Housing Bubbles and Busts: The Role of Supply Elasticity

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ABSTRACT. *Existing studies of the relationship between housing price dynamics and housing supply—which have relied upon questionable proxies for supply elasticity—have yielded highly mixed results. In this paper we provide new evidence on this relationship based on actual estimates of the price elasticity of supply for local markets in Florida. Intermarket differences in housing supply elasticity are found to explain much of the variation in housing price movements and new construction during the most recent boom-bust cycle. Additionally, we find that variance in the elasticity of housing supply can be attributed to differences in land availability and the local fiscal and regulatory environment.* (JEL R14)

I. INTRODUCTION

In recent years there has been considerable interest in the role that housing supply has played in explaining differences in housing prices and their changes across U.S. housing markets. Unfortunately, research on this issue has been limited by the nonavailability of estimates of local areas' supply price elasticities. In the absence of supply price elasticities, studies have resorted to using two proxy variables; these proxies have been used by Glaeser, Gyourko, and Saiz (2008), Huang and Tang (2012), and Davidoff (forthcoming). Cox (2011) provides a detailed critique of these proxies and finds that both have important limitations that may result in biased estimates.[1] This may help explain why studies that have used these proxies have yielded highly mixed results, with some evidence suggesting that the elasticity of supply does not affect housing market dynamics (Davidoff, forthcoming), but other evidence showing a strong relationship between the cross-sectional variation in housing supply conditions and the severity of housing price and construction cycles (Huang and Tang 2012).

The primary purpose of this paper is not to revisit the appropriateness of extant proxies for the price elasticity of housing supply, but the concerns raised by Cox and the mixed nature of the extant evidence suggest some new evidence is needed that is not dependent on the existing proxy variables for the supply price elasticity.[2] Hence, we take a completely

[1] One proxy variable is the measure of natural geographical constraints constructed by Saiz (2010). He estimates the amount of developable land within a metropolitan area, which is defined as the share of land within the area that has a slope of less than 15 degrees, after excluding land lost to bodies of water. The other proxy variable is the Wharton Residential Land Use Regulatory Index (Gyourko, Saiz, and Summers 2008), which is a measure of the regulatory restrictiveness within the metropolitan area. Cox offers two main criticisms of Saiz. First, Cox (2011, p. 14) notes that "Saiz uses an invariant 50 kilometer radius from the urban focal point of the metropolitan area to analyze geographic constraints. Given the huge differentials in the geographical sizes among the principal urban areas in the sample of the metropolitan areas over 500,000 population, the invariant 50 km radius is blunt in the extreme." Cox's second criticism pertains to how the Saiz measure treats developed land: "The second problem is that the Saiz natural constraint area does not take into consideration the area of existing development (urban area), which by virtue of it being largely occupied by buildings, also represents a geographical constraint (a development geographic constraint)" (2011, p. 18). Cox's principal criticism of the Wharton index is that "it is based partially on loose 'black box' opinions and judgments, and does not include responses from private housing industry participants" (2011, p. 16).

[2] However, later in the paper we do compare the ability of Saiz's measure of the share of undevelopable land (as an elasticity proxy) and our estimated supply price elasticities to explain housing price and quantity changes during Florida's housing boom and subsequent crash.

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different methodological approach toward investigating whether, at the local level, supply conditions affect price and quantity changes that occur in response to a demand shock. Specifically, we conduct a two-step longitudinal and cross-sectional analysis: First, we make use of a unique longitudinal database of single-family housing units and their prices to obtain estimates of the short-run supply price elasticity for each Florida county. Second, we use these estimates to conduct a cross-sectional investigation of the role that interarea differences in housing supply conditions played in explaining differences in housing price and construction changes across Florida communities during the most recent housing market boom and bust periods. The supply elasticities are found to strongly influence price and quantity changes, which provides strong confirmation of the importance of the role that housing supply plays in explaining housing market dynamics.

Having found that supply price elasticities strongly influenced price and quantity changes, we added a second equally important objective to this paper; namely, to investigate the factors that account for differences in supply price elasticities across markets. Apart from our findings, this has been recognized in the literature as an important issue, but it has received limited attention because the data necessary to undertake such studies are generally unavailable.

II. LITERATURE REVIEW

In this study we examine how differences in the supply price elasticity of housing (henceforth, supply elasticity) across markets correlate with changes in housing prices and quantities that result from positive and negative demand shocks and the factors that explain differences in the supply elasticity across markets.[3] Previous studies that address these two issues are reviewed below in separate sections that discuss the causes of variance in supply elasticity and the influence that this variance has on the housing market. There are also two types of studies in the housing literature that are related, albeit tangentially, to our current inquiry: those that estimate the supply price elasticity of housing and those that relate housing prices to land use regulations. These studies have been thoroughly reviewed elsewhere and therefore are not discussed here.[4]

Supply Elasticity and the Housing Market

There are four recent papers that study the effects that supply elasticity has on housing prices and the amount of new construction during booms and busts. The pioneering study is by Glaeser, Gyourko, and Saiz (2008). The other three studies, which build on the work of Glaeser, Gyourko, and Saiz, are by Grimes and Aitken (2010), Huang and Tang (2012), and Davidoff (forthcoming). The first two studies support Glaeser, Gyourko, and Saiz's conclusion that supply elasticity effects are important, while the last study contradicts all previous studies by concluding that these effects are unimportant.

Glaeser, Gyourko, and Saiz's inquiry is both theoretical and empirical. Their theoretical model describes housing bubbles that arise from irrational overoptimism and adaptive expectations. The model predicts what supply and demand analysis would tell us, namely, that inelastic supply leads to greater price increases and less new construction in a boom period. While a price increase in response to a demand shock may seem more damaging to social welfare than a spurt in new construction, Glaeser, Gyourko, and Saiz note that this may not be the case if excessive new construction distorts otherwise efficient migration decisions. The most novel prediction from their model is that the relationship between the supply elasticity and the size of the postbubble price correction is theoretically indeterminate. The qualitative relationship between supply elasticity and the price correction is ambigu-

[3] The source of a demand shock may either be a fundamental change in housing demand (e.g., as caused by a change in income) or a housing bubble (where overoptimistic price expectations generate greater demand), which is then followed by a bust.

[4] For a review of the estimates of housing supply elasticities, see DiPasquale (1999). For a review of the impacts that regulations have been found to have on housing prices, see Quigley and Rosenthal (2005).

ous because of two diametrically opposed forces: (1) a higher supply elasticity results in greater overbuilding during the boom, which causes greater excess inventory and larger price declines in the postboom period; and (2) a higher supply elasticity results in less of a price rise during the boom, so there is a smaller postboom price correction.

To test their model empirically, Glaeser Gyourko, and Saiz proxy for the supply elasticity using Saiz's (2010) measure of developable land within a metropolitan area. They study three distinct periods: two booms (1982–1988 and 1996–2006) and one bust (1989–1996). Price and construction models are estimated for each period, and the results from these models support the hypothesis that more inelastic areas have greater price appreciation in boom periods. Although weaker, the results also support the hypothesis that there is less construction in inelastic areas in boom periods. The results from the bust period "are not nearly as robust as those from the boom period" (Glaeser, Gyourko, and Saiz 2008, p. 208). Modest evidence suggests that there is no relationship between the supply elasticity and the fall in housing prices and that more elastic markets experienced a larger quantity response during the boom period.

Grimes and Aitken (2010) estimate a housing supply model to obtain supply elasticity and a model that registers the responsiveness of prices to a short-run excess demand disequilibrium for each of 73 regions in New Zealand, using panel data for 53 quarters. Their key result is obtained using the region as the observational unit and by regressing their price adjustment estimate on the estimated elasticity. The estimated coefficient on the latter variable is negative and statistically significant, indicating that a faster supply response moderates price spikes following a demand shock. While Grimes and Aitken do not employ the supply elasticity proxies criticized by Cox (2011), their analysis also suffers from an important measurement issue, namely, they were able to measure the existing housing stock within each region only at 5-year intervals, which forced them to obtain their quarterly estimates via linear interpolation. Obviously, the interpolation procedure introduces a substantial amount of noise into their housing data, which casts considerable doubt on their estimated elasticities.

