

Oliner, Stephen D.; Nichols, Joseph B.; Mulhall, Michael R.

Working Paper

Swings in commercial and residential land prices in the United States

AEI Economics Working Paper, No. 2012-03

Provided in Cooperation with:

American Enterprise Institute (AEI), Washington, DC

Suggested Citation: Oliner, Stephen D.; Nichols, Joseph B.; Mulhall, Michael R. (2012) : Swings in commercial and residential land prices in the United States, AEI Economics Working Paper, No. 2012-03, American Enterprise Institute (AEI), Washington, DC

This Version is available at:

<https://hdl.handle.net/10419/280505>

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.



AMERICAN ENTERPRISE INSTITUTE

Swings in Commercial and Residential Land Prices in the United States

Joseph B. Nichols
Federal Reserve Board

Stephen D. Oliner
AEI and UCLA Ziman Center for Real Estate

Michael R. Mulhall
J.P. Morgan Chase

Swings in Commercial and Residential Land Prices in the United States

Joseph B. Nichols
Federal Reserve Board

Stephen D. Oliner*
American Enterprise Institute and UCLA Ziman Center for Real Estate

Michael R. Mulhall
J.P. Morgan Chase

June 2012

Forthcoming in the Journal of Urban Economics

Abstract: We use a large dataset of land sales dating back to the mid-1990s to construct land price indexes for 23 MSAs in the United States and for the aggregate of those MSAs. The price indexes show a dramatic increase in both commercial and residential land prices over several years prior to their peak in 2006-07 and a steep descent since then. These fluctuations have exceeded those in well-known indexes of home prices and commercial real estate prices. Because those indexes price a bundle of land and structures, this comparison implies that land prices have been more volatile than structures prices over this period. This result is a key element of the land leverage hypothesis, which holds that home prices and commercial property prices will be more volatile, all else equal, in areas where land represents a larger share of real estate value.

Acknowledgements: We received helpful comments from Stuart Rosenthal (the editor), three anonymous referees, numerous Federal Reserve staff, and participants at the 2009 AREUEA meetings, the 2011 USC-UCLA-UC-Irvine Urban Economics Research Symposium, and the 2012 Barcelona Institute of Economics Workshop on Urban Economics. We thank Christopher Reynolds and Katherine Hayden for producing the location maps in the paper. The data for this study were provided by the CoStar Group, Inc. The views expressed herein are those of the authors alone and should not be attributed to any of the organizations with which the authors are affiliated.

*Corresponding author. Contact information:

Address: American Enterprise Institute, 1150 17th Street, NW, Washington, DC 20036

Email: stephen.oliner@aei.org

Phone: (202) 419-5205

1. Introduction

All types of economic activity require land. This requirement is obvious in industries such as farming and construction, but all other forms of commerce ultimately need land as well because workers, equipment, and buildings must be located somewhere. Even a high-tech company like Google has a corporate campus and more than 70 other facilities.

A broad measure of the value of land in the United States can be derived from the Flow of Funds (FOF) accounts published by the Federal Reserve Board. The FOF-based estimate covers land owned by households, nonprofit organizations, and most nonfarm businesses, and is calculated as the market value of real estate in those sectors minus the value of structures from the U.S. National Income and Product Accounts (NIPAs). This estimate of U.S. land value exceeded \$17 trillion in 2006 at the height of the real-estate boom.¹

With such a large aggregate value, changes in land prices can have a substantial effect on the net worth of businesses and households. In this regard, Davis and Heathcote (2007) estimate that swings in residential land prices accounted for most of the variation in house prices over 1975-2006 for the United States as a whole. Davis and Palumbo (2008) reach the same conclusion for a large set of metropolitan areas, as do Bostic, Longhofer, and Redfearn (2007) in their detailed analysis of home price changes in a single metropolitan area (Wichita, Kansas).

Land also serves as a form of collateral for loans, especially for construction loans. If the borrower defaults before completing a construction project, the lender's recovery will depend largely on the value of the land pledged as collateral. Commercial banks in the United States

¹ The data for this estimate are in tables B.100, B.102, and B.103 of the *Flow of Funds Accounts of the United States* (www.federalreserve.gov/release/z1). Barker (2007), Case (2007), and Davis (2009) estimated aggregate land value from earlier vintages of the FOF data. We should note that all of these estimates are subject to substantial measurement error. Indeed, the FOF accounts stopped publishing series for land value in 1995, at least in part because of concerns about the accuracy of the estimates. For estimates of aggregate land values that do not use the FOF data (but that rely on a variety of strong assumptions), see Davis and Heathcote (2007), the Bureau of Labor Statistics (2007), and the earlier work by Goldsmith (1951), Manvel (1968), and Milgram (1973).

have substantial exposure to land prices through their lending activities. Indeed, domestic banks held more than \$500 billion in construction and land development loans in 2007, and the subsequent delinquency rates on these loans exceeded those on all other major types of bank loans.²

Despite the importance of land as a component of wealth, as a source of variation in real estate prices, and as collateral for loans, only a handful of studies have calculated land price indexes for the United States as a whole or for a broad set of U.S. cities.³ Davis and Heathcote (2007) and Davis and Palumbo (2008) estimate price indexes for residential land in the United States, while Davis (2009) estimates indexes for both residential and commercial land. These indexes, however, are not based on transaction prices. Instead, Davis and his coauthors infer land prices as a residual in a measurement framework that is similar to the Flow of Funds methodology described above. Notably, their estimates rely on the assumption that the market value of existing structures is well approximated by the capital stock series in the NIPAs. Given that these series are constructed from limited information on depreciation rates, this assumption may not be appropriate.

In contrast, Sirmans and Slade (2011) use transaction prices to calculate national land price indexes. However, they do not estimate price indexes for individual metropolitan statistical areas (MSAs), an important limitation given the substantial local variation in real estate markets. Our study addresses this limitation by providing the first transaction-based indexes of land prices for a broad swath of MSAs across the United States.

² These figures represent aggregates from the Consolidated Reports of Condition and Income (Call Reports) filed by domestic commercial banks. See Lee and Rose (2010) for a full analysis of the Call Report data.

³ That said, there are numerous studies of land prices for narrow geographic areas within the U.S. For residential land, see Bryan and Sarte (2009), Downing (1970), Greenlees (1980), Ihlanfeldt (2007), and Voith (2001). For commercial land or a combination of commercial and residential land, see Brownstone and DeVany (1991), Colwell and Munneke (1997, 2003), Guntermann and Thomas (2005), Haughwout, Orr, and Bedoll (2008), Kowalski and Paraskevopoulos (1990), McMillen (1996), McMillen and McDonald (1991), Munneke (2005), Peiser (1987), Wallace (1988), and Wieand and Muth (1972).

For each MSA, we calculate separate price indexes for commercial land and residential land. Standard models of spatial equilibrium (see Glaeser, 2008, and Anas, Arnott, and Small, 1998) allocate land to its highest-value use, subject to regulatory constraints. This allocation process implies that expected risk-adjusted returns should be equal across alternative land uses. We are able to shed light on this prediction by comparing the price movements between the two broad types of land.⁴

Using source data from the CoStar Group, Inc., we construct a dataset that includes 180,000 land transactions in 23 MSAs from the mid- or late 1990s through mid-2011.⁵ These MSAs include the major population centers in the United States and some smaller cities. We estimate hedonic equations in which residential and commercial land prices depend on property-level characteristics, a flexible specification of property location, and half-yearly dummy variables to capture the changes in land prices over time after controlling for the other factors. The specification of locational effects includes both the property's distance from the central business district (CBD) and the semiparametric specification introduced in Colwell (1998) that can accommodate a much wider range of spatial patterns.

Our primary focus is on the land price indexes implied by the coefficients on the half-yearly dummy variables. For each MSA, we present price indexes for a composite of residential and commercial land along with separate indexes for these two broad types of land. We also report the analogous price indexes for the aggregate of the 23 MSAs taken together.

⁴ Note that our regressions generate price indexes for these two broad categories and do not permit a comparison of their absolute price levels. However, we are able to compare price levels among the different types of residential land and separately among the different types of commercial land.

⁵ In addition to our paper and Sirmans and Slade (2011), other studies that have used the CoStar land price data include Albouy and Ehrlich (2011), Kok, Monkkonen, and Quigley (2010), Haughwout, Orr, and Bedoll (2008), and Colwell and Munneke (2003). Albouy and Ehrlich focus on the differences in the level of land prices across MSAs and use these differentials to estimate cost functions and productivity for housing and tradable goods. The other three studies analyze land prices for San Francisco, New York City, and Chicago, respectively.

The results show that land prices trended up at a moderate pace from 1995 until about 2002 and then accelerated sharply. From the second half of 2002 to the second half of 2006, our composite index of residential and commercial land prices for the 23 MSAs as a group jumped nearly 135 percent, with even larger increases in the MSAs along the East Coast and in the Far West. However, prices tumbled over the next few years. As of mid-2011, the composite index had fallen more than 50 percent from its peak. Separate indexes for commercial and residential land prices display the same broad pattern, with especially large swings for the residential index. Residential land prices likely had greater gyrations because the excesses that led to the financial crises were rooted in the housing market (see the Economic Report of the President, 2009). Nonetheless, over the full sample period, the price indexes for residential and commercial land changed about the same amount on net, consistent with the expectation of similar returns across land uses in equilibrium.

The ups and downs in the aggregate land price indexes since 2002 outstrip those for well-known national indexes of home prices and commercial real estate prices. Because those indexes price a bundle of land and structures, this comparison implies that land prices have been more volatile than structures prices over this period. We examine whether the greater volatility of land prices is also evident at the MSA level using data for the housing sector.⁶ Consistent with the aggregate results, we find that the swings in residential land prices since 2002 have been far wider than those for home prices and structures prices in every MSA we examined. In addition, a standard decomposition of home prices assigns a much larger role to land than to structures in explaining the home-price cycle since 2002.⁷

⁶A parallel analysis for the commercial sector could not be undertaken because MSA-level price indexes for commercial real estate are not available.

⁷We do not test whether the jump in land prices from 2002 to the subsequent peak constituted a formal asset-price bubble. As discussed in Flood and Hodrick (1990), Case and Shiller (1989), Meese and Wallace (1994) and others,

Our finding that land prices in the United States have been more volatile than structures prices lines up with the results in Davis and Heathcote (2007), Davis and Palumbo (2008), and Sirmans and Slade (2011). It is also consistent with the available international evidence (Bourassa et al., 2009, 2010; Wu, Gyourko, and Deng, 2010; and Schulz and Werwatz, 2011). The difference in volatility likely arises because the supply of structures is much more elastic than the supply of land.⁸ As a result, shocks to demand for housing or commercial space have a greater effect on the price of land than on the price of structures.

This regularity supports the central assumption in the land leverage hypothesis of Bostic, Longhofer, and Redfearn (2007). The land leverage hypothesis holds that home prices and commercial property prices will be more volatile, all else equal, in areas where land represents a relatively large share of real estate value. This hypothesis can only be valid if land prices are more volatile than structures prices.

The rest of the paper is organized as follows. The next section discusses our dataset. Section 3 describes how we model the effects of location on land prices, and section 4 lays out all other aspects of our methodology. Section 5 presents the estimation results and the price indexes for land. Section 6 compares the land price indexes to the prices for other types of real estate. The final section summarizes our conclusions and suggests avenues for future research.

it is difficult to distinguish between bubbles and large price changes that result from shifts in fundamentals. Despite this difficulty, several recent studies have attempted to assess whether the run-up in house prices through the latest peak was a bubble (Coleman, LaCour-Little, and Vandell, 2008; Goodman and Thibodeau, 2008; Mikhed and Zemčik, 2009; Clark and Coggin, 2011; and Mayer, 2011). While not conclusive, the evidence is consistent with a house price bubble, at least in some cities. Given that land was the volatile component of house prices over the latest cycle, any such bubbles in house prices likely were due to the underlying land.

⁸ Gyourko and Saiz (2006) estimate that the supply of physical housing structures in the United States is extremely elastic. Because structures are built with labor and materials, the elastic supply of structures reflects the ability to vary the quantity of those inputs without large changes in their costs. In contrast, Saiz (2010) shows that land-use regulation and geographic constraints limit the supply of available land, resulting in a much lower supply elasticity for new homes than for the structure per se. See Gyourko (2009) for further discussion.

2. Data

The data for our analysis come from the CoStar Group, Inc., a leading provider of real estate information in the United States. Among its various products, CoStar maintains a database on sales of commercial property and land, which it compiles from public records, interviews with parties to the transactions, and field inspection of the properties. Currently, the CoStar database includes more than one million transactions.

We analyze the transactions that CoStar identifies as sales of residential or commercial land. CoStar defines land sales as transactions that involve vacant property or property with structures that are slated for demolition. These criteria ensure that the value of any existing structures should be incidental to the total value of the property. To be included in the CoStar database, a residential land parcel must consist of at least five single-family lots or be large enough to support multifamily buildings with at least five units. There is no lower size limit for commercial land parcels. Each transaction record contains the parcel's sales price and location, along with a series of text fields describing its characteristics. We use these text fields to create a number of indicator variables for our hedonic price regression.

Although the CoStar database includes transactions throughout the United States, we focused on 23 MSAs with the richest data. These MSAs include five cities in the Northeast (Boston, New York, Philadelphia, Baltimore, and Washington DC); three areas in Florida (Orlando, Tampa/St. Petersburg, and South Florida); six cities on the West Coast (Los Angeles, San Diego, San Francisco, Sacramento, Seattle, and Portland); and nine cities in the interior of the country (Atlanta, Chicago, Dallas, Denver, Detroit, Houston, Las Vegas, Phoenix, and Tucson). Prior to 1995, the transaction data for most of these MSAs are either sparse or nonexistent. Accordingly, we standardized the starting point of the dataset in 1995, except for

six MSAs for which the data begin between 1997 and 1999. The dataset runs through mid-2011. As described in Appendix A, we eliminated transactions with missing data and applied a variety of screens to improve the quality of the dataset. For example, we removed about 14,000 observations that appeared not to be market sales. This screen eliminated all government takings of land through eminent domain and all transfers of land from property owners to their lenders through foreclosure proceedings. Foreclosed parcels subsequently sold by lenders to new owners passed this screen, but there were too few such sales to generate meaningful results, and so we removed all foreclosure transactions from the dataset. We also removed all observations for an MSA in a given half-year when we judged the sample size to be insufficient to generate a reliable price index for that period.⁹

As shown in table 1, we have 180,155 sales in total, 55 percent of which are for commercial land. Summing the number of commercial and residential land sales yields a figure slightly less than the number of land sales in the “Total” column. This difference arises because about 800 sales are in MSAs that have more than 40 transactions in a given half-year when the commercial and residential land sales are pooled, but do not have more than 20 transactions for both types of land. Given our rules for minimum sample sizes, we exclude these observations when we estimate separate price indexes for commercial land and residential land.

The sample sizes vary widely at the MSA level, ranging from about 2,800 for Sacramento to more than 18,000 for Phoenix. This variation owes at least partly to differences in the amount of development activity across the MSAs and does not correlate closely with the relative shares of the MSAs in national stocks of residential or commercial land. As described in section 4, we

⁹ In particular, we included the residential or commercial transactions for an MSA in a given half-year only when we had more than 20 observations for that property type. For regressions in which we estimated an MSA-level price index that covered both property types, we included a given half-year only when we had more than 40 observations in total. We settled on these minimum half-yearly sample sizes after observing the volatility in the price indexes generated without regard to sample size.

employ MSA-level weights that correct for this discrepancy whenever we estimate aggregate results that pool the MSAs or test cross-MSA restrictions.

