THE GEOGRAPHIC DETERMINANTS OF HOUSING SUPPLY*

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I process satellite-generated data on terrain elevation and presence of water bodies to precisely estimate the amount of developable land in U.S. metropolitan areas. The data show that residential development is effectively curtailed by the presence of steep-sloped terrain. I also find that most areas in which housing supply is regarded as inelastic are severely land-constrained by their geography. Econometrically, supply elasticities can be well characterized as functions of both physical and regulatory constraints, which in turn are endogenous to prices and demographic growth. Geography is a key factor in the contemporaneous urban development of the United States.

I. INTRODUCTION

The determinants of local housing supply elasticities are of critical importance in explaining current trends in the shape of urban development and the evolution of housing values.[1] The existing literature on this topic has focused on the role that local land use regulations play in accounting for differences in the availability of land. The large variance in housing values across locales can indeed be partially explained by man-made regulatory constraints. However, zoning and other land-use policies are multidimensional, difficult to measure, and endogenous to preexisting land values. In this context, it is uncontroversial to argue that predetermined geographic features such as oceans, lakes, mountains, and wetlands can also induce a relative scarcity of developable land. Hence their study merits serious consideration: to what extent, if at all, does geography determine contemporaneous patterns of urban growth?[2]

This paper gives empirical content to the concepts of land scarcity and abundance in urban America. Using geographic information system (GIS) techniques, I precisely estimate the area that is forgone to the sea within 50-kilometer radii from metropolitan

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1. Glaeser, Gyourko, and Saks (2006); Saks (2008).

2. An important step in this direction has been taken by Burchfield et al. (2006), who relate terrain ruggednes and access to underground water to the density and compactness of *new* real estate development.

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central cities. I then use satellite-based geographic data on land use provided by the United States Geographic Service (USGS) to calculate the area lost to internal water bodies and wetlands. Using the USGS Digital Elevation Model (DEM) at 90–square meter cell grids, I also create slope maps, which allow me to calculate how much of the land around each city exhibits slopes above 15%. Combining all the information above, the paper provides a precise measure of exogenously undevelopable land in cities. I then turn to studying the links between geography and urban development.

To do so, I first develop a conceptual framework that relates land availability to urban growth and housing prices. Using a variation of the Alonso–Muth–Mills model (Alonso 1964; Mills 1967; Muth 1969), I show that land-constrained cities not only should be more expensive *ceteris paribus,* but also should *display lower housing supply elasticities* with respect to citywide demand shocks, a somewhat *ad hoc* claim in the existing literature. I also show that, in equilibrium, consumers in geographically constrained metropolitan areas should require higher wages or higher amenities to compensate them for more expensive housing.

Empirically, all of these facts are corroborated by the data. I find that most areas that are widely regarded as supply-inelastic are, in fact, severely land-constrained by their geography. Rose (1989b) showed a positive correlation between coastal constraints and housing prices for a limited sample of forty-five cities. Here I show that restrictive geography, including the presence of mountainous areas and internal water, was a very strong predictor of housing price levels and *growth* for all metropolitan statistical areas (MSA) during the period 1970–2000, even after controlling for regional effects. This association was not solely driven by coastal areas, as it is present even *within* coastal markets. I next deploy the Wharton Residential Urban Land Regulation Index recently created by Gyourko, Saiz, and Summers (2008). The index is constructed to capture the stringency of residential growth controls. Using alternate citywide demand shocks, I estimate metropolitan-specific housing supply functions and find that housing supply elasticities can be well characterized as functions of both physical and regulatory constraints.

These associations, however, do not take into account feedback effects between prices and regulations. Homeowners have stronger incentives to protect their housing investments where land values are high initially. The homevoter hypothesis (Fischel

2001) implies a reverse causal relationship from initially high land values to increased regulations. Empirically, I find that antigrowth local land policies are more likely to arise in growing, land-constrained metropolitan areas and in cities where preexisting land values were high and worth protecting. Hence, I next endogeneize the regulatory component of housing supply elasticity. I posit and estimate an empirical model of metropolitan housing markets with endogenous regulations. As exogenous land-use regulatory shifters, I use measures shown to be associated with local tastes for regulation. Both geography and regulation are important to account for housing supply elasticities, with the latter showing themselves to be endogenous to prices and past growth.

Finally, I use the results to provide operational estimates of local supply elasticities in all major U.S. metropolitan areas. These estimates, based on land-availability fundamentals, should prove useful in calibrating general equilibrium models of interregional labor mobility and in predicting the response of housing markets to future demand shocks. Housing supply is estimated to be quite elastic for the average metropolitan area (with a population-weighted elasticity of 1.75). In land-constrained large cities, such as cities in coastal California, Miami, New York, Boston, and Chicago, estimated elasticities are below one. These elasticity estimates display a very strong correlation of .65 with housing prices in 2000. Quantitatively, a movement across the interquartile range in geographic land availability in an average-regulated metropolitan area of 1 million is associated with shifting from a housing supply elasticity of approximately 2.45 to one of 1.25. Moving to the ninetieth percentile of land constraints (as in San Diego, where 60% of the area within its 50-km radius is not developable) pushes average housing supply elasticities down further to 0.91. The results in the paper ultimately demonstrate that geography is a key factor in the contemporaneous urban development of the United States.

II. Geography and Land in the United States: A New Data Set

The economic importance of geography for local economic development is an underexplored topic. Previous research has examined the correlation between housing price levels and proxies for the arc of circle lost to the sea in a limited number of cities (Rose 1989a, 1989b; Malpezzi 1996; Malpezzi, Chun, and Green 1998) but the measures proved somewhat limited. Recent papers in urban economics, such as Burchfield et al. (2006), Rosenthal and Strange (2008), and Combes et al. (2009), underline the relevance of geographic conditions as economic fundamentals explaining local population density.

Here, I develop a comprehensive measure of the area that is unavailable for residential or commercial real estate development in MSAs. Architectural development guidelines typically deem areas with slopes above 15% severely constrained for residential construction. Using data on elevation from the USGS Digital Elevation Model (DEM) at its 90-m resolution, I generated slope maps for the continental United States. GIS software was then used to calculate the exact share of the area corresponding to land with slope above 15% within a 50-km radius of each metropolitan central city.

Residential development is effectively constrained by the presence of steep slopes. To demonstrate this, I focus on Los Angeles (LA). Median housing values there are among the highest in the United States and the incentives to build on undeveloped land are very strong. Using GIS software to delineate the intersection between steep-slope zones and the 6,456 census block groups (as delimited in 2000) that lie within a 50-km radius of LA's city centroid, I calculated the share of the area in each block group with slope above 15%. Then I defined steep-slope block groups as those with a share of steep-sloped terrain of more than 50%. Steep-slope block groups encompassed 47.62% of the land area within 50 km of LA's geographic center in year 2000. However, only 3.65% of the population within this 50-km radius lived in them. These magnitudes clearly illustrate the deterrent effect of steep slopes on housing development.

The next step to calculate land availability involved estimating the area within the cities' 50-km radii that corresponds to wetlands, lakes, rivers, and other internal water bodies. The 1992 USGS National Land Cover Dataset is a satellite-based GIS source containing information about land cover characteristics at 30 by 30–m cell resolutions. The data were processed by the Wharton GIS lab to produce information on the area apportioned to each of the land cover uses delimited by the USGS by census tract. Next, the distance from each central city centroid to the centroid of all census tracts was calculated, and Census tracts within 50 km were used to compute water cover shares.

Last, I used digital contour maps to calculate the areas within the 50-km radii that are lost to oceans and the Great Lakes. The final measure combines the area corresponding to steep slopes, oceans, lakes, wetlands, and other water features. This is the first comprehensive measure of truly undevelopable area in the literature. The use of a radius from the city centroid makes it a measure of original constraints, as opposed to one based on *ex post* ease of development (e.g., density).

Table I displays the percentages of undevelopable area for all MSAs with population over 500,000 in the 2000 Census for which I also have regulation data (those included in the later regressions). Of these large metro areas, Ventura (CA) is the most constrained, with 80% of the area within a 50-km radius rendered undevelopable by the Pacific Ocean and mountains. Miami, Fort Lauderdale, New Orleans, San Francisco, Sarasota, Salt Lake City, West Palm Beach, San Diego, and San Jose complete the list of the top 10 most physically constrained major metropolitan areas in the United States. Many large cities in the South and Midwest (such as Atlanta, San Antonio, and Columbus) are largely unconstrained.

Table II studies the correlates of the newly constructed land unavailability variable. To do so, I run a number of independent regressions. The variables in Table II's rows appear on the left-hand side in each sequential regression, and the geographic-unavailability variable is always the main right-hand side control. Regional fixed effects (Northeast, South, Midwest, West) are included in all regressions. Each column shows the coefficient of the variable of reference on the unavailable land share, and its associated standard error appears in parentheses. A second set of regressions (2) also controls for a coastal status dummy, which identifies metropolitan areas that are within 100 km of the ocean or Great Lakes. The significant coefficients reveal that geographically land-constrained areas tended to be more expensive in 2000, to have experienced faster price growth since 1970, to have higher incomes, to be more creative (higher patents per capita), and to have higher leisure amenities (as measured by the number of tourist visits).[3] Observed metropolitan population levels were largely orthogonal to natural land constraints.

Interestingly, note that none of the major demand-side drivers of recent urban demographic change (immigration, education,

3. Carlino and Saiz (2008) demonstrate that the number of tourist visits is strongly correlated with other measures of quality of life and a strong predictor of recent city growth.

TABLE I
PHYSICAL AND REGULATORY DEVELOPMENT CONSTRAINTS (METRO AREAS WITH POPULATION > 500,000)

Rank	MSA/NECMA name	Undevelopable area (%)	WRI	Rank	MSA/NECMA name	Undevelopable area (%)	WRI
1	Ventura, CA	79.64	1.21	26	Portland–Vancouver, OR–WA	37.54	0.27
2	Miami, FL	76.63	0.94	27	Tacoma, WA	36.69	1.34
3	Fort Lauderdale, FL	75.71	0.72	28	Orlando, FL	36.13	0.32
4	New Orleans, LA	74.89	−1.24	29	Boston–Worcester–Lawrence, MA–NH	33.90	1.70
5	San Francisco, CA	73.14	0.72	30	Jersey City, NJ	33.80	0.29
6	Salt Lake City–Ogden, UT	71.99	−0.03	31	Baton Rouge, LA	33.52	−0.81
7	Sarasota–Bradenton, FL	66.63	0.92	32	Las Vegas, NV–AZ	32.07	−0.69
8	West Palm Beach–Boca Raton, FL	64.01	0.31	33	Gary, IN	31.53	−0.69
9	San Jose, CA	63.80	0.21	34	Newark, NJ	30.50	0.68
10	San Diego, CA	63.41	0.46	35	Rochester, NY	30.46	−0.06
11	Oakland, CA	61.67	0.62	36	Pittsburgh, PA	30.02	0.10
12	Charleston–North Charleston, SC	60.45	−0.81	37	Mobile, AL	29.32	−1.00
13	Norfolk–Virginia Beach–Newport News, VA–NC	59.77	0.12	38	Scranton–Wilkes-Barre–Hazleton, PA	28.78	0.01
14	Los Angeles–Long Beach, CA	52.47	0.49	39	Springfield, MA	27.08	0.72
15	Vallejo–Fairfield–Napa, CA	49.16	0.96	40	Detroit, MI	24.52	0.05
16	Jacksonville, FL	47.33	−0.02	41	Bakersfield, CA	24.21	0.40
17	New Haven–Bridgeport–Stamford, CT	45.01	0.19	42	Harrisburg–Lebanon–Carlisle, PA	24.02	0.54
18	Seattle–Bellevue–Everett, WA	43.63	0.92	43	Albany–Schenectady–Troy, NY	23.33	−0.09
19	Milwaukee–Waukesha, WI	41.78	0.46	44	Hartford, CT	23.29	0.49
20	Tampa–St. Petersburg–Clearwater, FL	41.64	−0.22	45	Tucson, AZ	23.07	1.52
21	Cleveland–Lorain–Elyria, OH	40.50	−0.16	46	Colorado Springs, CO	22.27	0.87
22	New York, NY	40.42	0.65	47	Baltimore, MD	21.87	1.60
23	Chicago, IL	40.01	0.02	48	Allentown–Bethlehem–Easton, PA	20.86	0.02
24	Knoxville, TN	38.53	−0.37	49	Minneapolis–St. Paul, MN–WI	19.23	0.38
25	Riverside–San Bernardino, CA	37.90	0.53	50	Buffalo–Niagara Falls, NY	19.05	−0.23

