23354.13	-11408	258612
ΔWHITE	338	26445.37	87052.08	-249040	548701
ΔROADPERCAP	294	0.00112	0.00146	-0.00445	0.01045
ΔTRANSITMI	338	262563.30	1892745.00	-12600000	12700000
ΔPOP	338	42644.46	112415.70	-256669	741359
POP2008	338	577360.00	1236573.00	79173	15300000
TRANSITMI2008	338	11400000.00	56500000.00	183016	928000000

Δ refers to change over the 2008-2013 period

On average, annual ridership increased by about 367,000 unlinked trips across the urban areas. In terms of annual ridership numbers, Atlanta, GA experienced the biggest decline (24 million unlinked trips or 15% of 2008 ridership) and the New York-Newark urban area saw the

[6] Out-migration estimates, however, also include counts of deaths and persons who shift from under to over 10 years in the US and hence leave the "recent immigrant" cohort.

biggest increase (about 102 million unlinked trips or 2.5% of 2008 ridership). Table 3 lists the top-5 (gainers) and bottom-5 (losers) urban areas in terms of ridership change.

Table 3: Top-5 and bottom-5 urban areas, ridership change (2008-2013)

Urban area	Change in unlinked passenger trips (2008-2013)
Top-5	
New York-Newark, NY-NJ-CT	102,000,000
Boston, MA-NH-RI	25,700,000
Philadelphia, PA-NJ-DE-MD	20,000,000
Athens-Clarke County, GA	11,100,000
New Orleans, LA	10,900,000
Bottom-5	
Milwaukee, WI	-9,459,318
Baltimore, MD	-11,000,000
Los Angeles-Long Beach-Anaheim, CA	-15,500,000
Houston, TX	-15,600,000
Atlanta, GA	-24,200,000

The estimated average annual in-migration of recent immigrants range from 0 (Uniontown-Connellsville, PA) to 121,157 (New York-Newark urban area) with a mean of about 3,700. Table 4 lists the top-5 and bottom-5 urban areas in terms of average annual in-migration, measured as 1) absolute numbers and 2) percentages relative to baseline total adult population. The table shows that, along with New York, Los Angeles, Miami, Washington DC and Chicago urban areas experienced the highest inflow of recent immigrants over the study period. When inflow is expressed as a percent of baseline population, relatively small (2008 urban area population under 200,000) college towns such as Champaign, IL, Ithaca, NY, College Station, TX and Lafayette, IN that attract new students in large volumes every year from all over the world rank high. The high-tech services capital of the US, San Jose, CA, ranks fifth.

Table 4: Top-5 and bottom-5 urban areas, recent immigrant in-migration (2008-2013)

Average annual in-migration of recent immigrants (2008-2013), rounded	Average annual in-migration of recent immigrants (2008-2013) as percent of 2008 population
Top-5	***Top-5***
New York-Newark, NY-NJ-CT (121,157)	Champaign, IL (2.67)
Los Angeles-Long Beach-Anaheim, CA (69,786)	Ithaca, NY (2.32)
Miami, FL (57,203)	College Station-Bryan, TX (2.00)
Washington, DC-VA-MD (44,249)	Lafayette, IN (1.83)
Chicago, IL-IN (43,636)	San Jose, CA (1.80)
Bottom-5	***Bottom-5***
Decatur, AL (41)	Wheeling, WV-OH (0.04)
Danville, IL (40)	Decatur, AL (0.03)
Monessen-California, PA (38)	Altoona, PA (0.03)
Altoona, PA (35)	Danville, IL (0.03)
Uniontown-Connellsville, PA (0)	Uniontown-Connellsville, PA (0.00)

On average, the number of recent immigrants across urban areas decreased by about 4,900. Among the million-plus (16+ population in 2013) urban areas, Los Angeles-Long Beach-Anaheim, New York-Newark, Chicago, Phoenix-Mesa, and Dallas-Fort Worth-Arlington urban areas each experienced decreases of more than 100,000 recent immigrants, whereas San Antonio, Seattle and Orlando urban areas each recorded increases of over 100,000. Road miles increased only by about a mile per 1000 persons on average, which is expected. There is a large variation in the change in annual transit revenue miles in our dataset. Among million-plus (16+ population in 2013) urban areas, annual revenue service was reduced by over 5 million miles in New York-Newark, San Francisco-Oakland, Atlanta, Chicago, and Pittsburg urban areas, whereas service was increased by over 10 million miles annually in Los Angeles-Long Beach-Anaheim and Washington, DC urban areas. The aggregate statistics on the changes in numbers of persons in poverty, college-educated persons, unemployed persons and White persons are for reference only – the changes in these variables should be interpreted at the disaggregate urban area level, relative to changes in populations.

