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# LOCAL ECONOMIC CONDITIONS AND THE NATURE OF NEW HOUSING SUPPLY

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## Abstract

We explore the effects of local economic conditions on the type and size of newly constructed housing units in a city. Exploiting the metro area samples of the American Housing Survey from 1984 to 2004, we find that positive local income shocks (i) increase a city's share of multi-family housing in new construction and (ii) trigger the construction of smaller units. These responses are driven by migration. Our findings are consistent with a modified open monocentric city model that more realistically assumes land is available for conversion into new housing throughout the city.



**JEL classification:** R11, R21, R31, R52.

**Keywords:** Local economic conditions, open monocentric city model, land conversion, housing supply, housing type, housing consumption, land use regulation, migration.

# 1 Introduction

The composition of the existing housing stock in a city does not only determine the character of a city – its skyline – but affects a host of important attributes such as the city's household composition, its homeownership propensity, or its population density and hence, conceivably, productivity. As a consequence of the durability of housing, if the nature of the existing housing stock is important for a location's fortunes, then so should be the nature of *new* housing supply. Housing units built in a certain period – reflecting the demand and supply conditions at that point in time – last for several decades and sometimes centuries, continuing to exert positive and negative externalities associated with their characteristics.

Little is known to date about the determinants of the characteristics of new housing supply. In particular, little is known about how local economic shocks affect the nature of the newly built housing stock. In this paper we explore empirically how local economic conditions, at the time when new housing developments are being planned and built, affect the composition of new housing supply, i.e., whether new housing is of single-family or multi-family type and the size of newly built housing units in square feet.

To guide our empirical analysis, we first derive predictions from a modified version of the open monocentric city model. The model assumes that some land – either already developed or still undeveloped – is available for conversion into new housing *throughout* the city. Unlike the standard open monocentric city model that implicitly assumes that the city is rebuilt from scratch in each period, our implicit assumptions are that (i) the building stock is durable and depreciates slowly and at some point can become obsolete and is replaced and (ii) during the outward development process of a city, some pockets of land may remain undeveloped but are subject to infill at a later stage.[1]

Formally, we assume that some (re)development takes place all over the urban area and that the share of land that is converted into new housing may vary between more central and more peripheral locations, consistent with stylized facts derived from the American Housing Survey (AHS) and empirical evidence presented by e.g. Brueckner and Rosenthal (2009), Burchfield *et al.* (2006), or McDonald and McMillen (2000). We also assume that conversion rates depend on local income levels, capturing the idea that the opportunity cost of keeping land open or of not redeveloping existing defunct property stock increases with local income. To keep our model tractable and provide a simple benchmark case for our empirical analysis, we assume that a local income shock brings about the same percentage change in the conversion rate for both types of housing. Given these assumptions, our model predicts that positive local income shocks (i) increase the share of multi-family housing in the city and (ii) cause the construction of smaller units.

[1] Land may not be uniformly developed at each distance to the center for a number of reasons: Undeveloped land varies in soil quality and topography and consequently in the development cost. Heterogeneous owners of undeveloped land may differ in the reservation price, at which they are willing to develop their land. Undeveloped land also possesses a real option (to wait and develop at a later point in time) and the valuation of this option may too vary across heterogeneous owners. Certain sites may be off limits to developers because of zoning, historic preservation or other types of government intervention. Finally, certain undeveloped sites may be awkward and costly to develop because of their unusual shapes so not viable at the time of general outward development. But these sites may become viable for development at a later stage because of increased demand pressures.

The economic intuition for these predictions is, at a first glance at least, not straightforward: Housing is a normal good and hence one might expect that local economic booms induce the construction of more single-family units and of larger units, whereas local economic crises might have the opposite effect. However, this view ignores the important insight that, in response to a local shock, residents can relocate across cities and that such relocation may equalize differences in living standards across cities. Consider an open monocentric city, where households can relocate freely between cities and housing is assumed to be perfectly malleable. In such a setting, a positive city-level income shock temporarily increases the utility of its residents compared to the utility of the outside option. This attracts more migrants into the city and thereby increases land and house prices and reduces the quantity of housing consumed *at each distance from the Central Business District (CBD)*. In the new spatial equilibrium, household migration exactly equalizes the utility of households across cities.

The higher land prices also invoke a substitution away from land to capital in the housing production process, implying more capital-intensive housing in spatial equilibrium, again, at each distance from the CBD. But this does not necessarily imply that in the metro area as a whole the capital intensity of new construction must increase. This is because a positive income shock also generates more new housing development with low capital intensity at the urban fringe. Hence, the aggregate effect is a priori ambiguous. One contribution of our theoretical analysis is that we derive predictions *at the aggregate metro area level*: a positive income shock increases the share of construction of multi-family housing in a city and, on average, leads to the construction of smaller units.

In order to test our model predictions, we turn to the AHS. We gather information on over 700,000 housing units, including their year of construction, from all AHS Metropolitan Statistical Area (MSA)-samples between 1984 and 2004. A key assumption in our empirical strategy is that after a unit is built, the type (single-family vs. multi-family) and the size of the unit (floor area in square feet) remain unchanged for a few years. In our baseline specification we assume that the fraction of units that increase their floor size within the first ten years after construction is small. (In a robustness check we narrow down this time window to five years.) Building on this plausible assumption, measures of the type and size of new housing supply can be created for each MSA by taking means conditional on the year of construction. These measures are subsequently related to one-year lagged MSA-level income per capita and construction industry-wages in a fixed effects panel data analysis that fully controls for time-invariant spatial heterogeneity and trends at the national level. Ultimately, we end up with a panel dataset consisting of 47 MSAs and nearly 2000 observations.

The fixed effects baseline estimates provide strong support for our two model predictions; that positive local income shocks are associated with the construction of more multi-family housing and smaller units. To illustrate the magnitude of the effects, consider the case where income is constant everywhere in the country except in one metro area. (Alternatively we could assume that income grows more strongly in the focal metro area than nationally without loss of generality.) This metro area receives a one-time productivity shock that raises local income by 10 percent from year $\tau$-1 to year $\tau$. From year $\tau$ onwards income remains again unchanged. Our baseline result implies that such a shock *permanently* raises the share of multi-family units in *new* construction by 6.4 percentage points from period $\tau$ onwards.

To put this into context, say a city consists of 1 million housing units, 30 percent of which are of the multi-family type prior to the shock. Say 40,000 new units are being built each year in equilibrium, either replacing existing run-down housing stock, converting brownfield land, or being built on open land. This implies that prior to the shock 12,000 new multi-family units are built each year. Our estimates imply that the positive income shock will induce a permanent increase in construction of multi-family units from 12,000 to 14,560 in each subsequent period. A one-time positive income shock of 10 percent will thus after 10 years generate an additional 25,600 multi-family units and correspondingly fewer single-family ones.[2] While this illustrative example suggests that our estimated annual effects are not enormous, because the effects are permanent in nature and cumulative over time, they are quantitatively quite meaningful in the long-run. Moreover, our derived quantitative effects are consistent with the observation that the built housing stock in a city typically only changes gradually and slowly due to the extreme durability and thus slow depreciation of housing.

Our other base line results are similarly meaningful and plausible: A one-time 10 percent increase in local income, holding national level income constant, reduces the square footage of an average single-family house by 119 square feet (4.8 percent) and that of an average multi-family unit by 350 square feet (16.8) percent.

The mechanism in our theoretical setting that drives our findings is *migration across cities* in response to local income shocks. Using MSA-level migration data, we first provide evidence that in MSAs with lax land use regulation[3] positive local income shocks are indeed fairly strongly positively related to incoming migration into MSAs, and, that local employment growth is a good proxy for incoming migration. We then use a Bartik (1991)-type identification strategy, to show that local employment growth in response to changing local economic conditions (i.e., local labor demand shocks) can explain our empirical findings in a causal sense. Put differently, our instrumental variable estimates are indicative that local labor demand shocks, via inducing migration, *cause* changes in the composition of new housing supply.

Our paper makes two contributions to the existing literature. First, we shed light on the link between local economic conditions and the nature of new housing supply. Second, we propose a modified version of the open monocentric city model that explicitly distinguishes between multi-family and single-family units and reconciles the theoretical framework with observed regularities with respect to new residential development.

Our paper is structured as follows. In Section 2 we briefly review the related literature. Section 3 provides a theoretical framework that guides our empirical analysis. Section 4 describes the data and empirical strategy in more detail. In Section 5 we present results. Conclusions are offered in the final section.

[2] This back of the envelope calculation assumes that the income differential does not converge. Blanchard and Katz (1992) however document a tendency for regional income differentials to converge. To the extent that there is convergence over time, the effect will get smaller over time and the aggregate effect will be less pronounced.

[3] Our theoretical setting implicitly assumes that the nature of new housing supply is purely driven by market forces. It thus *abstracts from land use controls*, which may hamper MSA-level labor market adjustment through limits on the quantity of newly supplied housing units (Glaeser *et al.*, 2006; Saks, 2008).

## 2 Background and related literature

Our theoretical framework builds upon the seminal work on the monocentric city model (MCM) by Alonso (1964), Mills (1967, 1972), and Muth (1969), and, in particular on the Muth model that incorporates housing construction (see e.g. Brueckner, 1987; Fujita, 1989). One limitation of the standard MCM is the extreme assumption that the conversion rate of the existing housing stock into new stock is 1 throughout the city; housing capital is assumed to be perfectly malleable. Put differently: the city is rebuilt from scratch in every single period. The reverse extreme assumption is perfectly durable housing capital. In such a setting new construction of (single-family) housing can only take place within small concentric rings at the edge of the city, where agricultural land is converted into housing.