TABLE I
(CONTINUED)

Rank	MSA/NECMA name	Undevelopable area (%)	WRI	Rank	MSA/NECMA name	Undevelopable area (%)	WRI
51	Toledo, OH	18.96	−0.57	74	Dallas, TX	9.16	−0.23
52	Syracuse, NY	17.85	−0.59	75	Richmond–Petersburg, VA	8.81	−0.38
53	Denver, CO	16.72	0.84	76	Houston, TX	8.40	−0.40
54	Columbia, SC	15.23	−0.76	77	Raleigh–Durham–Chapel Hill, NC	8.11	0.64
55	Wilmington–Newark, DE–MD	14.67	0.47	78	Akron, OH	6.45	0.07
56	Birmingham, AL	14.35	−0.23	79	Tulsa, OK	6.29	−0.78
57	Phoenix–Mesa, AZ	13.95	0.61	80	Kansas City, MO–KS	5.82	−0.79
58	Washington, DC–MD–VA–WV	13.95	0.31	81	El Paso, TX	5.13	0.73
59	Providence–Warwick–Pawtucket, RI	13.87	1.89	82	Fort Worth–Arlington, TX	4.91	−0.27
60	Little Rock–North Little Rock, AR	13.71	−0.85	83	Charlotte–Gastonia–Rock Hill, NC–SC	4.69	−0.53
61	Fresno, CA	12.88	0.91	84	Atlanta, GA	4.08	0.03
62	Greenville–Spartanburg–Anderson, SC	12.87	−0.94	85	Austin–San Marcos, TX	3.76	−0.28
63	Nashville, TN	12.83	−0.41	86	Omaha, NE–IA	3.34	−0.56
64	Louisville, KY–IN	12.69	−0.47	87	San Antonio, TX	3.17	−0.21
65	Memphis, TN–AR–MS	12.18	1.18	88	Greensboro–Winston–Salem–High Point, NC	3.12	−0.29
66	Stockton–Lodi, CA	12.05	0.59	89	Fort Wayne, IN	2.56	−1.22
67	Albuquerque, NM	11.63	0.37	90	Columbus, OH	2.50	0.26
68	St. Louis, MO–IL	11.08	−0.73	91	Oklahoma City, OK	2.46	−0.37
69	Youngstown–Warren, OH	10.52	−0.38	92	Wichita, KS	1.66	−1.19
70	Cincinnati, OH–KY–IN	10.30	−0.58	93	Indianapolis, IN	1.44	−0.74
71	Philadelphia, PA–NJ	10.16	1.13	94	Dayton–Springfield, OH	1.04	−0.50
72	Ann Arbor, MI	9.71	0.31	95	McAllen–Edinburg–Mission, TX	0.93	−0.45
73	Grand Rapids–Muskegon–Holland, MI	9.28	−0.15				

Note. WRI = Wharton Regulation Index.

TABLE II
PARTIAL CORRELATES OF UNAVAILABLE LAND SHARE (50-KM RADIUS)

	Share of area unavailable for development	
	OLS-regional FE β	Adds coastal dummy β
	(1)	(2)
Log population in 2000	0.443	−0.01
	(0.336)	(0.364)
Log median house value in 2000	0.592	0.41
	(0.081)***	(0.085)***
ΔLog median house value (1970–2000)	0.240	0.122
	(0.054)***	(0.057)**
Log income in 2000	0.233	0.164
	(0.056)***	(0.060)***
ΔLog income (1990–2000)	−0.002	0.006
	(0.020)	(0.022)
ΔLog population (1990–2000)	−0.027	−0.043
	(0.027)	(0.029)
Immigrants (1990–2000)/population (1990)	0.009	−0.007
	(0.011)	(0.012)
Share with bachelor's degree (2000)	0.006	−0.004
	(0.020)	(0.022)
Share workers in manufacturing (2000)	−0.01	0.005
	(0.021)	(0.023)
Log(patents/population) (2000)	0.762	0.771
	(0.260)***	(0.287)***
January monthly hours of sun (average 1941–1970)	−3.812	−12.047
	(11.252)	(12.318)
Log tourist visits per person (2000)	0.493	0.719
	(0.261)*	(0.286)**

Notes. Standard errors in parentheses. Rows present the coefficients (β) and standard errors of separate regressions, where the variable described in the row is the dependent variable on the left-hand side and the unavailable land share (geographic constraint) is the explanatory variable on the right-hand side. The regressions in column (1) include regional fixed effects as controls, whereas those in column (2) also include a coastal dummy for metropolitan areas within 100 km of the oceans or Great Lakes (as defined in Rappaport and Sachs [2003]). * significant at 10%; ** significant at 5%; *** significant at 1%.

manufacturing orientation, and hours of sun) was actually correlated with geographic land constraints.

All results hold after controlling for the coastal dummy, indicating that the new land-availability variable contains information above and beyond that used in studies that focus on coastal status (Rose 1989a, 1989b; Malpezzi 1996). Taking into account the standard deviations of the different components of land unavailability, mountains contribute 42% of the variation in this variable, whereas coastal and internal water loss account for

31% and 26% of the variance in land constraints, respectively. After controlling for region fixed effects, as I do throughout the paper, there is no correlation in the data between coastal area loss and the extent of land constraints begotten by mountainous terrain. The loss of developable land due to the presence of large bodies of internal water (70% of which is attributable to wetlands, as in the Everglades) tends to be positively associated with coastal area loss and, not surprisingly, negatively associated with mountainous terrain.

The other major data set used in the paper is obtained from the 2005 Wharton Regulation Survey. Gyourko, Saiz, and Summers (2008) use the survey to produce a number of indexes that capture the intensity of local growth control policies in a number of dimensions. Lower values in the Wharton Regulation Index, which is standardized across all municipalities in the original sample, can be thought of as signifying the adoption of more laissez-faire policies toward real estate development. Metropolitan areas with high values of the Wharton Regulation Index (WRI henceforth), conversely have zoning regulations or project approval practices that constrain new residential real estate development. I process the original municipal-based data to create average regulation indexes by metropolitan area using the probability sample weights developed by Gyourko, Saiz, and Summers (2008).[4]

Table I displays the average WRI values for all metropolitan areas with populations greater than 500,000 and for which data are available. A clear pattern arises when the regulation index is contrasted with the land-availability measure. Physical land scarcity is associated with stricter regulatory constraints to development. Of the twenty most land-constrained areas, fourteen have positive values of the regulation index (which has a mean of −0.10 and a s.e. of 0.81 across metro areas). Conversely, sixteen of the twenty least land-constrained metropolitan areas have negative regulation index values.

Other data sources are used throughout the paper: the reader is referred to Appendices I–III for descriptive statistics and the meaning and provenance of the remaining variables.

4. Note that, because of different sample sizes across cities, in regressions where the WRI is used on the left-hand side (Table IV), heteroscedasticity could be an issue, and therefore Feasible Generalized Least Squares (FGLS) are used. In fact, however, the results in Table IV are very robust to all reasonable weighting schemes and the omission of metro areas with smaller number of observations in the WRI.

III. Geography and Local Development: A Framework

Why should physical or man-made land availability constraints have an impact on housing supply *elasticities*? How does geography shape urban development? To characterize the supply of housing in a city, I assume developers to be price takers in the land market. Consumers within the city compete for locations determining the price of the land input. Taking land values and construction outlays as given, developers supply housing at cost. All necesary model derivations and the proofs of propositions are in the mathematical appendix, Appendix I.

The preferences of homogeneous consumers in city k are captured by the utility function $U(C_k) = (C_k)^{\rho}$. Consumption in the city (C_k) is the sum of the consumption of city amenities (A_k) and private goods. Private consumption is equal to wages in the city minus rents, minus the (monetized) costs of commuting to the central business district (CBD), where all jobs are located. Each individual is also a worker and lives in a separate house, so that the number of housing units equals population ($H_k = \text{POP}_k$). Utility can be expressed as $U(C_k) = (A_k + w_k - \gamma \cdot r' - t \cdot d)^{\rho}$, where w_k stands for the wage in the city, γ for the units of land/housing-space consumption (assumed constant), r' for the rent *per unit of housing-space consumption*, t for the monetary cost per distance commuted, and d for the distance of the consumer's residence to the CBD. As in conventional Alonso–Muth–Mills models (Brueckner 1987), a nonarbitrage condition defines the rent gradient: all city inhabitants attain utility $\overline{U_k}$ via competition in the land markets. Therefore the total rent paid by an individual ($r = \gamma \cdot r'$) takes the functional form $r(d) = r_0 - td$.

Consider a circular city with radius Φ_k. Geographic or regulatory land constraints make construction unfeasible in some areas: only a sector (share) Λ_k of the circle is developable.[5] The city radius is thus a function of the number of households and land availability: $\Phi_k = \sqrt{\gamma H_k / \Lambda_k \pi}$.

Developers are price takers and buy land at market prices. They build and sell homes at price $P(d)$. The construction sector is competitive and houses are sold at the cost of land, $\text{LC}(d)$, plus construction costs, CC, which include the profits of the builder: $P(d) = \text{CC} + \text{LC}(d)$. In the asset market steady state equilibrium

5. This feature appears in conventional urban economic models that focus on a representative city (Capozza and Helsley 1990). Here, I add heterogeneity in the land availability parameter across cities and derive explicit housing supplies elasticities from it.

there is no uncertainty and prices equal the discounted value of rents: $P(d) = r(d)/i$, which implies that $r(d) = i \cdot \text{CC} + i \cdot \text{LC}(d)$. At the city's edge there is no alternative use for land so, without loss of generality, $\text{LC}(\Phi_k) = 0$. Therefore $r(\Phi_k) = i \cdot \text{CC}$, which implies that $r_0 = i \cdot \text{CC} + t \cdot \sqrt{\gamma H_k / \Lambda_k \pi}$.