Another key dimension of the sample is the number of observations over time. Figure 1 shows one measure of this time pattern – the median sample size across the 23 MSAs for each half-year since 1995 – for both commercial and residential transactions.¹⁰ The median MSA has at least 80 residential land transactions in every half-year through the first half of 2006, but transaction volume then plunged with the collapse in housing activity. For commercial land, transaction volume was well maintained through 2007 but dropped sharply thereafter. Given the sample-size cutoffs we impose at the MSA level, the reduced volume of sales implies that we cannot calculate land price indexes in every period for some MSAs.

Tables 2 and 3 provide summary information about the characteristics of the land parcels in our dataset. As shown in table 2, the parcels range in size from less than half an acre to more than 70 acres, with a median size of about 7 acres for residential land and 2.5 acres for commercial land. Table 2 also indicates that the median distance from the CBD is almost 20 miles, as most opportunities for land development occur beyond the urban core.

Table 3 shows the shares of the sample observations classified by the type of property, condition, and intended use. Almost half of the residential land is intended for single-family homes, while 12 percent is slated for multifamily rental projects. The large “other residential” category consists principally of condominium development. The commercial land spans a wide range of property types, with 8 percent of the parcels planned for office buildings, 22 percent for industrial sites, 16 percent for retail stores, and the remaining 53 percent for “other commercial”

¹⁰ We show the median sample size across MSAs for each period, rather than the average sample size or the total number of observations, to reduce the distorting effect from the increase in the number of MSAs in the sample from 1995 to 1999. Note also that the figure presents the median sample size across MSAs *before* applying the cutoff rules described in the previous footnote. We do this to display the underlying time patterns in transaction volume.

projects, which include mixed-use development as well as such buildings as hospitals, hotels, and convention centers. Regarding property condition, about half the land in the dataset is unimproved, while smaller shares of the observations have identified improvements, shown by each row down through the row labeled “Structure present.” About 9 percent of the residential land and 19 percent of the commercial land has insufficient information to classify the extent of improvements. A very small share of the parcels have environmental problems of some type, the severity of which are unknown. Finally, as shown at the bottom of the table, about three-quarters of the land in the sample is intended for private development, while roughly 10 percent will be held for investment purposes with no immediate development plans. Only a small fraction of the land falls into the other categories for known uses.

3. Locational Effects on Land Prices

There is a vast literature on the spatial structure of urban areas and the locational factors that influence property prices. Numerous studies have documented that large cities tend to have multiple centers of employment (see, for example, Giuliano and Small, 1991; McMillen and McDonald, 1998; Craig and Ng, 2001; McMillen and Smith, 2003; and Redfeard, 2007). In particular, McMillen and Smith (2003) found strong evidence of employment subcenters for nearly every MSA that appears in our dataset. Moreover, the proximity to these subcenters tends to be an important determinant of land prices (see Peiser, 1987; Sivitanidou, 1996; and McMillen, 1996). A separate line of research has shown that local characteristics such as the quality of schools and the amount of crime affect house prices, and ultimately the prices of the underlying land.¹¹

¹¹See Ries and Somerville (2010), Brasington and Haurin (2009), and Clapp, Nanda, and Ross (2007) for recent studies of school quality and house prices. Ihlanfeldt and Mayock (2010) and Pope (2008) provide recent analyses of the connection between crime and house prices.

To capture spatial effects, empirical studies of land prices typically use a variety of locational variables. The primary variable is usually distance from the CBD, which is augmented by other variables that include distance from major roads, rail lines, and airports; distance from suburban business nodes; distance from the coastline; dummy variables for local topography; dummy variables for location within the city limits and within specific counties; the amount of street frontage; and demographic information for the surrounding area.

For our study – which covers 23 separate MSAs – it is not practical to specify a vector of locational variables for each land transaction. Instead, we use the semiparametric approach in Colwell (1998) to capture locational effects on real estate prices over and above those related strictly to distance from the CBD. Colwell’s method superimposes a grid over the transactions in the dataset. The grid can be sized to fit the boundaries of this area, and it can be divided into as many component squares as the researcher wishes to use. Colwell’s method estimates the price level at each of the grid vertices, which correspond to the corners of the component squares.

To carry out the estimation, each vertex is treated as a separate variable in the hedonic regression. Let V_1, \dots, V_n denote the set of n vertices. Prior to estimation, values must be assigned to V_1, \dots, V_n for each transaction in the dataset. This is done by measuring the nearness of a transaction to the four corners of the square in which it is located; all other grid vertices have a value of zero for this observation. For a transaction located exactly at the center of a given square, each of the four surrounding vertices receives a value of 0.25. In this case, the transaction contributes equal to the estimated price level for each of the surrounding vertices. For transactions located elsewhere in the square, the values assigned to the four vertices are weights that sum to one and that reflect an area-based measure of closeness.

Figure 2 illustrates this weighting scheme for a single square in the grid, with an observation at point A. The value assigned to vertex V_1 for this observation equals the area of the rectangle formed by A and the opposite vertex (the shaded area), divided by the total area of the square. This ratio of areas will converge to one as point A approaches V_1 . The values for V_2 , V_3 , and V_4 associated with point A are calculated in the same manner as for V_1 : Form the analogous opposite rectangles and compute the ratio of the area of each rectangle to the area of the square. All other vertices in the grid have a value of zero for this observation.

After the values for V_1, \dots, V_n have been assigned in this manner for each observation, a hedonic price regression can be estimated with the vertices included as explanatory variables.¹² The estimated coefficient for a given vertex represents the height of the price surface at that point on the grid. Colwell shows that the surface defined by these grid points is continuous, piecewise linear along the edges of the individual squares, and parabolic along slices within each square.¹³

In implementing this method, we laid out grids for the individual MSAs to conform to the spatial pattern of the land transactions. For Atlanta, we were able to encompass the vast majority of transactions with a grid that has five squares across and five squares down. With this grid, we estimate coefficients for 36 vertices (6x6) on the spatial price surface. For Dallas and Tucson, we used a five-by-five grid from which we omitted any vertex on the outside edge that had very few nearby transactions. As an example, figure 3 shows the grid for Dallas superimposed on a scatter plot of the transactions in our dataset. For the other MSAs, the spatial distribution of transactions did not fit well within a square outline – generally because of the presence of a body of water or mountains – and we drew the outline of the grid to fit these patterns. In each case,

¹² If the regression includes a constant term, one vertex must be omitted to avoid perfect colinearity among the explanatory variables. The colinearity arises because the vertex weights for any observation sum to one.

¹³ An alternative to the Colwell grid procedure would be to simply include a fixed effect for each grid square. However, the fixed-effect specification would force the estimated price surface to be flat for each square, with jumps where the squares meet, which is less appropriate than the flexible surface that we estimate.

the grid consists of between 22 and 28 individual squares (with between 33 and 42 vertices), arranged in a non-rectangular shape. As an illustration, figure 4 presents the grid used for South Florida, where the land transactions occupy a band along the coastline.¹⁴

In summary, we take a hybrid approach to modeling locational effects on land prices. We include distance from the CBD as an explanatory variable in the regressions, which we augment with the grid method described above to capture the features of the spatial price surface that do not lie on a constant gradient away from the CBD.

4. Specification and Estimation of the Hedonic Price Equation

We use the data described in section 2 to estimate a flexible hedonic regression for land prices. The dependent variable in the regression is the natural log of the price of land per square foot. The explanatory variables can be broken into three broad categories: property characteristics other than location (X_1, \dots, X_J), measures of location (Z_1, \dots, Z_K), and a set of half-yearly time dummies (D_1, \dots, D_T). We estimate 46 separate OLS regressions, one for residential land and another for commercial land in each of the 23 MSAs. Each regression can be written as

$$(1) \quad \ln P_{m,l} = \alpha + \sum_j \beta_{m,l,j} X_j + \sum_k \theta_{m,l,k} Z_k + \sum_t \gamma_{m,l,t} D_t + \varepsilon_{m,l}$$

where $P_{m,l}$ is the vector of observations of price per square foot in the m^{th} MSA for the l^{th} broad land type (l equals either residential or commercial), and $\varepsilon_{m,l}$ is the corresponding vector of iid errors assumed to be distributed $N(0, \sigma_{m,l}^2)$. The subscripts on the β , θ , γ , and σ^2 coefficients show explicitly that they vary across MSAs and property types.¹⁵

¹⁴ The position of the grid for each MSA was dictated by the desire to encompass as many observations as possible given the constraint that the blocks of the grid had to be equal-size squares. This constraint meant that it was impossible to place a specific vertex at the exact center of any MSA.

¹⁵ Technically, we stack the 46 equations and estimate the stacked system using the SAS maximum likelihood procedure PROC MIXED. Under the assumed structure for the error terms, this procedure is equivalent to equation-by-equation OLS. We use the stacked system because it provides a convenient framework for testing cross-equation restrictions.

The vector of property characteristics (X_1, \dots, X_J) includes the natural log of the size of the parcel in order to test the so-called “plattage effect”. Plattage refers to the common finding that the price of a land parcel rises less than proportionally with its size. This relationship likely arises because there is an optimal scale for buildings of a given type, which implies that parcels larger than the size needed to support the optimal building scale earn a lower return.

The other variables in X_1, \dots, X_J are the indicator variables mentioned in section 2. These variables are largely the same as those used by Haughwout, Orr, and Bedoll (2008) and include:

- Type of property: Dummy variables in the residential land regressions for single-family housing, multifamily rental housing, and other residential uses (the multifamily dummy is omitted to avoid perfect colinearity) and separate dummies in the commercial land regressions for industrial, office, retail, and other commercial parcels (industrial land is the omitted property type).
- Condition of the property: Dummy variables for whether the land has been graded, paved, finished, fully improved, platted and engineered, previously developed, has an existing structure, or if information on improvements is not reported. The omitted condition is unimproved land. This group of variables also includes a dummy for whether environmental problems exist as defined by reported soil or building contamination.
- Intended use of the property: Dummy variables for property to be privately developed, to be used in the public sector, to be kept as open space, to be held for investment, or with an unknown use; the omitted category is private development. We also include a dummy for whether the land was purchased by an entity that intended to expand an adjacent property.

The vector of location characteristics (Z_1, \dots, Z_K) includes the natural log of distance of each land parcel from the geographic center of its MSA as determined by Google Earth, along

with the grid vertices described in the previous section. We specify a separate grid for residential land and commercial land in each MSA. To avoid perfect colinearity, each regression omits one vertex variable located close to the center of the MSA. Finally, we include a set of MSA fixed effects, with New York City as the omitted MSA.

The coefficients on the half-yearly dummies in the 46 regressions yield MSA-level price indexes for residential and commercial land. We also estimate a composite price index for each MSA by constraining the coefficients on the half-yearly dummies to be the same for both types of land in that MSA.

In addition, we estimate price indexes for the sample as a whole by constraining the coefficients on the half-yearly dummies to be the same in all MSAs, and we test a variety of other cross-MSA restrictions. As discussed in section 2, the number of land sales by MSA in our dataset does not match the MSA shares of commercial and residential land. Accordingly, in estimating the regressions with cross-MSA constraints, we weight the data so that each MSA contributes to the results roughly in proportion to its share of land. For example, if Phoenix's share of residential land sales in our dataset is five times greater than its share of residential land in the 23 MSAs, we apply a weight of 0.2 to those sales. Ideally, the weights would be based on the available land area devoted to commercial and residential real estate by MSA. Such data are not available. Instead, we weighted the commercial land observations using estimates from Torto Wheaton Research of the total space in commercial and industrial buildings by MSA and the residential land observations using the number of occupied single-family and multifamily housing units from the 2000 Census. We re-adjust these weights on a period-by-period basis to account for the exclusion of MSAs with insufficient observations during particular time periods.

Our dataset is not a random sample of all existing land parcels, which raises the possibility of sample selection bias. However, the standard Heckman (1979) procedure to correct for selection bias is not feasible in our case, as we would need data on land parcels that were not sold during the sample period. Such data would be very difficult to assemble for a large number of MSAs. Moreover, it is unclear whether any such effort would be worthwhile. The results of studies that have applied the Heckman procedure to real estate prices have been mixed to date. Although there is some evidence of selection bias in house prices (see, for example, Gatzlaff and Haurin, 1997 and 1998; Jud and Seaks, 1994; and Rosenthal and Helsley, 1994), the few available studies for commercial real estate and land prices have found that selection effects were generally small and insignificant.¹⁶

5. Results

We discuss the estimation results in three steps, focusing first on the estimated coefficients for the characteristics of the land parcels, then on the spatial price effects, and finally on the estimated variation in land prices over time.

Property characteristics

Table 4 summarizes the estimated coefficients for property characteristics across the 23 MSAs for residential and commercial land. Each row reports the median and range of coefficient values across the MSAs, along with the number of MSAs for which the coefficients were either negative and significant or positive and significant at the five-percent level.

Starting in the first row, the coefficient on the log of parcel size is negative and significant in all 23 MSAs for both residential land and commercial land. The median value for residential land indicates that doubling the size of a parcel reduces price per square foot by 54

¹⁶ See Colwell and Munneke (1997) for commercial and residential land prices, Rosenthal and Helsley (1994) for residential land prices, and Fisher, Geltner, and Pollakowski (2007) and Munneke and Slade (2000, 2001) for commercial real estate prices.

percent; the median result for commercial land is very similar. These results confirm the plottage effect – that the price of a land parcel rises less than one-for-one with its size.

The next block of the table shows the differences in price per square foot across the different types of land. Office, retail, and other commercial land sell at a significant price premium to industrial land (the omitted category) in almost every MSA. On the residential side, land for single-family housing and land in the residual grouping generally sell at a discount to land for multifamily housing (the omitted category). These price differentials are consistent with those found in other studies (Albouy and Ehrlich, 2011; Haughwout, Orr, and Bedoll, 2008; Kok, Monkkonen, and Quigley, 2010; and Sirmans and Slade, 2011).¹⁷

Most of the indicators of the condition of a property have the expected effects on price. Property that has been graded, paved, finished, fully improved, platted and engineered, or previously developed sells at a price per square foot that is 10 to 27 percent above that for unimproved land in the median MSA; these price differentials are statistically significant in a large number of MSAs for some variables, but not for others.¹⁸ The presence of an existing structure has a small positive effect on land prices in the median MSA. The positive coefficients suggest that an existing structure may proxy for unobserved factors that made the land well suited for development in the first place. Unobserved factors also may account for the

¹⁷ Zoning rules are a potential source of these price differentials. The key question is whether zoning alters the supply of different types of land away from the market allocation or whether it simply ratifies market outcomes. Zoning can induce the observed price differentials only if the former is true. Studies that attempt to distinguish these two hypotheses (Wallace, 1988; McMillen and McDonald, 1991 and 2002; Pogodzinski and Sass, 1994; and Munneke, 2005) have found that zoning follows the market for single-family housing but that it interferes with market outcomes for other land uses. Thus, zoning can result in persistent price differences across property types. The Wharton survey on land-use regulation, conducted by Gyourko, Saiz, and Summers (2008), provides additional evidence that zoning could account for the price differentials in our estimates. According to the survey responses, zoning rules restrict the supply of land more for multifamily housing than for single-family housing; the responses also show somewhat tighter supply of land for commercial uses relative to land for industrial uses. These results are consistent with the pattern of coefficients on our property-type dummies.

¹⁸ If we pool the data for all 23 MSAs, the estimated price effects are similar to those in the median MSA, but the coefficients become significant at the five-percent level for almost every variable for both residential and commercial land. The pooled results are more precise because of the substantial increase in sample size.

inconsistent price differentials for parcels that lack information on land improvements; these parcels tend to sell at a premium to unimproved land in the residential sector but at a discount in the commercial sector. Finally, the dummy for environmental problems is almost never significant, which may reflect both the small number of parcels with any noted problems and the absence of information about the severity of those problems.