Neither extreme squares well with empirical evidence. The American Housing Survey (AHS) – the data source underlying our empirical analysis – for example reveals that the rate of new construction relative to the existing housing stock is only about twice as high in the suburbs than in the city center. Brueckner and Rosenthal (2009) document the percentage of housing stock in 2000 that is under 10 years old as a function of distance to the city center. While they find that newer housing is disproportionately located in suburban areas, they also document that except for the largest cities, the percentage hardly varies between 10 and 40 miles from the city center. In a similar vein, Albouy and Ehrlich (2013) document the geo-locations of transactions of undeveloped land for Chicago, Houston, Los Angeles, and New York. The transactions are not found to be more frequent at the boundaries of urban areas.

New residential construction is not confined to greenfield sites but also occurs on brownfield land or to replace older housing, typically at higher density. Such conversion takes place especially in older parts of cities, in or close to the center (Brueckner and Rosenthal, 2009). In these areas there is often a considerable amount of brownfield development as previously industrial sites are converted into housing.[4]

Both, the standard MCM and models with perfect durability ignore the important fact that the housing stock is durable but depreciates (see e.g., Brueckner, 2000). In models with durability and depreciation the possibility that multi-family and single-family housing are not strictly separated in different concentric zones, as is the case when housing capital is malleable, arises naturally. Upward sloping and even discontinuous building height contours can result. Moreover, in the face of uncertainty, urban development does not necessarily occur from the city center outwards. Instead there may be leapfrog development, in which some land within the city boundary remains vacant (Capozza and Helsley, 1990).[5]

The theoretical prediction of leapfrogging is consistent with an important stylized fact: pockets of land are open and developable within most US cities. Remote-sensing information on land use dynamics points to the importance of infill: Burchfield *et al.* (2006) document that scattered

[4] Geocoded data on residential development by year of construction for New York City – plotted in Figure A1 (Web-Appendix A) – allows us to illustrate this stylized fact. Recent residential development – between 2000 and 2017 – is scattered throughout the metro area, including the most central parts of the city. Whereas residential development in the most central parts of the metro area is almost exclusively of the multi-family type, in more peripheral areas it is mostly single-family housing.

[5] An alternative explanation for leapfrogging is that people have a preference for housing that is close to open space (Turner, 2005). See also Burchfield *et al.* (2006) for a summary of the related theoretical literature.

and incomplete residential development is the rule rather than the exception. Suburban developments tend to leave substantial amounts of space open, which may become an important source of new construction within the urban fringe at a later stage. They also provide evidence that areas that were about half developed in 1976 were subject to the most intense residential development between 1976 and 1992. In a similar vein, McDonald and McMillen (2000) examine the location of residential and commercial real estate development in the Chicago metro area between 1990 and 1996. Their findings indicate, among other things, that new residential development did not just take place at the edge of, but, throughout the Chicago metro area, forming clusters of their own between major highways.

In our model we attempt to reconcile the open monocentric city framework with the stylized facts that (i) housing is durable but depreciates and is ultimately redeveloped and (ii) there are pockets of open land even close to the city center and these pockets often are developed at a later stage. Specifically, we assume that in each period some fraction, $\alpha$, of all land (developed *or* undeveloped) is redeveloped *or* newly developed, respectively. We assume that this fraction depends on the local income in the city (relative to the national average) – reflecting opportunity cost considerations – and we allow it to differ between the core and the periphery of the city.

Rosenthal and Helsley (1994) developed a structural model of redevelopment in the context of a monocentric city and show that the probability that a property will be redeveloped depends on the ratio of the price of the current house and that of undeveloped land. They assume demolition costs to be zero, but the logic of their model implies that conversion rates decrease in such costs. A positive local income shock tends to make all existing housing suboptimal, which implies a decrease in the value of existing property relative to that of undeveloped land, and hence an increase in redevelopment probability throughout the city. Their approach can also be applied to conversion from non-residential (e.g. agricultural) to residential use, which would imply a model for such conversion rates with similar characteristics. Our main assumptions about conversion rates thus appear to be consistent with a structural model of redevelopment behavior in urban areas.

Dye and McMillen (2007) follow Rosenthal and Helsley (1994) in their empirical study of teardowns in Chicago and find that the specific characteristics of properties affect their teardown-probability. Since the presence of idiosyncratic characteristics is difficult to include in stylized models of the urban economy, our assumption that a constant share of land is redeveloped in each year appears to be a reasonable approximation, at least for the purposes of the present paper.

Overall, the evidence discussed above suggests that a positive local income shock does not lead to a sharp discontinuity in the conversion rate at the boundary of the city. A gradual decrease towards zero near the urban fringe of the city might be a more appropriate formulation. However, in the interest of simplicity and to stay close to the conventional monocentric model we adopt a formulation in which the conversion rate is constant up to the new (post-shock) boundary of the city.[6]

[6] A gradual decrease in the conversion rate nearer the edge of the city would complicate the derivations, but should strengthen our conclusions.

Our paper also relates to the housing supply literature. This literature has either focused on new housing supply in units or on the "volume" of residential investment at the national level, thus aggregating all composition and quality aspects into one single variable and ignoring the spatial dimension (e.g., DiPasquale, 1999). Studies in the former category generally focus on the single-family sector, thus ignoring the supply of multi-family housing. Heterogeneity within the single-family-sector is ignored as well, even though the hedonic literature suggests that the value of single-family housing units varies widely depending on their attributes.

Another strand of the urban economics literature considers the link between the housing stock (supply) and the corresponding household composition. Affluent households in the United States tend to sort into communities that predominately consist of spacious and expensive single-family homes. As a consequence, such communities have higher local property tax revenue per capita and therefore can offer better local public schools and other local public services. In contrast, low income households prefer to sort into inexpensive lower "quality" housing in decaying areas (Rosenthal, 2008)[7] or into areas where government programs have contributed to "affordable housing" (Baum-Snow and Marion, 2009). Minimum lot size restrictions, imposed by affluent households in order to keep less well-off households at bay, tend to reinforce such sorting by income based on the underlying built environment.

Our study also ties into the literature on the determinants of homeownership. This is because single-family (multi-family) housing is strongly positively associated to homeownership (renter-occupation). According to the national AHS, only about one in seven multi-family units in the US is owner-occupied. Roughly the reverse is the case for single-family units. Coulson and Fisher (2014) and Linneman (1985) provide different theoretical explanations, suggesting that the housing type causally affects the optimal tenure. Hilber (2005 and 2014) provide empirical evidence for the US and Europe, respectively, that even conditional on location and occupant characteristics, the housing type is the key determinant explaining the homeownership status of a property. Hence, if a positive local income shock increases the share of multi-family housing in new construction, all else equal, this implies a reduction of the homeownership propensity in the city, possibly offsetting the positive direct effect of income on the propensity that a household owns.

To the extent that the construction of more multi-family housing indeed causes a decrease in the homeownership propensity, our study has direct relevance for the voluminous literature on the social and economic consequences of homeownership. This literature suggests that homeownership is linked to housing maintenance (Galster, 1983), investment in local public goods such as public schools (Hilber and Mayer, 2009), investment in social capital (DiPasquale and Glaeser, 1999; Hilber, 2010; Hoff and Sen, 2005), labor market outcomes and entrepreneurship (Blanchflower and Oswald, 2013; Bracke *et al.*, 2018; Harding and Rosenthal, 2017; Oswald, 1996), or local political participation and land use regulation (e.g., Ahlfeldt and

[7] Glaeser and Gyourko (2005) show that the same mechanism is at play at the geographical level of MSAs. In declining cities where labor demand is weak, house prices are low but through decay, the housing stock adjusts only slowly to these conditions. This leads to a sorting process in which people with lower human capital levels stay in declining cities in order to benefit from relatively cheap housing.

Maennig, 2015; Fischel, 2000 and 2001; Hilber and Robert-Nicoud, 2013; Ortalo-Magné and Prat, 2014).

# 3 Theoretical framework: Single and multi-family housing in a monocentric city

In this section we present a simple modified version of the open monocentric city model to guide our empirical analysis. Our model differs from the standard model in that (i) we explicitly distinguish between two types of housing; single-family and multi-family, (ii) we assume that in each period a small fraction of all land – already developed or undeveloped – is converted into new residential development throughout the city, and (iii) the conversion rate can differ between more central and more peripheral locations.

The literature about the monocentric model with housing production (see, for instance, Brueckner, 1987) typically assumes that the structure to land ratio is continuous, whereas we take it to be discrete and distinguish only two housing types. The reason for this is that in our empirical work we make heavy use of the qualitative distinction between single-family and multi-family housing and we prefer to have a theoretical model that corresponds closely to the regressions. However, the mechanism through which income shocks translate into changes in housing supply in our model is entirely similar to that in the more conventional set-up.

Our model is static in that we compare the income $y$ in a focal city with the outside income $y^*$. In the steady-state $y_0$ equals $y^*$. Our model predicts that if $y_0$ increases to $y$ such that $y>y^*$, this will increase the share of multi-family housing in new construction and will lead to new construction of smaller units. By implication, if income *grows* more strongly in our focal city than in the nation as a whole, the prediction remains that the share of multi-family housing in new construction will increase in the focal city and newly constructed units will be smaller compared to the counterfactual with equal income growth rates.

The empirical implication is that – absent of land use regulations and other restrictions – cities that receive a positive income shock, controlling for national-level shocks and city specific unobserved characteristics (captured through the inclusion of year and metro area fixed effects), will observe an increase in the share of construction of capital-intensive multi-family housing and will see smaller units being built.

We proceed discussing the various components and specific features of our model.