Huang and Tang (2012) proxy the supply elasticity of 200 U.S. cities using the Saiz (2010) measure of developable land and the Wharton Residential Land Use Regulatory Index (Gyourko, Saiz, and Summers 2008). They find that both proxy variables are linked to larger booms and busts in housing prices. Huang and Tang measure changes in housing prices using data from Zillow.com. While Zillow has been shown to yield reasonably accurate estimates of the market value of individual homes, it is not clear that these Zillow estimates can be transformed into a reliable intertemporal price index. The value of the Zillow Home Value Index for a given year is the median of Zillow's value estimates for all homes in an area. The major drawback to using this index is that the Zillow value estimates are only as good as the "comparison sales" on which they are based. Because the quality of these comparable sales can vary greatly over time as market conditions change, the Zillow Home Value Index may be subject to considerable time period–dependent measurement error. More importantly, the median value of homes in an area can change as a result of changes in the stock (e.g., demolition followed by new construction can raise housing values). As the composition of the housing stock can be affected by booms and busts, Huang and Tang's results may reflect changes in either the price of housing services or the quantity of housing services, which makes it difficult to use their results to draw conclusions about the relationship between price changes and the elasticity of housing supply.

Davidoff's results are strongly inconsistent with those of the previous three studies, despite the fact that he also proxies the supply elasticity with the Saiz measure and the Wharton index. He conducts three tests of the hypothesis that the supply elasticity affected cross-sectional variation in the severity of the housing price cycles during the 2000s. He acknowledges that two of the tests are weak tests because of "troubling" assumptions.[5] The re-

[5] The first test assumes a common national demand

sults from these first two tests are primarily of interest because they are consistent with those obtained from his final test. The latter involved regressing measures of a metropolitan area's housing price cycle severity on state fixed effects and three proxies for the area's supply elasticity: (1) whether the metropolitan area is a coastal market (which implies a low supply elasticity), (2) Saiz's measure of developable land, and (3) the Wharton Residential Land Use Regulatory Index. The intent behind the inclusion of the state fixed effects is that they capture a "critical source of demand heterogeneity" (Davidoff, forthcoming, p. 16). The results indicate that variation in the supply elasticity proxies is not associated with increased cycle severity. Davidoff, therefore, rejects the hypothesis that differences in supply elasticity caused cross-sectional variation in the severity of the 2000s housing cycle among U.S. housing markets.

Determinants of Supply Elasticity

Two papers have investigated the factors that explain differences in supply elasticity across areas (Green, Malpezzi, and Mayo 2005; Saiz 2010). Although they take very different empirical approaches, both studies share a common conclusion, namely, that the elasticity is lower in metropolitan areas where land use regulations are more restrictive.

Green, Malpezzi, and Mayo (2005) first estimate supply elasticities for 45 metropolitan statistical areas (MSAs) in the United States using annual data for the years 1979–1996. To obtain their elasticity estimates, Green, Malpezzi, and Mayo regress a measure of the percentage change in the housing stock, which is based on new building permits, on the lagged first difference in natural logarithms of the Fannie Mae repeat-sales index of house prices for the MSA.[6] In the second stage of their study, the estimated elasticity (which is the estimated coefficient on the home price index from stage one) is regressed on a land use regulation index, population density, and controls. The regulation index, which the authors admit is extremely crude, is the unweighted sum of seven variables that describe the regulatory environment, which were constructed using answers to survey questions. Results from the second stage of their analysis show that elasticities are lower where land use regulation is more stringent and population density is higher, with the latter result presumably capturing the availability of developable land.

Saiz uses decennial price and quantity data at the MSA level from the U. S. Census Bureau to regress the 30-year changes (1970–2000) in the natural logarithm of housing prices on the same 30-year changes in the natural logarithm of construction costs, the natural logarithm of the housing stock, and the natural logarithm of the housing stock interacted with (1) his measure of developable land (the Saiz measure, as described above), and (2) the Wharton Residential Land Use Regulation Index. A higher value of the Wharton index is interpreted as evidence that land use regulation in the market is more restrictive. The possible endogeneity of changes in the stock and the Wharton index is addressed through the use of an instrumental variables estimator. The results from Saiz's empirical models show that the supply elasticity varies directly with the Saiz measure of land availability and inversely with the Wharton index.[7]

To summarize the literature, a common limitation of the studies of both the consequences and causes of supply elasticity is that the data used to measure many of the key variables have been suspect. As described below, the database utilized in this study allows for far more precise measurements of housing prices, the housing stock, and the local regulatory environment than those used in previous studies. In lieu of using median values,

shock and the second test assumes that demand and supply parameters have not changed since the 1980s.

[6] Green, Malpezzi, and Mayo's (2005) justification for lagging their price change variable is that this reduces the possibility of reverse causation bias. But builders obtaining permits incorporate current prices into their price expectations, so lagging may have biased their estimates by introducing measurement error.

[7] As with Huang and Tang, a possible weakness of Saiz's empirical methodology is his use of median values to measure housing price changes. Additionally, the Saiz measure pertains to the geography as it existed in 2000, and the Wharton index captures the regulatory environment in 2005. To ensure their exogeneity, these variables should have been measured at the beginning of the period (i.e., 1970).

we measure changes in housing prices by constructing a quality-controlled house price index for each market area. To measure new home construction, we actually count the number of finished homes. Prior studies have used building permits to measure additions to the housing stock; these permit-based measures may inaccurately measure new construction, particularly during the collapse of the housing market, when builders began to revise their expectations of future housing demand. To ensure their exogeneity, our land use regulation and available land variables are measured prior to the start of our price and construction sample. As precise measurements of all of these variables are required to reliably estimate supply elasticities and their effects on housing prices and quantity changes during boom and bust periods, the empirical results advanced below are also more robust and readily interpretable than what is currently found in the literature.

III. DATA

There are three parts to our study: (1) estimating county supply elasticities, (2) investigating the role these elasticities played in explaining housing price and quantity changes within counties during boom and bust periods, and (3) determining the factors that underlie differences in supply elasticity across counties. Each part of the analysis required assembling a different database. For the first part of the analysis, we assembled a 21-year panel of housing prices and stocks for 63 of Florida's 67 counties that runs from 1990 to 2010.[8] In the second part of the study, we constructed cross-county databases that included, in addition to the estimated supply elasticities and housing price and quantity changes, a set of demand-side variables matching those employed by Glaeser, Gyourko, and Saiz (2008). The third data set includes county-level variables that are hypothesized to affect the supply elasticity, such as measures of the regulatory and fiscal environment and the amount of developable land.[9] Each of these databases is described in turn below.

Panel Data Sets

The primary data sources used to construct the panels are the standardized property tax rolls that each county must submit annually to the Florida Department of Revenue (FDOR).[10] Tax rolls for the years 1995–2011 are made available to the public by FDOR.[11] The rolls contain two pieces of information that make possible our analysis of the causes and consequences of variance in supply elasticity. First, each tax roll includes a data field indicating the year in which each structure was built. This variable is used to estimate the size of the single-family housing stock and single-family new construction for each year of our panel, 1990–2010.[12] We utilize the 2011 tax roll to generate this supply series. Though the 2011 data is essentially just a cross-sectional database, we can utilize the information on the timing of development to analyze the evolution of the housing stock over time. More specifically, to estimate the stock of single-family housing in year X, we utilize the 2011 tax roll to generate a count of all of the properties found on the 2011 roll that were constructed prior to X. The amount of new construction in year X is defined as all of the properties that were constructed in year $X-1$.

[8] In 3 of Florida's 67 counties, the number of sales is too limited or the field reporting the year of construction was too unreliable for the county to be used in our analysis. In one county (Pinellas), severe errors in the construction of the digital map made the calculation of the land and regulatory variables impossible.

[9] In order to compare the explanatory power of our estimate elasticity and Saiz's proxy for elasticity in models of housing price and quantity changes, we also built a database that involved replicating Saiz's methodology for calculating the percentage of undevelopable land for each county.

[10] These tax rolls are used by FDOR to monitor the performance of the county tax assessors, who must abide by certain state statutes in assessing properties within their jurisdiction. The tax assessors, who are titled "property appraisers" in Florida, are elected to office.

[11] The year on the tax roll corresponds to the state of the housing stock in January of that year. For instance, the count of all single-family properties found on the 1996 tax roll represents the stock of single-family units as of January 1, 1996.

[12] A county-level regression of our single-family stock estimate for the year 2000 on the number reported by the 2000 census gives us an R-squared of 99.2, a slope of 0.997, and an intercept that is not significantly different from zero.