The last block of the table presents the results for intended land use. Land to be held for investment or kept as open space generally sells at a significant discount to land purchased for immediate private development, as expected. Land with an unknown intended use also tends to sell at a discount. For the other land-use categories – land purchased for public use or to facilitate the expansion of an adjacent property – fewer of the MSAs show significant price effects. We did not have strong priors about the signs of any such price effects.

The results in table 4 indicate that the coefficient estimates for a given variable can span a wide range across MSAs and property types. As detailed in Appendix B, we find strong evidence that these differences are statistically significant, implying that the property types and MSAs should not be aggregated when estimating the price effects of the variables in table 4.¹⁹

Locational effects

As discussed above, our regression estimates both a log-linear distance gradient from the CBD and a grid to capture additional spatial effects. The estimated distance gradient is negative and significant at the five-percent level for both residential and commercial land in all but a few MSAs. The median value of the gradient across MSAs is -0.33 for residential land and -0.41 for

¹⁹ We also tested whether the coefficients on the variables in table 4 are stable over time. Specifically, we split the sample at the end of 2002 – roughly the beginning of the boom-bust cycle in land prices – and tested for the equality of individual sets of coefficients and of all coefficients before and after that point. The results rejected the hypothesis of constant coefficients. However, allowing the coefficients to vary across the two subperiods had almost no effect on the land price indexes calculated from the coefficients on the half-yearly dummies, which represent the core contribution of the paper. In light of this result, we decided not to add time variation to what was already a regression with many estimated coefficients. Accordingly, the coefficient estimates in table 4 should be regarded as the average price effects of these variables over the full sample period.

commercial land, so that doubling the distance from the CBD, all else equal, reduces land value by one-third or more in the median MSA.

Distance from the CBD does not fully characterize locational effects, as indicated by the significant coefficients on many grid vertices. In 20 of the 23 MSAs, at least ten of the grid vertices for residential land are statistically significant at the five-percent level; the same result holds for commercial land in 21 MSAs. These results are consistent with the prior literature that has shown the classic monocentric urban model does not adequately describe property prices in most cities. As an example, figure 5 displays the estimated grids for South Florida. The back edge of the figure lies along the Atlantic coast, with the highest part of the contour representing Miami Beach. The figure clearly shows a price premium for land near the coast after controlling for distance from the CBD.

Price indexes

Given our controls for spatial price effects and key characteristics of the land parcels, the coefficients on the half-yearly dummy variables trace out the indexes of land prices. The price index in a given half-year equals the exponentiated coefficient for that period divided by the exponentiated coefficient for the base period of 2002:H2.

The top panel in figure 6 presents the land price index from the restricted version of the regression that includes a single set of half-yearly dummies. The resulting price index represents a composite index for residential and commercial land in all 23 MSAs. This index trended up from 1995 to 2002 at an average annual rate of about 7 percent. Prices then surged, increasing almost 135 percent in total from the second half of 2002 to the series peak in the second half of 2006. After holding about steady for a year, the price index plummeted, and by the first half of 2009, it had unwound nearly its entire post-2002 run-up.

The lower panel shows separate price indexes for commercial and residential land. To obtain these indexes, we re-estimated the constrained regression with two sets of time dummies, one for commercial land in all MSAs and another for residential land in all MSAs. Both indexes display the same broad patterns as the composite index in the upper panel – a moderate uptrend through 2002, followed by a sharp rise that lasted until the first half of 2006 for residential land and into 2007 for commercial land. The earlier peak for residential land is consistent with the housing market having been the catalyst for the broader downturn in the economy. Prices for both residential and commercial land then plunged through the first half of 2009 and have changed little on net since then. During both phases of the cycle, the index for residential land displays greater amplitude than the index for commercial land, and a likelihood ratio test decisively rejects the hypothesis that the two price indexes are equal.

The basic features of the land price indexes estimated by Sirmans and Slade (2011) are similar to those in figure 6. Both sets of indexes indicate that land prices surged after 2002 and have fallen sharply in recent years. There are, however, some differences in the indexes. The total increase in the Sirmans-Slade indexes from 2002 to the peak, while substantial, is not as large as in our indexes, and their indexes peak somewhat earlier than ours. In particular, their composite index of commercial, industrial, and residential land prices peaked in late 2005, when real estate markets were still booming, which seems less plausible than the late 2006 peak in our composite index.

Figure 7 presents composite land price indexes for the individual MSAs. These price indexes cover commercial and residential land and are calculated from regressions that have single set of time dummies for all land in a given MSA. The upper row of the figure shows the indexes for MSAs on the East Coast while the middle row presents the indexes for the West

Coast MSAs and the bottom row shows the indexes for MSAs located in the interior of the country. All the series are indexed to equal 100 in 2002:H2 and every panel has the same scale, so the magnitude of the post-2002 run-up in land prices and the subsequent decline can be compared across MSAs.

The most striking feature of figure 7 is that the swing in land prices generally has been much larger on the coasts and in Las Vegas and Phoenix than elsewhere in the country, a pattern that mirrors the boom-bust cycle in the housing market in recent years. Another notable result is that land prices in most MSAs have unwound a large share of the earlier price jump.

Table 5 provides additional information on the MSA-level indexes. As shown, the price indexes for 18 of the 23 MSAs reached a peak in 2006 or 2007, with the median peak date across the MSAs occurring in the second half of 2006. The total price increase from the second half of 2002 to the peak ranged from a low of 52 percent in Denver and 57 percent in Dallas to nearly 300 percent in Las Vegas; the median rise across MSAs was 145 percent. Although the range is extremely wide, even the increases at the low end are not small in any absolute sense. The table also shows the extent of the price decline from the MSA-specific peaks to the series low. Thirteen of the MSAs hit their post-peak low in 2009, while prices in the others continued to drop into 2010 or 2011. The cumulative peak-to-low decline ranged from 33 percent in Denver to 70 percent or more in Atlanta, Detroit, Las Vegas, and South Florida, with a median price drop of 60 percent across the MSAs.

We present analogous results for commercial land in figure 8 and table 6 and for residential land in figure 9 and table 7. The underlying price indexes are calculated from the unrestricted version of the regression equation that allows the coefficients of the half-yearly dummies to vary across both MSAs and property type.

Three main conclusions can be drawn from these exhibits. First, the tables show that the MSA-level indexes generally have the same timing pattern seen in the aggregate indexes – namely that residential land prices peaked earlier than commercial prices but that the subsequent lows were closer together. In particular, for the median MSA, the residential peak preceded the commercial peak by a full year (2006:H1 versus 2007:H1), while the low occurred only a half-year ahead (2009:H2 versus 2010:H1). Second, for most MSAs, the price swings were greater for residential land than for commercial land, which also comports with the aggregate measures in figure 6. Taking the median across MSAs, residential land prices increased 167 percent from the second half of 2002 to the peak, nearly 30 percentage points more than for commercial land prices. Similarly, since the peak, the index for residential land has dropped 69 percent in the median MSA, 10 percentage points more than the median decline for commercial land. Finally, as shown in figures 8 and 9, the price swings for commercial and residential land generally have been wider on the East Coast and in the Far West than elsewhere in the country, consistent with the regional differences for the composite indexes in figure 7.

6. Further Analysis of Land Prices

An important question is how these movements in land prices compare with the price changes for housing and commercial real estate over the same period. In their assessments of this issue, Davis and Heathcote (2007), Davis and Palumbo (2008), and Sirmans and Slade (2011) found that the price swings for residential land have been wider than those for home prices over various periods. Sirmans and Slade obtained the same result when comparing their price index for industrial land to standard price indexes for industrial real estate.²⁰

²⁰ However, their comparison of relative price movements for the non-industrial part of the commercial sector was inconclusive because of differences in the sectoral coverage of the indexes.

These results have important implications for the land leverage hypothesis of Bostic, Longhofer, and Redfearn (2007). As noted in the introduction, this hypothesis holds that home prices and commercial property prices will be more volatile, all else equal, in areas in which land represents a relatively large share of real estate value. Clearly, this hypothesis can only be valid if land prices are more volatile than structures prices.

To see whether our results validate this condition, table 8 compares our land price indexes to well-known indexes of home prices and commercial real estate prices. The land price indexes are the aggregate indexes for residential and commercial land that were plotted in the lower panel of figure 6. We compare the index of residential land prices to the S&P/Case-Shiller 20-city index of home prices and the index of commercial land prices to the CoStar commercial repeat-sales index.²¹

As shown in the table, the peaks and subsequent lows in our land price indexes do not systematically lead or lag those for the comparison price indexes. The lack of a lead-lag relationship between land prices and other real estate prices may reflect the fact that all such prices are forward-looking assessments of the discounted profits anticipated from holding these assets.

Despite the similar timing, the price swings are substantially larger for land prices than for home prices or commercial property prices. On the residential side, the index of land prices rose three times as much as home prices from 2002:H2 to the peak in each series and then fell twice as much from the peak.²² The greater amplitude of residential land prices relative to home

²¹ Information about the S&P/Case-Shiller and the CoStar indexes can be found at www.standardandpoors.com and www.costar.com/ccrsi/index.aspx, respectively. Table 8 shows an aggregation of the CoStar sub-indexes for office, retail, and industrial properties. We exclude apartment properties to match the coverage of the commercial land price index.

²² The S&P/Case-Shiller index excludes multifamily properties, while the residential land index includes multifamily land parcels. This coverage difference, however, does not appear to distort the comparison between the residential land and home price indexes. CoStar's price index for apartment properties closely tracks the changes in

prices over real estate cycles is consistent with the results from earlier studies. Because the Case-Shiller index and other home price indexes cover a bundle of land and structures, the results in this paper and elsewhere imply that residential land prices have been more variable than the prices of housing structures. The same pattern holds qualitatively for the commercial sector, where land prices rose and then fell roughly one and half times as much as the CoStar measure of property prices. As discussed in the introduction, the wider swings for land prices than for property prices and, by implication, for structures prices likely owe to the relatively inelastic supply of land, which amplifies price movements in the face of shifts in demand.²³

A formal test of the statistical significance of these differences would require information on the standard errors of all the series in table 8. Although standard errors for the Case-Shiller and CoStar series have not been published, we can make some headway based solely on the standard errors for our land price indexes. In particular, we calculated the percent change between the top of the two-standard-error band in 2002:H2 for the residential land price index and the bottom of this band at the peak date; we did the same for the price index for commercial land. The results should provide a generous downside bound for the actual increase in land prices from 2002:H2 to the peak. This calculation yields a rise of 144 percent for residential land prices and 106 percent for commercial land prices – figures that exceed the increases in the Case-Shiller index and CoStar indexes, respectively, by a comfortable margin. We calculated an analogous lower bound for the absolute percentage drop in land prices from the peak to the subsequent low. The resulting bounded declines were considerably larger than those in the Case-

the S&P/Case-Shiller index since 2002. Thus, a broader measure of home prices that included apartments would still be considerably less volatile than the index for residential land prices.

²³ The NIPAs provide additional evidence that structures prices varied less than land prices over the latest real estate cycle. From 2002:H2 to the 2006:H1 peak for residential land prices, the price index for residential structures (in NIPA table 5.3.4) rose about 25 percent; the index then edged down about 2 percent from the peak to 2009:H1. The percent changes are not greatly different for commercial structures. For both residential and commercial structures, these price changes are dwarfed by those for land prices shown in table 8.

Shiller and CoStar indexes. This assessment suggests that the differences in price volatility in table 8 are statistically significant unless the standard errors for the Case-Shiller and CoStar indexes are very large.

The results in table 8 are based on aggregate price indexes that differ somewhat with respect to geographic coverage and the weights used to aggregate across MSAs. These differences in index construction could influence the magnitude of the estimated swings in the series. Table 9 addresses this issue by examining the relative price movements at the MSA level. For each MSA, we compare the changes in three series: the Case-Shiller home price index, our residential land price index, and a price measure for residential structures from Davis and Palumbo (2008).²⁴ The table shows the cumulative percent change in each series from 2002:H2 to the MSA-specific peak in land prices and the change from that point to the MSA-specific low in land prices; the 14 MSAs included in the table are those for which all three price series are available. This MSA-level test cannot be conducted for the commercial sector because of the absence of price indexes for commercial real estate for individual MSAs.

For every MSA, table 9 shows that the appreciation in residential land prices from 2002:H2 to the MSA-specific peak substantially exceeded the price appreciation for both homes and residential structures. As a summary measure, the median rise in land prices across the MSAs was 168 percent, three times the median increase in home prices (55 percent) and more than five times the median price increase for structures (31 percent). This same pattern holds in reverse after the peak. Residential land prices fell more sharply than home prices in every MSA,

²⁴ The Davis-Palumbo series were obtained from their online data files: (<http://www.lincolnst.edu/subcenters/land-values/>). As described in Davis and Palumbo (2008), these indexes use the change in construction costs for new residential structures as a proxy for the change in the prices of these structures. Rosenthal (1999) provides evidence of a tight linkage between structures prices and construction costs, which is consistent with the high degree of competition in the residential construction industry. However, we are unaware of other research on the tightness of this relationship.

while the price index for structures actually continued to rise in all but one MSA. Thus, the wider swings in land prices than in home prices at the aggregate level were not an artifact of aggregation but instead reflect a robust feature of relative price movements at the MSA level.

A further question of interest is the degree to which the documented volatility in land prices can account for the movements in home values. We address this issue with a standard decomposition of the growth in home prices into the contributions from residential structures prices and land prices (see Davis and Palumbo, 2008):

$$(2) \quad g_{hi} = w_{si}g_{si} + w_{li}g_{li} = (1 - w_{li})g_{si} + w_{li}g_{li}$$

where g_{hi} , g_{si} , and g_{li} denote the growth in the prices of homes, structures, and land, respectively, in MSA i , w_{si} and w_{li} are the shares of home values represented by structures and land, and $w_{si} = 1 - w_{li}$. The data for g_{hi} , g_{si} , and g_{li} are the same as in table 9, while the land share of home value is from Davis and Palumbo (2008). For each MSA, we apply the decomposition first to the period from 2002:H2 to the peak in land prices and then from the peak to the low for land prices; the land shares used in the decomposition are the MSAs-specific average values over each of the two periods. In each period, the contribution from land prices to the growth of home prices is $w_{li}g_{li}$, and the contribution from structures prices is $(1 - w_{li})g_{si}$. The sum of these contributions represents the implied growth in home prices from the decomposition. Note that this implied growth rate will not match the growth in the Case-Shiller index, except by chance, because the data employed in the decomposition are not drawn from a fully consistent measurement framework. That is, equation 2 implicitly includes a residual term.

Table 10 reports the results of the decomposition over the two periods. The first column for each period (labeled “CS”) shows the percent change in the Case-Shiller index, while the other columns represent terms from the right side of equation 2. Focusing first on the period of

price appreciation, the estimated contribution from land prices far outstrips the contribution from structures prices in 12 of the 14 MSAs, while the contributions are about the same for the two MSAs in which home prices rose the least (Atlanta and Denver). The relative contribution from land prices is even greater in the period after the peak. In every MSA except Atlanta, land prices fully account or more than account for the implied drop in home prices, as structures prices continued to rise. Overall, the decomposition assigns a much larger role to land prices than to structures prices in explaining the cycle in home prices since 2002.

However, the residual in the decomposition – the difference between the implied change in home prices and the change in the Case-Shiller index – is often large. This is especially true in the period from 2002:H2 to the peak, when the implied increase in home prices outpaced the rise in the Case-Shiller index in every MSA. The sizable residuals reflect some combination of definitional differences among the series in equation 2 and measurement error in these series. Future work to calculate the various series in a unified framework would be valuable.²⁵

Finally, we examine the contributions of land prices and structures prices to explaining the cross-MSA variation in home price changes. To do this, we regress the change in the Case-Shiller home price index on the contributions from structures prices and land prices from table 10; the regression also includes a constant term. We estimate the regression for the period from 2002:H2 to the peak and then for the period from the peak to the low. If the accounting decomposition in equation 2 held exactly, the constant term would equal zero and the coefficient for each price contribution would equal one.

