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**HOUSING PRICES AND SINGLE-FAMILY PERMITS:
THE CASE OF CALIFORNIA CITIES IN THE 1990s**

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Housing Prices and Single-Family Permits:
The Case of California Cities in the 1990s

An Undergraduate Honors Thesis

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ABSTRACT

This paper analyses the role of local land use policies in determining the levels and changes of housing prices in California cities between 1990 and 2000. Three models are developed. First, a cross-sectional model fits price levels to income, demographic, and regulatory variables that characterize the extent to which a given city's land use policy reflects exclusivity or growth-hospitality. Second, a "before and after" transformation of the cross-sectional model compares housing price changes with variables that relate the extent to which local policy favors single-family, detached housing construction through the permits process. Third, the "initial conditions" model relates changes in housing prices to the permits variables. The models address potential endogeneity built in the permits variables with a two-stage least squares procedure. For each sampled municipality, the mentioned exclusivity and growth-hospitality variables, along with state-proposition voting outcomes, serve as instrumental variables. On the whole, the regressions substantiate the hypotheses that both regulations and demographics influence housing prices.

HOUSING PRICES AND SINGLE-FAMILY PERMITS:
THE CASE OF CALIFORNIA CITIES IN THE 1990S

I. Introduction

A considerable body of literature addresses the impact of local land use regulation on housing prices. Upon review, one could speculate that, since California's cities categorically and historically bear more dramatic population and income growth than cities in any other state, researchers have often looked there to examine the local public sector's impact on housing markets. Furthermore, many policy analysts have noted that California municipalities, on the whole, exert a particularly strong range of regulations, including specific zoning requirements, residential and commercial development incentives, and growth management schemes. In this respect, California provides an apt laboratory for the effects of land use regulation. Appropriately, this analysis uses California to develop a model for inter-municipal housing price change with particular attention to the effect of a land use bias towards single-family housing.

Aside from the direct effect of increasing the cost of construction and limiting the supply of developable land, land use regulations have the potential to confer monopoly power on builders and to reorient the market toward higher income housing consumers who ostensibly demand fewer services from the local public sector. Additionally, a well-zoned community may simply appear more attractive and, therefore, may generate higher housing demand.[1] Thus, with all else equal, regulation increases housing prices. The economic justification for regulation cites the potential for spillover effects throughout

[1] Brueckner and Lai (1996) refer to these two explanations as the "supply-restriction" model and the "amenity-creation" model, respectively.

the urban economy.[2] These benefits include alleviating congestion and reducing environmental and infrastructural costs. By extension, regulation may alter a community's economic and social situation by influencing transportation costs, housing tenure choice, and the sorting of different economic and demographic groups.[3] These wide-ranging effects are complex and often difficult to measure.

Concerning more direct market effects, existing research has shown that, unsurprisingly, land use regulation increases housing prices. Empiricists have proposed many ways to measure regulation, e.g. dummy variables for specific construction requirements or indices constructed from public policy surveys. This paper uses two indices developed by Quigley, Raphael, and Rosenthal (2002) as explanatory variables to measure inter-jurisdictional housing policy variations in 1990s California. Specifically, these variables quantify the extent of local bias in favor of single-family detached housing construction. The census counts a net increase of 1,031,667 housing units in California between 1990 and 2000 – the following indices simply characterize the local distribution of housing stock growth in terms of permits for single-family detached housing.

The first, the "Deviations Index," measures the extent to which the number of newly issued permits for single-family detached housing deviates from expectations. Expectations are drawn from the statewide issuance of permits and the local proportion of

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[2] For a historical example, city officials initially proposed zoning as a remedy to the dense, noxiously mixed-usage (and the accompanying volatility of land values) symptomatic of the unfettered privatism that characterized land use decisions in 19th Century American cities. Zoning remains one of the most prominent manifestations of modernity's influence over urban spaces.

[3] Malpezzi (1996) examines externalities associated with regulation. Quigley, Raphael, and Rosenthal (2002) measure the effect on racial and ethnic composition.

single-family units at the start of the decade. Higher values of the Deviation Index could reflect either a public policy bias toward low-density development or the relative inclination of local land use decision-makers to favor growth in the housing stock. The second variable, the "Proportions Index," measures the proportion of all new residential permits that are allotted to single-family detached housing. Unlike the Deviations Index, higher values of the Proportions index reflect only policy bias towards low-density development and not the overall expansion of single-family housing stock.