Because of demolition, some of the houses that were constructed prior to 2011 may be missing in our stock and construction counts, especially in the early years of the panel. To adjust for this removal, we inflated the pre-2010 construction and stock estimates using depreciation rates provided by DiPasquale and Wheaton (1994).[13] In the analysis below, models utilizing the depreciation-corrected data are labeled " corrected," whereas models utilizing the uncorrected data series are labeled "standard."

Properly estimating the price elasticity of supply necessitates measuring the movement in the price of housing services. To do this, we utilize the repeat-sales index approach pioneered by Bailey, Muth, and Nourse (1963). The first step in constructing the traditional repeat-sales index is the estimation of the following equation:

$$\ln(P_r) - \ln(P_s) = \sum_{t=0}^{N} D_t\beta_t + u_{r,s}, \quad [1]$$

where r denotes the period of the most recent sale, s denotes the period of the second most recent sale, P_k represents the sales price in period k, and N denotes the last period in the sample. D_t is a variable that takes on a value of 1 if the most recent sale occurred in period t, a value of -1 if the second most recent sale occurred in period t, and a value of 0 otherwise. Equation [1] can be estimated using ordinary least squares (OLS), and the estimated coefficients can be used to construct a repeat-sales index as follows:

$$RSI_t = RSI_0 \exp(\beta_t), \quad [2]$$

where RSI_t is the index value in period t, β is the estimated period-t coefficient from equation [1], and RSI_0 is the value of the index in the base period, which is traditionally set equal to 100.

One of the great strengths of the repeat-sales approach is the limited amount of data that is needed for its implementation. Because the influence of property characteristics on housing values is essentially differenced away in equation [1], the only information required to estimate a repeat-sales index is the date and transaction price of the paired sales. The FDOR rolls include all of this information, as well as an indicator of whether the sale was arm's length, which allows us to mitigate the possibility of nonmarket transactions biasing our estimates of housing price movements.

Because repeat sales thin as we go backward in time from 1995 (the year of our first tax roll), we estimate a variant of equation [1] that preserves degrees of freedom by assuming that prices change continuously, rather than discretely, over time; this assumption allows us to estimate the price index more efficiently in time periods with few sales. This more flexible sales price index, developed by McMillen and Dombrow (2001), utilizes the Fourier flexible form estimator of Gallant (1981) to estimate a repeat-sales model.[14] Specifically, we estimate

$$\ln(P_r) - \ln(P_s) = g(T_r) - g(T_s) + u_{r,s}, \quad [3]$$

where $g(T_i)$ is defined as

$$g(T_i) = \alpha_1 + \alpha_2 z_i^2 + \sum_{q=1}^{Q} (\gamma_q sin(qz_i) + \lambda_q cos(qz_i) - 1), \quad [4]$$

where z_i denotes a time variable that is transformed to lie between 0 and 2π.[15] If T_i equals 1 in the first sale month in the first year of the sample and $\max(T)$ equals the number of

[13] DiPasquale and Wheaton (1994) utilize removal rates of 0.35%, 0.3%, and 0.2% to construct their national series. Because the housing stock in Florida is quite new by national standards and the length of our panel is relatively short, we utilize the 0.2% removal rate for our corrected housing supply series.

[14] The Fourier flexible form is, in essence, a highly flexible smoothing technique that can be used to capture trends even when sales are thin within time intervals, so long as there are sales on each side of the interval.

[15] Although T_i can, in principle, vary continuously, the FDOR data contain fields that report the month and year—but not the day—of transactions. Consequently, we utilize the month of sale to rescale the time variable using equation [5].

months of sales price observations included in the sample, z_i is defined as[16]

$$z_i = \frac{2\pi T_i}{\max(T)}. \quad [5]$$

After transforming the time variables, we estimate the parameters from equation [3] using OLS, estimating separate models for each county in the sample. Following the estimation of equation [3], we utilize the estimated parameters to trace out price movements over time. To do this, for each county we first find the median sales price for single-family units in the year 2000; we then use the estimated coefficients from equation [3] to estimate the value of this representative property in January of each year between 1990 and 2010.[17] These price variables serve as our measure of the price of housing services in our housing supply models.

As noted above and described below, housing stock–adjustment models are estimated to obtain a supply elasticity for each county. These models, which are based on those estimated by DiPasquale and Wheaton (1994), require as controls the price of undeveloped residential lots and construction costs.[18] The former is obtained by estimating price indexes using the Fourier flexible form in the same manner as described above, using turnover in the residentially zoned, vacant lots that appear on the tax rolls. To estimate the physical cost of construction, we first calculate the median value of physical improvements to land for single-family properties using the replacement cost variable from the 2000 FDOR tax roll. We then utilize the construction cost series published by the RS Means Company, a commercial vendor that researchers in the past have used to investigate the cost of construction, to estimate the value of this representative structure annually from 1990 to 2010.[19] Lastly, the price, land, and construction cost series are expressed in 2010 dollars using the CPI-less-shelter index published by the Bureau of Labor Statistics.

County Cross-Sectional Database Used to Study Consequences of Supply Elasticity

In addition to boom and bust period housing price and quantity changes calculated from our panel data and our estimated elasticities, our data set includes the same demand variables used by Glaeser, Gyourko, and Saiz (2008): the percentage of college graduates, average January temperature, personal income, income growth, and precipitation.[20] They measure these variables at the MSA level, while we measure them for each Florida county.

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[16] The number of expansion terms (Q) is chosen to minimize the Bayesian information criterion statistic, where the number of expansion terms was permitted to range from 0 to 5.

[17] Although we allow for housing prices to evolve in continuous time when estimating the repeat-sales index, we must measure time in discrete units in order to estimate the housing supply equations described below because the FDOR date-of-construction field indicates only the year in which the structure was built.

[18] Another approach to obtaining estimates of the supply price elasticities has been provided by Mayer and Somerville (MS) (2000). Their approach expresses new construction as a function of changes in home prices and costs rather than as a function of levels of these variables. However, they argue that the superiority of their model over DiPasquale and Wheaton's (DW) (1994) stock-adjustment model results solely from the need of DW to construct a housing stock measure, which is likely mismeasured in noncensus years because building permits (which are not always acted upon) are used to measure additions to the housing stock. Because we actually count the number of completed homes off of our tax rolls and do not rely upon permit data, our use of the DW model is not subject to MS's criticism. While both the DW and MS models are well grounded theoretically, the estimation of the MS model is much more data intensive. The MS model includes current changes in prices and costs, along with lagged changes, as explanatory variables. Because they are endogenous, current changes must be instrumented, which puts heavy demands on MS's data that go beyond what is possible with our data.

[19] For instance, if the construction value in 2000 is \$100,000, and the RS Means index increases 10% between 2000 and 2001, our 2001 construction cost estimate will be \$110,000.

[20] To quote Glaeser, Gyourko, and Saiz (2008, p. 207), "While these variables [temperature and precipitation] do not change over time, demand for them might have, so they provide a natural way of controlling for changes in demand." Later in the same paragraph they say, "The actual level of educational achievement and income are static, of course, and should only matter if they are correlated with changes in local economies."

County Cross-Sectional Database Used to Study Determinants of Supply Elasticities

To estimate models explaining differences in supply elasticities across counties, we need data on factors that influence a developer's ability to produce more housing in response to increases in the price of housing. Such factors include the local regulatory environment, the availability of developable land, and the current fiscal situation of the county.

We hypothesize that supply elasticity is higher in counties that have less restrictive land use regulations and a larger amount of potentially developable land. Additionally, because local governments have an incentive to block developments that could strain the public coffer, we expect the elasticity of housing supply to be higher in locales where the additional tax revenue from development is more likely to offset the concomitant increase in the demand for public services than to generate a fiscal deficit.

In Florida, land use regulation is primarily the responsibility of county governments. In some other states, especially those in the Northeast and Midwest regions, land use regulation is chiefly a municipal government function. This raises the issue of how relevant our results may be to other states. Within the South, Florida's structure is not unique. County governments in the Southern states tend to have more power and service responsibility than counties in other regions. In many ways, counties in the South are more like municipalities than counties in other regions of the United States. The land use powers of Florida counties emanate from the Growth Management Act of 1985 and do not result from constitutional provisions. Hence, as in the case of cities, Florida counties regulate land on the basis of local legislation. Also of interest is that a majority of Florida's population lives within the unincorporated areas of counties.