²⁵ Additional research on the land share of home value would be especially useful, as Albouy and Ehrlich (2011) and Kok, Monkkonen, and Quigley (2010) both estimate land shares that are lower than the Davis-Palumbo shares. Pending such future research, we would note that the contribution of land prices to the change in home prices in table 10 would still greatly exceed the contribution of structures prices even if the land shares were scaled down considerably from the Davis-Palumbo values.

As shown in table 11, the constant term is not significantly different from zero in either period, consistent with equation 2. However, the coefficient on the contribution from structures prices is also not significant, and in the latter period it is not even positive, contrary to the prediction from the price decomposition. The regression might not work for structures because the variation in the structures contribution across MSAs is quite small and therefore could have been swamped by measurement error in the decomposition. The cross-MSA variation in the land contribution is much wider than for structures, and the results for land are somewhat better. In both periods, the coefficient on the land contribution is positive and close to significance at the 10 percent level, though the estimated values are well below one. These cross-MSA results provide some modest additional evidence of a link between changes in land prices and changes in home prices.

7. Conclusions and Directions for Future Work

This paper constructs land price indexes for a broad set of metropolitan areas in the United States. To calculate the indexes, we estimate a hedonic regression for land prices in 23 large MSAs with a sample of 180,000 land transactions from the mid-1990s through the first half of 2011. The regressions control for a variety of characteristics of the land parcels and for spatial price patterns within an MSA. Given these controls, the half-yearly dummy variables in the regressions trace out the implied land price indexes.

The resulting indexes show a dramatic increase in both residential and commercial land prices over several years prior to their peaks in 2006-07 and a steep descent since then. The magnitude of the run-up and the subsequent decline differs across the MSAs, with the largest movements in MSAs on the East Coast and in the Far West. Another key result is that the swings in land prices for the 23 MSAs as an aggregate have been considerably larger than those

in well-known indexes of commercial real estate and home prices. We confirm this finding at the MSA level by comparing our residential land price indexes to the Case-Shiller home price indexes. Moreover, a decomposition of the changes in home prices into the contributions from construction costs and land prices shows that land prices were by far the more important driver of the recent boom-bust cycle in home prices.

The CoStar data on land transactions open up many possibilities for future research. One fruitful avenue would be to investigate the sources of the wide swings in land prices. In particular, to what degree have these movements reflected the availability and cost of financing, the use of leverage in property transactions, supply and demand fundamentals in local real estate markets, or broader economic conditions? A second avenue would be to develop additional evidence on the share of land in the prices of homes and commercial real estate. This would be especially valuable for commercial property, for which no estimates are currently available.

References

- Albouy, David, and Gabriel Ehrlich. 2011. "Metropolitan Land Values and Housing Productivity." Working paper, University of Michigan.
- Anas, Alex, Richard Arnott, and Kenneth A. Small. 1998. "Urban Spatial Structure." *Journal of Economic Literature* 36: 1426-64.
- Barker, David. 2007. "Urban Land Rents in the United States". In *Land Policies and Their Outcomes*, edited by Gregory K. Ingram and Yu-Hung Hong. Cambridge, MA: Lincoln Institute of Land Policy.
- Bostic, Raphael W., Stanley D. Longhofer, and Christian L. Redfearn. 2007. "Land Leverage: Decomposing Home Price Dynamics." *Real Estate Economics* 35(2): 183-208.
- Bourassa, Steven C. et al. 2010. "Land Leverage and House Prices." Swiss Finance Institute Research Paper Series no. 10-48.
- Bourassa, Steven C. et al. 2009. "House Price Changes and Idiosyncratic Risk: The Impact of Property Characteristics." *Real Estate Economics* 37(2): 259-78.
- Brasington, David M., and Donald R. Haurin. 2009. "Parents, Peers, or School Inputs: Which Components of School Outcomes are Capitalized into House Value?" *Regional Science and Urban Economics* 39(5): 523-9.
- Brownstone, David, and Arthur De Vany. 1991. "Zoning, Returns to Scale, and the Value of Undeveloped Land." *Review of Economics and Statistics* 73(4): 699-704.
- Bryan, Kevin A., and Pierre-Daniel G. Sarte. 2009. "Semiparametric Estimation of Land Price Gradients Using Large Data Sets." *Federal Reserve Bank of Richmond Economic Quarterly* 95(1): 53-74.
- Bureau of Labor Statistics. 2007. "Technical Information About the BLS Multifactor Productivity Measures." (September 26). Posted at www.bls.gov/mfp/mprtech.pdf
- Case, Karl E. 2007. "The Value of Land in the United States: 1975-2005". In *Land Policies and Their Outcomes*, edited by Gregory K. Ingram and Yu-Hung Hong. Cambridge, MA: Lincoln Institute of Land Policy.
- Case, Karl E., and Robert J. Shiller. 1989. "The Efficiency of the Market for Single-Family Homes." *American Economic Review* 79(1): 125-37.
- Clapp, John M., Anupam Nanda, and Stephen L. Ross. 2008. "Which School Attributes Matter? The Influence of School District Performance and Demographic Composition on Property Values." *Journal of Urban Economics* 63(2): 451-66.

- Clark, Steven P., and T. Daniel Coggin. 2011. "Was There a U.S. House Price Bubble? An Econometric Analysis Using National and Regional Panel Data." *The Quarterly Review of Economics and Finance* 51(2): 189-200.
- Coleman, Major IV, Michael LaCour-Little, and Kerry D. Vandell. 2008. "Subprime Lending and the Housing Bubble: Tail Wags Dog?" *Journal of Housing Economics* 17(4): 272-90.
- Colwell, Peter F. 1998. "A Primer on Piecewise Parabolic Multiple Regression Analysis via Estimations of Chicago CBD Land Prices." *Journal of Real Estate Finance and Economics* 17(1): 87-97.
- Colwell, Peter F., and Henry J. Munneke. 2003. "Estimating a Price Surface for Vacant Land in an Urban Area." *Land Economics* 79(1): 15-28.
- Colwell, Peter F., and Henry J. Munneke. 1997. "The Structure of Urban Land Prices." *Journal of Urban Economics* 41(3): 321-36.
- Craig, Steven G., and Pin T. Ng. 2001. "Using Quantile Smoothing Splines to Identify Employment Subcenters in a Multicentric Urban Area." *Journal of Urban Economics* 49(1): 100-20
- Davis, Morris A. 2009. "The Price and Quantity of Land by Legal Form of Organization in the United States." *Regional Science and Urban Economics* 39(3): 350-9.
- Davis, Morris A., and Jonathan Heathcote. 2007. "The Price and Quantity of Residential Land in the United States." *Journal of Monetary Economics* 54(8): 2595-2620.
- Davis, Morris A., and Michael G. Palumbo. 2008. "The Price of Residential Land in Large US Cities." *Journal of Urban Economics* 63(1): 352-84.
- Downing, Paul B. 1970. "Estimating Residential Land Value by Multivariate Analysis." In *The Assessment of Land Value*, edited by Daniel M. Holland. Madison, WI: University of Wisconsin Press.
- Economic Report of the President. 2009. Washington DC: Government Printing Office.
- Fisher, Jeff, David Geltner, and Henry Pollakowski. 2007. "A Quarterly Transactions-based Index of Institutional Real Estate Investment Performance and Movements in Supply and Demand." *The Journal of Real Estate Finance and Economics* 34: 5-33.
- Flood, Robert P., and Robert J. Hodrick. 1990. "On Testing for Speculative Bubbles." *The Journal of Economic Perspectives* 4(2): 85-101
- Gatzlaff, Dean H., and Donald R. Haurin. 1998. "Sample Selection and Biases in Local House Value Indices." *Journal of Urban Economics* 43(2): 199-222.

- Gatzlaff, Dean H., and Donald R. Haurin. 1997. "Sample Selection Bias and Repeat-Sales Index Estimates." *The Journal of Real Estate Finance and Economics* 14: 33-50.
- Giuliano, Genevieve, and Kenneth A. Small. 1991. "Subcenters in the Los Angeles Region." *Regional Science and Urban Economics* 21(2): 163-82.
- Glaeser, Edward L. 2008. *Cities, Agglomeration, and Spatial Equilibrium*. New York: Oxford University Press.
- Goldsmith, Raymond W. 1951. "A Perpetual Inventory of National Wealth." In *Studies in Income and Wealth, vol. 14*, edited by Martin R. Gainsbrugh and Elwyn T. Bonnell: 5-61. New York: National Bureau of Economic Research.
- Goodman, Allen C., and Thomas G. Thibodeau. 2008. "Where are the Speculative Bubbles in US Housing Markets?" *Journal of Housing Economics* 17(2): 117-37.
- Greenlees, John S. 1980. "Residential Land Price Indices Using Multiple Regression." In *Urban Land Markets: Price Indices, Supply Measures, and Public Policy Effects*, edited by J. Thomas Black and James E. Hoben. Washington DC: The Urban Land Institute.
- Guntermann, Karl L., and Gareth Thomas. 2005. "Parcel Size, Location, and Commercial Land Values." *Journal of Real Estate Research* 27(3): 343-54.
- Gyourko, Joseph. 2009. "Housing Supply." *Annual Review of Economics* 1: 295-318.
- Gyourko, Joseph, and Albert Saiz. 2006. "Construction Costs and the Supply of Housing Infrastructure." *Journal of Regional Science* 46(4): 661-80.
- Gyourko, Joseph, Albert Saiz, and Anita Summers. 2008. "A New Measure of the Local Regulatory Environment for Housing Markets: The Wharton Residential Land Use Regulatory Index." *Urban Studies* 45(3): 693-729.
- Haughwout, Andrew, James Orr, and David Bedoll. 2008. "The Price of Land in the New York Metropolitan Area." Federal Reserve Bank of New York, *Current Issues in Economics and Finance* 14(3): 1-7
- Heckman, James. 1979. "Sample Selection Bias as a Specification Error." *Econometrica* 47: 153-61
- Ihlanfeldt, Keith R. 2007. "The Effect of Land Use Regulation on Housing and Land Prices." *Journal of Urban Economics* 61(3): 420-35.
- Ihlandeldt, Keith, and Tom Mayock. 2010. "Panel Data Estimates of the Effects of Different Types of Crime on Housing Prices." *Regional Science and Urban Economics* 40(2-3): 161-72.

Jud, G. Donald, and Terry G. Seaks. 1994. "Sample Selection Bias in Estimating Housing Sales Prices." *Journal of Real Estate Research* 9(3): 289-98.

Kok, Nils, Paavo Monkkonen, and John M. Quigley. 2010. "Economic Geography, Jobs, and Regulations: The Value of Land and Housing." Working paper.

Kowalski, Joseph G., and Christos C. Paraskevopoulos. 1990. "The Impact of Location on Urban Industrial Land Prices." *Journal of Urban Economics* 27(1): 16-24

Lee, Seung Jung, and Jonathan D. Rose. 2010. "Profits and Balance Sheet Developments at U.S. Commercial Banks in 2009." *Federal Reserve Bulletin* 96 (May): A1-37. Posted at www.federalreserve.gov/pubs/bulletin/2010/default.htm.

Manvel, Allan D. 1968. "Trends in the Value of Real Estate and Land, 1956 to 1966." In *Three Land Research Studies*, prepared for the consideration of the National Commission on Urban Problems. Washington DC: Government Printing Office.

Mayer, Christopher. 2011. "Housing Bubbles: A Survey." *Annual Review of Economics* 3: 559-77.

McMillen, Daniel P. 1996. "One Hundred Fifty Years of Land Values in Chicago: A Nonparametric Approach." *Journal of Urban Economics* 40(1): 100-24.

McMillen, Daniel P., and John F. McDonald. 2002. "Land Values in a Newly Zoned City." *The Review of Economics and Statistics* 84(1): 62-72.

McMillen, Daniel P., and John F. McDonald. 1998. "Suburban Subcenters and Employment Density in Metropolitan Chicago." *Journal of Urban Economics* 43(2): 157-80.

McMillen, Daniel P., and John F. McDonald. 1991. "A Simultaneous Equations Model of Zoning and Land Values." *Regional Science and Urban Economics* 21(1): 55-72.

McMillen, Daniel P., and Stefani C. Smith. 2003. "The Number of Subcenters in Large Urban Areas." *Journal of Urban Economics* 53(3): 321-38.

Meese, Richard, and Nancy Wallace. 1994. "Testing the Present Value Relation for Housing Prices: Should I Leave My House in San Francisco?" *Journal of Urban Economics* 35(3): 245-66.

Mikhed, Vyacheslav, and Petr Zemčík. 2009. "Do House Prices Reflect Fundamentals? Aggregate and Panel Data Evidence." *Journal of Housing Economics* 18(2): 140-9.

Milgram, Grace. 1973. "Appendix II, Estimates of the Value of Land in the United States Held by Various Sectors of the Economy, Annually, 1952 to 1968." In *Institutional Investors and Corporate Stock – A Background Study*, edited by Raymond W. Goldsmith: 343-77. New York: National Bureau of Economic Research.

- Munneke, Henry J. 2005. "Dynamics of the Urban Zoning Structure: An Empirical Investigation of Zoning Change." *Journal of Urban Economics* 58(3): 455-73.
- Munneke, Henry J., and Barrett A. Slade. 2001. "A Metropolitan Transaction-Based Commercial Price Index: A Time-Varying Parameter Approach." *Real Estate Economics* 29(1): 55-84.
- Munneke, Henry J., and Barrett A. Slade. 2000. "An Empirical Study of Sample-Selection Bias in Indices of Commercial Real Estate." *The Journal of Real Estate Finance and Economics* 21(1): 45-64.
- Peiser, Richard B. 1987. "The Determinants of Nonresidential Land Values." *Journal of Urban Economics* 22(3): 340-60.
- Pogodzinski, J. M., and Tim R. Sass. 1994. "The Theory and Estimation of Endogenous Zoning." *Regional Science and Urban Economics* 24: 601-30.
- Pope, Jaren C. 2008. "Fear of Crime and Housing Prices: Household Reactions to Sex Offender Registries." *Journal of Urban Economics* 64(3): 601-14.
- Redfearn, Christian L. 2007. "The Topography of Metropolitan Employment: Identifying Centers of Employment in a Polycentric Urban Area." *Journal of Urban Economics* 61(3): 519-41.
- Ries, John, and Tsur Somerville. 2010. "School Quality and Residential Property Values: Evidence from Vancouver Rezoning." *The Review of Economics and Statistics* 92(4): 928-44.
- Rosenthal, Stuart S. 1999. "Residential Buildings and the Cost of Construction: New Evidence on the Efficiency of the Housing Market." *The Review of Economics and Statistics* 81(2): 288-302.
- Rosenthal, Stuart S., and Robert W. Helsley. 1994. "Redevelopment and the Urban Land Price Gradient." *Journal of Urban Economics* 35(2): 182-200.
- Saiz, Albert. 2010. "The Geographic Determinants of Housing Supply." *Quarterly Journal of Economics* 125(3): 1253-96.
- Schulz, Rainer, and Axel Werwatz. 2011. "Is There an Equilibrating Relationship between House Prices and Replacement Cost? Empirical Evidence from Berlin." *Journal of Urban Economics* 69(3): 288-302.
- Sirmans, C.F., and Barrett A. Slade. 2011. "National Transaction-based Land Price Indices." *The Journal of Real Estate Finance and Economics*. Published online. Available online at www.springerlink.com/content/

Sivitanidou, Rena. 1996. "Do Office-Commercial Firms Value Access to Service Employment Centers? A Hedonic Value Analysis within Polycentric Los Angeles." *Journal of Urban Economics* 40(2): 125-49.