## 3.1 Demand for floor space

We consider a monocentric city with a homogeneous population and two types of housing: single-family (*sf*) and multi-family (*mf*). The generalization to an arbitrary number of *mf* dwelling types is discussed in Web-Appendices B and C. Utility is characterized by the function $u = u(c, s, i)$, where $c$ is a composite consumption good, $s$ is the amount of floor space, and $i$ (= *sf, mf*) indicates the dwelling type. Households can switch between dwelling types, however, in each period they can only inhabit one type. Utility is continuously differentiable, increasing and quasi-concave in $c$ and $s$. Floor space in *mf* units is assumed to be *inferior*, which is reflected in the assumption that $u(c,s,sf) > u(c,s,mf)$ for all $c$ and $s$.[8] This assumption may appear to be at odds with the existence of luxury apartments with nice views that would be preferred over modest single family houses by many. However, note that we have assumed a homogeneous population to keep our model tractable, whereas the luxury apartments are inhabited by families that are probably substantially different in income and other characteristics from those living in modest single family houses. Within the housing segments that are relevant, it seems reasonable to assume both groups prefer single-family over multi-family housing.

The inverse of the utility function with respect to $c$, $z = z(u,s,i)$, may be interpreted as the amount of composite consumption good that has to be offered to a household that lives in a housing unit of type $i$ with an amount $s$ of floor space, in order to guarantee utility level $u$. Its partial derivative with respect to $s$ equals minus the willingness to pay for floor space. By assumption, this willingness to pay is always larger for floor space in $sf$ housing: $-\partial z(u,s,sf)/\partial s > -\partial z(u,s,mf)/\partial s$ for all $u$ and $s$.

For simplicity we assume that all housing is rented and that rent levels adjust fully to changes in market conditions.[9] Let $p_i$ denote the rent of a square unit of floor space in housing of type $i$. Normalizing the price of the composite consumption good to unity, we can then write the household budget constraint as $y - tx = c + p_i s_i$, where $y$ denotes income, $x$ is the distance to the CBD, and $t$ the transportation cost per distance unit. Equalization of utility within the city requires that $p_i$ is equal to the bid rent for floor space. This is the maximum amount a household can afford to pay for a unit of floor space, while still being able to reach utility level $u$:

$$\Psi(u,x,y,i) = \max_s \frac{y - tx - z(u,s,i)}{s} \tag{1}$$

For the floor size $s$ that solves this optimization problem, it holds that:

$$-\frac{\partial z(u,s,i)}{\partial s} = \frac{y - tx - z(u,s,i)}{s}. \tag{2}$$

This equation states that the marginal willingness to pay for floor space equals the amount of money per unit of floor space available to a household that has to reach utility level $u$. Hence, the equilibrium rent level for each type of housing is equal to the corresponding marginal willingness to pay. These bid rents are decreasing convex functions of the distance to the city center, and it can be shown that the bid rent function for floor space in $mf$ units lies below that for floor space in $sf$ units.[10]

[8] We treat floor space in *mf* and *sf* units as inherently different and, specifically, *mf* floor space as inferior because apartments are much more prone to negative externalities such as noise from neighbors. In the extension to an arbitrary number of *mf* housing types in Web-Appendices B and C, we also assume that the floor space of *mf* units is inferior, but we assume that consumers are indifferent to building height within the *mf* sector.

[9] Owner-occupied housing could be dealt with by modeling user costs rather than rents. However, in that case we would need to take into account the wealth effect of house price changes (capital gains), which can be ignored in the standard setting with absentee landowners.

[10] Suppose this is not the case. If the bid rents for both types of housing would be the actual prices for floor space, the budget line for *mf* housing would lie entirely below that for *sf* housing. With the indifference curve referring to *mf* housing lying entirely above that for *sf* housing the consumer would be unable to reach the given utility level on which the bid rents are based in both dwelling types. This implies a contradiction.

### 3.2 Demand for developable land

Developers build and rent out *sf* and *mf* units in perfectly competitive markets, implying that all their profits disappear into bid rents for residential land. The two types of housing are distinguished by the number of square feet of floor space per unit of land. In our model, buildings containing *mf* units, have $F$ square units of floor space per square unit of land, whereas *sf* housing has one square unit of floor space per unit of land. We assume that the floor size $s$ for new housing is determined by the prevailing market circumstances at the time of construction and remains fixed at that level until the building is demolished. Our data indicate that the average number of floors roughly equals 2 in *sf* housing and 3 in *mf* housing. Our model would capture this proportion by assuming an $F$ of 1.5. However, *sf* housing often uses more land for gardens instead of floor space relative to the *mf* sector, implying $F > 1.5$.

Profits per square unit of land, $\pi_i$, are given by either $\pi_{mf} = F\Psi(u,x,y,mf) - C_{mf} - p_l$ or $\pi_{sf} = \Psi(u,x,y,sf) - C_{sf} - p_l$, where $p_l$ denotes the land rent and $C_{mf}$ and $C_{sf}$ the (annualized) construction costs per square unit of land. While this is not strictly necessary for our model, it is conventional to assume that $C_{mf} > C_{sf}$.[11] Setting these profits to zero, we obtain the bid rent functions for land:

$$\Pi(u,x,y,sf) = \Psi(u,x,y,sf) - C_{sf}\,, \tag{3a}$$

$$\Pi(u,x,y,mf) = F\Psi(u,x,y,mf) - C_{mf}\,. \tag{3b}$$

Developers choose floor sizes optimally, implying that they choose the floor sizes that solve the consumer problem (1).

Whether *mf* or *sf* units are constructed is determined by the highest bid; *mf* units are built when $\Pi(u,x,y,mf) > \Pi(u,x,y,sf)$, or:

$$\Psi(u,x,y,mf) > \frac{\Psi(u,x,y,sf) - C_{sf} + C_{mf}}{F}\,. \tag{4}$$

We would expect to see *mf* housing close to the city center and single-family housing in the suburban ring around the center. This pattern emerges when inequality (4) is satisfied in the center and when, at the intersection of the two sectors, the bid rent curve for land in *mf* housing is steeper than the bid rent curve for land in *sf* housing, or equivalently, when the profits associated with the construction of *mf* housing decrease faster than the profits associated with the construction of *sf* housing:

$$\frac{\partial\Pi(u,x,y,mf)}{\partial x} < \frac{\partial\Pi(u,x,y,sf)}{\partial x}\,. \tag{5}$$

To provide some intuition for this condition, we substitute the expressions for bid rents for floor space (1) into condition (5), and use the 'Muth condition' that $\partial\Psi(u,x,y,i)/\partial x = -t/s_i$. We

[11] The extended model discussed in Web-Appendices B and C makes the realistic assumption that the construction cost per unit of land is convex in the number of floors.

can now rewrite this condition as $F/s_{mf} > 1/s_{sf}$ . Hence, (5) is equivalent to the requirement that the household density in the area with *mf* housing adjacent to the boundary between the two sectors exceeds the corresponding density in the area that contains the *sf* housing. If we assume that this mild condition holds, multi-family housing will be nearer to the center and single-family housing nearer to the edge of the city. In what follows we assume (4) and (5) to be true.

**Figure 1** illustrates bid rents for floor space (left) and land (right) in both sectors as a function of distance to the CBD. The black and grey lines refer to *mf* and *sf* housing respectively. Bid rents for floor space are highest for *sf* housing everywhere, yet close to the center, building *mf* housing is more profitable because of the higher unit density. The boundary $x^*$ between the *mf* and *sf* sectors occurs at the intersection of the two bid rent curves for land, where equation (4) is solved with equality. Bid rent curves referring to the *sf* type are therefore dashed to the left of $x^*$, whereas *sf* units are built beyond this boundary. The reverse is the case for *mf* units; bid rent curves referring to the *mf* type are dashed to the right of $x^*$, *mf* units are built to the left. The urban fringe $x^b$ is determined by the condition that land in *sf* housing is equally valuable as land in agricultural use. The bid rent curve for land is steeper to the left of $x^*$ than to the right of it.

**Figure 2** illustrates the effect of an increase in income. It shows the bid rent for land in highest use, i.e. the bid rent referring to the housing type that maximizes land rents at this location. The dashed line in this figure refers to the situation before and the solid line to the situation after the increase in income. An increase in local income shifts all bid rent curves outwards by the same amount, and hence, the boundary between the *mf* and *sf* sectors and the urban fringe shift outwards by this amount as well. More precisely, if income rises by an amount $\Delta y$, then bid rents shift outward by an amount $\Delta y / t$. This is seen for the bid rents for floor space by revisiting expression (1):

$$\begin{aligned} \Psi\left(u, x+\Delta y/t, y+\Delta y, i\right) &= \max_s \frac{y+\Delta y - t\left(x+\Delta y/t\right) - z\left(u,s,i\right)}{s} \\ &= \Psi\left(u,x,y,i\right), \end{aligned} \tag{6}$$

and a similar derivation can be made for bid rents for land in expressions (3a) and (3b). As a consequence, the nature of new housing supply in terms of type and floor space shifts outwards from the CBD by an amount $\Delta y / t$ as well.

### 3.3 Conversion rates and testable predictions

In this subsection we derive predictions about the nature of new housing supply in situations where housing construction takes place everywhere in the city, consistent with stylized facts from the AHS and empirical evidence discussed in Section 2. For ease of exposition we assume that in the part of the city where *mf* housing is constructed, a percentage $\alpha_{mf}$ of all land (developed or open and developable) is converted into new housing in each period, while in the part of the city where *sf* housing is constructed, a percentage $\alpha_{sf}$ of all land is converted into new development in each period. The new development can be thought of as either redevelopment of defunct housing stock or as construction of new housing on greenfield or brownfield land.[12] The two percentages $\alpha_{sf}$ and $\alpha_{mf}$ may be different, and in particular $\alpha_{sf}$ may be larger than $\alpha_{mf}$, as suggested by the empirical evidence discussed in Section 2.