To measure the restrictiveness of land use regulation, we make use of two sources of data. First, following Ihlanfeldt (2009), we use the county's expenditures on comprehensive planning. These expenditures are reported by the county and each city within the county annually to the Florida Department of Financial Services. These expenditure measures include spending related to land use planning and the enforcement of regulations designed to implement the plans. One of the advantages of this variable is that it responds to the concern—expressed by Fischel (1989) and Quigley and Rosenthal (2005)—that a local government's statutory land use regulations may correlate only weakly with the actual level of restrictiveness, given that planners' enforcement and interpretation of the regulations are highly discretionary.[21] What is actually spent on land use regulations may better reveal the government's impact on land markets in practice. Thus, we hypothesize that supply elasticity is lower in those counties where the local government is spending more to affect land use.

The second measure of land use regulation restrictiveness is from Mayock (2009). He employs the future land use maps that jurisdictions must submit to the Florida Department of Community Affairs in order to comply with Florida's Growth Management Act. These maps specify the number of housing units that can be constructed on the land within the jurisdiction. For each county, a measure akin to a minimum lot size requirement is calculated as the total acreage of undeveloped land that is classified as "residential land" by FDOR in the 2011 tax roll divided by the total number of housing units allowed under the future land use map.[22] Minimum lot size requirements are generally enforced and their interpretation is straightforward. Moreover, these requirements are the most frequently cited type of regulation that local jurisdictions use to exclude unwanted development (Quigley and Rosenthal 2005).

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[21] Planners' discretion is empirically examined by Cheung, Ihlanfeldt, and Mayock (2009), whose evidence suggests that developers' regulatory costs are best described as negotiated outcomes between them and planners.

[22] In the future land use maps, each parcel in the state of Florida is assigned a range of allowable housing units. To construct our minimum lot size measure, we first calculate the maximum number of units that are allowed on each parcel using the midpoint of the allowable range that is indicated on the future land use map. The minimum lot size measure is then calculated by dividing the total area of undeveloped land in each county by the maximum number of housing units allowed on undeveloped land.

Our measures of land use regulation are different from those employed in previous studies. Almost without exception, these studies have used a count variable to measure regulation, where the number of specific policies or regulations (out of a given maximum possible) is summed, with a higher total interpreted as implying greater restrictiveness. The issue, of course, with count indexes is how to weight the importance of various regulations in forming the sum. There is little basis for the weights that have been used in the existing literature.

To measure the quantity of undeveloped land within each county that is potentially developable, we start with the universe of parcels in the county and first eliminate all land on which major improvements (e.g., homes, commercial property) are located. We then eliminate all parcels that are not classified by FDOR as residential land.[23] Lastly, we eliminate all parcels that are held by public entities (e.g., national parks). The total acreage on the parcels remaining after this successive elimination process is summed to generate our aggregate measure of undeveloped land.

It should be emphasized that our definition of developable land is much more narrow than what has previously been adopted in the literature. Although all land can, in principle, be developed or redeveloped, it is much easier for developers to construct new housing on parcels that do not require any demolition work and have already been approved for residential use.[24] Moreover, if developers target residentially zoned land for new construction projects, the land use regulations on such parcels should influence the pattern and timing of development more so than regulations that fall on other types of land that are less amenable to new development. Because we seek to explain the variance in the *short-run* supply elasticity, we focus on the availability of land—and the density restrictions on such land—on which developers can build most rapidly.

We use two variables to capture the fiscal environment. The first variable is the countywide millage rate that is used to calculate ad valorem property tax bills. We hypothesize that counties with higher millage rates will have higher supply elasticities because local officials are more likely to approve development that generates tax revenue that exceeds the increase in expenditures on public services associated with the development. For instance, consider a proposed development that will be assessed at V dollars and will consume B additional dollars in public services. If the current tax rate is τ, this proposed development will generate a fiscal surplus if $\tau V > B$. This surplus would allow local officials to either increase expenditures on public services or lower the tax rate; neither of these actions would be likely to draw political backlash from incumbent citizens and NIMBY groups, and rational politicians and planners would have strong incentives to support the new development. On the other hand, if $\tau V < B$ and the new development generates a fiscal deficit, local officials will be faced with the politically unpalatable choice between raising property taxes on both new and existing residents or reducing the level of public service provision. Thus, in the case where a new development generates a fiscal deficit, local officials have strong incentives to block the new development. For a fixed level of public services, it is clear that the probability that a new development generates a fiscal surplus increases with the millage rate (τ). Because the short-run supply response to price increases is a function of the rate at which new developments

[23] Although the FDOR classification does not correspond directly to municipal-level zoning classifications, FDOR officials informed us that these parcels are generally unimproved lots that are currently zoned for residential use. The residential classification is composed of 10 subclassifications: vacant residential, single-family, mobile homes, multifamily with 10 units or more, condominiums, cooperatives, retirement homes, miscellaneous residential, multifamily with less than 10 units, and an "undefined" classification that is used at FDOR's discretion.

[24] For instance, many of the central business districts in the counties in our sample have very high allowable densities, and we expect a higher supply elasticity in areas that have less stringent land use regulations. As would be expected, however, very large proportions of the land in the urban core have already been developed. It is thus highly unlikely that the land use regulations that apply to such parcels would influence the supply of new single-family housing, because profit-maximizing developers have a strong incentive to eschew the costly demolition of high-value structures and to target development toward parcels without major improvements.

can be approved, we expect the short-run supply elasticity to be higher in locations with higher millage rates.

Another variable that influences the fiscal impact of new development is the value of the county's existing property. It is straightforward to show that if the marginal cost of providing public services is constant and the average assessed value of the new development is lower than that of existing property, the millage rate must rise in order to provide the same level of public services per household. For reasons described above, politically astute local officials will be inclined to prohibit or delay development projects that they believe will reduce the average value of property in the jurisdiction. This screening out of low-valued projects results in an inverse relationship between the supply elasticity of housing and the value of existing property. To test this hypothesis, we include among the regressors the average estimated market value of existing single-family structures. We focus on the value of single-family units as opposed to the value of all real estate because the demand for public services, which is the primary determinant of the millage rate, is driven by the existing residents of the community, many of whom own single-family properties.

Because the regulatory and local fiscal variables will generally change as new development occurs, the inclusion of such variables that are measured in the same time period that is used to calculate the elasticity measures could introduce simultaneity bias in the regressions that investigate the determinants of elasticity. To avoid such bias, our available land measure, minimum lot size measure, and fiscal variables are calculated in 1990, the year the panel used to calculate the elasticity of supply starts.[25] Because the municipal expenditure data is unavailable before 1993, our measure of comprehensive planning expenditures is measured in 1993.

[25] Because the FDOR tax rolls are not available in digital format before 1995, to calculate the average value of single-family housing in 1990, we first calculate the average value of single-family housing using the 1995 tax roll. We then use the single-family repeat-sales index to back forecast this mean to its 1990 value.

Constructing the Saiz Proxy for Florida Counties

To construct the Saiz measure at the county level in Florida, we first used GIS technology to overlay a 60 m by 60 m grid on each county in the state.[26] Each square on the grid is assigned a slope value using the 2011 U.S. Geological Survey (USGS) Digital Elevation Model map and a land use code from the 2001 USGS National Land Cover Dataset.[27] A grid point is classified as undevelopable if its slope is greater than 15 degrees or the land use code indicates that the grid point is water, woody wetlands, or emergent herbaceous wetlands.[28] Following the classification of each grid point, we define the percentage of undevelopable land as the area of all undevelopable grid points in the county divided by the area of all grid points (developable and undevelopable) in the county.

IV. METHODOLOGY AND RESULTS

To estimate a supply elasticity for each county, housing supply stock–adjustment models similar to that of DiPasquale and Wheaton (1994) are estimated. These models, along with the results from their estimation, are described below. We then describe the models and results that comprise our analysis relating boom and bust changes in housing prices and quantities to measures of supply elasticity. Lastly, we explore factors accounting for intercounty differences in elasticity and estimate models in which the Saiz proxy for the supply elasticity is used in lieu of our elasticity estimate.