Voith, Richard. 2001. "How Responsive Is the Demand for Residential Land to Changes in Its Price?" *Federal Reserve Bank of Philadelphia Business Review*, Q3: 33-9.

Wallace, Nancy E. 1988. "The Market Effects of Zoning Undeveloped Land: Does Zoning Follow the Market?" *Journal of Urban Economics* 23(3): 307-26.

Wieand, Kenneth, Jr., and Richard F. Muth. 1972. "A Note on the Variation of Land Values with Distance from the CBD in St. Louis." *Journal of Regional Science* 12(3): 469-73.

Wu, Jing, Joseph Gyourko, and Yongheng Deng. 2010. "Evaluating Conditions in Major Chinese Housing Markets." National Bureau of Economic Research Working Paper no. 16189.

Appendix A

This appendix details the construction of the dataset used in the paper. We apply a sequence of filters to remove transactions that lacked sales prices or supporting documentation, were not at arms length, contained apparent data recording errors, fell too far outside the MSA-specific grid, or were from half-year periods with an insufficient number of observations for a given MSA. The CoStar database contains several detailed fields with notes on the property and transaction. We scan these fields for key phrases to identify observations that should be excluded from our analysis. Table A.1 reports the initial sample size by year and the sequential effects of each filter.

The first screen removed transactions for which the sale price was missing or for which CoStar collected the transaction information solely from public records without contacting the participants to confirm the transaction details. Given the lack of confirmation, CoStar staff recommended that we exclude these transactions from our analysis.

Although the CoStar notes state whether a transaction has been classified as non-arms-length, we constructed our own, more comprehensive definition of these and other transactions that do not represent the sale of land at current market prices. We removed transfers of deeds in lieu of foreclosure, foreclosed properties seized by lenders, and other properties acquired by lenders that lack any notes providing additional clarification.²⁶ We also exclude any transaction in which the buyer and seller have either the same name or the same address, transfers within a company or family, section 1031 transfers, direct exchanges, transfers of partial interest, gifts, eminent domain acquisitions, the exercise of an existing option, and transactions that include a ground lease. Finally, we excluded any transactions that passed these screens but that CoStar

²⁶ As noted in section 2, foreclosed parcels subsequently sold by lenders would pass this screen, but there were too few such sales to generate meaningful results. Accordingly, we removed all foreclosure transactions from the dataset.

classifies as a non-arms-length or non-market sale or as a sale in which the price does not represent the true value of the land.

We also screened the data for observations where the price per square foot was an outlier that appears to reflect data errors in two separate ways. First, transactions were excluded that had an unusually high or low price per square foot and for which the notes indicate that reported values could not be confirmed, were misleading, or reflected considerations other than the land itself. Second, we excluded observations where the reported gross square footage of the site was dramatically different from the reported net square footage. The difference between the two represents the part of the site that is not buildable, but we were concerned that very large differences could indicate a data recording error.

We controlled for spatial variation within each MSA using Colwell's (1998) semiparametric approach. This technique, described in section 3, superimposed a grid over a map of each MSA, where the grid consists of equal-size squares. As required by Colwell's method, observations that were more than one square outside this grid – those on the periphery of the MSA – were excluded from the analysis.

Our final filter excluded observations from half-yearly periods that lacked sufficient observations to reliably estimate the value of the corresponding time dummy for that particular MSA. As described in section 2, we set this threshold at 20 observations for residential land and commercial land separately when we estimate MSA-level price indexes for each property type and at 40 observations for the two property types taken together when we estimate an aggregate MSA-level price index. The 20-observation threshold applied to each property type is the more restrictive of the two tests, and the table reports the effects of this filter.

A final issue concerns the time lag before a transaction appears in the CoStar database. Our analysis of the historical data indicates that 66 percent of sales are recorded in the database within three months of their sale date, 86 percent after six months, and 93 percent after one year. Based on these results, we judged that sufficient data were available to estimate preliminary price indexes for a given half-year period six months after the end of the period.

Appendix B

This appendix presents the results of a large set of likelihood ratio tests for the equality of coefficients in our hedonic regression. For each hypothesis, the test statistic is $-2 \ln(\lambda^R/\lambda^U)$, where λ^R is the value of the likelihood function after the equality restrictions have been imposed and λ^U is the value of the unrestricted likelihood function. This statistic is distributed $\chi^2(\rho)$, where ρ equals the number of restrictions imposed by the null hypothesis.

Each column in table B.1 reports the test results for a subset of independent variables. The column labeled “Condition of property,” for example, includes the nine dummy variables that fall under this heading, for which the coefficient estimates were summarized in table 4. Similarly, the column labeled “Time effects” includes the full set of half-yearly dummy variables. For any column in the table, the first row shows the results for the most restrictive null hypothesis – that every variable in that column has a single coefficient across all 23 MSAs and both property types. The second row tests a less restrictive null hypothesis – that each variable has a single coefficient across all MSAs for residential land and a separate coefficient across all MSAs for commercial land. The next row tests the analogous null within MSAs – that each variable has a single coefficient across residential and commercial land in every MSA, while allowing this coefficient to differ across MSAs. The remaining rows test the hypothesis that each variable has the same coefficient across the two property types for MSAs taken one at a time.

Each entry in table B.1 presents the p-value for a particular hypothesis test. p-values smaller than 0.05 indicate that the null hypothesis is rejected at the five-percent level of significance, while p-values smaller than 0.01 indicate rejection at the one-percent level. The p-

values greater than 0.05 have been shaded to highlight the hypotheses that cannot be rejected at the standard five-percent level.

The results in the first three lines of the table provide strong evidence in favor of the highly disaggregated regression model that we estimate instead of a model that aggregates the individual MSAs, the two broad property types, or both. As shown, all of these aggregation hypotheses are rejected at the one-percent level of significance.²⁷

The remaining rows of the table show that some of the coefficient restrictions cannot be rejected for individual MSAs. For example, in 18 of the 23 MSAs we cannot reject the hypothesis that distance from the CBD has the same price effect on residential and commercial land. Similarly, in six MSAs we cannot reject the hypothesis that parcel size has the same price effect on both types of property. However, in every MSA at least one of the equality restrictions is rejected at the one-percent level. Accordingly, it is not appropriate to estimate an aggregated model for residential and commercial land prices in any of the MSAs.

²⁷ Note that we do not test the equality of the locational grids across MSAs. Given the differences in geographic features across MSAs, we would expect the price surfaces to differ as well.

Table 1
Sample Size

MSA	Starting period	Number of sales		
		Total	Commercial	Residential
Total	---	180,155	99,152	80,231
Atlanta	1995:H1	17,004	8,626	8,378
Baltimore	1995:H2	3,616	1,952	1,608
Boston	1995:H1	3,841	2,028	1,724
Chicago	1995:H1	14,610	8,707	5,903
Dallas	1995:H2	4,911	4,039	728
Denver	1995:H2	8,492	4,702	3,775
Detroit	1999:H2	3,347	2,223	1,116
Houston	1999:H2	2,920	2,467	330
Las Vegas	1995:H1	10,449	4,861	5,571
Los Angeles	1995:H1	13,970	9,631	4,339
New York	1998:H2	7,023	4,181	2,842
Orlando	1995:H1	7,819	4,539	3,242
Philadelphia	1998:H1	5,311	2,729	2,488
Phoenix	1995:H1	18,249	8,472	9,777
Portland	1995:H1	6,589	2,601	3,951
Sacramento	1995:H1	2,782	2,010	746
San Diego	1995:H1	3,261	2,152	1,118
San Francisco	1995:H1	4,584	3,153	1,385
Seattle	1995:H1	9,609	3,893	5,716
South Florida	1997:H2	9,507	6,234	3,257
Tampa/St. Petersburg	1997:H2	6,928	4,355	2,565
Tucson	1995:H1	5,312	2,009	3,283
Washington DC	1995:H1	10,021	3,588	6,389

Note. Los Angeles is defined to include Orange County and the Inland Empire; New York is defined to include northern New Jersey, Westchester County, and Long Island; and San Francisco is defined to include Marin/Sonoma, East Bay/Oakland, and South Bay/San Jose. The number of observations in the “Total” column does not equal the sum of the observations in the “Commercial” and “Residential” columns because of the sample construction rules we applied. See the text for details.

Source. Authors’ analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 2
Summary Statistics for Parcel Size and Distance from CBD

Variable	10 th pct.	Median	90 th pct.
Parcel size (acres)			
Residential	.6	7.2	77.7
Commercial	.4	2.5	20.1
Distance from CBD (miles)			
Residential	6.7	19.6	42.8
Commercial	4.8	17.6	44.2

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 3
Shares of Sample by Indicator Variable
(percent)

Variable	Residential	Commercial
<i>Type of property</i>		
Single-family	45.2	
Multifamily rental	12.0	
Other residential	42.8	
Office		8.1
Industrial		22.3
Retail		16.3
Other commercial		53.2
<i>Condition of property</i>		
Unimproved	57.3	49.3
Graded	4.4	8.2
Paved	.6	2.7
Finished	18.6	6.9
Fully improved	2.9	2.0
Platted and engineered	2.8	1.4
Previously developed	4.0	6.2
Structure present	33.4	30.8
Improvements unknown	8.9	19.1
Environmental problems	.1	.4
<i>Intended use</i>		
Private development	76.7	74.6
Hold for investment	11.1	8.8
Open space	1.2	.5
Public use	1.1	2.2
Unknown	10.0	13.8
Expansion of adjacent property	.3	1.6

Note. The property types are mutually exclusive within the residential and commercial samples. The condition of property categories are not mutually exclusive and may overlap. The intended use categories, with the exception of “Expansion of adjacent property”, are mutually exclusive.

Source. Authors’ analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 4
Coefficient Estimates Across MSAs

Variable	Residential				Commercial			
	Median	Range Min. Max.	# Significant ¹ Neg. Pos.		Median	Range Min. Max.	# Significant ¹ Neg. Pos.	
Log of parcel size	-.54	-.71 -.19	23	0	-.50	-.65 -.33	23	0
<i>Type of property</i>								
Office					.44	.22 .69	0	23
Retail					.61	.28 .81	0	23
Other commercial					.26	.04 .42	0	22
Single-family	-.13	-.66 .04	12	0				
Other residential	-.37	-.74 -.09	22	0				
<i>Condition of property</i>								
Graded	.14	-.11 .91	0	10	.10	-.07 .31	0	9
Paved	.20	-.29 1.48	1	5	.11	-.05 .51	0	7
Finished	.26	-.15 1.84	1	17	.11	-.03 .35	0	8
Fully improved	.13	-.23 1.34	1	6	.27	.02 .60	0	15
Platted and engineered	.14	-.34 .70	0	7	.17	-.09 .45	0	7
Previously developed	.17	-.24 1.20	0	10	.27	.01 .46	0	19
Structure present	.06	-.81 .30	1	8	.01	-.11 .15	0	3
Improvements unknown	.08	-.15 .49	1	8	-.11	-.27 .05	12	0
Environmental problems	.15	-.98 .92	1	0	-.07	-.51 .26	1	0
<i>Intended use</i>								
Hold for investment	-.25	-.47 .00	13	0	-.14	-.26 .09	16	0
Open space	-.33	-.94 .21	12	0	-.31	-1.65 .75	12	0
Public use	.02	-.37 .50	1	3	.13	-.31 .40	1	9
Unknown	-.05	-.55 .19	5	3	-.09	-.17 .03	11	0
Expansion of adjacent property	-.05	-.63 .48	1	0	-.12	-.35 .01	7	0

1. At the five-percent level.

Note. The omitted property type for the residential observations is "multifamily rental" and the omitted property type for the commercial observations is "industrial". The omitted condition of property is "unimproved", and the omitted intended use is "private development". Some MSAs lacked observations to estimate every coefficient. For example, six MSAs had no residential land sales reported with environmental problems; the results in that line of the table are based on estimates for the other 17 MSAs.

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 5
Price Indexes for Composite of Commercial
and Residential Land by MSA

MSA	Date of		Percent change	
	Peak	Post-peak low	2002:H2 to peak	Peak to low
Atlanta	2007:H1	2011:H1	91	-70
Baltimore	2006:H2	2009:H1	108	-56
Boston	2005:H2	2009:H1	145	-69
Chicago	2005:H2	2009:H1	106	-60
Dallas	2006:H2	2009:H1	57	-40
Denver	2008:H1	2009:H1	52	-33
Detroit	2005:H1	2011:H1	82	-73
Houston	2006:H1	2009:H2	100	-47
Las Vegas	2007:H2	2011:H1	287	-76
Los Angeles	2006:H2	2011:H1	165	-63
New York	2007:H2	2009:H1	228	-49
Orlando	2006:H2	2010:H1	146	-60
Philadelphia	2007:H2	2009:H1	145	-47
Phoenix	2006:H1	2011:H1	195	-68
Portland	2006:H1	2010:H2	178	-53
Sacramento	2006:H2	2010:H2	148	-61
San Diego	2005:H2	2011:H1	145	-42
San Francisco	2006:H2	2009:H1	164	-62
Seattle	2006:H2	2009:H1	158	-62
South Florida	2007:H2	2009:H1	205	-71
Tampa	2005:H2	2011:H1	132	-60
Tucson	2007:H1	2009:H1	116	-57
Washington DC	2007:H1	2009:H1	241	-46
Median across MSAs	2006:H2	2009:H1	145	-60

Note. See table 1 for definitions of selected MSAs.

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 6
Price Indexes for Commercial Land by MSA

MSA	Date of		Percent change	
	Peak	Post-peak low	2002:H2 to peak	Peak to low
Atlanta	2007:H1	2011:H1	94	-68
Baltimore	2007:H2	2009:H1	117	-59
Boston	2005:H2	2009:H1	167	-69
Chicago	2006:H2	2011:H1	121	-59
Dallas	2006:H2	2009:H1	64	-40
Denver	2008:H1	2009:H1	66	-34
Detroit	2005:H2	2010:H1	55	-73
Houston	2006:H1	2009:H2	111	-47
Las Vegas	2007:H2	2011:H1	264	-76
Los Angeles	2006:H2	2011:H1	162	-63
New York	2007:H2	2009:H1	202	-42
Orlando	2007:H2	2010:H1	155	-65
Philadelphia	2007:H2	2009:H1	143	-49
Phoenix	2007:H2	2011:H1	131	-65
Portland	2007:H1	2011:H1	95	-38
Sacramento	2006:H2	2010:H2	163	-61
San Diego	2005:H2	2010:H2	142	-44
San Francisco	2007:H2	2009:H1	138	-60
Seattle	2007:H2	2009:H1	161	-49
South Florida	2007:H2	2009:H1	225	-67
Tampa	2007:H1	2011:H1	118	-61
Tucson	2007:H1	2011:H1	139	-58
Washington DC	2007:H1	2010:H1	218	-57
Median across MSAs	2007:H1	2010:H1	139	-59

Note. See table 1 for definitions of selected MSAs.

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 7
Price Indexes for Residential Land by MSA

MSA	Date of		Percent change	
	Peak ¹	Post-peak low ²	2002:H2 to peak ¹	Peak to low
Atlanta	2006:H2	2011:H1	93	-74
Baltimore	2006:H1	2009:H1	129	-55
Boston	2005:H2	NA	133	NA
Chicago	2005:H2	2009:H1	128	-76
Dallas	NA	NA	NA	NA
Denver	2006:H2	2011:H1	42	-38
Detroit	2005:H1	2009:H2	138	-83
Houston	NA	NA	NA	NA
Las Vegas	2007:H1	2011:H1	334	-77
Los Angeles	2006:H2	2009:H1	166	-70
New York	2007:H2	2009:H2	336	-68
Orlando	2005:H2	2011:H1	218	-76
Philadelphia	2005:H2	2009:H2	180	-55
Phoenix	2006:H1	2009:H1	292	-80
Portland	2006:H1	2010:H2	266	-69
Sacramento	2005:H2	2009:H2	151	-55
San Diego	2005:H2	2007:H2	167	-52
San Francisco	2006:H2	2009:H2	248	-64
Seattle	2006:H2	2009:H1	169	-71
South Florida	2006:H1	2009:H1	183	-83
Tampa	2005:H2	2009:H2	197	-77
Tucson	2006:H1	2009:H1	129	-60
Washington DC	2006:H1	2008:H2	145	-47
Median across MSAs	2006:H1	2009:H2	167	-69

1. Calculated only for MSAs for which the price index is available through at least 2006:H2 or for which the available data prior to 2006:H2 indicate an earlier peak.