A local income shock implies that the city expands (the boundary shifts from $x^b$ to $x^{b\prime}$) but that only a fraction $\alpha_{sf}$ of the new urban land will be converted immediately. This is in line with the empirical evidence that shows that large amounts of undeveloped land remain in suburban areas.

We assume that the conversion rates $\alpha_i$ ($i = mf, sf$) are increasing functions of local income $y$: $\alpha_i = \alpha_i(y)$. The rationale for this assumption is that the opportunity cost of keeping land open or in existing (no longer-optimal) use increases with income. We further assume that $\alpha_i$ is always positive, which means that there will always be some new construction, even if local income (growth) is below the national income (growth). This assumption is consistent with stylized facts; there has been some new construction of *mf* and *sf* housing even during Detroit's worst crisis period.

Our predictions are derived under a further simplifying assumption on the relationship between income and conversion rates. This *proportionality assumption* is that a local income shock brings about the same percentage change in the conversion rate for both types of housing, i.e. $\alpha_{mf}(y) = k\alpha_{sf}(y)$ for some positive constant $k$. It reflects the fact that we do not have a clear prior on whether an income shock should bring about a larger or smaller relative change in the conversion rate for single-family as opposed to multi-family housing. There are several possible considerations. A potentially important one is that an increase in income should raise the demand for living close to the center through its impact on the value of commuting time, which would suggest a larger conversion rate of multi-family housing. Another is that conversion of multi-family housing may be more difficult due to higher demolition costs. Rosenthal and Helsley (1994) argue that demolition costs are small, which might suggest that the forces that increase conversion rates close to the center are strongest. If this is true, our proportionality assumption would lead to under-estimation of the true effect of an income shock on the share of *mf* housing in new construction. The proportionality assumption also has the significant advantage that it helps us to keep the model tractable and provides a simple benchmark case for our empirical analysis.

The assumption is illustrated in **Figure 3**, where the black bars refer to the situation before and the grey bars to the situation after the change in income. The figure illustrates the conversion rates when $k$ is assumed to be markedly smaller than 1, i.e. the bulk of conversion occurs in the *sf* sector. The proportionality assumption then imposes that the rise in conversion rates after an increase in income is also markedly larger in the *sf* sector.

[12] The conversion rate can be defined formally as the amount of land that is either developed or redeveloped in a certain area in each period relative to the total amount of all land (developed and open and developable) in the area, both at a given distance from the CBD. So the denominator does not distinguish between open land, already developed land (with depreciating capital), or brownfield land. In reality, more redevelopment or brownfield development takes place in more central locations. In our model, since empirically we cannot distinguish between redevelopment and development of open land, we are only interested in the nature of new housing supply. Hence, in our model there is also no need to distinguish between the fraction $\alpha_i$ that is due to redevelopment of depreciated stock or due to new development of open land.

The empirical evidence discussed in the background section suggests that $k < 1$. This is also consistent with theoretical reasoning: Since more central locations are more developed than more remote ones and since *mf* housing requires larger sites, fewer suitable parcels of open land will typically be available for construction of *mf* housing. Thus, building new tall apartment blocks in central locations normally implies redevelopment of depreciated properties rather than development on an open plot of land. Such redevelopment is typically highly involved – much more complex and costly than (re-)development of low density *sf* housing in more peripheral locations.

With $k < 1$ there will be a discontinuity in the conversion rate, which arises by assumption, at the boundary $x^*$ (i.e., $\alpha_{mf} < \alpha_{sf}$ ). The discontinuity in the conversion rate implies that a positive local income shock extends the part of the city where the conversion rate is comparably lower, that is, where *mf* housing is constructed (i.e., the boundary between the *mf* and the *sf* sector shifts outwards). To the extent that the conversion rate does not drop in the part of the city where the local income shock causes a change in the predominant housing type (from *sf* to *mf*) but instead remains constant or increases, the positive impact of a positive income shock on the share of new *mf* housing construction will be understated (i.e., Prediction 1 below holds *a fortiori*).

We take the city as it has been developed in previous periods as given and consider what happens in a single period, say period 1, when the local income level is $y$ and utility, which is determined at the national level, is $u^*$. If local income had grown at the national average, its value would be $y^*$. In what follows we refer to the situation, in which $y > y^*$ as a *local increase in income* (holding national income constant).

Let $N_i(y)$ denote the number of newly built units of type $i$ when local income equals $y$. The quantity $N_i(y)$ is computed by multiplying the unit density that solves the consumer problem in (1) with the amount of newly converted land at each distance from the CBD, and then integrating this product over $x$. Our first prediction about the composition of new urban housing supply can be expressed as:

> **Prediction 1**: *A local increase in income, holding national income constant, raises the share of mf units in new construction. More formally; y>y* implies:*
>
> $$\frac{N_{mf}(y)}{N_{mf}(y)+N_{sf}(y)} > \frac{N_{mf}(y^*)}{N_{mf}(y^*)+N_{sf}(y^*)} \tag{7}$$
>
> *and an analogous implication holds for the opposite inequality.*

Proof. See Web-Appendix C.

We already discussed the intuition for this result above: a rise in local income pushes up land prices everywhere in the city, which in turn leads to substitution away from land in the housing production process, that is, from the production of single-family to that of multi-family housing. This also implies that the amount of floor space in newly constructed units decreases with local income. At the aggregate city level, this will be the case if the number of newly constructed units increases more strongly than the amount of land devoted to these new units. To state this formally, let $A_i(y)$ denote the amount of land converted to use for new construction of units of type *i* when income rises from $y_0$ to $y$. The average amount of floor space in newly built units in the *mf* and *sf* sector is then given by $\overline{s}_{mf}(y) = FA_{mf}(y)/N_{mf}(y)$ and $\overline{s}_{sf}(y) = A_{sf}(y)/N_{sf}(y)$ . We can express our second prediction as:

> **Prediction 2**: *A local increase in income, holding national income constant, lowers the average amount of floor space of newly built units for both types. More formally; y>y* implies:*
>
> $$\overline{s}_i(y) < \overline{s}_i(y^*) \tag{8}$$
>
> *and an analogous implication holds for the opposite inequality.*

Proof. See Web-Appendix C.

Our model makes heavy use of the assumption that the city is *open*—in-migration is the driver of our predictions. A closed city would also expand in response to a positive income shock, but the reason would be that households demand *more* floor space. The increase in housing consumption flattens the bid rent curve, and house prices close to the city center will decrease, which makes it unlikely that the share of *mf* housing in new construction increases. See Brueckner (1987) and Wheaton (1977). A counteracting force may be that demand for living close to the city center increases due to the higher value of commuting time implied by the income shock. It is therefore unlikely that our two predictions will hold in a closed city. We consider the empirical validity of our two predictions in Section 5.

# 4 Data, empirical strategy, and baseline specification

## 4.1 The data

Our main outcome measures of interest (share *mf* units, square footage of *sf* and *mf* units) come from the 1984 to 2004 AHS metropolitan area datasets, obtained through HUD User. The US Census conducted these AHS metro surveys annually between 1984 and 1993 and at irregular dates after that. In each year, a different set of MSAs was surveyed. In total, we have information for 47 MSAs and the average number of times that an MSA is surveyed is 3.6. See Table A1 (in Web-Appendix D) for a list of all MSAs and the years they were surveyed in the AHS. For our period of observation, definitions of the variables of interest were overall consistent, though a few minor adjustments

We also obtained Census building permit data for the years 1980 to 2018 from the US Census Bureau. While this dataset allows us to only derive one measure of interest (share *mf* units), it covers all MSAs (rather than just a small subset of larger MSAs) and data is available for a longer period (until 2018). Another advantage of this data is that it measures permit issuance. There should not be nearly as long of a lag between economic conditions and permits than between economic conditions and realized construction.

Our main measure of local economic conditions, the MSA-level income per capita, is derived from the Regional Economic Information System of the BEA.[13] From this dataset we also construct a proxy for construction wages by dividing total earnings in the construction industry by employment. The County Business Patterns (CBP) dataset provides employment data. We use this data to generate two additional variables: the employment growth in the MSA and a measure indicating a labor demand shock (our instrument to identify employment growth). In computing this labor demand shock variable, we follow Saks (2008), using the same underlying data and methodology.

Finally, in order to distinguish MSAs with more stringent land use controls from those with less stringent ones, we use two indices of regulatory restrictiveness. The first index, developed by Saks (2008) is the simple average of six independent surveys conducted during the late 1970s and the 1980s. The method of index construction and the underlying surveys are described in detail in Saks (2008). The second index, the Wharton Residential Land Use Regulatory Index (WRLURI), is derived from a survey conducted during the early 2000s (see Gyourko *et al.*, 2008, and Saiz, 2010, for details). Both indices aggregate information on many different types of land use regulation at the level of municipalities. Since our main data spans the period from the early 1980s up to the early 2000s, we create a new 'combined index' by averaging the Saks and WRLURI indices and we proceed by using this combined index in our empirical analysis.[14] We note however that our results remain virtually unchanged if we use either the Saks index or the WRLURI instead of the combined index.