Neither of these variables measures any specific regulatory characteristic or zoning enactment *per se*. Rather, the permits indices, as quantifiable products of a range of land use decisions, provide a proxy for the ultimately imperceptible variable "regulation," a unique political quality that exerts a complex influence over housing supply. In particular, the permits indices represent regulation in terms of its hypothesized influence over housing densities. This allows for a new empirical question: Is a bias toward single-family development associated with higher housing prices?

The models presented in this paper relate the mentioned indices to changes in first, second, and third quartile owner-occupied values and rents between 1990 and 2000 for a sample of California municipalities. Two base specifications are developed to estimate the determinants of housing price change. The first uses a "before and after" transformation of simple cross-sectional housing price models to control for unobserved, municipality-specific fixed variables and for changes in other determinants of price. The second controls for an array of initial conditions in the local housing market as measured by the 1990 Census. The permits indices are then added to each model to draw out an association between regulation and housing price change.

However, regulations are not exogenously determined in these models. While permits hypothetically affect housing prices, a range of market conditions, certainly including housing prices themselves, influences the demand for and supply of permits. A two stage least squares procedure (2SLS), however, mitigates this simultaneous causality bias. Quigley, Raphael, and Rosenthal (2002) use a survey of local officials to construct two instrumental variables, one to measure the extent to which municipal land use policies are "pro-growth" and another to quantify the degree of "exclusivity."[4] To strengthen the first stage estimates, this study adds selected municipal ballot box responses to state propositions.[5] The chosen propositions concern issues that often polarize liberal/conservative political divisions. Assuming that these distinctions carry over into similarly opposed attitudes toward local land use, the ballot box instruments function as bottoms-up political measurements to accompany the top-down political survey indices as first stage estimators of the single-family permits variables.[6] Ideally, these instrumental variables only affect housing price changes through their associated permits outcomes.

The regressions presented in this paper provide a general characterization for the determinants of housing prices in California. Overall, the permits indices provide mixed results. On the contrary, the regressions suggest that racial and ethnic changes have a

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[4] For elaboration on the determinants of the "pro-growth" and "exclusionary" variables, see the original paper.

[5] The ballot box instruments rely on a dataset and procedure put forth in Dunn, Quigley, and Rosenthal (2004) to examine the effect of prevailing wage requirements on affordable housing construction.

[6] Fischel, noting that "local government actions . . . are based, ultimately, on the results of the ballot box," bolsters the assumption that voters dictate land use policy (1985, p. 95).

clearer dynamic effect on housing prices. Still, when controlling for initial conditions, the Proportions Index negatively correlates with housing price inflation.

II. A Review of Literature on Land Use Controls

As stated earlier, a considerable body of research has established a connection between land use regulation and housing prices. Many of these studies feature a hedonic housing model with prices as a function of amenities, demand parameters, regulations, and other supply constraints. Empirical work on regulation's efficiency for internalizing externalities associated with non-conforming uses remains significantly less developed.

In critiquing one such study, Fischel (1990) highlights a common problem in the literature. Mark and Goldberg (1986) use a sample of single-family housing units sold in Vancouver from 1957 to 1980 to test three hypotheses associated with land use regulation: (1) zoning classification affects sale price, (2) allowing non-single-family uses in a neighborhood lowers sales price, and (3) re-zoning to allow higher densities and different uses increases sale price. Running separate regressions for each sample year, Mark and Goldberg note that regulation coefficients are often insignificant and change signs from year to year. They conclude that zoning produces ambiguous price effects that fail to satisfy the externalities justification. Fischel, however objects on the grounds that Mark and Goldberg mistakenly assume land use decisions are exogenous to prices. Under the assumption that zoning both limits the negative impact of non-conforming uses and provides such uses where they are most desired by residents, a political economy model suggests that a well-functioning zoning scheme should, in fact, demonstrate ambiguous price effects over time.[7] This assumption actually reverses Mark and Goldberg's interpretation and supports zoning's capacity for internalizing external costs. The effects of zoning may simply be difficult to detect under a rational, participatory political system. Moreover, accepting the fact that political conditions vary over time and space provides an economic reason to suspect simultaneous causality bias in models that strictly regard price as a function of regulations.