In this setup, average housing rent in the city, $\widetilde{r}_k$, can be shown to be equivalent to the rent paid by the household living two-thirds of the distance from the CBD to the city's edge: $\widetilde{r}_k = r(\frac{2}{3}\Phi_k)$ (see Derivation 1 in Appendix II). The final housing supply equation in the city has average housing values ($\widetilde{P}_k^S$) expressed as a function of the number of households:

$$\widetilde{P}_k^S = \text{CC} + \frac{1}{3i} t \cdot \sqrt{\frac{\gamma H_k}{\Lambda_k \pi}}. \tag{1}$$

I next define the aggregate demand function for housing in the city. In a system of open cities, consumers can move and thus equalize utility across locations, which I normalize to zero (i.e., the spatial indifference condition is $\overline{U_k} = 0 \ \forall k$). Furthermore, in all cities, w_k and A_k are functions of population. I model the level of amenities as $A_k = \widetilde{A_k} - \alpha\sqrt{\text{POP}_k}$. The parameter α mediates the marginal congestion cost (in terms of rivalry for amenities, traffic, pollution, noise, social capital dilution, crime, etc.). α could also be interpreted in the context of an alternative but isomorphic model with taste heterogeneity: people with greater preferences for the city are willing to pay more and move in first, but later marginal migrants display less of a willingness to pay for the city (e.g., Saiz [2007]). Labor demand is modeled as $w_k = \widetilde{w_k} - \psi\sqrt{\text{POP}_k}$ and is assumed to be downward sloping; marginal congestion costs weakly increase with population ($\psi, \alpha \geq 0$).[6] Recalling that $H_k = \text{POP}_k$, substituting into the intercity spatial equilibrium equation, and focusing w.o.l.o.g. on the spatial indifference condition of consumers living in the CBD, I obtain the demand schedule for housing in the city:

$$\sqrt{H_k} = \frac{\widetilde{A_k} + \widetilde{w_k}}{(\psi + \alpha)} - \frac{i}{(\psi + \alpha)} P(0). \tag{2}$$

6. Of course, cities may display agglomeration economies up to some congestion point (given predetermined conditions, these may be captured by $\widetilde{A_k} + \widetilde{w_k}$). It is necessary only that, in equilibrium, the *marginal* effect of population on wages and amenities be (weakly) negative. This is a natural assumption that avoids a counterfactual equilibrium where all activity is concentrated in one single city with $\Lambda_k = 1$.

Note that relative shocks to labor productivity or to amenities $(\widetilde{A}_k + \widetilde{w}_k)$ shift the city's demand curve upward, *which I will use to identify supply elasticities later*.

I can now combine the expression for home values in the CBD via the supply equation and the city-demand equation (2) to obtain the equilibrium number of households in each city,

$$H_k^* = \left(\frac{\widetilde{A}_k + \widetilde{w}_k - i \cdot \mathrm{CC}}{(\psi + \alpha) + t \cdot \sqrt{\frac{\gamma}{\Lambda_k \cdot \pi}}} \right)^2 \quad \text{(Derivation 2).}$$

Note that amenities and wages have to at least cover the annuitized physical costs of construction for a potential site to be inhabitable.

Within this setup, I first study the supply response to growth in the demand for housing that is induced by productivity and amenity shocks. Its is clear that $\partial \widetilde{P}_k^S / \partial \Lambda_k < 0$. *Other things equal*, more land availability shifts down the supply schedule. Do land constraints also have an effect with respect to supply *elasticities*? Defining the city-specific supply inverse elasticity of average housing prices as $\beta_k^S \equiv \partial \ln \widetilde{P}_k^S / \partial \ln H_k$ one can demonstrate

PROPOSITION 1. The inverse elasticity of supply (that is, the price sensitivity to demand shocks) is decreasing in land availability. Conversely, as land constraints increase, positive demand shocks imply stronger positive impacts on the the growth of housing values.

Proposition 1 tells us that land-constrained cities have more inelastic housing supply and helps us understand how housing prices react to exogenous demand shocks. In addition, two interesting further questions arise from the general equilibrium in the housing and labor markets: Why is there any population in areas with difficult housing supply conditions? Should these areas be more expensive *ex post* in equilibrium? Assume that the covariance between productivity, amenities, and land availability is zero across all locales. Productivity–amenity shocks are *ex ante* independent of physical land availability, which is consistent with random productivity shocks and Gibrat's Law explanation for parallel urban growth (Gabaix 1999). Assume further that the relevant upper tail of such shocks is drawn from a Pareto distribution. I can now state

PROPOSITION 2. Metropolitan areas with low land availability tend to be more productive or to have higher amenities; in the observable distribution of metro areas the covariance between land availability and productivity–amenity shocks is negative.

The intuition for Proposition 2 is based on the nature of the urban development process. As discussed by Eeckhout (2004), existing metropolitan areas are a truncated distribution of the upper tail of inhabited settlements. In order to compensate for the higher housing prices that are induced by locations with more difficult supply conditions, consumers need to be rewarded with higher wages or urban amenities. Although costly land development reduced *ex ante* the desirability of marshlands, wetlands, and mountainous areas for human habitation, those land-constrained cities *that thrived ex post* must be more productive or attractive than comparable locales. Observationally, this implies a positive association between attractiveness and land constraints, conditional on metropolitan status. Conversely, land-unconstrained metropolitan areas must be, on average, observationally less productive and/or amenable.

Note that because the spatial indifference condition has to hold, this implies that expected home values are also decreasing in land availability: metropolitan areas with lower land availability tend to be more expensive in equilibrium. These conclusions are reinforced if the *ex ante* covariance between productivity/amenities and land availability is negative, albeit this is not a necessary condition.[7]

Although, due to a selection effect, land-constrained metropolitan areas have higher amenities, productivity, and prices, they are not necessarily larger. In fact, if productivity–amenity shocks are approximately Pareto-distributed in the upper tail (consistent with the empirical evidence on the distribution of city sizes in most countries), one can posit

PROPOSITION 3. Population levels in the existing distribution of metropolitan areas should be independent of the degree of land availability.

Proposition 3 tells us that population levels in metropolitan areas are expected to be orthogonal to initial land availability. In equilibrium, higher productivity and/or amenities are required

7. Glaeser (2005a, 2005b) and Gyourko (2005) emphasize the importance of access to harbors (a factor that limits land availability) for the earlier development of some of the larger oldest cities in the United States: Boston, New York, and Philadelphia.

in more land-constrained cities, which further left-censors their observed distribution of city productivities. With a Pareto distribution of productivity shocks, this effect exactly compensates for the extra costs imposed by a difficult geography.

In sum, the model tells us that one should expect those geographically constrained metropolitan areas *that we observe in the data* to be more productive or to have higher amenities (Proposition 2) and the correlation between land availability and population size to be zero (Proposition 3), precisely the data patterns found in the preceding section. In addition, due to Proposition 1, one should expect metropolitan areas with lower land availability not only to be more expensive in equilibrium, but also to display *lower housing supply elasticities*, as I will demonstrate in the next sections.

IV. Geography and Housing Price Elasticities

I now move to assessing how important geographic constraints are in explaining local housing price elasticities. Recall from the model that, on the supply side, average housing prices in a city are the sum of construction costs plus land values (themselves a function of the number of housing units): $\widetilde{P}_k = \mathrm{CC} + \mathrm{LC}(H_k)$. Totally differentiating the log of this expression, and manipulating, I obtain

$$d\ln\widetilde{P}_k = \frac{d\mathrm{CC}}{\widetilde{P}_k} + \frac{d\mathrm{LC}(H_k)}{dH_k}\cdot\frac{H_k}{\widetilde{P}_k}\cdot\frac{dH_k}{H_k}.$$

For now, I assume changes in local construction costs to be exogenous to local changes in housing demand: the prices of capital and materials (timber, cement, aluminum, and so on) are determined at the national or international level, and construction is an extremely competitive industry with an elastic labor supply. The assumption is consistent with previous research (Gyourko and Saiz, 2006), but I relax it later. Defining $\sigma_k = \mathrm{CC}/\widetilde{P}_k$ as the initial share of construction costs on housing prices, and assuming that $d\widetilde{P}_k/dH_k = d\mathrm{LC}(H_k)/dH_k$, one obtains $d\ln\widetilde{P}_k = \sigma_k\cdot d\mathrm{CC}/\mathrm{CC} + \beta_k^S\cdot dH_k/H_k$. As defined earlier in the model, β_k^S is the inverse elasticity of housing supply with respect to average home values. I can reexpress this as the empirical log-linearized supply equation: $d\ln\widetilde{P}_k = \sigma_k\cdot d\ln\mathrm{CC} + \beta_k^S\cdot d\ln H_k$. Note that by considering changes in values and quantities, initial scale differences across cities are differenced out (Mayer and Somerville

2000). Throughout the rest of the paper I use long differences (between 1970 and 2000) and hence focus on long-run housing dynamics, as opposed to high-frequency volatility.[8] However, I will also later briefly discuss results at higher (decadal) frequencies. The empirical specification also includes region fixed effects (R_k^j, for $j = 1, 2, 3$) and an error term (ε_k), and estimates the supply equation in discrete changes:

$$(3) \qquad \Delta \ln \widetilde{P}_k = \sigma_k \cdot \Delta \ln \mathrm{CC}_k + \beta_k^S \cdot \Delta \ln H_k + \sum R_k^j + \varepsilon_k.$$

$\widetilde{P}_k$ is measured by median housing prices in each decennial Census.[9] The city-specific parameter σ_k (construction cost share in 1970) is calculated using the estimates in Davis and Heathcote (2007) and Davis and Palumbo (2008) and data on housing prices. Combined with existing detailed information about the growth of construction costs in each city from published sources, the city-specific intercept $\sigma_k \cdot \Delta \ln \mathrm{CC}$ is thus known and calibrated into the model. Changes in the housing stock are, of course, endogenous to changes in prices via the demand side. Therefore, I instrument for $\Delta \ln H_k$ using a shift-share of the 1974 metropolitan industrial composition, the log of average hours of sun in January, and the number of new immigrants (1970 to 2000) divided by the population in 1970. The first variable, as introduced by Bartik (1991) and recently used by Glaeser, Gyourko, and Saks (2006) and Saks (2008), is constructed using early employment levels at the two-digit SIC level and using national growth rates in each industry to forecast city growth due to composition effects. Hours of sun capture a well-documented secular trend of increasing demand for high-amenity areas (Glaeser, Kolko, and Saiz 2001; Rappaport 2007). Finally, previous research (Saiz 2003, 2007; Ottaviano and Peri 2007) has shown international migration to be one of the strongest determinants of the growth in housing demand and prices in a number of major American cities. Immigration inflows

8. Short-run housing adjustments involve considerable dynamic aspects, such as lagged construction responses and serial correlation of high-frequency price changes (Glaeser and Gyourko 2006).

9. A long literature, summarized by Kiel and Zabel (1999), demonstrates that the evolution of self-reported housing prices generally mimics that of actual prices (for a recent confirmation of this fact, see Pence and Bucks [2006]). The correlation between the change in log median census values and the change in the log of the Freddie Mac repeat sales index between 1980 and 2000 is 0.9 across the 147 cities for which the measures were available. The repeat sales index, obtained from Freddie Mac, is unavailable in 1970, and its coverage in our application is limited to the 147 aforementioned cities. Therefore, in this context, I prefer to use the higher coverage of the Census measure.

have been shown to be largely unrelated to other citywide economic shocks, and very strongly associated with the predetermined settlement patterns of immigrant communities (Altonji and Card 1989).

The instruments for demand shocks prove to be strong, with an F-test 47.75 compared to the critical 5% value in Stock and Yogo (2005) of 13.91. The instruments also pass conventional exogeneity tests (with a p-value of .6 in the Sargan–Hansen J test). Note that the specification explicitly controls for all factors that drive physical construction costs. Equation (3) is estimated using 2SLS, with the assumptions $E(\varepsilon_k \cdot Z_k) = 0$, and with Z_k denoting the exogenous variables: the demand instruments, evolution of construction costs, the constant, and regional fixed effects in (3).

In Table III, column (1), I start exploring the data by imposing a common supply inverse-elasticity parameter for all cities ($\beta_k^S = \beta^S \ \forall k$). The estimates of β^S suggest a relatively elastic housing supply on average, with an elasticity of 1.54 (1/0.65). This is well within the range of 1 to 3 proposed by the existing literature at the national level (for a review see Gyourko [2008]). Importantly, unreported regressions where I use each of the demand IV separately always yield similar and statistically significant results.