[26] The main difference between our construction of the developable land proxy and that of Saiz is one of granularity; whereas we utilize an overlay where each piece of the grid is a 60 m by 60 m square, Saiz (2010, p. 1254) utilizes a grid of 90 m by 90 m squares. Hence, our sampling is at a finer level than that of Saiz.

[27] The raw land cover and elevation data can be downloaded using the USGS Seamless Data Warehouse, which is available at http://seamless.usgs.gov/.

[28] The water, woody wetlands, and emergent herbaceous wetlands classifications correspond to land use codes 11, 90, and 95, respectively, on the 2006 USGS National Land Cover Dataset.

Housing Supply Stock–Adjustment Models

To estimate a supply elasticity for each county, the following stock-adjustment model was estimated:[29]

$$C_t = \alpha_1 + \alpha_2 P_t + \alpha_3 I_t + \alpha_4 LAND_t + \alpha_5 COST_t - \alpha_6 S_{t-1} + \varepsilon_t, \quad [6]$$

where C_t is the number of single-family homes completed in year t, P_t is the housing price in year t, I_t is the average construction financing interest rate in year t, $LAND_t$ is the land price in year t, $COST_t$ is the contruction cost in year t, S_t is the stock of single-family housing in year t, and ε_t is the error term. Following DiPasquale and Wheaton (1994), our method of estimation was OLS. It should be emphasized that because of the durable nature of housing, in equilibrium it is the price of housing and the *total stock of housing*—not the level of construction—that are simultaneously determined; OLS estimates of equation [6] should thus not be plagued by simultaneity bias.[30]

The short-run elasticity is equal to

$$\varepsilon_{SR} = \alpha_2 = \frac{\overline{P}_t}{\overline{C}_t}, \quad [7]$$

where $\overline{P}_t$ and $\overline{C}_t$ are the average annual housing price and average annual level of construction over the course of the panel.[31,32]

Summary measures for the estimated supply elasticities, along with summary measures for the variables used in the rest of our analysis, are reported in Table 1. The average short-run price elasticity in our sample was approximately 2, with a maximum of approximately 8.[33] A comparison of the corrected and uncorrected elasticity terms suggests that, at least on average, the estimated short-run elasticities are relatively insensitive to whether the removal-rate correction is utilized in the construction of our housing supply series. Regarding price movements, the average county in our sample experienced real single-family price appreciation of approximately 89% between 2000 and 2006; this dramatic

[29] Our estimated supply models are an improvement over those estimated by DiPasquale and Wheaton in that we are able to better measure a number of the variables: (1) their dependent variable is permits, while ours is actual homes built; (2) they use the average price of farmland as their land cost measure, while we use a quality-controlled residential land price index; and finally (3) their stock measure is estimated by updating census data with permits, while ours is the actual number of single-family homes on the tax roll.

[30] Topel and Rosen (1988) and Mayer and Somerville (2000) estimate housing construction models in which the housing price term is treated as an endogenous variable. Neither of these studies, however, provide estimates from a benchmark OLS model or perform any statistical tests that provide evidence that construction and housing prices are simultaneously determined. Furthermore, Topel and Rosen argue that any such bias should be minimal: "Endogeneity [in a housing construction model] is unlikely to be a serious problem because investment is such a small fraction of existing stock" (1988, p. 728). Because of the well-known poor performance of instrumental variables models in small samples and the absence of empirical evidence indicating that simultaneity bias is a problem, we utilize OLS to estimate equation [6]. Another possible concern raised by equation [6] is that falling prices may cause construction to fall to zero, in which case our dependent variable would be censored, possibly biasing our estimator. However, while construction did substantially fall in many counties as prices plummeted, in no case did new construction totally cease.

[31] The long-run supply elasticity, which measures the responsiveness to changes in the total stock of housing to price changes, is equal to $\varepsilon_{LR} = (\alpha_2/\alpha_6)/(\overline{P}_t/\overline{C}_t)$. As the focus of this paper is on construction activity and short-run price changes, our analysis focuses on the short-run elasticity of supply.

[32] Previous work (e.g., Rosenthal 1999; Mayer and Somerville 2000) has found that housing starts are stationary, whereas a number of the independent variables in our stock-adjustment equation are nonstationary. Should this be the case, equation [6] could possibly be an unbalanced regression "in which the regressand is not of the same order of integration as the regressors, or any linear combination of the regressors" (Banerjee et al. 1993, p. 164). In such unbalanced regressions, standard distributions cannot be used to conduct inference. Researchers wishing to conduct inference on unbalanced regressions have generally resorted to one of two approaches that allow for the use of standard distribution theory: transforming the independent variables so that they are integrated of the same order as the regressand or constructing a set of cointegrated regressors. Mayer and Somerville (2000) adopt the former approach, whereas Grimes and Aitken (2010) adopt the latter. As we make no attempt to conduct inference on the parameters of the stock-adjustment equation, we do not adopt either of those approaches here.

[33] An appendix is available upon request that contains the estimated price coefficient, its estimated standard error, and the implied supply price elasticity obtained from estimating the stock-adjustment model for each of the 64 counties.

TABLE 1
Summary Statistics

Variable	Description	Mean
Bubble appreciation	Real price appreciation: 2000–2006	89.00 (32.66)
Crash appreciation	Real price appreciation: 2007–2010	− 36.94 (18.12)
Short-run elasticity	Short-run price elasticity of supply	2.07 (2.02)
Corrected short-run elasticity	Depreciation-corrected short-run price elasticity of supply	2.03 (1.98)
Bubble construction	Single-family construction: 2000–2006	14,797.62 (17,753.75)
Corrected bubble construction	Depreciation-corrected single-family construction: 2000–2006	14,994.59 (17,991.13)
Crash construction	Single-family construction: 2007–2010	3,209.00 (3,666.87)
Corrected crash construction	Depreciation-corrected single-family construction: 2007–2010	3,222.21 (3681.89)
Total housing units	Total housing units in 2000	108,083.00 (167,058.30)
Land area	Land area in square miles	825.45 (386.41)
% College graduates	Percentage of residents with a college degree	17.08 (8.04)
Income per capita	Real income per capita in 2000	24,355.79 (7,132.74)
Income growth	Real income growth between 2000 and 2006	3,240.60 (2,708.66)
Rainfall	Average annual rainfall	53.79 (5.24)
January temperature	Mean January temperature	58.16 (4.54)
Planning expenditures	Comprehensive planning expenditures in 1993	1,913,983.00 (3,368,927.00)
Minimum lot size	Minimum lot size on undeveloped residential land in 1990	0.89 (0.69)
Undeveloped land	Acres of undeveloped residential land in 1990	42,788.55 (27,948.03)
Millage	Countywide millage rate in 1990	16.89 (2.41)
Average housing value	Average value of single-family housing in 1990	58,435.30 (28,818.90)

Note: Standard deviations in parentheses. Observations = 63.

price swing was accompanied by, on average, about 15,000 new single-family homes (an annualized rate of approximately 2,100 units). After reaching a peak around 2007, housing prices in Florida began a precipitous decline, with a mean rate of appreciation between 2007 and 2010 of −37%. As expected, the collapse of housing prices was met with a concomitant reduction in construction activity, with the mean annualized rate of new single-family construction between 2007 and 2010 falling to just above 800.

The comovement of single-family housing prices and construction and land costs for six of the most populous counties in Florida is displayed in Figure 1.[34] These plots, which are constructed using the cost and price series described above, reveal that, prior to 2000 (the year that we use to mark the start of the housing bubble, which is delineated with a vertical bar), the price of single-family housing roughly tracked movements in the physical costs of construction. From 2000 through 2006, however, there was an incredibly sharp divergence in housing values and construction costs. Movements in land prices, which are also susceptible to speculation, followed housing price movements closely over this time period.

Starting around the year 2007 (the year that we define as the end of the bubble, which is marked with a vertical bar), housing prices and land prices began to fall sharply. Con-

[34] Similar graphs for all of the other counties in our sample are available upon request. Pinellas County is actually the fifth-largest county in Florida by population. Because of some serious problems in the way that the digital parcel map was formatted in Pinellas, data from this county could not be used for the sample.