2. Calculated only for MSAs for which the price index is available in 2009:H1 or later.

Note. See table 1 for definitions of selected MSAs. NA indicates that the price index is not available.

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

Table 8
Price Indexes for Land, Homes, and Commercial Real Estate

Index	Date of		Percent Change	
	Peak	Post-peak low	2002:H2 to peak	Peak to low
<i>Residential</i>				
Land price index	2006:H1	2009:H1	164	-63
S&P/Case-Shiller 20-city home price index	2006:H2	2011:H1	55	-32
<i>Commercial</i>				
Land price index	2007:H2	2011:H1	125	-52
CoStar commercial repeat-sales index	2007:H2	2009:H2	80	-35

Note. The land price indexes are those calculated for the aggregate of all 23 MSAs, which were shown in the lower panel of figure 6. The CoStar commercial repeat-sales index is a value-weighted index produced for the Federal Reserve Board that includes office, retail, and industrial properties for the U.S. as a whole and is not limited to the 23 MSAs in our sample. The S&P/Case-Shiller index covers single-family homes; the 20 cities in the home price index do not map exactly into the 23 MSAs in our land price index. All figures in the table are calculated from data that are not seasonally adjusted.

Sources. Land price indexes: Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

CoStar commercial repeat-sales index: CoStar. S&P/Case-Shiller index: Standard and Poor's.

Table 9
Comparing Residential Land, Structures, and Home Prices by MSA

MSA	Percent change 2002:H2 to peak			Percent change peak to low		
	Land	Structures	Home	Land	Structures	Home
Atlanta	93	31	17	-74	-7	-24
Boston	133	20	25	NA	NA	NA
Chicago	128	22	29	-76	21	-22
Denver	42	30	11	-38	17	-12
Los Angeles	166	28	96	-70	14	-41
New York	336	34	46	-68	8	-15
Phoenix	292	40	93	-80	21	-52
Portland	266	27	51	-69	20	-16
San Diego	167	23	66	-52	14	-12
San Francisco	248	26	51	-64	24	-38
Seattle	169	33	58	-71	16	-18
South Florida ¹	183	37	99	-83	19	-46
Tampa	197	37	66	-77	27	-34
Washington DC	145	33	75	-47	19	-25
Median across MSAs²	168	31	55	-70	19	-24

1. The Miami Case-Shiller index is used to proxy for homes in South Florida.

2. Non-integer median values have been rounded up to the nearest integer.

Note. NA indicates that the land price index is not available in any period after 2008:H2. The table includes all MSAs in the Case-Shiller 20-city home price index for which we calculate a land price index that extends far enough through the sample to designate a peak date and for which there was an estimate of structure prices from Davis and Palumbo (2008). The dates for the peak and low differ across MSAs and are defined from the series for land prices in each MSA. For a given MSA, the same peak and low dates are used for land, structures, and home prices. All results are calculated from data that are not seasonally adjusted.

Source. Land prices based on authors' analysis of data from the CoStar Group, Inc. (www.costar.com). Home prices based on S&P/Case-Shiller data from Standard & Poor's. Structures prices are from the online data files for Davis and Palumbo (2008) (<http://www.lincolinst.edu/subcenters/land-values/>).

Table 10
Decomposition of Home Price Changes by MSA

MSA	2002:H2 to peak						Peak to low					
	Home prices			Contributions from			Home prices			Contributions from		
	CS	Implied	Land	Struct.	Land	Avg. land share	CS	Implied	Land	Struct.	Land	Avg. land share
	(percent change)	(percent change)	(percentage points)	(percentage points)	(percentage points)	(percent)	(percent change)	(percent change)	(percentage points)	(percentage points)	(percentage points)	(percent)
Atlanta	17	48	26	22	26	28	-24	-15	-9	-6	-9	12
Boston	25	102	97	5	97	73	NA	NA	NA	NA	NA	NA
Chicago	29	74	63	11	63	49	-22	-20	-32	12	-32	42
Denver	11	35	19	17	19	45	-12	0	-11	12	-11	30
Los Angeles	96	136	129	6	129	78	-41	-49	-53	4	-53	75
New York	46	233	222	12	222	66	-15	-37	-40	3	-40	59
Phoenix	93	181	164	18	164	56	-52	-26	-38	11	-38	47
Portland	51	161	149	12	149	56	-16	-28	-37	9	-37	54
San Diego	66	137	132	5	132	79	-12	-38	-41	3	-41	79
San Francisco	51	219	216	3	216	87	-38	-48	-52	4	-52	82
Seattle	58	117	105	13	105	62	-18	-38	-44	6	-44	62
South Florida ¹	99	139	128	11	128	70	-46	-48	-55	6	-55	66
Tampa	66	115	97	19	97	49	-34	-16	-32	16	-32	41
Washington DC	75	107	96	11	96	66	-25	-23	-30	7	-30	63
Median across MSAs²	55	126	116	11	116	64	-24	-28	-38	6	-38	59

1. The Miami Case-Shiller index is used to proxy for homes in South Florida.

2. Non-integer median values have been rounded up to the nearest integer.

Note. CS stands for the S&P/Case Shiller index. NA indicates that the land price index is not available in any period after 2008:H2. The table includes all MSAs in the Case-Shiller 20-city home price index for which we calculate a land price index that extends far enough through the sample to designate a peak date and for which Davis and Palumbo (2008) provide estimates of structures prices and the land share of home value. The dates for the peak and low differ across MSAs and are defined from the series for land prices in each MSA. For a given MSA, the same peak and low dates are used for all series shown. The contributions from land and structures may not sum to the implied change in home prices due to rounding. All results are calculated from data that are not seasonally adjusted. Source. Land prices based on authors' analysis of data from the CoStar Group, Inc. (www.costar.com). Home prices based on S&P/Case-Shiller data from Standard & Poor's. Structure prices and average land shares are from the online data files for Davis and Palumbo (2008) (<http://www.lincolninst.edu/subcenters/land-values/>).

Table 11
Cross-MSA Regressions of Home Prices on Contributions from
Structures and Land Prices

Variable	2002:H2 to peak	Peak to low
Constant	0.252 (0.317)	-0.100 (0.112)
Structures price contribution	0.248 (1.567)	-0.251 (0.674)
Land price contribution	0.238 (0.147)	0.429 (0.266)
R ²	0.223	0.218

Note: Dependent variable is the S&P Case-Shiller home price index. Standard errors in parentheses.

Table A.1
Construction of Sample by Year

Year	(1) Initial sample	(2) Price missing or public record	(3) Non-arms length	(4) Outliers	(5) Too far outside grid	(6) Fails MSA obs. test	(7) Final sample
1995	8,230	584	231	0	547	23	6,845
1996	9,912	682	258	2	441	37	8,492
1997	11,598	874	370	0	580	66	9,708
1998	13,972	777	633	1	240	52	12,269
1999	16,338	858	860	8	180	30	14,402
2000	15,798	697	862	5	106	39	14,089
2001	13,792	585	802	7	101	0	12,297
2002	15,231	498	860	6	115	28	13,724
2003	16,410	392	805	9	115	0	15,089
2004	18,190	702	818	28	164	66	16,412
2005	16,103	340	747	31	249	60	14,676
2006	15,399	1,094	702	29	864	97	12,613
2007	13,841	2,671	860	18	1,143	196	8,953
2008	12,033	3,311	1,147	21	456	198	6,900
2009	14,799	7,145	1,931	18	361	180	5,164
2010	12,605	5,218	1,690	13	342	110	5,232
2011	6,261	2,842	634	2	190	75	2,518
Total	230,512	29,270	14,210	198	6,194	1,257	179,383

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com).

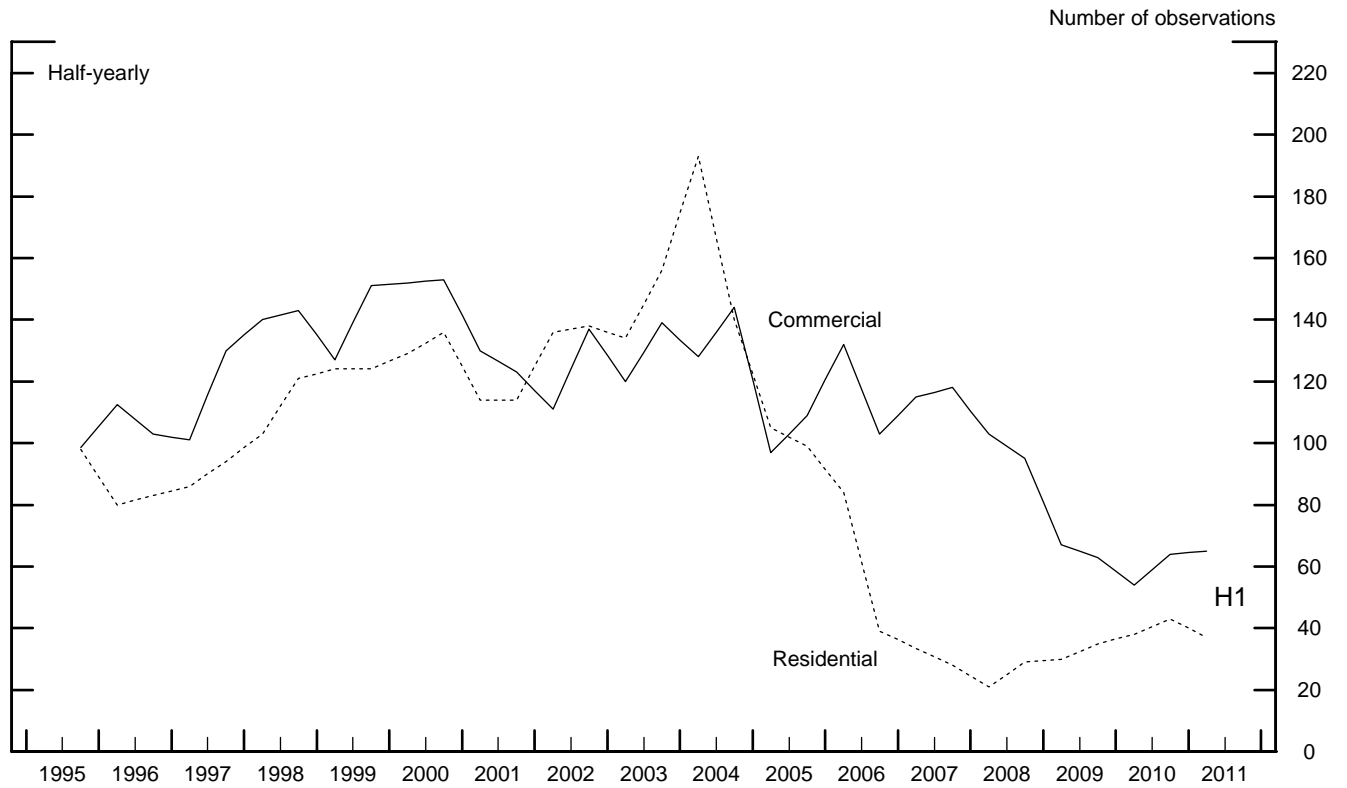
Table B.1
P-values from Hypothesis Tests for Equivalence of Coefficient Estimates

	Parcel size	Condition of property	Intended use	Distance from CBD	Locational grids	Time effects
Across MSAs and property type	<.001	<.001	<.001	<.001	NA	<.001
Across MSAs, by property type	<.001	<.001	<.001	<.001	NA	<.001
Across property type, by MSA						
All MSAs	<.001	<.001	<.001	<.001	<.001	<.001
Atlanta	<.001	.020	.235	.006	<.001	<.001
Baltimore	<.001	.033	.662	.603	.022	<.001
Boston	<.001	.021	.442	.528	.001	<.001
Chicago	.038	.013	.042	.290	<.001	<.001
Dallas	.256	.001	.012	.909	.101	.429
Denver	<.001	<.001	<.001	<.001	<.001	<.001
Detroit	.205	.008	.669	.170	.016	.003
Houston	<.001	<.001	.051	.002	<.001	.003
Las Vegas	<.001	<.001	.002	.116	<.001	<.001
Los Angeles	<.001	<.001	.030	.914	<.001	<.001
New York	<.001	<.001	.103	.107	<.001	<.001
Orlando	.453	.370	.004	.939	<.001	<.001
Philadelphia	.005	<.001	.797	.471	.258	.003
Phoenix	.001	<.001	<.001	<.001	<.001	<.001
Portland	<.001	<.001	<.001	.597	.007	<.001
Sacramento	<.001	.022	.403	.066	.082	<.001
San Diego	<.001	.154	.045	.192	<.001	.004
San Francisco	.921	.112	.405	.366	.085	<.001
Seattle	<.001	<.001	.005	.202	<.001	<.001
South Florida	.004	<.001	.001	.989	<.001	<.001
Tampa	.909	.114	.034	.120	<.001	<.001
Tucson	.189	<.001	.099	.875	<.001	<.001
Washington DC	<.001	.009	<.001	<.001	<.001	<.001

Note. See table 1 for definitions of selected MSAs. The p-values are based on likelihood ratio tests. The likelihood ratio has a chi-squared distribution with degrees of freedom equal to the number of restrictions imposed for that test. NA indicates that we did not perform this hypothesis test. All p-values greater than 0.05 are shaded.

Source. Authors' analysis of data from the CoStar Group, Inc. (www.costar.com)

Figure 1
Median Half-Yearly Sample Size Across MSAs



Source. Authors' analysis of data from the CoStar Group, Inc.