From the model in Section III, I know that the inverse of supply elasticities should be a function of land availability with $\partial\beta_k/\partial\Lambda_k < 0$. A first-degree linear approximation to this relationship can be posited as $\beta_k^S = \widetilde{\beta}^S + (1 - \Lambda_k) \cdot \beta^{\text{LAND}}$.[10] The supply equation becomes

$$\text{(4)} \qquad \Delta \ln \widetilde{P}_k = \sigma_k \cdot \Delta \ln \text{CC}_k + \widetilde{\beta}^S \cdot \Delta \ln H_k + \beta^{\text{LAND}} \cdot (1 - \Lambda_k) \cdot \Delta \ln H_k + \sum R_k^j + \varepsilon_k.$$

In Table III, column (2), as in all specifications thereafter, $(1 - \Lambda_k)$—the share of area *unavailable* for development—is considered predetermined and exogenous to supply-side shocks in the period 1970–2000. Of course, mountains and coastal status could potentially be drivers for increased housing demand in the period under consideration. Note, however, that equation (4) is

10. Nonlinear versions of the functional relationship between β_k^{LAND} and Λ_k did not add any improvement of economic or statistical significance to the fit of the supply equation in this small sample of 269 cities. Note that the specific functional form of $\partial\beta_k/\partial\Lambda_k$ in the model is driven by the assumptions on the nature of Ricardian land rents: these are solely due to commuting to the CBD, and commuting costs are linear.

TABLE III
HOUSING SUPPLY: GEOGRAPHY AND LAND USE REGULATIONS

	$\Delta\log(P)$ (supply): 1970–2000					
	(1)	(2)	(3)	(4)	(5)	(6)
$\Delta\log(Q)$	0.650	0.336	0.305	0.060		
	(0.107)***	(0.116)***	(0.146)***	(0.215)		
Unavailable land $\times\Delta\log(Q)$		0.560	0.449	0.511	0.516	−5.329
		(0.118)***	(0.140)***	(0.214)***	(0.116)***	(0.904)***
Log(1970 population) $\times$ unavailable land $\times\Delta\log(Q)$						0.481
						(0.117)***
log(WRI)$\times\Delta\log(Q)$				0.237	0.268	0.301
				(0.130)*	(0.068)***	(0.066)***
$\Delta\log(Q)$ $\times$ ocean			0.106			
			(0.065)			
Midwest	−0.099	−0.041	−0.022	−0.015	−0.009	0.002
	(0.054)*	(0.052)	(0.054)	(0.055)	(0.050)	(0.049)
South	−0.236	−0.170	−0.163	−0.129	−0.116	−0.115
	(0.065)***	(0.062)***	(0.062)***	(0.069)*	(0.050)**	(0.048)**
West	0.016	0.057	−0.022	0.059	0.069	0.035
	(0.076)	(0.072)	(0.054)	(0.072)	(0.063)	(0.046)
Constant	0.550	0.594	0.594	0.528	0.601	0.061
	(0.055)***	(0.052)***	(0.052)***	(0.058)***	(0.046)***	(0.045)***

Notes. Standard errors in parentheses. The table shows the coefficient of 2SLS estimation of a metropolitan housing supply equation. On the left-hand side, I try to explain changes in median housing prices by metro area between 1970 and 2000, adjusted for construction costs (see theory and text). On the right-hand side, the main explanatory endogenous variable is the change in housing demand [the log of the number of households − log(Q)] between 1970 and 2000. Some specifications interact that endogenous variable with the unavailable land share (due to geography) and the log of the Wharton Regulation Index (WRI), which we treat as exogenous in this table. The instruments used for demand shocks are a shift-share of the 1974 metropolitan industrial composition, the magnitude of immigration shocks, and the log of January average hours of sun. The identifying assumptions are that the covariance between the residuals of the supply equations and the instruments are zero. *significant at 10%; **significant at 5%; ***significant at 1%.

consistently estimated even if demand shocks $\Delta \ln H_k$ are also correlated with $(1 - \Lambda_k)$. Intuitively, land unavailability can be safely included in both the supply and demand equations insofar as there are enough exclusion restrictions specific to the supply equation.

The results in Table III, column (2), strongly suggest that the impact of demand on prices is mediated by physical land unavailability. Moving within the interquartile range of land unavailability (9% to 39%), the estimates show the impact of demand shocks on prices to increase by about 25%.

Are the results simply capturing the fact that cities with less land availability tend to be coastal? Table III, column (3), allows the impact of demand shocks to vary for coastal and noncoastal areas. Coastal areas are defined as MSAs within 100 km of the ocean (as calculated by Rappaport and Sachs [2003]). Formally $\beta_k^S = \widetilde{\beta}^S + (1 - \Lambda_k) \cdot \beta^{\text{LAND}} + \text{COAST}_k \cdot \beta^{\text{COAST}}$, where COAST is a coastal status dummy. The results show the coastal variable not to be significant. Land unavailability is important within coastal (and noncoastal) areas.

In column (4) of Table III, the inverse elasticity parameter is approximated by a linear function of land use regulations and geographic constraints: $\beta_k^S = \widetilde{\beta}^S + (1 - \Lambda_k) \cdot \beta^{\text{LAND}} + \ln \text{WRI}_k \cdot \beta^{\text{REG}}$. In this specification, $\ln \text{WRI}_k$ stands for the natural log of the WRI.[11] The supply equation becomes

$$\begin{aligned} \Delta \ln \widetilde{P}_k = \sigma_k \cdot \Delta \ln \text{CC}_k + \widetilde{\beta}^S \cdot \Delta \ln H_k + \beta^{\text{LAND}} \cdot (1 - \Lambda_k) \cdot \Delta \ln H_k \\ + \beta^{\text{REG}} \cdot \ln \text{WRI}_k \cdot \Delta \ln H_k + \sum R_k^j + \varepsilon_k. \end{aligned} \tag{5}$$

For now, ln WRI is assumed to be predetermined and exogenous to changes in housing prices through the period 1970–2000. As in all specifications hereafter, I cannot reject that $\widetilde{\beta}^S = 0$: the impact of demand shocks on prices is solely mediated by geographic and regulatory constraints, which is the assumption that I carry forward. In Table III, column (5), I explicitly present results of the model with the constraint $\widetilde{\beta}^S = 0$, which largely leaves the coefficients of interest unchanged.

It is important to remark that independent regressions that consider changes in prices and housing units in the three decades

11. I added three to the original index to ensure that log(WRI) always has positive support, which is consistent with the theoretical predictions of a positive supply parameter across the board. Alternative (unreported) normalizations never had major quantitative impacts on the estimates.

separately (1970s, 1980s, 1990s) cannot reject the coefficients on geography and regulations to be statistically equivalent across decades.[12]

It is apparent that the elasticity of housing supply depends critically on both regulations and physical constraints. However, standard errors on the land unavailability parameter are larger. This can be explained by heterogeneity in how binding physical constraints are. Whereas regulatory constraints matter regardless of the existing level of construction, physical constraints may not be important until the level of development is high enough to render them binding. Using the model in the preceding section, it is straightforward to show that $\partial(\partial\beta_k/\partial\Lambda_k)/\partial \mathrm{POP}_k < 0$: the (negative) impact of land availability on inverse elastiticies should be stronger in larger metro areas. The most parsimonious way to capture this effect is to model the impact of physical constraints on elasticities as an interacted linear function of predetermined initial log population levels. In this specification $\beta_k^S = (1-\Lambda_k)\cdot\beta^{\mathrm{LAND}} + (1-\Lambda_k)\cdot\ln(\mathrm{POP}_{T-1})\cdot\beta^{\mathrm{LAND,POP}} + \ln\mathrm{WRI}\cdot\beta^{\mathrm{REG}}$. Hence the supply equation becomes

$$\text{(6)}\quad \begin{aligned}\Delta\ln\widetilde{P}_k = {} & [\beta^{\mathrm{LAND}} + \beta^{\mathrm{LAND,POP}}\cdot\ln(\mathrm{POP}_{T-1})]\cdot(1-\Lambda_k)\cdot\Delta\ln H_k \\ & + \sigma_k\cdot\Delta\ln\mathrm{CC}_k + \beta^{\mathrm{REG}}\cdot\ln\mathrm{WRI}_k\cdot\Delta\ln H_k + \sum R_k^j + \varepsilon_k.\end{aligned}$$

The results in Table III, column (6), strongly suggest that physical constraints matter more in larger metropolitan areas, consistent with the theory. Figure I depicts the difference in the inverse of β_k^S (that is, the supply elasticity) across the interquartile range of land availability as a function of initial population levels. In the graph, I assign the median level of regulation to all cities in order to create counterfactuals with respect to differences in land unavailability exclusively. At the lowest population levels supply elasticity is mostly determined by regulations: the difference between the seventy-fifth and twenty-fifth percentiles in the distribution of physical land constraints is not large. Nonetheless, geographic constraints become binding and have a strong

12. The average coefficients across decades are $\beta^{\mathrm{LAND}} = 0.29$ and $\beta^{\mathrm{REG}} = 0.21$. Due to the strong mean-reversion of prices at decadal frequencies, the topography coefficient is closer to zero in the 1990s, but larger in the 1980s, whereas the opposite pattern is apparent for the regulation coefficient. They are close to the mean in the 1970s.

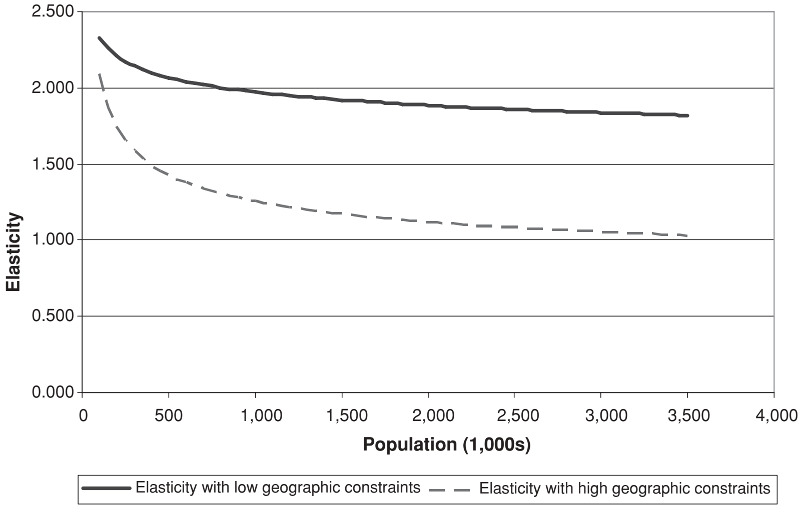

FIGURE I
Impact of Geography on Elasticities by Population

impact on prices as metropolitan population becomes larger. In metropolitan areas above 1,000,000 inhabitants, moving from the twenty-fifth to the seventy-fifth percentile of land unavailability implies supply elasticities that are 40% smaller.

V. The Indirect Effects of Geography

V.A. *Endogenous Regulations*

The previous results confirm the well-known empirical link between land use regulations and housing price growth. Recent examples in this literature include Glaeser, Gyourko, and Saks (2005a, 2005b), Quigley and Raphael (2005), and Saks (2008). However, the existing evidence has arguably not fully established a *causal* link: regulations may be endogenous to the evolution of housing prices.

In the theoretical literature, zoning and growth controls have long been regarded as endogenous devices to keep prices high in areas with valuable land (Hamilton 1975; Epple, Romer, and Filimon 1988; Brueckner 1995). In a review of much of this literature, Fischel (2001) develops the *homevoter hypothesis*, according to which zoning and local land use controls can be largely understood as tools for local homeowners to maximize land prices.