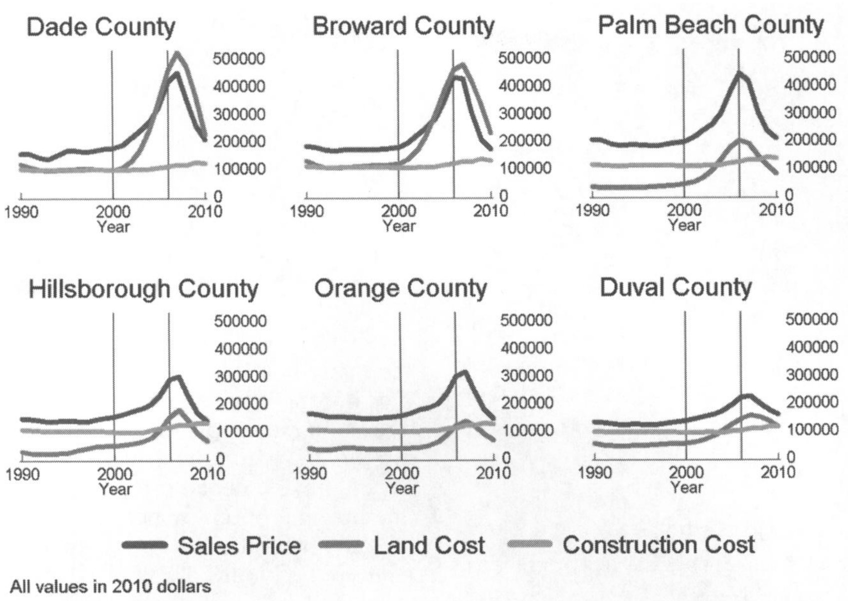

FIGURE 1
Home Values, Land Costs, and Construction Costs in Florida, 1990–2010

struction costs, on the other hand, continued the slow increase that they had experienced throughout the course of the last two decades.

Comparing the counties in Figure 1, we see that although the bubble began and ended around the same time in each of the six counties, the magnitudes of the price swings during these periods were not the same throughout the state. The three counties in southern Florida (Dade, Broward, and Palm Beach) saw much more dramatic price increases than did the counties in central Florida (Hillsborough and Orange) and north Florida (Duval), and the postbubble corrections have subsequently been more severe in southern Florida. The housing bubble thus clearly manifested itself in different forms in different housing markets, and one of the goals of this study is to investigate the role of supply elasticity in generating these differences. We turn now to this investigation.

The Relationship between Supply Elasticity and Changes in Housing Prices and Quantities

Like Glaeser, Gyourko, and Saiz (2008), to explore the effect that supply elasticity has on changes in housing prices during boom and bust periods, we begin by estimating simple bivariate cross-sectional regression equations. The dependent variables are the county change in (1) housing prices between 2000 and 2006, and (2) housing prices between 2007 and 2010. The single explanatory variable is the short-run estimated supply elasticity.[35] The variation that is used to estimate these models is displayed in Figure 2, which reveals that there was a large amount of variance in price appreciation over this period throughout the state. The most dramatic appreciation during these years was concentrated in the southern counties, with price growth particularly robust along the coast. In terms of housing supply, counties in southern Florida are also characterized by low supply elasticity relative to many counties in the central and northern parts of the state. The clustering of high appreciation rates in areas of low supply elasticity shown in Figure 2 and

[35] Our bivariate models are analogous to those estimated by Glaeser, Gyourko, and Saiz (2008), except that they measured price and quantity changes at the MSA level and, in lieu of using the supply elasticity, they use Saiz's (2010) proxy for supply elasticity.

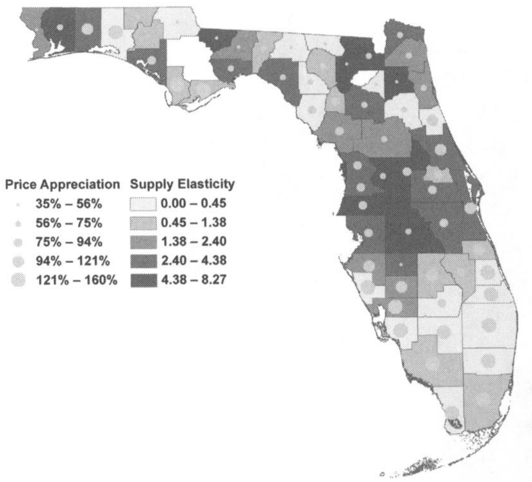

FIGURE 2
Price Appreciation and Supply Elasticity in Florida, 2000–2006

the bivariate regression results reported in Column 1 and Column 3 of Table 2 (showing the elasticity impact to be negative and statistically significant) bear out the theoretical predictions of Glaeser et al.[36]

As noted by Glaeser, Gyourko, and Saiz (2008), the biggest challenge to any implication that might be drawn from simple bivariate models is that they reflect demand-side, rather than supply-side, differences across markets. We therefore follow Glaeser, Gyourko, and Saiz by adding to our estimated models their different measures of housing demand as controls: mean January temperature, mean annual precipitation, share of population with college degrees in 2000, the level of real income in 2000 as measured by the Bureau of Economic Analysis, and the growth in that income between 1990 and 2000. An inspection of Column 2 and Column 4 of Table 2 provides strong confirmation that counties with higher short-run supply elasticities experienced substantially less price appreciation during the 2000–2006 period, even when controlling for demand-side shocks. The coefficients on the estimated elasticity terms in Table 2 suggest that a one-standard deviation increase in the short-run supply elasticity—an increase of about 2—leads to a decrease in real appreciation over this period of between 7% and 12%. The impact of supply elasticity on appreciation is far more pronounced for more dramatic changes in the supply environment. For instance, our estimates suggest that an increase in the supply elasticity from 0 to 8 (approximately the maximum in the sample) would have reduced price appreciation during the housing boom by between 29% and 48%.

As noted in Section 2, in the model of Glaeser , Gyourko, and Saiz, the qualitative relationship between supply elasticity and the postbubble price correction is ambiguous. The opposing effects that Glaeser, Gyourko, and Saiz identify appear to be off-setting in our sample. In Figure 3, which depicts the supply elasticity and evolution of housing prices between 2007 and 2010, there appears to be no relationship between supply elasticity and the postbubble price corrections, as some high-elasticity markets experienced price declines that exceeded the price correction in markets with much less elastic housing supply. Bivariate and multivariate regressions (reported in Table 3) in which price appreciation between 2007 and 2010 is the dependent variable also fail to reveal any meaningful statistical relationship between post-2006 housing price dynamics and housing supply conditions: the sign on the short-run elasticity coefficient varies across specifications in Table 3, and the coefficients are never statistically significant.

To explore the effect that supply elasticity has on changes in housing output during booms and busts, we again follow Glaeser, Gyourko, and Saiz (2008). The dependent variables are new construction measured over the periods (1) 2000–2006 and (2) 2007–2010. The simple model includes as explanatory variables the supply elasticity and the same controls as Glaeser, Gyourko, and Saiz included: the natural logarithm of land area and the natural logarithm of the housing stock

[36] There may appear to be an endogeneity bias in these results. Changes in prices are explained by estimated elasticities, which are obtained by relating new construction to changes in price levels. Indeed, if both the elasticity and price (quantity) change models were estimated using longitudinal county-level data, it would be hard to argue that our elasticity estimates are exogenous to the boom and bust price and quantity changes we seek to explain. However, our price and quantity change models are cross-sectional, which allays possible endogeneity concerns.

TABLE 2
Bubble Appreciation Regressions, Dependent Variable: Price Appreciation: 2000–2006

	Housing Supply Measure			
Variable	Standard (1)	Standard (2)	Corrected (3)	Corrected (4)
Short-run elasticity	− 5.897*** (1.652)	− 3.593*** (1.332)	− 6.010*** (1.684)	− 3.645*** (1.353)
% College graduates		− 0.1577 (0.3795)		− 0.1621 (0.3798)
Income per capita		0.000205 (0.000612)		0.000209 (0.000612)
Income growth		0.00360** (0.00140)		0.00360** (0.00140)
Rainfall		1.395** (0.578)		1.392** (0.578)
January temperature		4.040*** (0.695)		4.037*** (0.696)
Constant	101.2*** (6.566)	− 227.5*** (47.61)	101.2*** (6.570)	− 227.3*** (47.76)
Observations	63	63	63	63
R-squared	0.134	0.615	0.133	0.614

Note: Robust standard errors in parentheses.
** $p < 0.05$; *** $p < 0.01$.