However, other empirical work, in contrast with that of Mark and Goldberg, has suggested that zoning does have real, non-ambiguous price effects and externalities. In a review of the literature on growth controls, Fischel (1990) concludes that, for a given municipality, growth control measures reduce the value of undeveloped land subject to restrictions, increase housing prices in the restricted municipality, and create spillover effects in neighboring municipalities.[8] Much of this research relies heavily on the assumption that local governments act in the interest of the electorate – essentially, an extension of the Tiebout hypothesis that housing consumers choose their residential locations with regard to personal preferences for the outputs of local public services.[9] This conjecture positions a framework for understanding land use restrictions in relation to maximizing the utility of the median voter. Hamilton (1978), assuming that local housing demand slopes downward with labor demand, hypothesizes that restrictions on the quantity of housing generate gains for owner-occupants and landlords. Renters

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[7] Chapters 2-5 of Fischel (1985) develop this political model in much greater depth.
[8] Additionally, Fischel notes that studies focusing on undeveloped land (and in that way controlling for the political effects of ownership) should show the cleanest results for land-use constraints (1990, 21).
[9] Mills and Oates (1985) offer an extensive discussion of the Tiebout Model's relevance to fiscal zoning and other land use controls.

receive no net gain as wages rise to compensate for rent increases. Thus, in a jurisdiction with concentrated land use decisions, zoning authorities engage in monopolistic restriction of housing supply to the benefit of constituent landowners. In equilibrium, landowners favor the restricting of supply and renters are indifferent. Hamilton, finding that fragmentation of zoning authority reduces property values, offers support for his assertion that municipal governments may use land use powers to raise the price of housing to the benefit of the electorate. In a similar vein, Brueckner and Lai (1996) use a Nash equilibrium model to show how land use restrictions may benefit resident and absentee landowners at the expense of renters in a city that controls development (p. 142).

These political models logically extend into an explanation for "fiscal zoning," i.e. land use decisions targeted to limit fiscal impacts, often by excluding residents that pay less taxes and demand more public services, namely, the poor. To suggest an example relevant to this paper, a bias towards single-family housing construction may reflect these exclusionary tactics. If landowners associate lower average tax revenues and higher average demands on public services with poorer residents (renters and multi-family housing consumers), fiscally minded landowners may pursue single-family construction through the political process as a means to exclude less desirable residents.

As evidence of regulation's impact on housing construction, Thorson (1997) uses the example of agricultural downzoning in a rural Illinois county to assess the effects of exclusionary zoning on housing permits. He runs pooled time series regressions on a reduced form equation for change in the quantity of housing as a function of income, population, amenities, construction costs, and zoning. The results show that, in the long run, despite the potential for sidestepping by developers, zoning significantly reduced the number of permits issued by the county. In another paper, Thorson (1996) provides evidence in support of Hamilton's claim that urban areas with greater monopoly power over land use decisions tend to have higher housing prices. Similarly, Mayer and Sommerville (2000) find that metropolitan areas with more extensive regulation can have up to 45 percent fewer housing construction starts and price elasticities more than 20 percent lower than those in less-regulated markets, with different effects measured for different types of regulation.

Studies performed by Schwartz, Zorn, and Hansen (1986), Schwartz and Zorn (1988), and Katz and Rosen (1987) use hedonic price models to compare sale prices in growth-controlled municipalities with neighboring "control" groups. An included dummy variable indicates the "treatment" of growth control. Katz and Rosen find a price effect of 17 to 38 percent for San Francisco Bay Area communities, while Schwartz et al. find estimates of about 9 percent for Petaluma and Davis, California. The difference in estimates may be attributable to the Schwartz and Zorn sample's greater proximity to San Francisco (Fischel 1990, p. 32). Pogodzinski and Sass (1991), however, point out a problem with the use of dummy variables in hedonic specifications of zoning models. While most studies include only one dummy or one set of related dummies for specific regulations (e.g. growth control or large lot requirement), regulations actually consist of multidimensional bundles of regulations that could affect the choice of specific housing characteristics (p. 272). Additionally, treating zoning as a shift parameter ignores the potential capitalization of zoning restrictions into the implicit prices of other hedonic attributes in the model.