Results of the OLS regression model is summarized in Table 5. The findings are generally consistent with expectations and prior studies. The model has a high adjusted R-square value of 74%, indicating that the included variables explain a large portion of the variation in transit ridership across urban areas over the study period.

Table 5: OLS regression model of change in transit ridership (2008-2013)

Variable	Coef.	P>t
MIG	-156.89	0.05
ΔIMM10	68.99	0.00
ΔPOVERTY	-3.66	0.79
ΔCOLLEGE	-0.14	0.99
ΔUNEMP	36.17	0.32
ΔWHITE	-23.74	0.01
ΔROADMI	1.66E+08	0.24
ΔTRANSITMI	0.28	0.03
ΔPOP	12.22	0.29
POP2008	-2.63	0.11
TRANSITMI2008	0.17	0.00
Constant	4.72E+05	0.15

Adj. R-square=0.74; N=294

Among sociodemographic variables, recent immigrants and White persons have statistically significant but opposite associations – positive and negative respectively – with transit ridership, all else equal. Transit service increase is positively associated with ridership gain; between two urban areas that underwent equal increases in transit service, the one with a larger baseline transit supply experienced a larger gain all else equal, indicating economies of scale in transit investments.

The regression results suggest that after controlling for the size of the total and immigrant population, the in-migration rate[7] of recent immigrants is negatively associated with aggregate urban area transit ridership. The finding is contradicts our initial hypothesis, but there are several possible explanations. Since we control for the change in total population as well as the change in the recent immigrant population among other variables, the observed effect suggests that the average recent immigrant who migrated from one city to another within the US over the

[7] Rate refers to annualized average over the study period.

study period was not a transit patron in the destination city. This might be plausible because those who chose to move could have been relatively more educated and/or higher-income than others to begin with, and therefore were unlikely to be transit patrons on average in the origin city. Upward economic and social mobility in the destination city could further reduce their propensity of transit use. We, however, cannot test this using available data. Moreover, holding the change in the recent immigrant population constant, higher annual in-migration implies higher annual outflow as well – owing to both out-migration and cohort shifts (under to over 10 years in the US). This suggests that the outflow of recent immigrants negatively affects transit ridership in an urban area, all else equal.

The finding above does not imply that in-migration of immigrants, in general, is unproductive for the transit industry. Table 6 proves that in-migration of recent immigrants (from other urban areas within the US) and in-migration of new immigrants (from other countries) have opposite effects on transit ridership. We use the *FOREIGN* variable (mean=3074.54, SD=9663.49, min=10.60, and max=141053), an estimate of the annualized average number of foreign-born persons migrating into an urban area over the study period, to test the new immigrant effect. All else equal, increases (change in the positive direction, generally speaking, as the *FOREIGN* variable can take any value – positive or negative) in the number of new immigrants into an urban area is found to be associated with bigger positive change in transit ridership, suggesting that new immigrants drive, in part, urban transit use. Comparison of standardized (beta) coefficients indicate that the *MIG* (-0.47) and *FOREIGN* (0.38) variables have comparable (in magnitude) opposite effects on ridership.

Table 6: OLS regression model of change in transit ridership (2008-2013) – intra-urban migration of recent immigrants vs. inflow of new immigrants

Variable	Coef.	P>t
MIG	-292.29	0.00
FOREIGN	254.86	0.01
ΔIMM10	76.34	0.00
ΔPOVERTY	3.31	0.81
ΔCOLLEGE	1.90	0.88
ΔUNEMP	85.14	0.04
ΔWHITE	-15.76	0.11
ΔROADMI	1.49E+08	0.29
ΔTRANSITMI	0.26	0.04
ΔPOP	0.36	0.98
POP2008	-3.21	0.05
TRANSITMI2008	0.14	0.00
Constant	4.91E+05	0.13

Adj. R-square=0.75; N=294

Analysis reveals that the effect of in-migration rate on change in transit ridership is moderated by the magnitude of change in transit supply. We measure change in transit supply in terms of the change in per capita miles of transit service, denoted by: $\Delta PERCAPTRANSITMI = PERCAPTRANSITMI_{2013} - PERCAPTRANSITMI_{2008}$ (mean=0.28, SD=4.37, min=-27.66, and max=43.07). Table 7 shows that in urban areas where per capita miles of transit service increased (*ΔPERCAPTRANSITMI*>0), the association between in-migration rate of recent immigrants and change in transit ridership is positive. Conversely, places where per capita miles of transit service decreased significantly (*ΔPERCAPTRANSITMI* <-2.8, or in the bottom 10th percentile), the association is negative. Note that we use two new dummy variables *ΔPCT_POSITIVE (*=1 when *ΔPERCAPTRANSITMI*>0; 0 otherwise) and *ΔPCT_LOW (*=1 when *ΔPERCAPTRANSITMI*<-2.8; 0 otherwise) and interact them with the *MIG* variable in the regression model.