To discuss these issues, consider a stylized version of the supply equation:

$$(7)\qquad \Delta \ln \widetilde{P}_k = \beta_0 + \beta^{\mathrm{REG}} \cdot \ln \mathrm{WRI}_k \cdot \Delta \ln H_k + \overline{\beta} \cdot \Delta \ln H_k + \xi_k.$$

Housing supply inverse elasticities are modeled here as an invariant coefficient ($\overline{\beta}$) plus a linear function of regulatory constraints (the log of the WRI). Assume that, in fact, the local supply elasticity varies for other reasons than regulation that are uncontrolled for in the model,

$$(8)\qquad \Delta \ln \widetilde{P}_k = \beta_0 + \beta^{\mathrm{REG}} \cdot \ln \mathrm{WRI}_k \cdot \Delta \ln H_k + \overline{\beta} \cdot \Delta \ln H_k + \underbrace{\overline{\beta_k^{\delta}} \cdot \Delta \ln H_k + \eta_k}_{\xi_k},$$

where $\overline{\beta_k^{\delta}}$ is a local deviation from average supply elasticities unrelated to regulation. Even with suitable instruments for $\Delta \ln H_k$, consistent estimates will not be obtained if $\ln \mathrm{WRI}_k$ is correlated with ξ_k. Consider as a working hypothesis the following empirical equation describing the optimal choice of voters with regard to land use policies:

$$(9)\qquad \ln \mathrm{WRI}_k = \varphi_0 + \varphi_1 \cdot \overline{\beta_k^{\delta}} + \varphi_2 \cdot \overline{\beta_k^{\delta}} \cdot \Delta \ln H_k + \varphi_3 \cdot \ln \widetilde{P}_k + \mu_k.$$

What are the potential sources of regulation endogeneity in equation (9), which includes an independent error term denoted by μ_k? In Ortalo-Magné and Prat (2007), voters may explicitly restrict the supply of land in order to keep its value high, but only have an incentive to do so in areas where land was initially dear. The only source of supply constraints in Ortalo-Magné and Prat (2007) comes from regulation, but there are additional reasons that in areas that were initially land-constrained voters may want *further* limits on development (implying $\varphi_1 > 0$ in equation (9)). Consider the problem of a voter trying to maximize future land price growth. From the model in Section II, equilibrium housing prices in an initial steady state may be obtained as a function of local amenity–productivity levels. Assume now that we introduce some uncertainty about future amenity–productivity shocks, which are assumed to be uncorrelated with factors that condition initial population, such as geographic land availability (Gabaix 1999). In this context, expected changes to housing prices ($E(\Delta \widetilde{P}_k)$) are a function of expected productivity shocks ($E(\Delta \chi_k)$), as mediated by land availability. It is staightforward to show (see

Derivation 3 in Appendix II) that $dE(\Delta\widetilde{P}_k)/d\Lambda_k < 0$. Reduced land availability amplifies the effects of productivity shocks on home values. Conversely, productivity shocks largely translate into population growth in unconstrained cities.

Moreover, $d^2E(\Delta\widetilde{P}_k)/(d\Lambda_k)^2 > 0$: the *marginal impact* of additional land constraints on expected price growth is *larger in areas that already had lower land availability* initially. The intuition for this result comes from the geometry of land development. Recall from the model that the average city radius corresponds to $\Phi_k = \sqrt{\gamma \mathrm{POP}_k/\Lambda_k\pi}$; decreasing land availability has a stronger impact in pushing away the city boundary at low initial values, thereby further increasing Ricardian land rents. In the presence of positive marginal costs of restrictive zoning, voters in land-constrained regions have more of an incentive to pass such regulations. Conversely, marginal changes in zoning regulations do not have much of an expected impact on home values in areas where land is naturally abundant, thereby reducing their strategic value.

Furthermore, strategic growth-management considerations should be less of an issue in shrinking cities, where new constraints on growth are not binding, suggesting also that $\varphi_2 > 0$.

Restrictive land use policies are not exclusively enacted in order to limit the supply of housing, however. Citizens' demands for antigrowth regulations partially stem from the perceived nuisances of development, such as increased traffic, school congestion, and aesthetic impact on the landscape (Rybczynski 2007). These issues only arise in growing cities, and may be more salient in congested areas, where population densities are initially high. Therefore, restrictive nuisance zoning may be more prevalent in growing, land-constrained metro areas, which implies again that $\varphi_2 > 0$.

The existing literature offers additional reasons to expect reverse causality from growing prices to higher regulations ($\varphi_3 > 0$ in equation (9)). Recent examples include Fischel (2001) and Hilber and Robert-Nicoud (2006), who argue for a demand-side link from higher prices to increased growth controls. Several mechanisms have been identified that imply such a reverse causal link.

Rational voters may want to enact restrictive zoning policies in regions with valuable land *even when they do not aim to increase metropolitan housing prices*. Changes in the future local best-and-highest use of land are highly uncertain. Such uncertainty generates considerable wealth *risk* for homeowners who are unsure about the nature of future neighborhood change (Breton 1973). Therefore "since residents cannot insure against neighborhood change, zoning offers a kind of second-best institution" (Fischel 2001, p. 10). In regions with high land values, voters limit the scope and extent of future land development in their jurisdiction in order to reduce housing wealth risk. Because all jurisdictions in a region try to deflect risks and compete à la Tiebout, the equilibrium outcome at the metropolitan level implies stricter development constraints everywhere. Conversely, concerns about the variability of land values are absent in regions where home prices are close to, and pinned down by, structural replacement costs.

Similarly, voters have vested interests in fiscal zoning (Hamilton 1975, 1976). In areas with very cheap land, development usually happens at relatively low densities. However, as land values in a metropolitan area or jurisdiction increase, new entrants into the community want to consume less land. Simultaneously, in metropolitan areas where the land input is relatively expensive, developers want to use less of it and build at higher densities. However, existing homeowners do not want new arrivals to pay lower-than-average taxes, which may induce them to mandate large lot sizes on new development. According to the fiscal-zoning theories, land use regulations should become more restrictive in areas with expensive land.

In order to see whether the above theories have empirical content, I start by asking whether natural geographic constraints beget regulatory constraints. Table IV, column (1), displays regressions similar to equation (9) with the log of the WRI on the left-hand side. The main explanatory variable is the measure of undevelopable area. Geographic constraints were strongly associated with regulatory constraints in 2005, evidence consistent with $\varphi_1 > 0$ in equation (9). The regression includes other controls, such as regional fixed effects, the percentage of individuals older than 25 with a bachelor's degree, and lagged white non-Hispanic shares.[13]

Regardless of the evolution of local housing markets, there are regional differences in the propensity of local governments to regulate economic activity (Kahn 2002). As a proxy for preferences for

13. A previous working paper version (Saiz 2008) explored other potential correlates of land use regulations across metropolitan areas. Alternative hypotheses based on local politics, optimal regulation of externalities, and snob-zoning do not change the importance of reverse causation and original land constraints to account for regulations and are never quantitatively large.

governmental activism (as opposed to laissez-faire), regressions in Table IV control for the log of the public expenditure on protective inspection and regulation by local governments at the MSA level as a share of total public revenues. The government expenditure category "Protective inspection and regulation" in the Census of Governments includes local expenditures in building inspections; weights and measures; regulation of financial institutions; taxicabs; public service corporations; private utilities; licensing, examination, and regulation of professional occupations; inspection and regulation or working conditions; motor vehicle inspection and weighting; and regulation and enforcement of liquor laws and sale of alcoholic beverages. As expected, areas that tended to regulate economic activity in other spheres also regulated residential land development more strongly.

Regressions in Table IV also control for the share of Christians in nontraditional denominations in 1970, defined as one minus the Catholic and mainline protestant Christian shares.[14] Political scientists, economists, and historians of religion have claimed that the ethics and philosophy of nontraditional Christian denominations (especially those self-denominated Evangelical) are deeply rooted in individualism and the advocacy of limited government role.[15] Column (1) in Table IV (which controls for region fixed effects) finds that a one–standard deviation increase in the nontraditional Christian share in 1970 was associated with a -0.21-standard deviation change in land use regulations.

In column (2) of Table IV, I examine another source of endogeneity in equation (9), namely the possibility that $\varphi_2 > 0$. Land-constrained areas that have been declining or stagnating for a long time do not seem to display strong antigrowth policies. Consider the case of Charleston, West Virginia: 71% of its 50-km radius area is undevelopable according to our measure, yet the WRI's value is -1.1. Similar examples are New Orleans (LA), Asheville (NC), Chattanooga (TN), Elmira (NY), Erie (PA), and Wheeling (WV). In order to capture the fact that antigrowth regulations may not be important in declining areas, I interact the geographic-constraints variable with a dummy for MSA in the bottom quartile of urban growth between 1940 and 1970 (column (2) in Table IV).

14. Mainline Protestant denominations are defined as United Church of Christ, American Baptist, Presbyterian, Methodist, Lutheran, and Episcopal.

15. See Moberg (1972), Hollinger (1983), Magleby (1992), Holmer Nadesan (1999), Kyle (2006), Barnett (2008), and Swartz (2008). Crowe (2009) points to a negative correlation between housing price volatility and the Evangelical share, which could be explained by looser land use regulations in Evangelical areas.

TABLE IV
ENDOGENEITY OF LAND USE REGULATIONS

	Log WRI			
	(1)	(2)	(3)	(4)
Unavailable land, 50-km radius	0.134		−0.174	−0.241
	(0.067)**		(0.125)	(0.132)*
Unavailable land in growing cities (1940–1970)		0.165		
		(0.069)**		
Unavailable land in declining cities (1940–1970)		−0.054		
		(0.153)		
Declining cities dummy (1940–1970)		−0.076		
		(0.051)		
Unavailable land, 50-km radius × Δlog housing units (1970–2000)			0.451	0.375
			(0.158)***	(0.154)**
ΔLog housing price (1970–2000) = log housing price (1970)				0.198
				(0.088)**
Log (inspection expenditures/local tax revenues) (1982)	0.051	0.047	0.051	0.041
	(0.015)***	(0.015)***	(0.015)***	(0.015)***
Share of Christian “nontraditional” denominations (1970)	−0.308	−0.304	−0.314	−0.291
	(0.086)***	(0.084)***	(0.084)***	(0.090)***
Share with bachelor’s degree in 1970	0.983	0.538	0.867	0.089
	(0.332)***	(0.342)	(0.328)***	(0.404)
Non-Hispanic white share in 1980	0.036	0.08	−0.069	−0.017
	(0.113)	(0.110)	(0.116)	(0.120)

TABLE IV
(CONTINUED)

	Log WRI			
	(1)	(2)	(3)	(4)
Midwest	−0.266	−0.307	−0.289	−0.266
	(0.039)***	(0.039)***	(0.039)***	(0.044)***
South	−0.19	−0.222	−0.261	−0.196
	(0.054)***	(0.053)***	(0.058)***	(0.066)***
West	−0.029	−0.08	−0.096	−0.088
	(0.050)	(0.050)	(0.054)*	(0.056)
Constant	1.425	1.471	1.578	−0.759
	(0.137)***	(0.135)***	(0.144)***	(1.055)
Observations	269	269	269	269
R^2	.43	.46	—	—
Method	FGLS	FGLS	2SLS	3SLS

Notes. Standard errors in parentheses. The dependent variable in all regressions is the log of the WRI for each metro area. To deal with heterogeneous sample sizes (strong correlation of WRI values within MSA) columns (1) and (2) use a Feasible Generalized Least Squares (FGLS) procedure, where each observation is weighted proportionally to the inverse of the square error of OLS estimates (which are actually always very close in magnitude and significance). In columns (3) and (4), changes in the log of local housing prices and quantities are instrumented using the demand shocks in Table III (industry shift-share, hours of sun, and immigrant shocks), plus the land unavailability variable. *significant at 10%; **significant at 5%; ***significant at 1%.