Other studies concerned the inter-jurisdictional effects of the overall municipal regulatory environment. For a sample of San Francisco Bay Area municipalities, Dowall and Landis (1982) examine the effect of land use controls on average housing prices. Their results indicate that new home prices are higher in communities with low-density development policies, limited supplies of vacant land, and high development fees. Likewise, Elliott (1981) finds that, in a sample of California cities from 1969 to 1976, housing prices increased significantly faster in growth-controlled jurisdictions of extensively regulated housing markets than in non-growth-controlled jurisdictions. Pollakowski and Wachter (1990) similarly find that zoning and growth controls together produce a greater effect than either form of restriction taken individually. Again, their results indicate that regulations increase the price of housing and developable land.

Another group of studies, mostly authored by Malpezzi and Green, has paid special attention to the different effects of regulations at different ends of the market. Green (1999) regresses a reduced form equation of specific zoning restrictions on housing prices, rents, and tenure choices for a sample of municipalities in Waukesha County, Wisconsin. When the sample is limited to owner-occupied houses valued at less than $75,000, the effects of restrictions increase in both magnitude and significance. Green, speculating that land use restrictions are likely to be more binding at the lower end of the market, cites these results as evidence that land use regulations tend to fall more heavily on lower income households (p. 158).

Malpezzi (1996), in developing a simple model for estimating median housing values and rents across a sample of 48 American metropolitan areas, also pays particular attention to regulations. Like most studies on regulations, Malpezzi constructs a reduced form model of prices, which he then ports onto a range of measurements for regulatory externalities. The model includes controls for income level and change, population level and change, and geographical constraints. Regulatory variables include a state regulatory index, a measure of local restrictiveness, and a dummy for rent control. He gleans each of these variables from policy surveys. With all else held constant, the regression results suggest that moving from a "lightly" regulated environment to a "heavily" regulated environment increases rents by 17 percent and owner-occupied house values by 51 percent.[10] With a higher R-squared and greater significance for the regulatory variables, the value model slightly outperforms the rent model. Malpezzi and Green (1996) extend the model to first and third quartile house value and rents and find, like Green (1999), that the effects of regulations are more pronounced at the bottom of the market.

However, Malpezzi's drawing of observations from across state-lines confuses any attempt to disentangle the effect of state regulatory policies from local regulatory policies. Furthermore, he relies heavily on purely linear combinations of land use restrictions to index the regulatory environment. Noting the inherent simultaneity of a regulatory model, Malpezzi experiments with two and three stage least squares but loses too many observations to incomplete data. Still, by considering median housing prices, Malpezzi (1996), like Dowall and Landis (1982) and Elliot (1981), may draw from a broader sample of geographies for the effects of land use regulations.

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[10] For the purposes of this example, Malpezzi defines "lightly regulated" as first quartile state and local regulatory indices and a rent control dummy value of 0. He defines "heavily regulated" as third quartile state and local regulatory indices and a rent control dummy value of 1 (1996, p. 230).

Quigley, Raphael, and Rosenthal (2002) provide a blueprint for a different prospective on land use studies. To measure the effect of a single-family permits bias on demographic change, they draw from a large sample of California municipalities and thereby control for state-determined regulatory effects. Most importantly, unlike Malpezzi, when using instrumental variables to account for simultaneity, Quigley et al. retain a relatively large sample size, an advantage over control/treatment hedonic studies that may suffer selection bias. Another study that considers local building permits policy, Asabere and Huffman (2001), finds that a growth-conducive permits environment tends to increase prices of vacant land. If pro-growth municipalities tend to allow higher density development, vacant land is valued at higher rates per acre. However, applying a similar line of reasoning to housing prices rather than vacant land prices creates a greater potential for ambiguity. This problem is discussed in the next section.