Figure 2
Illustration of Vertex Weights

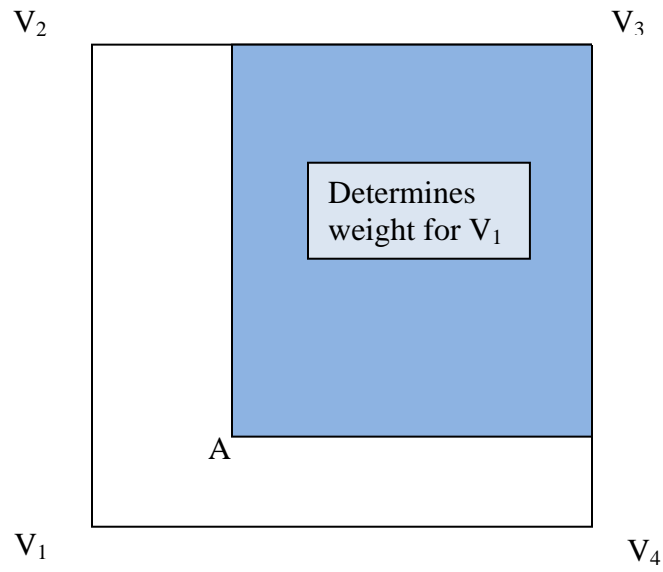


Figure 3
Locational Grid for Dallas

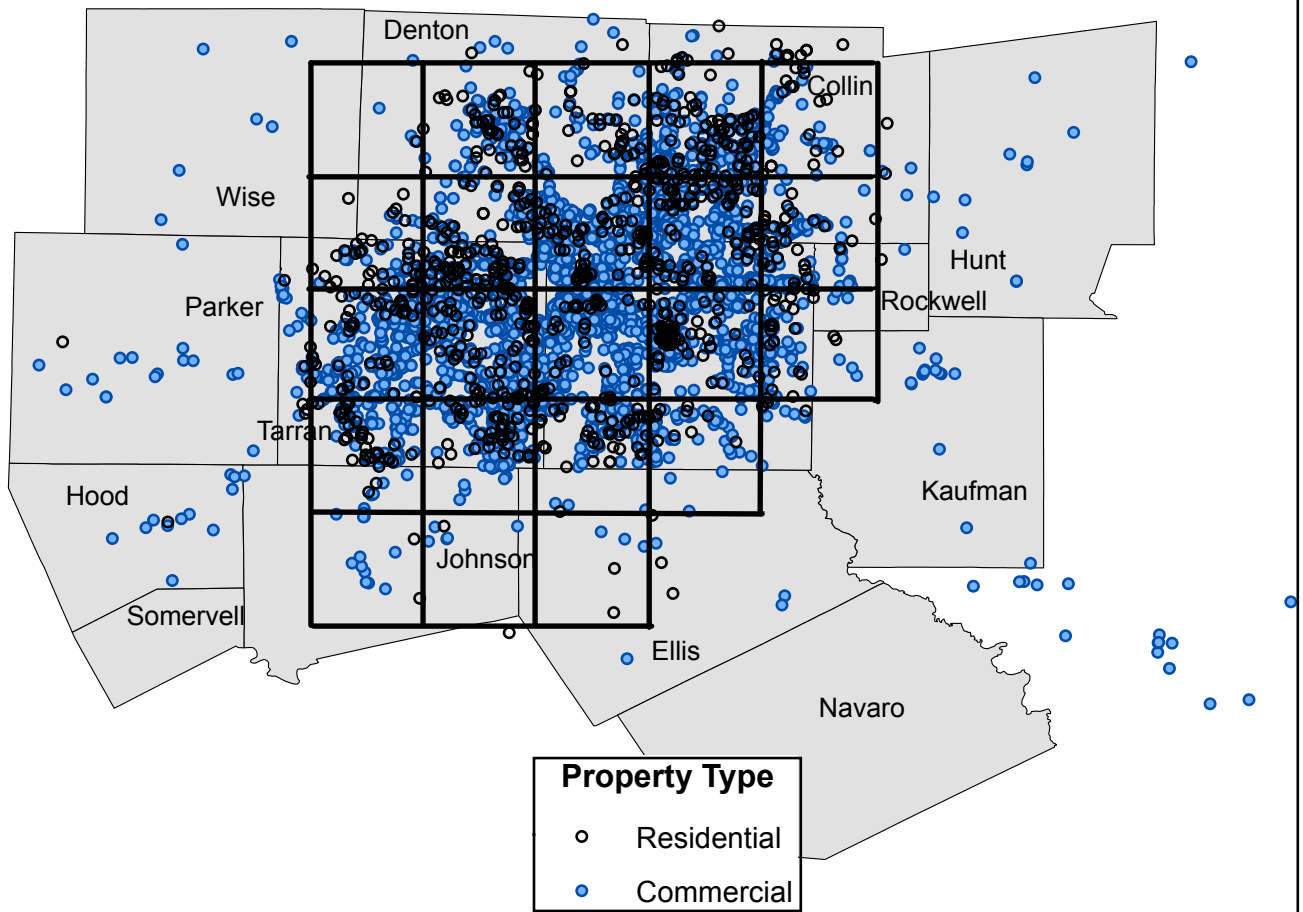


Figure 4
Locational Grid for South Florida

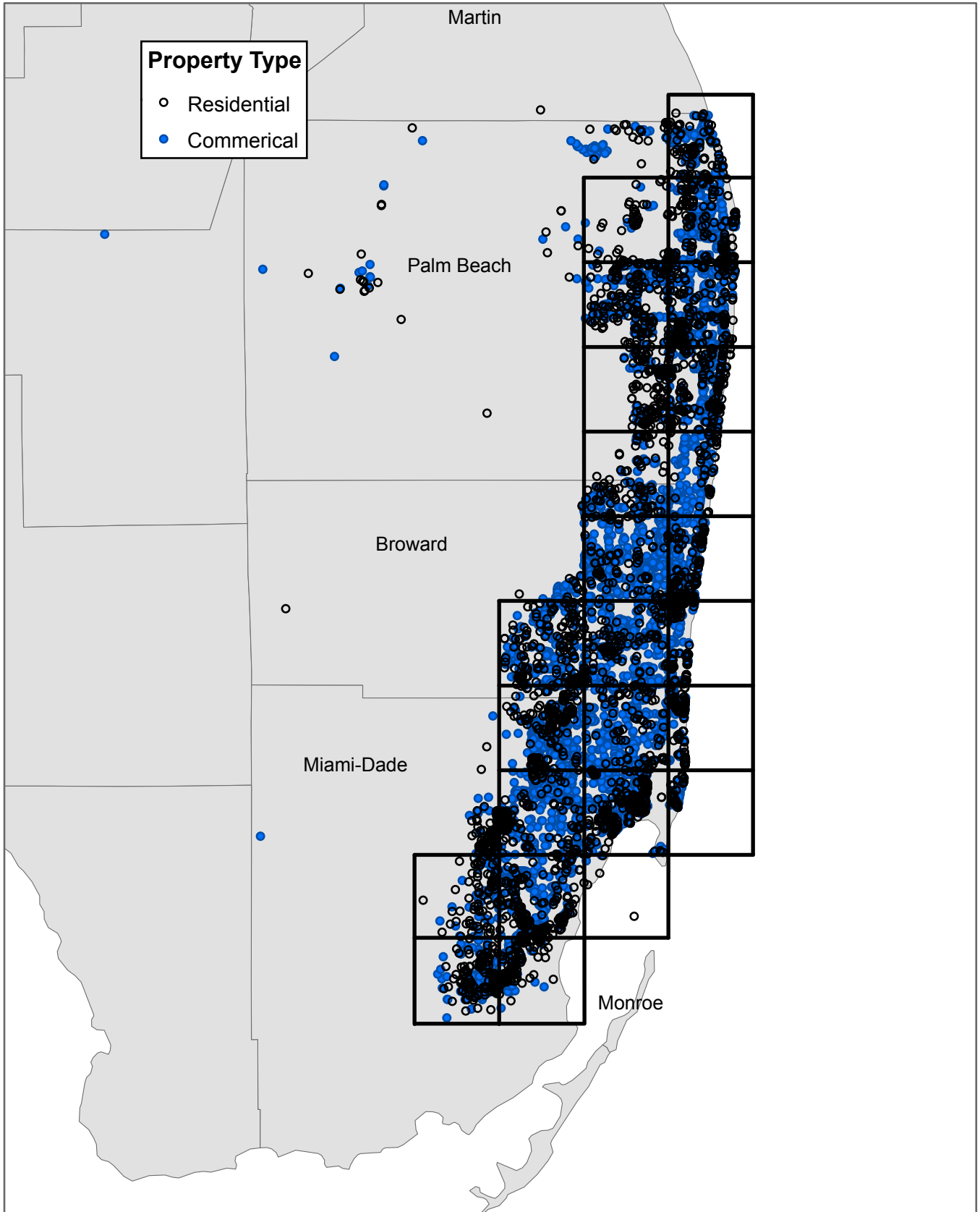
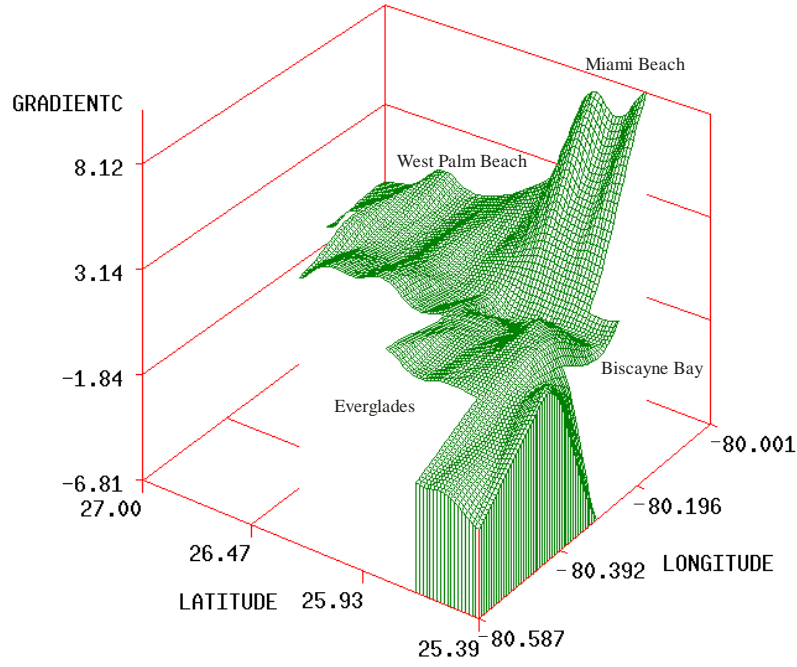


Figure 5
Land Price Surface for South Florida

Commercial



Residential

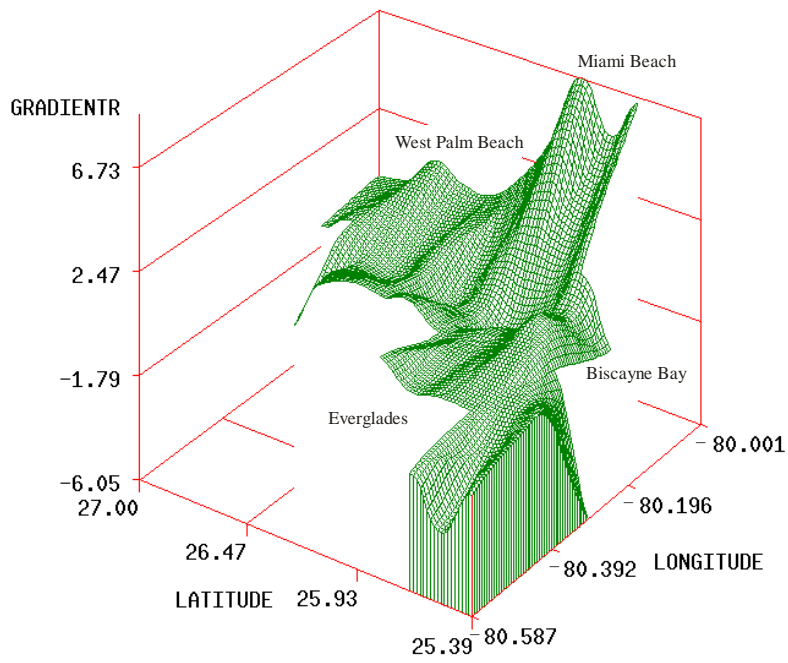
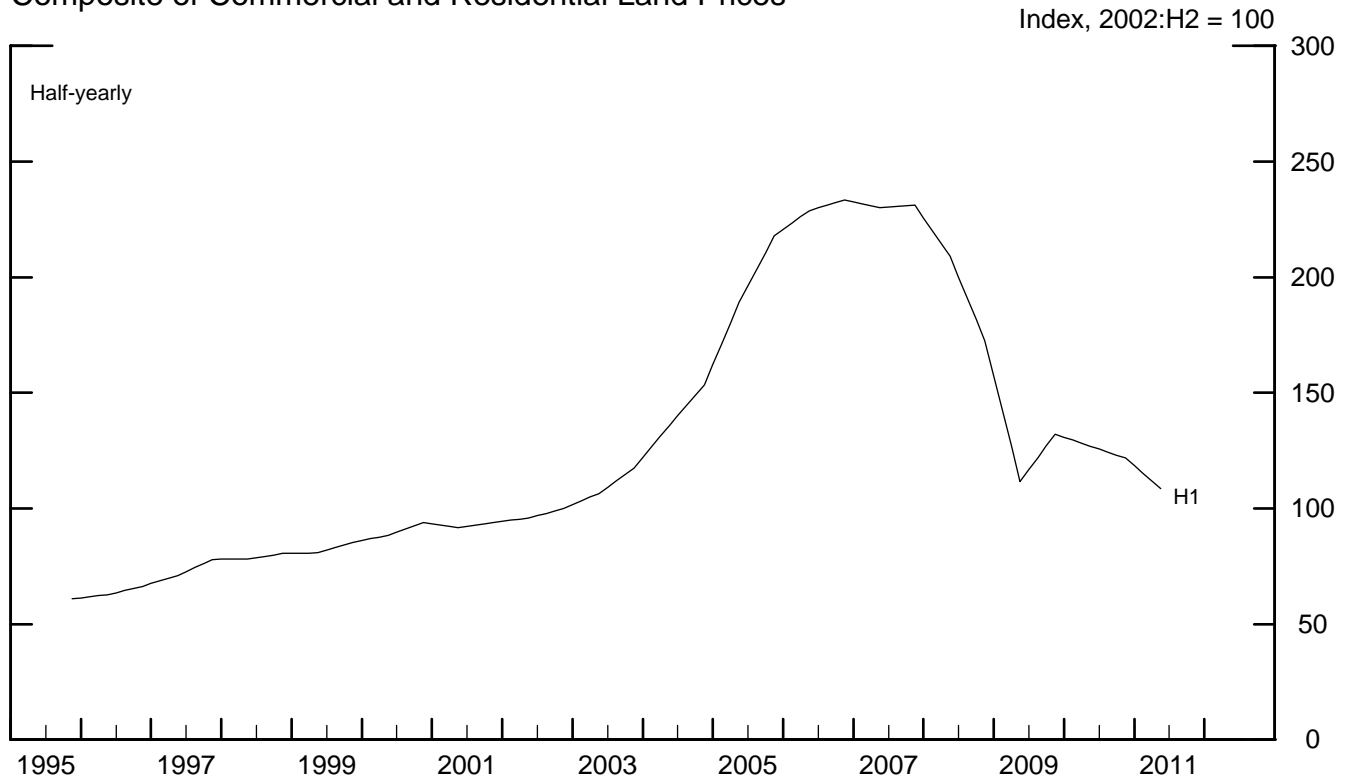
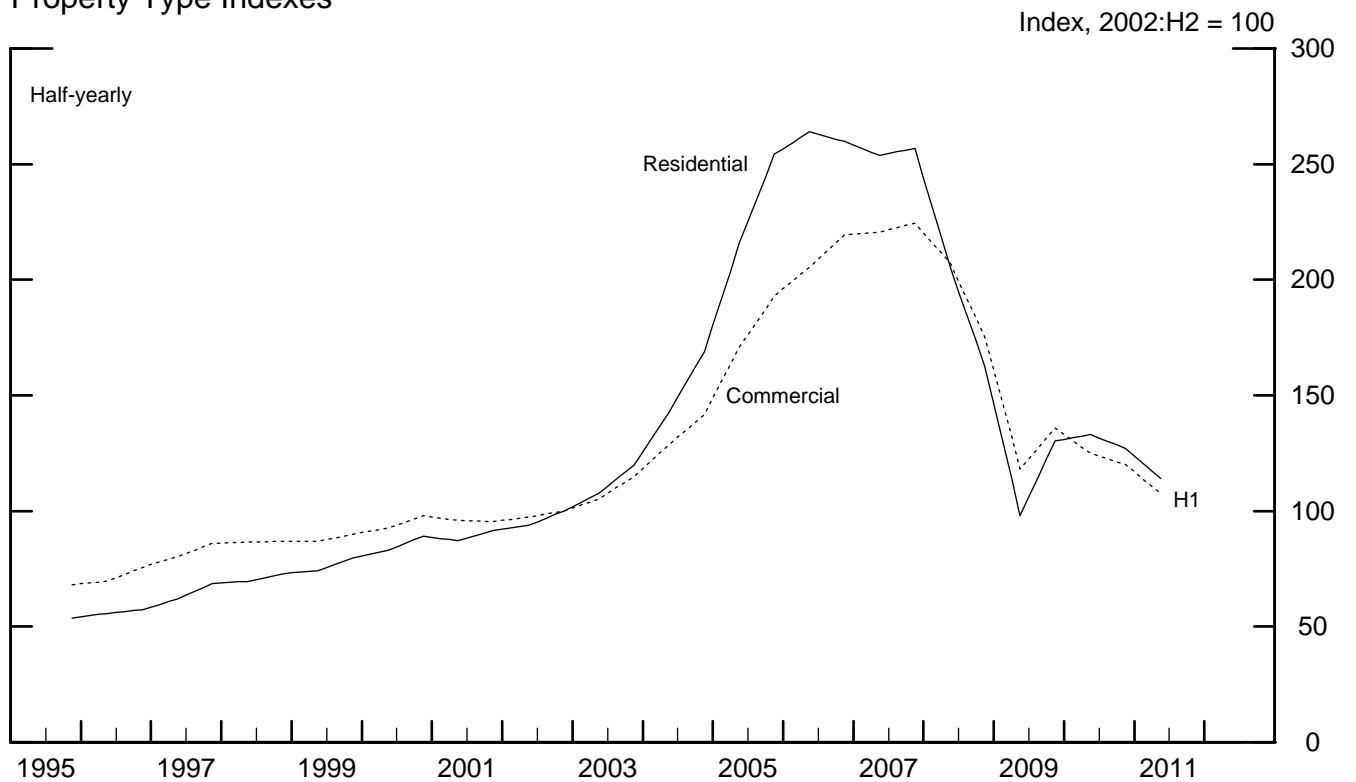