Lagged growth rates in a period that is, on the average, 45 years in the past are unlikely to be caused by the regulation environment in 2005. But they are likely to be good predictors of future growth, because of the permanence of factors that drove productivity during the second half of the 20th century, such as reliance on manufacturing or mining or relative scarcity of institutions of higher education. Similarly, in column (3) of Table IV, I interact the change in housing growth *between 1970 and 2000* with the geographic land-unavailability variable. Of course, housing construction is endogenous to regulations in this equation. Hence I use the demand shock instruments in Table III and interactions with geographic land unavailability as instrumental variables for the interacted endogenous variable. The results suggest that regulations are stricter in land-constrained metro areas that are thriving ($\varphi_2 > 0$). In declining cities, however, regulations are insensitive to previous factors that made housing supply inelastic.

Finally, in column (4) of Table IV, I test for reverse causation from price levels to higher regulation ($\varphi_3 > 0$ in equation (9)). Because $P_t = \Delta P_{t,t-n} + P_{t-n}$, I express the log of housing values in 2000 as the sum of the change in the log of prices plus the log of initial prices in 1970 (for comparability with Table III) and constrain the coefficient on both variables to be the same.[16] The instruments now are hours of sun, immigration shocks, and the Bartik (1991) employment shift-share and their interactions with geographic land unavailability. There are two endogenous variables: lagged changes in housing prices, and household growth interacted by the geographic constraints. The equation is estimated via 3SLS and strongly suggests that *both* a constraining geography *in growing cities* and higher housing prices led to a more regulated supply environment circa 2005.

In sum, the regulation equations in Table IV demonstrate that higher housing prices, demographic growth, and natural constraints beget more restrictive land-use regulations.

V.B. Endogeneizing Regulations in the Supply Equation

Because regulations are endogenous to ε_k in equations (5) and (6), one needs to use additional identifying exclusions to estimate housing supply elasticities. As suggested by the results in Table IV, the local public expenditure share in protective inspection and

16. In unconstrained equations, I cannot reject that the separate coefficients on $\Delta P_{2000,1970}$ and P_{1970} are statistically equivalent.

the nontraditional Christian share in 1970 can be used as instruments for the 2005 WRI: although they predict land use regulations, they are unlikely to impact land supply otherwise (note that the supply equation controls for the evolution of construction costs). As seen in Table IV, these variables prove also to be strong instruments.[17] Note that even if these variables were correlated with demand shocks, the regression have more supply-specific exclusion restrictions than endogenous variables and all parameters are fully identified. In fact, because the two endogenous variables appear in interacted form, I can now also include in the IV list the interactions of the instruments used for changes in quantities (hours of sun, employment shift-share, and immigration shocks) with those used for the regulation index (municipal inspections expenditure share and nontraditional Christian share). Importantly, *the results are very similar when I simply use each one of the regulation instruments separately.*

Column (1) in Table V reestimates the specification in Table III, column (5) (elasticities as linear functions of regulations and geographic constraints), this time allowing for endogenous regulations. The coefficient on the WRI declines to about 60% of its previous value. However, when the model in equation (6) is reestimated (land constraints matter more in large cities), the coefficient on the regulation index takes a value that is only 8% smaller than in the earlier estimates. Therefore, parameters from previous research are bound to somewhat overestimate the impact of regulations on prices, but it is still true that more regulated areas tend to be relatively more inelastic, and this impact is quantitatively large. In Table V, column (2), a move across the interquartile range in the WRI of a city of one million inhabitants with average land availability is associated with close to a 20% reduction in supply elasticity: from 1.76 to 1.38.

The impact of constrained geography is larger, especially in larger cities. For example, in a metro area with average regulations and a population of one million, the interquartile change in the share of unavailable land (from 0.09 to 0.38) implies a 50% reduction in supply elasticity (from 2.45 to 1.25).

In a separate Online Appendix, the interested reader can further see that endogeneizing construction costs (which could be themselves a function of geography) and immigration shocks does not change the main parameters of interest.

17. Partial R^2 of .074 in the first stage and F-test of 10.413, above the 20% maximal bias threshold (8.75) in Stock and Yogo (2005).

TABLE V
HOUSING SUPPLY: ENDOGENOUS REGULATIONS

	$\Delta\log(P)$ (supply)	
	(1)	(2)
Unavailable land $\times\Delta\log(Q)$	0.581	−5.260
	(0.119)***	(1.396)***
Log(1970 population) × unavailable land $\times\Delta\log(Q)$		0.475
		(0.119)***
Log(WRI) $\times\Delta\log(Q)$	0.109	0.280
	(0.078)*	(0.077)***
Midwest	−0.009	0.002
	(0.049)	(0.048)
South	−0.075	−0.109
	(0.049)	(0.049)**
West	0.149	0.059
	(0.063)	(0.065)
Constant	0.659	0.577
	(0.048)***	(0.048)***

Notes. Standard errors in parentheses. The table shows the coefficient of 2SLS estimation of a metropolitan housing supply equation. The specification and instruments used for demand shocks are as in Table III. Demand shocks are interacted with the unavailable land share (due to geography) and the log of the WRI. The latter variable is treated as endogenous using the share of local public expenditures on "protective inspections" and the share of nontraditional Christian denominations as instruments. Because we are instrumenting for log(WRI)$\times\Delta$ log(Q), I also include the interaction between the regulation and the demand instruments in the IV list. *significant at 10%; **significant at 5%; ***significant at 1%.

V.C. Estimated Elasticities

In this section, I use the coefficients in Table V, column (2), to estimate supply elasticities at the metro area level. Such estimates are simple nonlinear combinations of the available data on physical and regulatory constraints, and predetermined population levels in 2000. These elasticities are thus based on economic fundamentals related to natural and man-made land constraints and should prove useful in calibrating general equilibrium models of interregional labor mobility and in predicting the response of housing markets to future demand shocks.

The population-weighted average elasticity of supply is estimated to be 1.75 in metropolitan areas (2.5 unweighted). The results for metropolitan areas with population over 500,000 in 2000 can be found in Table VI. Estimated elasticities using only the geographic, regulatory, and initial population variables agree with perceptions about supply-constrained areas. Miami, Los Angeles, San Francisco, Oakland, New York, San Diego, Boston, Chicago, and Seattle are among the top fifteen in the list of the most inelastic cities. Houston, Austin, Charlotte, Kansas City, and Indianapolis are among the large metro areas with highly elastic housing supply.

Estimated elasticities (this time using predetermined 1970 population in order to avoid obvious endogeneity issues) also correlate very strongly with housing price levels in 2000 and changes over the 1970–2000 period. Figure II presents plots relating housing prices (Panel 1) or changes (Panel 2) on the vertical axis and the inverse of the estimated supply elasticity by metropolitan area on the horizontal axis. It is clear that a simple linear combination of physical and regulatory constraints goes very far to explain the evolution of prices, even without taking into account the differential demand shocks that cities experienced.

VI. Conclusion

The paper started by providing empirical content to the concept of land availability in metropolitan areas. Using satellite-generated data, I calculated an exact measure of land unavailable for real estate development in the metropolitan United States. This geographic measure can be used in future work exploring topics as diverse as housing and mortgage markets, labor mobility, urban density, transportation, and urban environmental issues.

I then developed a model for the impact of land availability on urban development and housing prices. In *ex post* equilibrium, land-constrained metro areas should have more expensive housing and enjoy higher amenities or productivity, as confirmed by the data. The model demonstrates that land constraints should also decrease housing supply *elasticities*, a somewhat *ad hoc* assumption in previous literature.

Empirically, most areas that are widely regarded as supply-inelastic were found, in fact, to be severely land-constrained by their geography. Deploying a new comprehensive survey on residential land use regulations, I found that highly regulated areas tend to be geographically constrained also. More generally, I found recent housing price and population growth to be predictive of more restrictive residential land regulations. The results point to the endogeneity of land use controls with respect to the housing market equilibrium.

Hence I next estimated a model where regulations are both causes and consequences of housing supply inelasticity. Housing

TABLE VI
SUPPLY ELASTICITIES (METRO AREAS WITH POPULATION > 500,000)

Rank	MSA/NECMA name	Supply elasticity	Rank	MSA/NECMA name	Supply elasticity
1	Miami, FL	0.60	26	Vallejo–Fairfield–Napa, CA	1.14
2	Los Angeles–Long Beach, CA	0.63	27	Newark, NJ	1.16
3	Fort Lauderdale, FL	0.65	28	Charleston–North Charleston, SC	1.20
4	San Francisco, CA	0.66	29	Pittsburgh, PA	1.20
5	San Diego, CA	0.67	30	Tacoma, WA	1.21
6	Oakland, CA	0.70	31	Baltimore, MD	1.23
7	Salt Lake City–Ogden, UT	0.75	32	Detroit, MI	1.24
8	Ventura, CA	0.75	33	Las Vegas, NV–AZ	1.39
9	New York, NY	0.76	34	Rochester, NY	1.40
10	San Jose, CA	0.76	35	Tucson, AZ	1.42
11	New Orleans, LA	0.81	36	Knoxville, TN	1.42
12	Chicago, IL	0.81	37	Jersey City, NJ	1.44
13	Norfolk–Virginia Beach–Newport News, VA–NC	0.82	38	Minneapolis–St. Paul, MN–WI	1.45
14	West Palm Beach–Boca Raton, FL	0.83	39	Hartford, CT	1.50
15	Boston–Worcester–Lawrence–Lowell–Brockton, MA–NH	0.86	40	Springfield, MA	1.52
16	Seattle–Bellevue–Everett, WA	0.88	41	Denver, CO	1.53
17	Sarasota–Bradenton, FL	0.92	42	Providence–Warwick–Pawtucket, RI	1.61
18	Riverside–San Bernardino, CA	0.94	43	Washington, DC–MD–VA–WV	1.61
19	New Haven–Bridgeport–Stamford–Danbury–Waterbury, CT	0.98	44	Phoenix–Mesa, AZ	1.61
20	Tampa–St. Petersburg–Clearwater, FL	1.00	45	Scranton–Wilkes-Barre–Hazleton, PA	1.62
21	Cleveland–Lorain–Elyria, OH	1.02	46	Harrisburg–Lebanon–Carlisle, PA	1.63
22	Milwaukee–Waukesha, WI	1.03	47	Bakersfield, CA	1.64
23	Jacksonville, FL	1.06	48	Philadelphia, PA–NJ	1.65
24	Portland–Vancouver, OR–WA	1.07	49	Colorado Springs, CO	1.67
25	Orlando, FL	1.12	50	Albany–Schenectady–Troy, NY	1.70

TABLE VI
(CONTINUED)