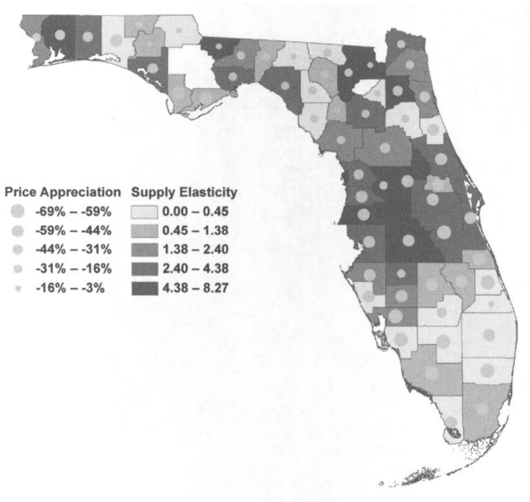

FIGURE 3
Price Appreciation and Supply Elasticity in Florida, 2007–2010

in 2000. The second specification of the model adds the demand variables. Results are reported in Table 4 and Table 5 for the bubble and crash regressions, respectively.

The regression results in Table 4 reveal that counties in which the price elasticity of supply is higher experienced a much more substantial level of building activity during the housing bubble than counties with lower price elasticities. Again, the regression results are robust to the inclusion of the demand variables. The results from estimating the crash construction regressions are presented in Table 5. These results show that there was a larger quantity response in counties with higher short-run supply elasticities. Once again, the results are highly insensitive to whether the models include the demand variables. All of these housing output results are consistent with the theoretical predictions of Glaeser, Gyourko, and Saiz (2008).

Explaining Differences in Supply Elasticities across Counties

The above results showing the important role played by the supply elasticity in explaining housing price and quantity changes across counties beg the question: What explains cross-county differences in the supply elasticity? We have hypothesized that the ability of the construction industry to respond to an increase in housing prices is hampered by land use regulations, a scarcity of developable land, and fiscal conditions that render local officials reluctant to grant approval for development projects. We therefore estimated models where the dependent variable is the county supply elasticity and the independent variables capture the aforementioned factors.

We have two alternative measures of land use regulation: the minimum lot size on vacant residential land in 1990 and comprehensive planning expenditures in 1993. Our measure of developable land is the amount of land, measured in acres, that had not been developed by 1990 and was classified by FDOR

TABLE 3
Crash Appreciation Regressions, Dependent Variable: Price Appreciation: 2007–2010

Variable	Housing Supply Measure Standard (1)	Standard (2)	Corrected (3)	Corrected (4)
Short-run elasticity	0.733 (1.082)	− 0.136 (0.896)	0.744 (1.102)	− 0.153 (0.913)
% College graduates		− 0.3741 (0.3791)		− 0.3737 (0.3790)
Income per capita		0.000110 (0.000603)		0.000110 (0.000604)
Income growth		0.000565 (0.000832)		0.000564 (0.000834)
Rainfall		− 0.466 (0.348)		− 0.467 (0.348)
January temperature		− 2.855*** (0.492)		− 2.856*** (0.492)
Constant	− 38.46*** (3.693)	156.3*** (32.92)	− 38.45*** (3.696)	156.5*** (32.91)
Observations	63	63	63	63
R-squared	0.007	0.473	0.007	0.473

Note: Robust standard errors in parentheses.
*** $p<0.01$.

TABLE 4
Bubble Construction Regressions, Dependent Variable: ln(New Construction: 2000–2006)

Variable	Housing Supply Measure Standard (1)	Standard (2)	Corrected (3)	Corrected (4)
Short-run elasticity	0.131*** (0.0471)	0.153*** (0.0561)	0.135*** (0.0475)	0.158*** (0.0564)
% College graduates		0.0075 (0.020)		0.00771 (0.0202)
Income per capita		2.32e–05 (2.59e–05)		2.31e–05 (2.57e–05)
Income growth		4.07e–06 (5.43e–05)		4.22e–06 (5.38e–05)
Rainfall		− 0.00792 (0.0207)		− 0.00759 (0.0206)
January temperature		0.00477 (0.0318)		0.00526 (0.0316)
ln(Total housing units)	1.261*** (0.0611)	1.132*** (0.110)	1.262*** (0.0608)	1.131*** (0.109)
ln(Land area)	− 0.498* (0.259)	− 0.417 (0.265)	− 0.499* (0.259)	− 0.418 (0.265)
Constant	− 1.970 (1.622)	− 1.735 (2.216)	− 1.959 (1.621)	− 1.755 (2.207)
Observations	63	63	63	63
R-squared	0.873	0.881	0.873	0.882

Note: Robust standard errors in parentheses.
* $p<0.10$; *** $p<0.01$.

TABLE 5
Crash Construction Regressions, Dependent Variable: ln(New Construction: 2007–2010)

Variable	Housing Supply Measure Standard (1)	Standard (2)	Corrected (3)	Corrected (4)
Short-run elasticity	0.154*** (0.0466)	0.165*** (0.0491)	0.159*** (0.0473)	0.170*** (0.0496)
% College graduates		0.00965 (0.0174)		0.00983 (0.01727)
Income per capita		− 2.31e–06 (2.63e–05)		− 2.42e–06 (2.61e–05)
Income growth		3.70e–05 (5.10e–05)		3.71e–05 (5.05e–05)
Rainfall		− 0.0180 (0.0158)		− 0.0177 (0.0157)
January temperature		− 0.000369 (0.0251)		9.72e–05 (0.0250)
ln(Total housing units)	1.009*** (0.0523)	0.952*** (0.0891)	1.009*** (0.0521)	0.952*** (0.0888)
ln(Land area)	− 0.360* (0.201)	− 0.324 (0.205)	− 0.361* (0.201)	− 0.325 (0.204)
Constant	− 1.479 (1.282)	− 0.383 (1.794)	− 1.477 (1.281)	− 0.408 (1.788)
Observations	63	63	63	63
R-squared	0.876	0.886	0.876	0.887

Note: Robust standard errors in parentheses.
* $p<0.10$; *** $p<0.01$.

TABLE 6
Elasticity Regressions, Dependent Variable: Short-Run Elasticity

	Housing Supply Measure	
Variable	Standard (1)	Corrected (2)
Minimum lot size	− 0.898*** (0.297)	− 0.878*** (0.290)
Undeveloped land	2.29e–05*** (8.40e–06)	2.25e–05*** (8.23e–06)
Planning expenditures	− 1.39e–07** (5.50e–08)	− 1.38e–07** (5.41e–08)
Average housing value	− 1.69e–05** (6.75e–06)	− 1.64e–05** (6.59e–06)
Millage	0.199* (0.103)	0.195* (0.101)
Constant	− 0.218 (2.034)	− 0.219 (1.993)
Observations	63	63
R-squared	0.217	0.217

Note: Robust standard errors in parentheses.
* $p<0.10$; ** $p<0.05$; *** $p<0.01$.

TABLE 7
Influence of Policy Variables on Price Appreciation during Bubble

A 1 Std. Dev. *Increase* in	Elasticity Change		Bubble Appreciation Change
Minimum lot size	0.62 unit decrease	⇒	3.7% increase
Undeveloped land	0.64 unit increase	⇒	3.8% decrease
Planning expenditures	0.46 unit decrease	⇒	2.8% increase
Average housing value	0.48 unit decrease	⇒	2.9% increase
Millage	0.48 unit increase	⇒	2.9% decrease

as residential land in 2011. The variables we use to capture local fiscal conditions are the 1990 countywide real property millage rate and the average value of single-family homes in 1990.

Results from these regression models are displayed in Table 6. The results show that counties with larger minimum lot size requirements are found to have lower short-run supply elasticities. The amount of developable land is also found to play an important role in determining the short-run supply response to price increases, as areas with more developable land have higher short-run elasticities. The estimated coefficients on the planning expenditure terms in Table 6 are all negative and are statistically significant. These results provide some evidence that counties that spend more money implementing and enforcing land use regulations have lower short-run supply elasticities. Turning to the fiscal variables, the positive and statistically significant coefficient on the millage rate term is consistent with the hypothesis that developers are less constrained by local officials when the tax revenue generated by new development is more likely to exceed the increase in expenditures on public services associated with the development. The negative and statistically significant coefficient on the housing value term is also consistent with the conjecture that the strategic approval of fiscally advantageous projects influences the elasticity of supply of housing, as local officials have strong incentives to prohibit or delay projects that may erode the average value of property in the jurisdiction.