Finally, this paper aims (1) to develop a simple cross-sectional model for city-level housing prices, (2) pool the cross-sectional model across two time periods to measure the dynamic effect of several housing price determinants over time, (3) to experiment with Quigley, Raphael, and Rosenthal's (2002) first stage estimates with the aim of isolating more relevant sets of instrumental variables, and (4) to asses the effect of a single-family permits regulatory bias on housing price change.

III. Speculating the Outcomes of a Single-Family Permits Bias

Definition of the Permits Indices

A discussion of the market effects of a single-family bias first requires a brief description of the permits indices that act as key explanatory variables in the subsequent regression analyses.[11] As stated earlier, the "Deviations Index" relates the proportion by which single-family residential building permits issued within a given jurisdiction deviate from expectations based on the city's initial proportion of single-family units in 1990 and the overall growth of single-family permits in the state between 1990 and 2000. Expectations, N_i, issued by city *i* is given by:

$$N_i = \Delta Single * Single_i / Single. \qquad (1)$$

$Single_i$ and *Single* define the number of single-family units for the given city and for the whole state, respectively. Likewise, $\Delta Single_i$ equals the number of new permits issued between 1990 and 2000, and $\Delta Single$ sums values of $\Delta Single_i$ over the entire state. As defined, expectations reflect the distribution of single-family permits according to a given city's initial 1990 share of the total state-wide single-family housing stock. The deviation from expectations, $Deviations_i$, equals the proportionate difference between the actual amount of permits issued by the city, $\Delta Single_i$, and the expected amount, N_i:

$$Deviations_i = (\Delta Single_i - N_i) / N_i. \qquad (2)$$

By the nature of its construction, the Deviations Index can reflect a variety of market and policy conditions. For instance, high values could reflect a local development bias toward single-family construction, at least with respect to the initial composition of

[11] Quigley, Raphael, and Rosenthal (2002) offer greater detail on the construction of the permits variables (pp. 10-13). This summary merely paraphrases the original paper where relevant to the explanation of housing prices.

the municipal housing stock in 1990. On the other hand, when a city experiences rapid growth, in-migrants likely demand more housing at all density levels, multi-family and single-family. In the latter case, high values of the Deviations Index may reflect a relatively dramatic expansion of total housing supply, not necessarily excluding multi-family and rental housing.

Quigley, Raphael, and Rosenthal construct the second permits variable, the Proportions Index, as the simple ratio of all new residential building permits issued between 1990 and 2000 that are specifically reserved for single-family detached units. Where T_i equals total residential permits issued in a given city i, $Proportions_i$ is defined:

$$Proportions_i = \Delta Single_i \,/\, T_i. \tag{3}$$

Unlike the Deviations Index, the Proportions Index does not account for initial composition of the housing stock. On the other hand, since the Proportions Index does not depend in any way on overall city growth, it avoids the ambiguities built into the Deviations Index.[12]

Applying the Permits Indices to a Simplified Housing Market

For the sake of simplification, assume the existence of a hypothetical city

[12] Quigley, Raphael, and Rosenthal (2002) also note a caveat associated with assigning single-family permits as proxies for a market oriented toward high-end housing construction. Many neighborhoods have favored high-end condominiums and rental units that likely draw a high-end income demographic but evade the specified permits indices. The existence of such neighborhoods, undetected by a measurement reliant on single-family permits, adds measurement error to the indices and biases OLS coefficients toward zero (p. 14).

populated by two types of housing consumers: low quality renters and high quality owners.[13] Both types interact with segmented but related housing submarkets such that changes in supply and demand in one submarket could affect price in the other. Assume that, regardless of quality or tenure, the standard demand determinants for housing quantity (Q_D) include housing value or rent (*P*), population (*pop*), a vector of income and human capital variables (*Y*), the cumulative monetary value of local amenities (amenities), and a vector of demographic variables (*demo*):

$$Q_D = f(P, pop, Y, amenities, demo) \quad (4)$$

Similarly, a city's housing supply (Q_S) depends on price, construction costs (*costs*), and a vector of regulatory variables (*reg*):

$$Q_s = f(P, costs, reg) \quad (5)$$

Housing is supplied as developers construct new housing or as residential consumers vacate existing housing. Logically, either by restricting the total supply of housing or by increasing the minimum quality of newly constructed units, local policies could significantly increase housing prices. On the other hand, a regulatory bias in favor of supplying one quality submarket (single-family detached housing for instance) may affect supply in other quality submarkets (low quality and rental housing).