Rank	MSA/NECMA name	Supply elasticity	Rank	MSA/NECMA name	Supply elasticity
51	Gary, IN	1.74	74	Atlanta, GA	2.55
52	Baton Rouge, LA	1.74	75	Akron, OH	2.59
53	Memphis, TN–AR–MS	1.76	76	Richmond–Petersburg, VA	2.60
54	Buffalo–Niagara Falls, NY	1.83	77	Youngstown–Warren, OH	2.63
55	Fresno, CA	1.84	78	Columbia, SC	2.64
56	Allentown–Bethlehem–Easton, PA	1.86	79	Columbus, OH	2.71
57	Wilmington–Newark, DE–MD	1.99	80	Greenville–Spartanburg–Anderson, SC	2.71
58	Mobile, AL	2.04	81	Little Rock–North Little Rock, AR	2.79
59	Stockton–Lodi, CA	2.07	82	Fort Worth–Arlington, TX	2.80
60	Raleigh–Durham–Chapel Hill, NC	2.11	83	San Antonio, TX	2.98
61	Albuquerque, NM	2.11	84	Austin–San Marcos, TX	3.00
62	Birmingham, AL	2.14	85	Charlotte–Gastonia–Rock Hill, NC–SC	3.09
63	Dallas, TX	2.18	86	Greensboro–Winston–Salem–High Point, NC	3.10
64	Syracuse, NY	2.21	87	Kansas City, MO–KS	3.19
65	Toledo, OH	2.21	88	Oklahoma City, OK	3.29
66	Nashville, TN	2.24	89	Tulsa, OK	3.35
67	Ann Arbor, MI	2.29	90	Omaha, NE–IA	3.47
68	Houston, TX	2.30	91	McAllen–Edinburg–Mission, TX	3.68
69	Louisville, KY–IN	2.34	92	Dayton–Springfield, OH	3.71
70	El Paso, TX	2.35	93	Indianapolis, IN	4.00
71	St. Louis, MO–IL	2.36	94	Fort Wayne, IN	5.36
72	Grand Rapids–Muskegon–Holland, MI	2.39	95	Wichita, KS	5.45
73	Cincinnati, OH–KY–IN	2.46			

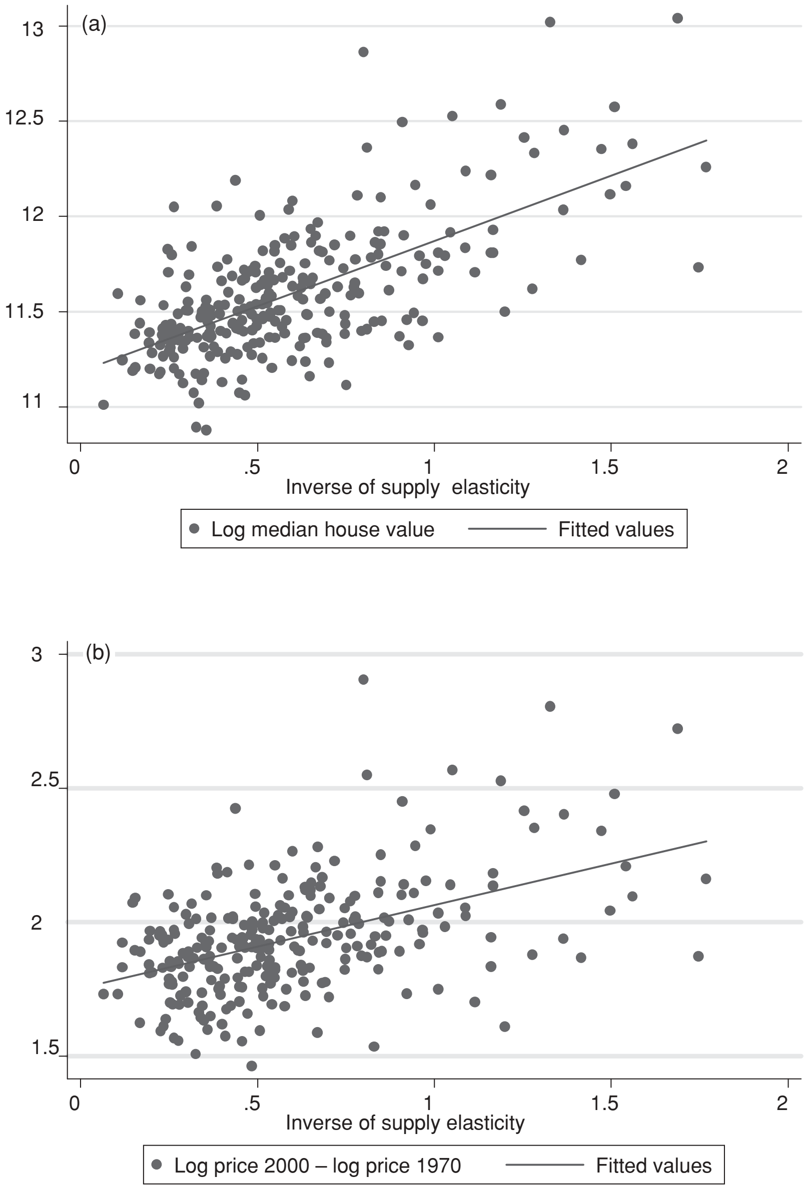

FIGURE II
Estimated Elasticities and Home Values (2000)
(a) Levels, (b) changes.

demand, construction, and regulations are all determined endogenously. Housing supply elasticities were found to be well characterized as functions of both physical and regulatory land constraints, which in turn are endogenous to prices and past growth.

Geography was shown to be one of the most important determinants of housing supply inelasticity: directly, via reductions in the amount of land availability, and indirectly, via increased land values and higher incentives for antigrowth regulations. The results in the paper demonstrate that geography is a key factor in the contemporaneous urban development of the United States, and help us understand why robust national demographic growth and increased urbanization has translated mostly into higher housing prices in San Diego, New York, Boston, and Los Angeles, but into rapidly growing populations in Atlanta, Phoenix, Houston, and Charlotte.

APPENDIX I
DESCRIPTIVE STATISTICS

	Mean (standard dev.)
Log population in 2000	12.893
	(1.060)
Log median house value in 2000	11.592
	(0.342)
Δlog median house value (1970–2000)	1.937
	(0.213)
Log income in 2000	10.200
	(0.184)
Δlog(income per capita) (1990–2000)	0.401
	(0.063)
Log population (1990–2000)	0.123
	(0.099)
Immigrants (1990–2000)/population (2000)	0.034
	(0.038)
Share with bachelor's degree (2000)	0.198
	(0.063)
Share workers in manufacturing (2000)	0.174
	(0.071)
Log(patents/population) (2000)	−8.978
	(0.866)
January monthly hours of sun (average 1941–1970)	151.342
	(38.199)
Log tourist visits per person (2000)	−12.679
	(0.830)

APPENDIX I
(CONTINUED)

	Mean (standard dev.)
Ocean dummy	0.331 (0.471)
Unavailable land, 50-km radius	0.261 (0.212)
Log(WRI)	1.025 (0.278)
Δlog housing units (1970–2000)	0.599 (0.319)
Log housing price (1970)	9.655 (0.228)
Log (inspection expenditures/local tax revenues)	−5.826 (0.971)
Share of Christian “nontraditional” denominations	0.351 (0.209)
Share with bachelors degree in 1970	0.111 (0.042)
Non-Hispanic white share in 1980	0.827 (0.138)
Midwest	0.264 (0.442)
South	0.383 (0.487)
West	0.201 (0.401)
Unionization in construction sector	0.208 (0.146)
Δlog(income per capita) (1970–2000)	1.965 (0.116)

APPENDIX II: DERIVATIONS AND PROOFS

Derivation 1. First, note that a share of $2\pi d\Lambda_k/\gamma \text{POP}_k$ households live in the sector of the circle at a distance d from the CBD. Average housing rents in the city, conditional on population, can thus be obtained as $\widetilde{r}_k = \int_0^{\Phi_k}(2\pi x\Lambda_k/\gamma\text{POP}_k)\cdot r(x)\cdot dx$, which implies that $\widetilde{r}_k = (2\pi\Lambda_k/\gamma\text{POP}_k)\int_0^{\Phi_k}(r_0x - tx^2)dx$, and so $\widetilde{r}_k = (2\pi\Lambda_k/\gamma\text{POP}_k)\cdot[\frac{1}{2}r_0x^2 - \frac{1}{3}tx^3]_0^{\Phi_k}$. Therefore $\widetilde{r}_k = (2\pi\Lambda_k/\gamma\text{POP}_k)\cdot[\frac{1}{2}r_0\Phi_k^2 - \frac{1}{3}t\Phi_k^3] = (\Phi_k^2\pi\Lambda_k/\gamma\text{POP}_k)\cdot[r_0 - \frac{2}{3}t\Phi_k] = \{[(\gamma\text{POP}_k/(\Lambda_k\pi))\cdot\pi\Lambda_k]/\gamma\text{POP}_k\}\cdot[r_0 - \frac{2}{3}t\Phi_k] = [r_0 - t\frac{2}{3}\Phi_k]$, which corresponds to rents in the location that is two-thirds of the way between the

CBD and the city's fringe. Substituting for the value of $r(\frac{2}{3}\Phi_k)$ yields $\widetilde{r}_k = i\text{CC} + \frac{1}{3}t\sqrt{\gamma \text{POP}_k/\Lambda_k \pi}$.

Derivation 2. Recall that

$$U(C_k) = (A_k + w_k - r - td)^\rho = 0. \tag{10}$$

Substituting into the intercity spatial equilibrium equation, I obtain $r(\text{POP}_k, d) = \widetilde{A}_k + \widetilde{w}_k - (\psi + \alpha)\sqrt{\text{POP}_k} - td$. Because all consumers are indifferent, I can focus w.o.l.o.g. on consumers living in the CBD. Recalling that $H_k = \text{POP}_k$ yields $r_0 = \widetilde{A}_k + \widetilde{w}_k - (\psi + \alpha)\sqrt{H_k}$. Defining $P(0) = r_0/i$, one obtains the demand schedule for housing in the city:

$$\sqrt{H_k} = \frac{\widetilde{A}_k + \widetilde{w}_k}{(\psi + \alpha)} - \frac{i}{(\psi + \alpha)} P(0). \tag{11}$$

Note also that changes in $P(0)$ shift all prices within a city vertically by the same amount and so, denoting $\widetilde{P}_k$ as the average housing price in city k, the city demand equation implies that $\partial \ln(H_k)/\partial \widetilde{P}_k = \partial \ln(H_k)/\partial P(0)$. Now recall the expression for rents in the CBD from the supply of land: $r_0 = i\text{CC} + t\sqrt{\gamma H_k/\Lambda_k \pi}$, which implies that $P(0) = \text{CC} + \frac{t}{i}\sqrt{\gamma H_k/\Lambda_k \pi}$.