Table 7: OLS regression model of change in transit ridership (2008-2013) – role of transit investments

Variable	Coef.	P>t
MIG	18.84	0.85
ΔPCT_POSITIVE	1.03E+05	0.86
ΔPCT_LOW	1.81E+05	0.83
MIG x ΔPCT_POSITIVE	282.86	0.04
MIG x ΔPCT_LOW	-184.82	0.03
ΔIMM10	67.67	0.00
ΔPOVERTY	2.27	0.87
ΔCOLLEGE	3.09	0.81
ΔUNEMP	42.61	0.25
ΔWHITE	-17.31	0.07
ΔROADMI	1.63E+08	0.25
ΔTRANSITMI	-0.08	0.68
ΔPOP	7.67	0.51
POP2008	-4.21	0.02
TRANSITMI2008	0.16	0.00
Constant	3.71E+05	0.44

Adj. R-square=0.75; N=294

The results suggest that recent immigrants positively contributed to transit ridership, on average, in urban areas that underwent significant transit investments leading to service increases that were more than proportional to population increases. It seems that those cities were on average able to attract incoming recent immigrants into transit efficiently. It is possible, although we cannot test using the datasets used in this study, that transit agencies in some cities of those cities increased service over the study period by observing (or in response to) in-migration trends of recent immigrants – and thereby boosted system-wide ridership. Sample cities in this group include Tucson, AZ, Columbus, OH, Champaign, IL, Raleigh, NC, New Orleans, LA, and Philadelphia, PA-NJ-DE-MD urban areas. Conversely, cities that lost per capita transit service and therefore riders despite high inflow of recent immigrants include, for example, Atlanta, GA, Urban Honolulu, HI, Austin, TX, Miami, FL, and Portland OR-WA urban areas.

5. LIMITATIONS

Our paper provides the first empirical evidence connecting the effect of in-migration of recent immigrants on the change in transit ridership across US urban areas using Census (ACS) sociodemographic and migration data, and federal (NTD) transit ridership data as the principal data sources. Data limitations, however, have governed the scope of our study, model specification, and consequently interpretation of results and takeaways for policy. Nonetheless, this research exercise has enabled us to suggest ways to improve existing surveys and conduct additional research to provide more useful information for urban transportation policy making.

Year-to-year out-migration rate cannot be determined from the ACS microdata, and hence the effect of out-migration of recent immigrants on ridership change cannot be directly modeled. Although we selected variables and specified our model strategically so that out-migration is indirectly captured, assignment of person-weights to each mover in the ACS sample in both the current year (available) and the previous year (unavailable) can help overcome this limitation.

The effect of in-migration of recent immigrants on transit ridership change can depend on the characteristics of the incomers, including: 1) personal-household-employment characteristics before and after the move, 2) prior mode use, auto ownership, and urban environment, 3) socioeconomic and transportation environments of the foreign country of origin, etc. Therefore, our average effect does not capture underlying heterogeneities. Explicit accounting of these differences can make demand analyses and forecasts better. Existing national datasets in the US provide limited information.

We have not accounted for two types of endogeneities in our model, which can render some of the observed effects to be associative rather than causal. First, transit service change over the study period in a given urban area could have been made in response to ridership change. Second, ridership change in an urban area, a reflection of transit investments all else equal, could have affected location choices of some recent immigrants who moved. The former is possible, but addressing the circular connection is not required in the present research context. The latter is plausible but difficult to address because of lack of data on how movers make relocation decisions. Even if causality cannot be reliably established, the associative effect itself is useful for transit agencies as they plan for the future considering various urban trends.

6. CONCLUDING COMMENTS AND IMPLICATION FOR TRANSPORT POLICY

This study was the first to explicitly model the relationship between transit use and migration flows. The results confirm various findings in the literature concerning the determinants of public transit use, including the positive role of increases in the transit supply, population density, and the size of the immigrant population. We also find that new immigrants from a foreign country are more likely to take public transit than other populations, and that immigrants migrating from other parts of the United States are less likely to take public. This fact may be due to the resources acquired by inter-metropolitan migrants during their stay since migrating from a country of origin.

The implications for transit agencies are that immigrant populations are not monolithic in their propensity to use public transit. Not only are recent immigrants more likely to use public transit, but those recent immigrants that have arrived directly from another country are even more likely. This suggests a focus on such data available from the American Community Survey can help identify trends in travel demand that are driven by the location of immigrant flows and the time that an immigrant has been in the county.

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