### 4.2 Empirical strategy, measures of new housing supply, and panel dataset

Our aim is to estimate, at the MSA-level, the impact of local economic conditions on the nature of new housing supply. We do this by regressing MSA-level measures of the type and size of newly built housing units on local income – our focal explanatory variable – and local construction wages. We include the latter variable as a control, to disentangle the effect of our focal variable from labor cost-induced changes in construction costs. The MSA-level measures that capture the type and size of newly built units are derived by averaging each characteristic over the MSA, year of construction, and year of observation (i.e., the survey year). Specifically, we aggregate up the following housing unit level measures from the AHS metro surveys: (1) an indicator that equals 1 if the unit is of the *mf* type and 0 otherwise, (2) the unit square footage if the unit is of the *mf* type, and (3) the unit square footage if the unit is of the *sf* type. Formally, we compute:

$$M^{l}_{t\times MSA\times\tau} = E\left(M^{l}_{h}\,\middle|\,t, MSA, \tau\right), \tag{10}$$

[13] In our baseline estimates we use income per capita rather than wages as our proxy for local economic conditions. This is because income per capita arguably more fully captures demand side shocks in the housing market. However, as a robustness check (not reported), we replicated our analysis using wages (derived from the County Business Pattern) and our findings are essentially unchanged.

[14] For two MSAs we do not observe the Saks index. We use instead the WRLURI. For one MSA we do not observe the WRLURI. For this MSA we use instead the Saks index. Thus we can assign a regulatory restrictiveness measure to each MSA in our sample.

where $M_h^l$ is the value that variable $M^l \ (l = 1,2,3)$ takes for housing unit *h*. We compute the expected value of this variable, for houses that are built in year $\tau$ and observed in an AHS survey of *MSA* in year *t*.

To illustrate our computation procedure, consider the share of housing built in Boston in 1994 that is of the *mf* type. First, we construct a dummy variable that is equal to one if a housing unit *h* is of the *mf* type (as defined in the AHS) and zero otherwise. The AHS metro sample for Boston in year *t* = 1998 (the earliest year after 1994 with a survey for Boston) provides us with information about the characteristics of a sample of housing units in this MSA, as well as the year in which these units were built. Hence, the value of this measure for $\tau$ = 1994 can be obtained by averaging the dummy variable over all housing units in this 1998-sample that were built in 1994. In this particular example, we assume that the housing units that were constructed in 1994 and observed in 1997, did neither change their type nor their square footage during the three-year time window.

The assumption that housing characteristics do not change between the year of construction $\tau$ and the year *t* in which the unit is observed in the AHS survey, is essential to our identification strategy. To ensure that we do not include any housing units that converted from *mf* to *sf* housing and vice versa, we drop all units, for which the AHS reports the construction year $\tau$ as a period of several years. This is the case for older houses; units that were built two decades or more before they are observed in the AHS. We maintain that conversions of units that are younger than 20 years are extremely rare.

Expansions of existing units – especially of *sf* housing – during renovations are more common. However, it would appear to be highly unlikely that such changes in unit size occur during the first ten or even fifteen years after construction. Hence, we include housing units in our analysis if the gap between *t* and $\tau$ is 10 years or less. In a robustness check, reported below, we narrow down this window further to 5 years and we also apply a time window of 10 years and 5 years, respectively, to compute the share of *mf* housing.

By computing measures of the nature of new housing supply according to (10) and subject to the conditions discussed above, we ultimately obtain a panel dataset in which the year of construction $\tau$ constitutes the time dimension and the cluster identifiers are AHS wave-MSA combinations $t \times MSA$. **Table 1** provides descriptive statistics of the resulting panel dataset. Apart from reporting the standard descriptive statistics (mean and standard deviation), we decompose the standard deviation into within and between cluster dimensions. This is relevant for our purposes because all estimates that are reported in the subsequent section are identified on variation within clusters only. Table 1 also reports overall minima and maxima, the number of clusters, and the number of observed cells.

As indicated in Panel A of Table 1, 167 $t \times MSA$ combinations are observed. The means in this panel are sensible and generally straightforward to interpret. 29.8 percent of newly constructed units are part of a *mf* structure. Units in the *sf* sector are on average significantly larger than in the *mf* sector. The average population size of an AHS-metro area is nearly 3 million, that is, our regression sample consists mainly of large MSAs. The variation of variables is usually larger between than within clusters, particularly for income per capita, population and construction wages. Only for the unit square footage of *mf* housing the variation within is larger than between clusters. Panels B and C of Table 1 document the descriptive statistics of our three measures characterizing the nature of new housing supply for the subset of MSA-cells, in which land use regulation is less restrictive or more restrictive than in the average cell. The mean of the three measures is similar for the sub-groups.

### 4.3 Econometric baseline model

Our main results are derived from the following specification:

$$M^{l}_{t\times MSA\times\tau} = C_{t\times MSA} + D_{\tau} + \beta_1 \log\left(Y_{MSA\times(\tau-1)}\right) + \beta_2 \log\left(W_{MSA\times(\tau-1)}\right) + \varepsilon_{t\times MSA\times\tau}, \qquad (11)$$

where $C_{t\times MSA}$ is a fixed effect that is specific to each *MSA* and to the year *t* in which it was surveyed in the AHS, $D_{\tau}$ is a time fixed effect that is specific to the construction year $\tau$, $Y_{MSA\times(\tau-1)}$ is the one-year-lagged per capita income, and $W_{MSA\times(\tau-1)}$ is the one-year-lagged wage level in the construction industry – a proxy for construction costs. The one-year lag in the latter two variables can be expected to capture the natural planning/development lag in the construction process. (We experimented with alternative lags and discuss the findings of these robustness checks below.)

Most metropolitan areas are surveyed several times between 1980 and 2004 (see Table A1 in Web-Appendix D for details), and for each time they appear in the AHS, we enter a separate fixed effect. These fixed effects control fully for all time-invariant heterogeneity at the MSA-level and for any heterogeneity across different AHS samples for the same MSA. The year fixed effects, $D_{\tau}$, control for all national level economic shocks at the time of construction. Hence, we can interpret the estimated coefficient $\beta_1$ as the impact of local income, holding national-level income constant. All remaining heterogeneity is absorbed by the error term $\varepsilon_{t\times MSA\times\tau}$. In the estimation of (11), we account for variation in the number of observations in the AHS on which each $t \times MSA \times \tau$ cell is estimated (and hence the precision of this estimate) by using Weighted Least Squares.[15] Furthermore, reported standard errors are clustered at the level of $t \times MSA$ cells.

## 5 Empirical results

### 5.1 Results for base line specifications

**Table 2** presents results for the base line specification with MSA × survey year plus construction year-fixed effects as in (11). The dependent variables are the share of *mf* housing in new construction (column 1) and the log square footage of *sf* and *mf* housing (columns 2 and 3). The focal coefficient in column 1 reveals that an increase in local income, holding national level income and unobserved time-invariant characteristics at the MSA-level constant, increases the share of *mf* housing in new construction, consistent with our Prediction 1. The relationship between one-year-lagged local income and the share of *mf* housing in new

[15] Time-varying weights are not allowed in a fixed effects estimator. Hence, we use as weights the number of observations, on which each cell is based, averaged over all observations in the same cluster that are based on at least one AHS observation.

construction is not only highly statistically significant but economically reasonably meaningful: an MSA that receives a one-time positive local income shock that raises local income 10 percent more than that at the national level, all else equal, will observe an increase in the share of *mf* housing in new construction of 6.4 percentage points. Local wages in the construction industry, interestingly, do not appear to have an independent effect, neither on the share of new *mf* housing in construction nor on the size of newly built *mf* or *sf* units.

Columns 2 and 3 reveal, consistent with our Prediction 2, that an increase in local income, holding national income constant, reduces the floor size of newly constructed housing units, in both the *mf* and the *sf* sector. Moreover, we find that the adverse effect is stronger in the *mf* sector. We have been unable to derive this as a prediction of our theory, but it is nevertheless interesting to note that it is consistently found in our data. An MSA that receives a one-time positive local income shock that raises local income 10 percent more than national level income, all else equal, will observe a reduction in the unit square footage by 4.8 percent in the *sf* sector and by 16.8 percent in the *mf* sector, respectively.

Since all our specifications reported in Table 2 include construction year fixed effects, all our effects of local income control for income at the US national level. The construction year fixed effects (time dummies) themselves reveal trends in the dependent variables at the national level that are unexplained by our other explanatory variables (i.e., by the MSA-level fixed effects and by the local income per capita and construction wage measures). Table A2 (in Web-Appendix D) reports the suppressed time dummies and reveals that aggregate trends in the dependent variables have been substantial. Notably, conditional on the time-varying and time-invariant local controls, there is a significant and continuous downward trend in the share of *mf* housing in aggregate construction and a significant upward trend in the unit surface of newly built units, particularly in the *mf* sector. The effect of raising income appears to have the opposite affect at the national level compared to the local level, consistent with theory.

### 5.2 Are the results driven by migration across cities?

Our empirical findings above are consistent with predictions derived from an open monocentric city model, in which demand for land is fully elastic as a consequence of costless migration across cities. In other words, our underlying theoretical framework suggests that migration is crucial to understanding why positive local income shocks lead to a greater share of construction of *mf* housing and of smaller units *at the local level*. In order to test for the appropriateness of this interpretation of the estimation results, we next explore whether positive income shocks are indeed correlated with in-migration into MSAs. To do this, we first directly relate MSA-level income growth, during a period with rapid growth (1995-2000), to in-migration from other MSAs during the same time period.[16] Our theoretical framework predicts that the two measures should be positively correlated, at least in MSAs with lax land use regulation. **Figure 4** reveals that this is indeed the case. Per capita income growth at MSA-level and in-migration into these MSAs are strongly positively correlated (with a correlation coefficient of 0.45). In MSAs (in our sample) with tight regulation, the correlation is still

[16] The Census provides MSA to MSA migration statistics for the period between 1995 and 2000. See https://www.census.gov/population/www/cen2000/migration/metxmet/index.html. For a further discussion of the role of internal migration in the United States and other potential data sources see Molloy *et al.* (2011).

positive but no longer statistically significant. While this evidence does not enable us to rule out that alternative channels may partially drive our results, it is suggestive that migration is an important driver of our findings.[17] Note, moreover, that these findings endorse our open city assumption.