The results from Table 6 can be used to investigate how characteristics of the local economic environment influenced price appreciation and new construction during the most recent housing boom and bust cycle. One such exercise is reported in Table 7, where we utilize the elasticity coefficients from Table 2 to study how local factors influenced price appreciation during the bubble through the supply elasticity channel. The results reveal that relatively small differences in the variables influencing supply elasticity result in nontrivial differences in price appreciation. From a public policy perspective, these results suggest that even moderate changes in a local community's regulatory environment and fiscal conditions can have a

TABLE 8
Appreciation Regressions with Saiz-Type Supply Measure, Dependent Variable: Price Appreciation

	Time Frame			
Variable	2000–2006 (1)	2000–2006 (2)	2007–2010 (3)	2007–2010 (4)
% Undevelopable land	118.8*** (16.41)	56.15** (25.76)	−48.68*** (9.477)	−14.67 (13.87)
% College graduates		−0.205 (0.327)		−0.402 (0.370)
Income per capita		0.000108 (0.000817)		0.000159 (0.000599)
Income growth		0.00347* (0.00176)		0.000711 (0.000864)
Rainfall		1.373** (0.568)		−0.424 (0.354)
January temperature		3.228*** (0.969)		−2.582*** (0.568)
Constant	41.91*** (8.263)	−205.3*** (54.04)		142.5*** (35.39)
Observations	63	63	63	63
R-squared	0.388	0.624	0.211	0.486

Note: Robust standard errors in parentheses.
* $p<0.10$; ** $p<0.05$; *** $p<0.01$.

significant impact on the magnitude of its price swings during a housing boom.

Comparison with Models Utilizing Saiz-Type Supply Measure

To date, authors wishing to investigate the relationship between housing supply conditions and price and construction dynamics during boom and bust cycles have relied heavily upon Saiz's (2010) measure of the share of undevelopable land within a metropolitan area to proxy for differences in the elasticity of housing supply. The popularity of the Saiz proxy stems from the presumption that it is likely exogenous to housing price and quantity changes. However, Cox (2011) describes it as "blunt in the extreme" because it is based on the availability of developable land in areas of the same geographic size (a circle with a 50 km radius) regardless of the size of the metropolitan area and ignores the land area that is already developed. Our data set, augmented by the Saiz proxy measured at the county level in Florida, allows us to investigate the relationship between the proxy and our estimated housing supply elasticities. Additionally, having both the proxy variable and the supply elasticity measure enables us to inspect to what extent the use of a proxy variable to test the hypotheses of Glaeser, Gyourko, and Saiz (2008) may result in spurious inference.

Results Using Saiz Proxy

Tables 8 and 9 report the results obtained from reestimating the price appreciation and new construction models for the boom and bust periods using the Saiz proxy (in lieu of our estimated supply elasticities). In the 2000–2006 price appreciation models, the estimated coefficient on the proxy is positive and statistically significant. In the 2007–2010 models, the proxy is negative but significant only in the bivariate model. These results parallel those obtained using our estimated supply elasticities (reported in Tables 2 and 3) and lend support to the use of the Saiz proxy. However, a reliable proxy variable for the supply elasticity must be able to explain both price and quantity changes. In the construction models, the Saiz proxy is not significant in any of the four models estimated (see Table 9), which lends support to Cox's criticism that the proxy may be too crude for the purposes at hand.

Finally, we estimated elasticity regressions that included the Saiz proxy as an explanatory variable using three different specifications: (1) a bivariate model including only the Saiz proxy, (2) a fully specified model in which the Saiz proxy is used in lieu of our measure of undeveloped land, and (3) a fully specified model that included both the Saiz proxy and our measure of undeveloped land. The results from these models are reported in Table 10.

TABLE 9
New Construction Regressions with Saiz-Type Supply Measure, Dependent Variable: ln(New Construction)

	Time Frame			
Variable	2000–2006 (1)	2000–2006 (2)	2007–2010 (3)	2007–2010 (4)
% Undevelopable land	0.119 (0.598)	0.202 (0.755)	− 0.163 (0.475)	0.0398 (0.598)
% College graduates		0.000301 (0.0203)		0.00230 (0.0173)
Income per capita		1.64e–05 (3.57e–05)		− 8.94e–06 (3.83e–05)
Income growth		− 5.42e–06 (6.82e–05)		2.79e–05 (6.77e–05)
Rainfall		− 0.0208 (0.0213)		− 0.0310* (0.0168)
January temperature		− 0.0269 (0.0289)		− 0.0305 (0.0207)
ln(Total housing units)	1.284*** (0.0638)	1.268*** (0.106)	1.042*** (0.0527)	1.092*** (0.0953)
ln(Land area)	− 0.515* (0.268)	− 0.425 (0.303)	− 0.365* (0.216)	− 0.328 (0.253)
Constant	− 1.884 (1.693)	− 0.0430 (2.557)	− 1.423 (1.451)	1.253 (2.229)
Observations	63	63	63	63
R-squared	0.853	0.859	0.836	0.847

Note: Robust standard errors in parentheses.
* $p<0.10$; ** $p<0.05$; *** $p<0.01$.

TABLE 10
Elasticity Regressions, Dependent Variable: Short-Run Elasticity

	Housing Supply Measure					
Variable	Standard (1)	Standard (2)	Standard (3)	Corrected (4)	Corrected (5)	Corrected (6)
% Undevelopable land	− 1.628	− 0.538	0.887	− 1.566	− 0.473	0.934
	(1.203)	(1.862)	(1.971)	(1.185)	(1.834)	(1.941)
Minimum lot size		− 1.007***	− 0.876***		− 0.984***	− 0.855***
		(0.332)	(0.321)		(0.325)	(0.314)
Undeveloped land			2.46e–05***			2.43e–05***
			(8.93e–06)			(8.74e–06)
Planning expenditures		− 6.97e–08	− 1.52e–07***		− 7.05e–08	− 1.52e–07***
		(5.32e–08)	(5.61e–08)		(5.23e–08)	(5.53e–08)
Average housing value		− 1.70e–05*	− 1.88e–05**		− 1.67e–05*	− 1.84e–05**
		(8.80e–06)	(8.06e–06)		(8.63e–06)	(7.91e–06)
Millage		0.114	0.210*		0.112	0.207*
		(0.111)	(0.111)		(0.109)	(0.109)
Constant	2.717***	2.384	− 0.714	2.655***	2.316	− 0.741
	(0.584)	(2.303)	(2.527)	(0.573)	(2.258)	(2.478)
Observations	63	63	63	63	63	63
R-squared	0.019	0.143	0.220	0.018	0.142	0.221

Note: Robust standard errors in parentheses.
* $p<0.10$; ** $p<0.05$; *** $p<0.01$.

In none of these models is the Saiz proxy statistically significant. These results, along with those obtained from using the Saiz proxy to explain new construction, suggest that researchers should be cautious in their reliance on Saiz's proxy for the supply elasticity.

V. CONCLUSION

Mixed evidence is provided in recent studies that have investigated the role played by interarea differences in supply elasticity in affecting the price and quantity changes that occur during a boom and bust housing cycle. In this paper, instead of using proxies for the supply elasticity, we use 21-year panels to estimate the short-run price elasticity of housing supply directly for 63 Florida counties. In areas where the supply elasticity is higher, we find that boom periods are characterized by greater construction and less price appreciation. During bust periods, there is more new construction in high supply-elasticity areas, but the magnitudes of price changes are unaffected by whether the supply elasticity is high or low.

The importance of supply elasticity as a determinant of price and quantity changes that occur in response to demand shocks begs the question: What factors account for differences in the supply elasticity across areas? We provide evidence that suggests that minimum lot size requirements, local comprehensive planning expenditures, the amount of developable land, and the local government's fiscal situation all affect the construction industry's response to higher prices.

An important unanswered question regarding the effect that supply has on housing markets during a boom and bust cycle is the extent to which overbuilding occurs on the upswing. More overbuilding translates into greater excess inventory as the cycle turns, which causes greater price deflation in the downturn. The relatively cheap housing units may affect household location decisions, resulting in significant welfare losses. Future research should attempt to develop reliable counterfactuals of the level of construction that would have occurred in the absence of the boom and bust phenomenon.

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