Figure 6
Aggregate Land Price Indexes

Composite of Commercial and Residential Land Prices

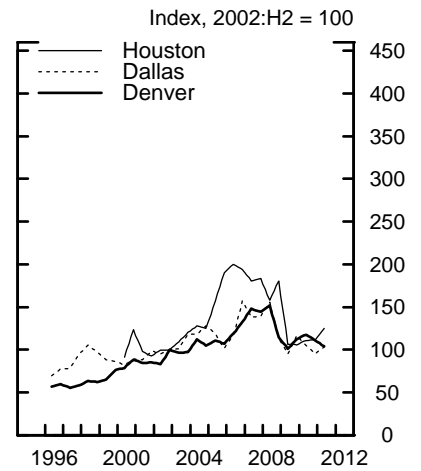
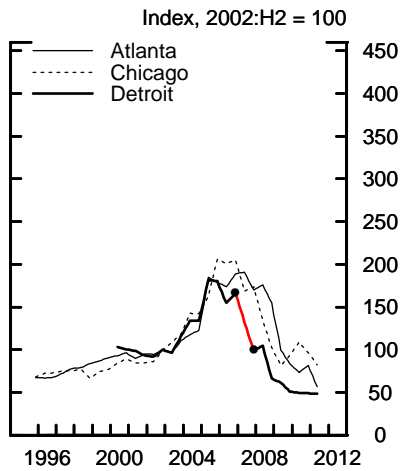
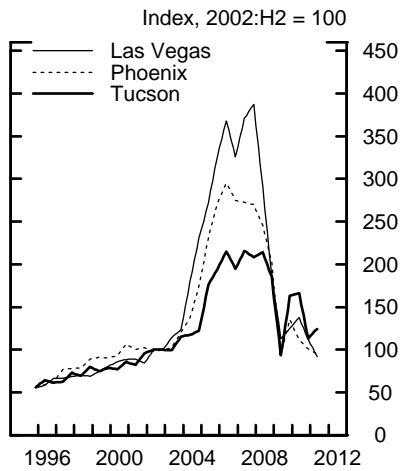
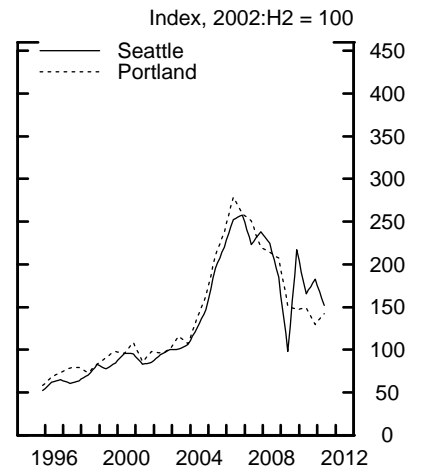
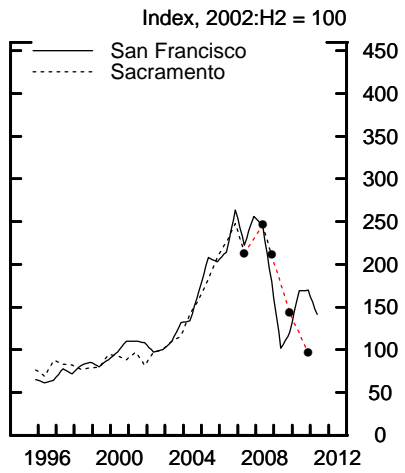
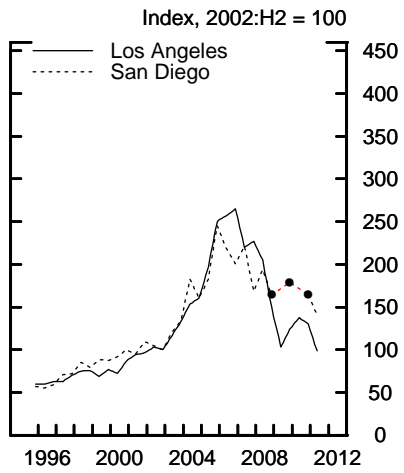
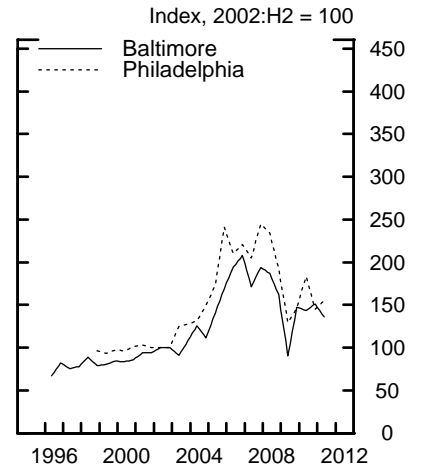
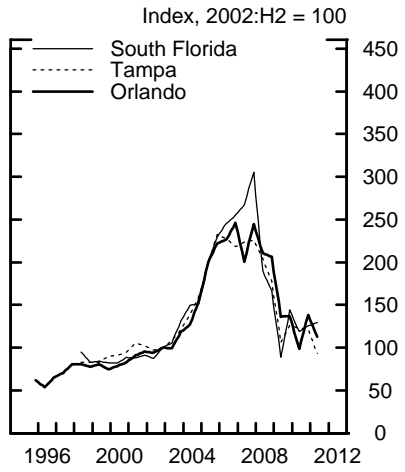
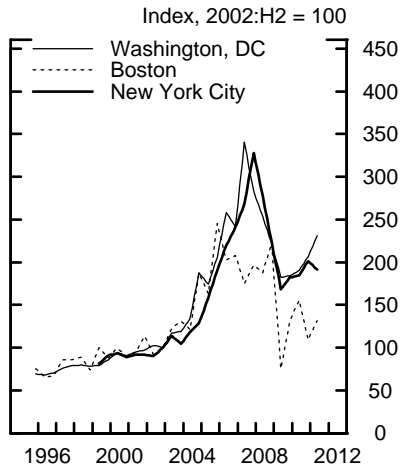


Property Type Indexes



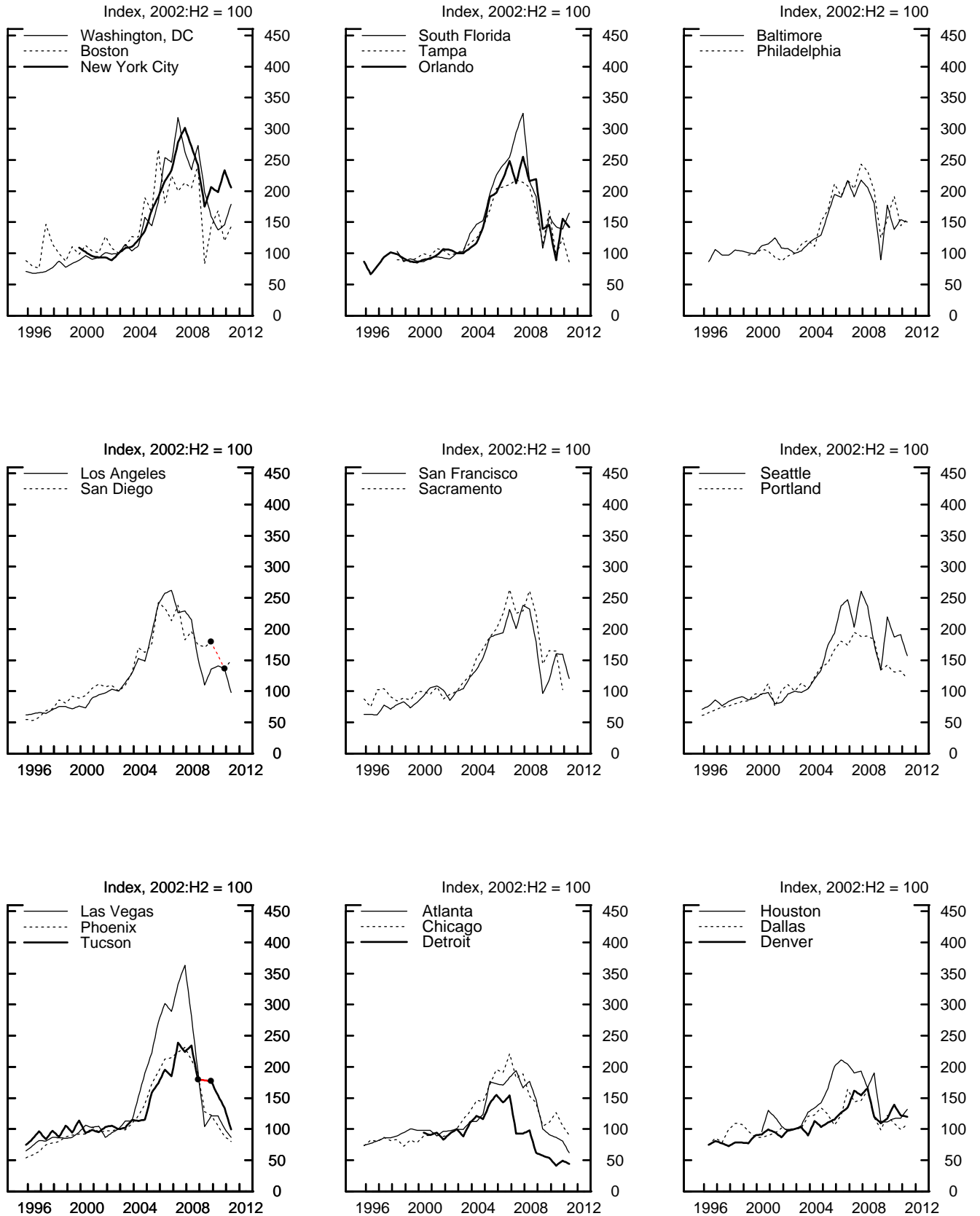
Source. Authors' analysis of data from the CoStar Group, Inc.

Figure 7
Composite Land Price Indexes by MSA



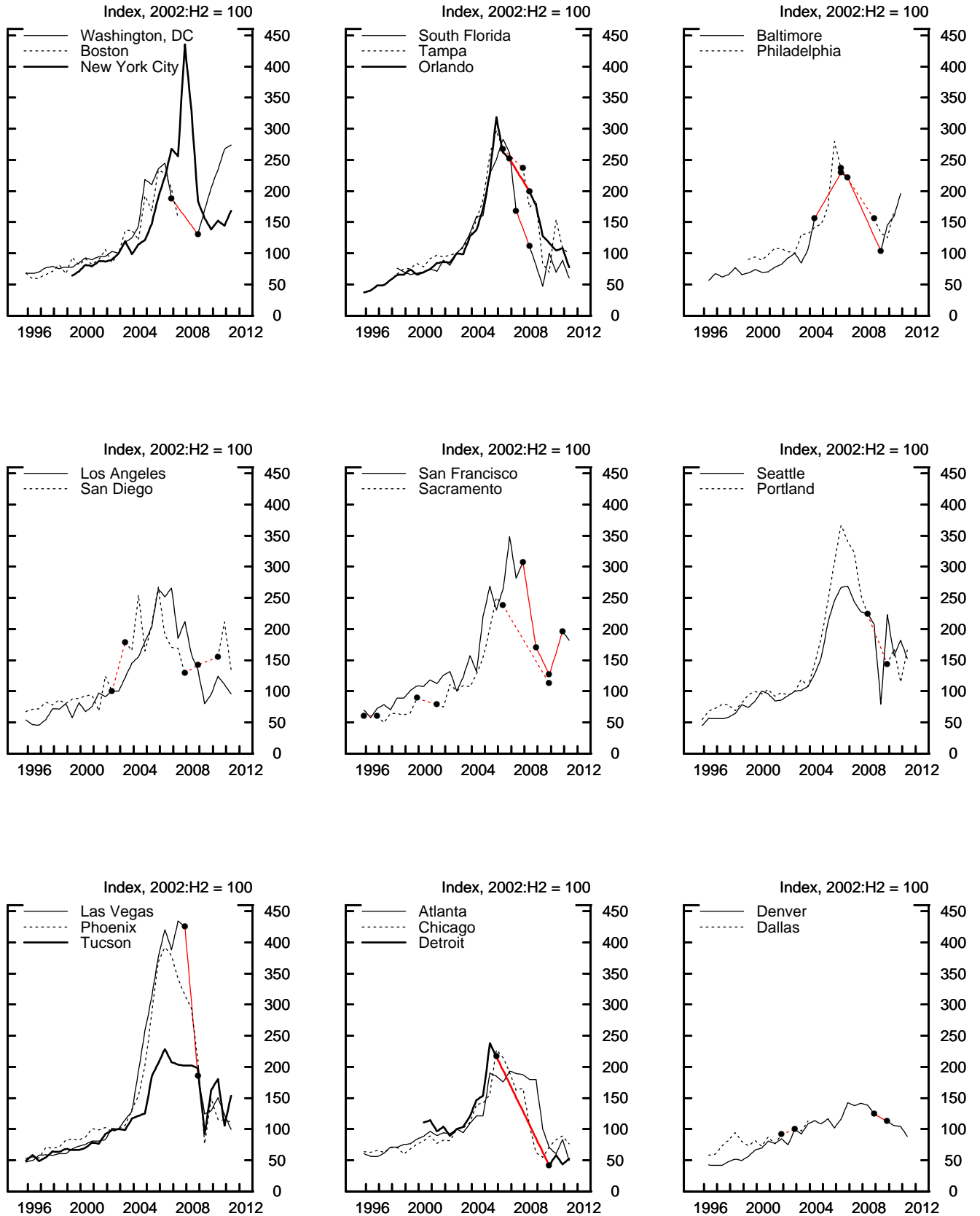
Note. Data plotted at half-year frequency. Red segments between dots represent interpolation for missing half-yearly estimates.

Figure 8
Commercial Land Price Indexes by MSA



Note. Data plotted at half-year frequency. Red segments between dots represent interpolation for missing half-yearly estimates.

Figure 9
Residential Land Price Indexes by MSA



Note. Data plotted at half-year frequency. Red segments between dots represent interpolation for missing half-yearly estimates. Houston not shown due to small number of half-yearly estimates.