We do not have access to annual migration data at MSA-level for our entire sample period and therefore are unable to rigorously *directly* test the migration channel within our empirical setting. We do, however, have data on a close proxy measure for in-migration: local employment growth. As is common in the literature (e.g., Blanchard and Katz, 1992 or Saks, 2008), we therefore use employment growth instead of in-migration statistics.

**Figure 5** indicates that for the period, for which we do have migration data (1995-2000), MSA-level employment growth is indeed closely related to MSA-level in-migration. In MSAs (in our sample) with lax regulation the correlation coefficient is 0.91. In MSAs with tight regulation it is still 0.63. Consistent with this observation, in **Figure 6** we confirm that in markets with fairly lax land use controls income growth is not only fairly strongly positively correlated with in-migration (Figure 4) but also with local employment growth (correlation coefficient of 0.62).

Employment growth or incoming migration are obviously endogenous. Migration depends not only on demand shocks, but also on the extent to which housing supply accommodates such shocks, as has been shown by Glaeser *et al.* (2006) and Saks (2008).

In order to identify the causal effect of employment growth on our measures that characterize the nature of new housing supply, we use an instrument proposed by Bartik (1991) and applied in empirical work, for example, by Blanchard and Katz (1992), Saks (2008), or Hilber and Vermeulen (2016). Specifically, we instrument for employment growth with a "labor demand shock variable" that equals the weighted average of national industry employment growth rates, where weights are equal to the lagged share of an industry's employment relative to total MSA employment. Intuitively, if an MSA has a large proportion of its jobs in an industry that is doing well at the national level, this MSA is predicted to have a high employment growth rate. The underlying idea is that both national industry specific demand shocks and the lagged industry composition of MSA employment are exogenous to local employment growth.

In **Table 3** we test for the alleged migration channel more rigorously. Panel A first reports simple OLS results of the effect of lagged local employment growth on our three measures capturing the nature of new housing supply. The results are qualitatively very similar to those reported in Table 2 for our baseline specification. Next we repeat this exercise but split our sample again into more and less regulated metro areas (Panels B and C). Again, we find strong effects with the expected signs in less regulated metro areas and the effects are again more pronounced than for the full sample. In more regulated areas we find that local employment growth is associated with a decrease rather than an increase in the share of *mf* housing and this effect is marginally statistically significant. There is no statistically significant effect of employment growth on the size of newly constructed housing units.

[17] An alternative interpretation of our findings could be that when local incomes rise, household formation increases, and newly formed households tend to demand more *mf* housing and smaller homes because they tend to be poorer and smaller. We do not have data to directly test this proposition.

Finally, in Panels D1 and D2 we report the findings of our instrumental variable approach: Panel D1 reports the 2nd stage of our TSLS estimates (along with a test statistic of the strength of the first stage) and Panel D2 reports the corresponding 1st stage results. The results are based on the sample of MSAs with comparably lax land use regulation only. We confine our sample to these MSAs because strict land use controls were demonstrated by Glaeser *et al.* (2006) and Saks (2008) to also prevent in-migration (i.e., house prices adjust rather than the composition of the housing stock), thus impairing the strength of our identification.[18] The findings provide further support for the proposition that the housing supply adjustments in metro areas with comparably lax land use regulation are driven by migration. Employment growth in these metro areas has a causal positive effect on the share of *mf* housing in new construction and a causal negative effect on the size of newly constructed housing, consistent with our Predictions 1 and 2. Moreover, the adverse causal effect on the size of new units is more pronounced for *mf* units, consistent with our theoretical Conjecture.

### 5.4 Robustness checks (using the AHS sample)

We carried out a number of robustness checks with our main sample. The results are reported in Tables A3 to A6 (in Web-Appendix D). In our baseline specification we explored the effect of an increase in lagged income (relative to the national average) on the *share* of all newly constructed units that are of the multi-family type (column 1 of Table 2). While focusing on the share of *mf*-construction allows us to directly test our theoretical Prediction 1, the reader may also be interested in the separate effects of an income shock on *mf*- and *sf*-construction.

In Table A3, we thus report the findings when we separately use the counts and log-counts of *mf* and *sf* construction as dependent variables, rather than the share of *mf* construction relative to all construction. Columns (1) and (2) report results for the counts. The estimates suggest that, not surprisingly, a one-year lagged income-increase (relative to the national average) induces an increase in both *mf* and *sf* construction. In absolute numbers, the increase is slightly larger for *mf* than for *sf* construction, which is remarkable given that, on average, slightly less than 30 percent of all existing housing units are of the *mf*-type and one might expect more sf-construction at the fringe going forward. Columns (3) and (4) report the results when we use the natural log of *mf* and *sf* unit counts. In doing so we confine our regression sample to MSA-year-observations with some construction (>0) of the type considered. The results imply that, consistent with our main findings, a 10% increase in lagged real income causes a significantly stronger increase in the construction of *mf* units (86%) than in the construction of *sf* homes (44%). The estimated coefficients suggest that the construction of *mf*-units relative to *sf*-units roughly doubles compared to the baseline. Column (5) finally reports the results for the log-ratio of *mf* to *sf* construction. The estimated coefficient implies that a 10% increase in lagged real income leads to a 38 percentage points stronger increase in *mf* compared to *sf* construction, consistent with the findings in columns (3) and (4).

Our baseline-specifications assume a *one-year lagged response* of construction to changes in real income. A one-year lag seemed most sensible to use given stylized facts about delays imposed by the planning and construction process and insights from the empirical literature on

[18] We also estimated the specifications for the full sample that includes MSAs with strict land use controls. As expected, the results are much less clear-cut.

the housing supply price elasticity in the United States (e.g., Mayer and Somerville 2000), which seems to suggest that construction generally responds fairly quickly (within a few quarters) to demand shock-induced price changes. In fact, developers' decisions when and what (and where) to build may not only be determined by contemporaneous or past changes in income but also by economic outlooks that may anticipate to some extent changes in economic conditions. However, one could also make a case for a longer lag based on the argument that especially multi-family projects are likely to take more than a year to plan and build.

To test the sensitivity of our findings to our lag-assumption, in Table A4 we report estimation results for our baseline model but we assume that the explanatory variables are either contemporaneous or lagged by 2 years. The findings of the sensitivity test suggest that the main effects are robust, although, they slightly decrease in strength if we use 2-year lagged income per capita.

In Table A5 we take this exercise one step further and first report findings for specifications with 3-year (Panel A) and even 4-year lags (Panel B). The estimated coefficients become significantly smaller and less strongly significant—and in the case of the effects on the size of *mf* units completely insignificant—when we use 3-year lags. The effects of 4-year lagged income cease to be statistically significant in all specifications and the estimated coefficients become very small. We have also experimented with specifications that jointly use several lags. We are somewhat wary about these specifications due to the likely auto-correlation in the construction figures. This caveat aside, when we use 1-year and 2-year lags, only the 1-year lag effects are statistically significant (with the predicted signs), while the two-year lagged effects are completely statistically insignificant. When we push things further and include 1-year, 2-year and 3-year lags, again only the 1-year lagged effects are statistically significant with the predicted signs. All in all, these findings seem to strongly support our assumption of a one-year lagged response.

As discussed in Section 4.2, the assumption that housing characteristics do not change between the year of construction and the year in which the unit is observed in the AHS, is crucial to our identification strategy. In our base specification we include, subject to some constraints, all units to compute the share *mf* housing and we apply a maximum time window of 10 years for the gap between the year of construction and the AHS survey year for the purpose of computing the *mf* and *sf* floor size indices. In Table A6 we report the findings of robustness checks, in which we impose even narrower time windows. Specifically, we check the sensitivity of our results for the share *mf* measure by introducing a time window and by limiting this to a maximum of 10 years and 5 years (columns 1 and 2), respectively, and we explore the robustness of our findings for the *mf* and *sf* floor size measures by limiting the time window to a maximum of 5 years (columns 3 and 4). Overall, the main results do not change significantly, even when these narrower windows are applied and the sample sizes, as a consequence, are significantly reduced.

### 5.5 Estimates using the Census building permit data

In a final step we replicate our baseline specification for the share *mf* units using the Census building permit data. We document the results in **Table 4**. Column (1) reports results for the full sample of 381 MSAs and the full sample period from 1980 to 2018. The estimated coefficient is highly statistically significant and suggests that an MSA that receives a one-time positive income shock, raising income 10 percent more than that at the national level, will observe an increase in the share of *mf* housing in new construction of 5.2 percentage points (coefficient of 0.52). This effect is smaller but of a similar magnitude compared to the corresponding effect for our AHS-sample (6.4 percentag points; coefficient of 0.64). The Census building permit sample differs from the AHS sample in three main respects: the sample period, geographical coverage, and the involved lag (permits vs. completed construction). Our AHS sample does not cover the years from 2005 to 2018. One potential explanation for the small difference in the estimated effects (0.52 vs. 0.64) could thus be that the effect may become weaker over time, plausibly because metro areas resemble less and less monocentric cities. However, column (2) of Table 4 suggests that the estimated effect, using the Census building permit data, is virtually unchanged for the sample period from 1980 to 2004 (0.524 vs. 0.521). Another potential explanation is the fact that the AHS sample consists of larger cities on average. Indeed, when we confine the Census building permit data to AHS-cities only (column 3), the effect becomes somewhat stronger with 0.56. Finally, when we combine the two sample restrictions, the estimated effect becomes 0.59, very close to our headline finding (0.64) from the main analysis. Overall, Table 4 is indicative that the estimated effects from the main analysis may be reasonably representative for all MSAs and time periods.