I can combine this supply-side price equation at the CBD with equation (11) to obtain $\widetilde{A}_k + \widetilde{w}_k - (\psi + \alpha)\sqrt{H_k} = i\text{CC} + t(\sqrt{\gamma/\Lambda_k \pi})\sqrt{H_k}$. Solving for housing yields

$$H_k = \left(\frac{\widetilde{A}_k + \widetilde{w}_k - i\text{CC}}{(\psi + \alpha) + t\sqrt{\frac{\gamma}{\Lambda_k \pi}}} \right)^2.$$

Proof of Proposition 1. The city-specific inverse elasticity of supply is $\beta_k^S = \partial \ln \widetilde{P}_k/\partial \ln H_k = \frac{1}{2}[\frac{1}{3i}t(\sqrt{\gamma H_k/\Lambda_k \pi})/\widetilde{P}_k]$, and therefore

$$\frac{\partial \beta_k}{\partial \Lambda_k} = \frac{\partial^2 \ln \widetilde{P}_k}{\partial \ln H_k \partial \Lambda_k} = -\frac{1}{4} \frac{\frac{1}{3i}t\sqrt{\frac{\gamma H_k}{\Lambda_k^3 \pi}}\text{CC}}{(\widetilde{P}_k)^2} < 0. \tag{12}$$

Proof of Proposition 2. I focus on relevant joint amenity and productivity shocks net of annuitized construction costs that are compatible with habitation: $\chi_k \equiv \widetilde{A}_k + \widetilde{w}_k - i\text{CC} > 0$. I further normalize the minimum city size that classifies a population center as metropolitan to one ($\text{POP}_k = H_k = 1$). The unit of population measurement could be, for instance, 50,000 people, which is the actual population level that qualifies an urban area for metropolitan status in the United States. The minimum necessary net wage–amenity shock observed in metropolitan areas ($\underline{\chi}$) is obtained with $\Lambda_k = 1$ (all land is developable) and therefore $\underline{\chi} = (\psi + \alpha) + t\sqrt{\gamma/\pi}$. Similarly, I denote the minimum amenity–productivity shock that a city with land availability (Λ_j) requires to reach metropolitan status as $\overline{\chi(\Lambda_j)} = \underline{\chi} + t\sqrt{\gamma/\pi}[(1/\sqrt{\Lambda_j}) - 1]$. Start by defining $\varepsilon(\Lambda_j) = \underline{\chi(\Lambda_j)} - \underline{\chi}$, to obtain $\varepsilon_j = t\sqrt{\gamma/\pi}[(1/\sqrt{\Lambda_j}) - 1]$. By assumption, conditional on qualifying as a metropolitan area, amenity–productivity shocks in land-unconstrained cities ($\Lambda_k = 1$) are drawn from the Pareto cdf: $f(\chi/\chi \geq \underline{\chi}, \lambda) = \lambda\underline{\chi}^{\lambda}/\chi^{\lambda+1}$, with $\lambda > 2$. Thus the expected value of shocks in such cities is $E(\chi/\chi \geq \underline{\chi}, \lambda) = \lambda\underline{\chi}/(\lambda - 1)$.

In turn, amenity–productivity shocks in land-constrained metropolitan areas with $\Lambda_j < 1$ will be drawn (*ex post*) from distributions with support $[\underline{\chi} + \varepsilon(\Lambda_j), \infty]$. The Pareto cdf implies that $F(\underline{\chi} + \varepsilon(\Lambda_j)) = 1 - (\underline{\chi}/\underline{\chi} + \varepsilon(\Lambda_j))^{\lambda}$, and so the upper tail truncated at $\underline{\chi} + \varepsilon(\Lambda_j)$ has mass $(\underline{\chi}/\underline{\chi} + \varepsilon(\Lambda_j))^{\lambda}$. Therefore

$$f(\chi/\chi \geq \underline{\chi} + \varepsilon(\Lambda_j)) = \frac{\lambda\underline{\chi}^{\lambda}}{\chi^{\lambda+1}} \Bigg/ \left(\frac{\underline{\chi}}{\underline{\chi} + \varepsilon(\Lambda_j)}\right)^{\lambda}$$
$$= \lambda(\underline{\chi} + \varepsilon(\Lambda_j))^{\lambda}/\chi^{\lambda+1},$$

which is itself Pareto distributed. Note that $E(\chi/\chi \geq \underline{\chi} + \varepsilon(\Lambda_j)) = E(\chi/\mathrm{POP}_j \geq 1\Lambda_j)$, and therefore $E(\chi/\mathrm{POP}_j \geq 1, \Lambda_j) = \lambda[\underline{\chi} + \varepsilon_j(\Lambda_j)]/[\lambda - 1]$, which is a decreasing function in land availability.

Proof of Proposition 3. Recall that $\mathrm{POP}_k = \{(\chi_k/[(\psi + \alpha) + t\sqrt{\frac{\gamma}{\Lambda_k \pi}})\}^2$. Using the relevant pdf:

$$E(\mathrm{POP}_j/\mathrm{POP}_j \geq 1, \Lambda_j)$$
$$= \int_{\underline{\chi}+\varepsilon_j(\Lambda_j)}^{\infty} \frac{\lambda[\underline{\chi} + \varepsilon_j(\Lambda_j)]^{\lambda}}{z^{\lambda+1}} \cdot \left(\frac{z}{(\psi + \alpha) + t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2 dz,$$

$$E(\text{POP}_j/\text{POP}_j \geq 1, \Lambda_j)$$
$$= \left(\frac{1}{(\psi+\alpha)+t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2 \int_{\underline{\chi}+\varepsilon_j(\Lambda_j)}^{\infty} \lambda[\underline{\chi}+\varepsilon_j(\Lambda_j)]^{\lambda} \cdot z^{1-\lambda} \cdot dz,$$
$$E(\text{POP}_j/, \text{POP}_j \geq 1, \Lambda_j)$$
$$= \left(\frac{1}{(\psi+\alpha)+t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2 \left[\frac{\lambda}{2-\lambda} \cdot (\underline{\chi}+\varepsilon_j(\Lambda_j))^{\lambda} \cdot z^{2-\lambda}\right]_{\underline{\chi}+\varepsilon_j(\Lambda_j)}^{\infty}$$
$$= \left(\frac{1}{(\psi+\alpha)+t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2 \left(0+\frac{\lambda}{\lambda-2} \cdot (\underline{\chi}+\varepsilon_j(\Lambda_j))^{\lambda} \cdot (\underline{\chi}+\varepsilon_j(\Lambda_j))^{2-\lambda}\right)$$
$$= \left(\frac{1}{(\psi+\alpha)+t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2 \left(\frac{\lambda}{\lambda-2} \cdot (\underline{\chi}+\varepsilon_j(\Lambda_j))^{2}\right)$$
$$= \left(\frac{(\underline{\chi}+\varepsilon_j(\Lambda_j))}{(\psi+\alpha)+t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2 \frac{\lambda}{\lambda-2}.$$

Because the first part of the equation defines the minimum population level normalized at one: $E(\text{POP}_j/\text{POP}_j \geq 1, \Lambda_j) = \lambda/(\lambda-2)$.

Derivation 3. Recall the equilibrium population level:

$$H_k^* = \left(\frac{\chi_k}{(\psi+\alpha)+t\sqrt{\frac{\gamma}{\Lambda_k \pi}}}\right)^2.$$

Substituting back into the supply equation, we obtain the equilibrium average price

$$\widetilde{P}_k^* = \text{CC} + \frac{1}{3i} t\sqrt{\frac{\gamma H_k^*}{\Lambda_k \pi}} = \text{CC} + \frac{1}{3i} \cdot \frac{\chi_k}{\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right)+1}.$$

Therefore changes in productivity–amenities imply

$$\Delta \widetilde{P}_k^* = \frac{1}{3i} \cdot \frac{\Delta \chi_k}{\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right) + 1}.$$

The expectation of changes in housing prices is therefore

$$E(\Delta \widetilde{P}_k^*) = \int \frac{1}{3i} \cdot \frac{\Delta \chi_k}{\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right) + 1} \cdot f(\Delta \chi_k) \cdot d\Delta \chi_k,$$

which, given the independence assumption of productivity shocks, implies that

$$E\left(\Delta \widetilde{P}_k^*\right) = \frac{1}{3i} \cdot \frac{E(\Delta \chi_k)}{\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right) + 1}.$$

Now, we can demonstrate that

$$\frac{dE\left(\Delta \widetilde{P}_k^*\right)}{d\Lambda_k} = -\frac{1}{3i} \cdot \frac{E(\Delta \chi_k)}{\left[\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right) + 1\right]^2} \cdot \frac{1}{2} \cdot \frac{(\psi+\alpha)\sqrt{\pi}}{t\sqrt{\gamma}} \cdot \Lambda_k^{-\frac{1}{2}} < 0.$$

More importantly,

$$\begin{aligned} \frac{d^2 E\left(\Delta \widetilde{P}_k^*\right)}{(d\Lambda_k)^2} = &+ \frac{1}{3i} \cdot \frac{E(\Delta \chi_k)}{\left[\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right) + 1\right]^3} \cdot \frac{1}{2} \cdot \frac{(\psi+\alpha)^2 \pi}{t^2 \gamma} \cdot \Lambda_k^{-\frac{1}{2}} \\ &+ \frac{1}{3i} \cdot \frac{E(\Delta \chi_k)}{\left[\left(\frac{(\psi+\alpha)\sqrt{\Lambda_k \pi}}{t\sqrt{\gamma}}\right) + 1\right]^2} \cdot \frac{1}{4} \cdot \frac{(\psi+\alpha)\sqrt{\pi}}{t\sqrt{\gamma}} \cdot \Lambda_k^{-\frac{3}{2}} > 0. \end{aligned}$$

Therefore the expected price change is a decreasing convex function of land availability: wherever land availability is high initially, further changes in land availability do not change expected price growth much. Conversely, in areas with initially low land availability, further constraints on land development have greater impacts on future prices.

Appendix III
Data Appendix

Variable	Source	Notes
Land unavailability	Calculated by author from elevation and land use GIS data from USGS	
Wharton Regulation Index	Gyourko, Saiz, and Summers (2008)	See *Data* section in text.
Declining metro area: 1950–1970	Calculated by author from data in Historical Census Browser—University of Virginia	A dummy that takes value 1 if growth is in the lowest quartile of the metro areas in our sample.
Non-Hispanic white share (1970)	HUD State of the Cities database (from the Census)	
BA/BS share (1970)	HUD State of the Cities database (from the Census)	
Foreign-born share (1970)	HUD State of the Cities database (from the Census)	
Tourist visits per person (2000)	Carlino and Saiz (2008)	
Log patents per capita	Glaeser and Saiz (2004)	
Immigration shock	HUD State of the Cities database (from the Census)	The difference in the number of foreign-born individuals between 1970 and 2000, divided by the metro area population in 1970.
Share of workers in manufacturing (1970)	HUD State of the Cities database (from the Census)	
Median housing price (1970, 2000)	HUD State of the Cities database (from the Census)	
Number of housing units (1970, 2000)	HUD State of the Cities database (from the Census)	
Percentage of Christians in “nontraditional” denominations, 1971	Churches and church membership in the United States, 1971—the Association of Religion data archives	Calculated as one minus the share of Catholic Church adherents and mainline Protestants (United Church of Christ, American Baptist, Presbyterian, Methodist, Lutheran, and Episcopal).
% voting for Carter (1980)	County and City Data Book 1983	

APPENDIX III
(CONTINUED)

Variable	Source	Notes
Local tax revenues (1982)	Census of Governments 1982	2000 county-based metropolitan definitions were used to aggregate at the metro level.
Inspection expenditures/local tax revenues (1982)	Census of Governments 1982	
Coastal metro area dummy	Rappaport and Sachs (2003)	A dummy that takes value 1 if the minimum distance in an MSA's county is below 100 km.
January monthly hours of sun (average 1941–1970)	Natural Amenities Scale—USDA Economic Research Service	
Construction costs (single-family, average quality)	Gyourko and Saiz (2006)—originally from Means et al.	
Housing price repeat sales index	Freddie Mac purchase-only conventional mortgage home price index	
Land value shares	Davis and Heathcote (2007); Davis and Palumbo (2008)	Davis and Heathcote (2007) calculate the average share of land for residential real estate in the United States in 1970 to be 20%. In 1984 (the first year for which their metropolitan data series is available) Davis and Palumbo (2008) suggest national and metropolitan land shares to be very similar. We therefore adapt an unweighted average 20% land share *across metropolitan areas* in 1970. We then calculate differences in the structure cost/value ratio by dividing the average construction cost in 1970 for a 2,000 sq. ft. home (the average home size) by the median home value in each metro area. The final metropolitan-level estimate of structural shares in 1970 (α_{it-1}) is proportional to the aforementioned ratio, and such that its unweighted mean across metro areas is 80%.
Central city areas (1970, 2000)	County and City Data Books 1972, 2002	

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