# 6 Conclusions

Our empirical analysis suggests that local economic conditions have a strong impact on the composition (type and size) of newly constructed housing units in a metro area. When one-year lagged local income rises, controlling for changes in income at the national level, more multi-family units and smaller units are being constructed in a metro area. We provide evidence in support of the proposition that these housing supply adjustments are driven by migration.

The standard urban economic model is a useful starting point for explaining these findings. In an open monocentric city where utility can be considered to be exogenous because of costless migration across cities, rising incomes should lead to higher land prices and therefore a higher capital intensity of land use. In this paper we propose a modified version of the open monocentric city model, in which this effect is brought about through substitution from single-family to multi-family construction and through a reduction of the square footage of dwellings, consistent with our main empirical findings. It suggests that the urban housing market responds to a positive income shock primarily by producing more apartments, rather than single family housing.

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# TABLES

Table 1
Metro area-level summary statistics

| Variable | Obs. | Mean | Std. Dev. | Between | Within | Min | Max | Clusters |
|---|---|---|---|---|---|---|---|---|
| *A. Full sample - no metro-dimension* | | | | | | | | |
| Share multi-family (*mf*) units | 1829 | 0.298 | 0.169 | 0.127 | 0.118 | 0 | 0.939 | 167 |
| Unit square footage, *sf* | 1548 | 2453 | 427 | 328 | 288 | 900 | 4500 | 167 |
| Unit square footage, *mf* | 1513 | 2093 | 1001 | 699 | 745 | 340 | 5000 | 167 |
| Income per capita (p.a.) | 1829 | 17855 | 6448 | 6377 | 3623 | 7616 | 42030 | 167 |
| Employment growth | 1829 | 0.0210 | 0.0241 | 0.0130 | 0.0209 | -0.0619 | 0.103 | 167 |
| Labor demand shock variable | 1829 | -0.0149 | 0.0207 | 0.00772 | 0.0194 | -0.0862 | 0.0426 | 167 |
| MSA population | 1829 | 2872300 | 2956012 | 2937453 | 203213 | 635318 | 17000000 | 167 |
| Wage per employee in construction sector (p.a.) | 1829 | 28597 | 7695 | 7465 | 4296 | 14546 | 56520 | 167 |
| Share of units in sample built between 1969-1979 | 1829 | 0.0689 | 0.253 | 0.0510 | .2495278 | 0 | 1 | 167 |
| Share of units in sample built during 1980s | 1829 | 0.599 | 0.490 | 0.354 | .3697993 | 0 | 1 | 167 |
| Share of units in sample built during 1990s | 1829 | 0.276 | 0.447 | 0.325 | .327798 | 0 | 1 | 167 |
| Share of units in sample built between 2000-2004 | 1829 | 0.0569 | 0.232 | 0.271 | .1281069 | 0 | 1 | 167 |
| Share developed residential land (excl. industrial) | 1829 | 0.142 | 0.104 | 0.102 | 0 | .0327 | 0.501 | 167 |
| Average regulatory restrictiveness | 1829 | 0.142 | 0.714 | 0.714 | 0 | -1.119 | 1.889 | 167 |
| *B. Metro areas with regulatory restrictiveness<average* | | | | | | | | |
| Share *mf* units | 918 | 0.285 | 0.165 | 0.120 | 0.121 | 0 | .826 | 83 |
| Unit square footage, *sf* | 769 | 2436 | 400 | 331 | 247 | 1525 | 4017 | 83 |
| Unit square footage, *mf* | 749 | 2021 | 987 | 690 | 737 | 403 | 5000 | 83 |
| *C. Metro areas with regulatory restrictiveness>average* | | | | | | | | |
| Share *mf* units | 911 | 0.311 | 0.173 | 0.134 | 0.114 | 0 | 0.939 | 84 |
| Unit square footage, *sf* | 779 | 2469 | 452 | 325 | 324 | 900 | 4500 | 84 |
| Unit square footage, *mf* | 764 | 2163 | 1009 | 701 | 753 | 340 | 5000 | 84 |

Table 2
Base specifications (weighted fixed effects models, full sample)
*Dependent variables:* Characteristics of newly built housing stock

| | (1) | (2) | (3) |
|---|---|---|---|
| | Share *mf* units | Log (unit sq. foot, *sf*) | Log (unit sq. foot, *mf*) |
| Log (Personal income per capita), 1-year lagged | 0.640*** | -0.484** | -1.676*** |
| | (0.130) | (0.223) | (0.556) |
| Log (Construction sector annual wage per employee), 1-year lagged | -0.117 | -0.00942 | 0.208 |
| | (0.105) | (0.112) | (0.316) |
| Metro area × AHS-year fixed effects | Yes | Yes | Yes |
| Year built-fixed effects† | Yes | Yes | Yes |
| Constant | -4.686*** | 12.58*** | 22.06*** |
| | (0.960) | (1.983) | (4.687) |
| Observations | *1829* | *1548* | *1513* |
| Number of AHS-year × MSA combinations | 167 | 167 | 167 |
| R-squared within | 0.242 | 0.178 | 0.063 |
| between | 0.307 | 0.047 | 0.0003 |
| overall | 0.234 | 0.043 | 0.003 |

*Notes:* Robust standard errors in parentheses. *** Significant at 1%; ** significant at 5%; * significant at 10%. † Coefficients and robust standard errors of year built-fixed effects are reported in Table A2 (Web-Appendix D).

Table 3
Are the Results Driven by Migration? (OLS- and TSLS-Estimates)

| | (1) | (2) | (3) |
|---|---|---|---|
| Dependent variables: | Share *mf* units | Log (unit sq. foot, *sf*) | Log (unit sq. foot, *mf*) |
| Panel A: Full sample (OLS) | | | |
| Employment growth, 1-year lagged | 0.660** | -0.392* | -2.254*** |
| | (0.309) | (0.233) | (0.779) |
| Log (Construction cost sector annual wage per employee), 1-year lagged | 0.150* | -0.197* | -0.425 |
| | (0.0764) | (0.105) | (0.282) |
| Fixed effects and controls | Yes | Yes | Yes |
| Observations | 1829 | 1548 | 1513 |
| R-squared within/between/overall | 0.23/0.50/0.31 | 0.17/0.15/0.11 | .060/.072/.060 |
| Number of AHS-year x metro area combinations | 167 | 167 | 167 |
| Panel B: More regulated metropolitan areas (OLS) | | | |
| Employment growth, 1-year lagged | -0.675* | 0.394 | 0.636 |
| | (0.399) | (0.297) | (1.088) |
| Log (Construction cost sector annual wage per employee), 1-year lagged | 0.180 | -0.195 | -0.340 |
| | (0.118) | (0.150) | (0.355) |
| Fixed effects and controls | Yes | Yes | Yes |
| Observations | 911 | 779 | 764 |
| R-squared within/between/overall | 0.22/0.36/0.26 | 0.19/0.13/0.10 | 0.085/0.24/0.12 |
| Number of AHS-year x metro area combinations | 84 | 84 | 84 |
| Panel C: Less regulated metropolitan areas (OLS) | | | |
| Employment growth, 1-year lagged | 0.979** | -0.769*** | -3.217*** |
| | (0.388) | (0.230) | (0.850) |
| Log (Construction cost sector annual wage per employee), 1-year lagged | 0.00280 | -0.105 | -0.365 |
| | (0.114) | (0.136) | (0.463) |
| Fixed effects and controls | Yes | Yes | Yes |
| Observations | 918 | 769 | 749 |
| R-squared within/between/overall | 0.32/0.63/0.39 | 0.23/0.24/0.17 | .093/.0001/.025 |
| Number of AHS-year x metro area combinations | 83 | 83 | 83 |
| Panel D1: Less regulated metropolitan areas (TSLS, 2$^{nd}$ stage) | | | |
| **Employment growth, 1-year lagged** [1)] | **1.614*** ** | **-1.665*** | **-3.539*** |
| | **(0.580)** | **(1.011)** | **(1.925)** |
| Log (Construction cost sector annual wage per employee), 1-year lagged | -0.00943 | -0.103 | -0.374 |
| | (0.115) | (0.135) | (0.459) |
| Fixed effects [2)] | Yes | Yes | Yes |
| Observations | 918 | 769 | 749 |
| Number of AHS-year x metro area combinations | 83 | 83 | 83 |
| Kleibergen-Paap rk Wald F-statistic (First-stage F) | 12.79 | 15.68 | 23.93 |
| Panel D2: Less regulated metropolitan areas (TSLS, 1$^{st}$ stage) | | | |
| Dependent variable: | | **Employment growth, 1-year lagged** | |
| Labor demand shock, 1-year lagged | 1.580*** | 1.540*** | 2.300*** |
| | (0.442) | (0.389) | (0.470) |
| Log (Construction cost sector annual wage per employee), 1-year lagged | -0.00104 | -0.00193 | -0.0221 |
| | (0.0233) | (0.0276) | (0.0278) |
| Centered/uncentered R-squared | 0.104 | 0.083 | 0.144 |

*Notes:* Robust standard errors in parentheses. *** Significant at 1%; ** significant at 5%; * significant at 10%. [1)] **Bold** variable is endogenous. Excluded instrument is labor demand shock variable. [2)] Year built-fixed effects and constant are partialled out.

Table 4
Estimates using Census building permit data (weighted fixed effects models, full sample)
*Dependent variable:* Share new multi-family units

| | | (1) | (2) | (3) | (4) |
|---|---|---|---|---|---|
| | | Full sample 1980-2018 | Full sample 1980-2004 | AHS-cities only 1980-2018 | AHS-cities only 1980-2004 |
| Log (Personal income per capita), 1-year lagged | | 0.521*** | 0.524*** | 0.564*** | 0.593*** |
| | | (0.0768) | (0.144) | (0.108) | (0.216) |
| Log (Construction sector annual wage per employee), 1-year lagged | | -0.0037 | 0.0621 | -0.0191 | 0.0628 |
| | | (0.0290) | (0.0597) | (0.0511) | (0.121) |
| Metro area fixed effects | | Yes | Yes | Yes | Yes |
| Year built-fixed effects | | Yes | Yes | Yes | Yes |
| Constant | | -4.734*** | -5.430*** | -4.949*** | -6.125*** |
| | | (0.670) | (1.148) | (0.968) | (1.628) |
| Observations | | 11,652 | 6,861 | 1,609 | 1,059 |
| Number of metro areas | | 381 | 346 | 43 | 43 |
| R-squared | within | 0.391 | 0.477 | 0.543 | 0.578 |
| | between | 0.092 | 0.128 | 0.390 | 0.297 |
| | overall | 0.159 | 0.210 | 0.453 | 0.450 |

*Notes:* Robust standard errors in parentheses. *** Significant at 1%; ** significant at 5%; * significant at 10%. Summary statistics for full sample (1980-2018; *N*=11,631; number of MSAs=381): Share *mf* units (mean: 0.264; std. dev.: 0.188), annual income per capita (mean: 27,052; std. dev.: 7,114), construction sector annual wage per employee (mean: 39,411; std. dev.: 9,451).

# FIGURES

*Figure 1:*
*Bid rent functions for floor space (left) and land (right)*

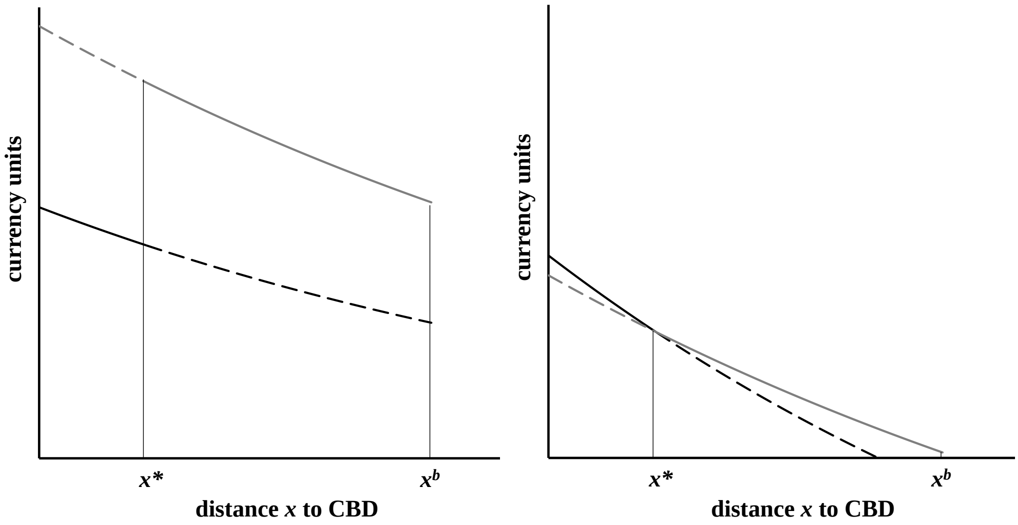


*Figure 2:*
*Effect of an increase in local income on the bid rent for land*

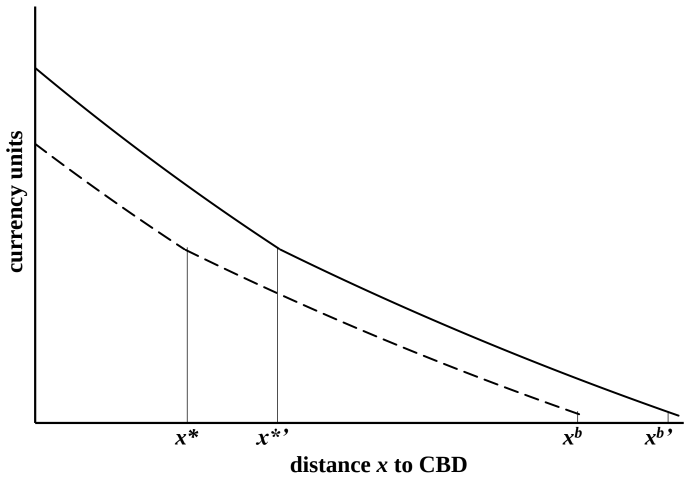


*Figure 3:*
*Change in conversion rates before and after an increase in income*

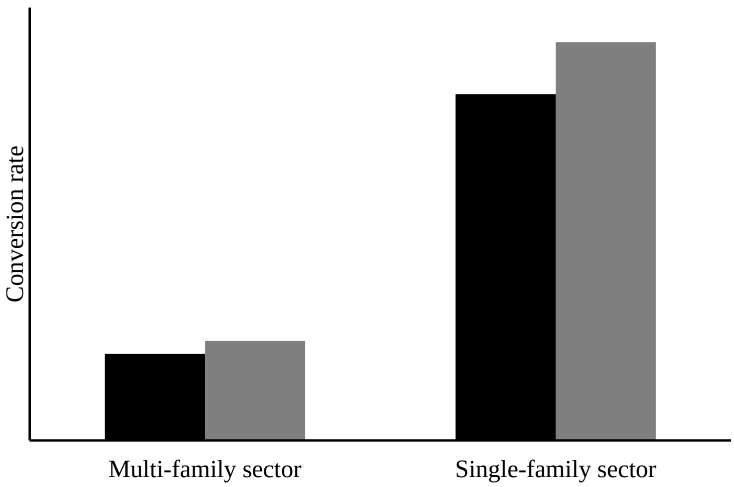

*Figure 4:*
*Income growth and in-migration, 1995-2000*

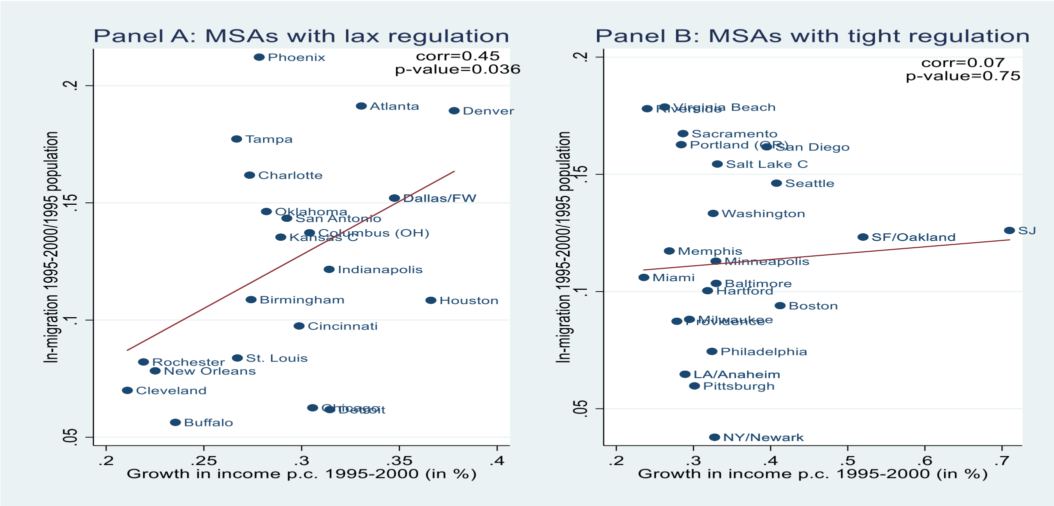


*Figure 5:*
*Employment growth and in-migration, 1995-2000*

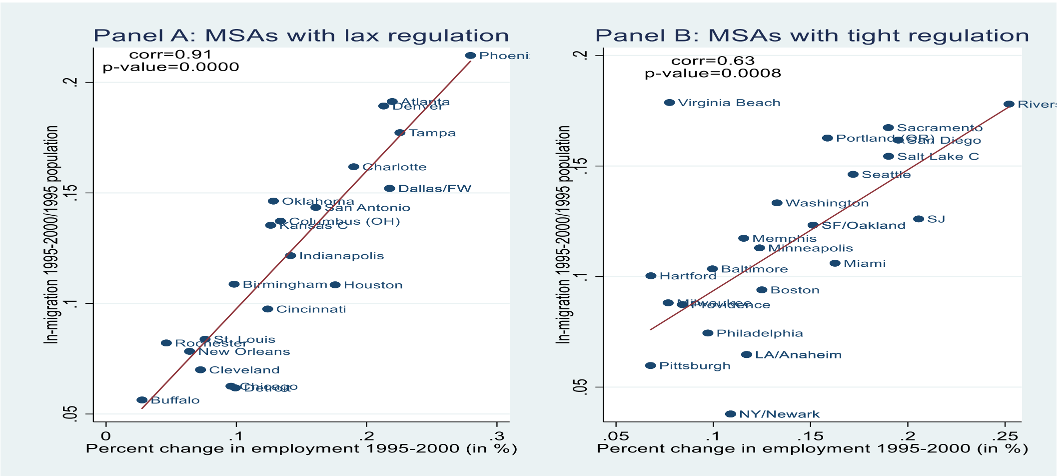


*Figure 6:*
*Income growth and employment growth, 1995-2000*