[13] Obviously, not every low-quality housing consumer rents, just as not every high-quality housing consumer owns. These assumptions simply facilitate the discussion of segmented but related markets.

In particular, high-end housing growth may affect the low-end market equilibrium through a quality-“filtering” process. The filtering model describes the interactions between housing quality submarkets: all else equal, as housing deteriorates over time, wealthier households move into newer and higher quality housing while poorer households come to occupy units vacated by mobile consumers. Because high-income, single-family housing is more profitable to develop, new construction directly increases the supply of high-quality housing. In an efficiently filtering environment, such a high-quality supply shift allows high-end consumers to vacate existing housing, thereby increasing the supply and lowering the price of lower quality housing. Still, for filtering to work, new units must not simply displace low quality units at the same price per unit of housing service.

Now, suppose that single-family housing starts account for all residential construction permits issued by the hypothetical city during an extended period, e.g. the 1990s decade. In other words, the Proportions Index equals 1. Meanwhile, assuming that rental units account for some proportion of the total housing stock in 1990, the Deviations Index is positive. In this case, both indices probably reflect an augmentation of the higher-end housing submarket, which may filter down into lower quality and rental submarkets by lowering prices.

However, housing consumers may favor such a single-family oriented municipality. For reasons explained earlier, high-end residents who pay more taxes and demand fewer public services may perceive single-family growth as more fiscally efficient. Additionally, low-density tracts of single-family detached housing may simply appeal to cultural norms and aesthetics: a symptom of the anti-urban ethic that runs throughout American geographical history. In this case, a single-family bias implies an increase in the monetary value of local amenities. In other words, the permits indices may reflect a demand shift. All else equal, favorability towards low-density, single-family areas may increase the demand for and the price of *all* qualities and types of housing.

In the context of segmented submarkets, this demand effect has the potential to crowd out lower income residents as high-quality consumers bid higher prices per unit of housing. Alternately, since the benefits of fiscal zoning and low-density development mostly accrue to homeowners, renter demand may not increase. In an extreme case, if single-family bias suggests that local policy favors homeowners over renters, demand may actually decrease, as renters may prefer either to own locally or to move to a more receptive environment. Independent of the volume of construction, the Proportions Index could reflect such a policy bias more than any actual supply effect. On the other hand, the Deviations Index somewhat more ambiguously could increase even if supply increases for all quality levels and tenure types.

This version of the single-family permits mechanism leads to *a priori* speculations of housing price outcomes. If the demand effect dominates, the Proportions Index should vary positively with housing values. In this case, values capitalize the amenities associated with a single-family bias. The Proportions Index should vary negatively with values if the supply effect dominates. Higher values of the Deviations Index could reflect either a deviation from historical proportions of single-family housing units or a dramatic growth of all housing types. In either case, if the supply effect dominates, value should increase with the Deviations Index. If housing units filter relatively efficiently through quality levels and the demand effects are nominal, both indices should vary negatively with rents and values at each quartile.

IV. Data Descriptions

The following empirical analysis draws data from several sources. First, Quigley, Raphael, and Rosenthal (2002) adapt the permits indices from data recorded by the California Industry Research Board (CIRB) and summed over the decade in question. Second, the Census Summary File 1 provides place-level data on population, race, and ethnicity for 1990 and 2000.[14] Similarly, the third source of data, Census Summary File 3, relates place-level estimates for median income, education, and housing price based on five percent samples. The demographic, income, and population data act as demand determinants in the simple housing price models presented below.

The census offers owner-occupied housing values and contract rents, the dependent variables, for the median, first, and third quartiles in each designated place.[15] Natural log transformations allow coefficients to be interpreted as elasticities. Compared to the rest of the country in 2000, California-level observations for each of the six housing price measures exceed those of every other state except Hawaii. At each

[14] Unlike the 1990 Census, the 2000 Census allowed respondents to choose more than one racial descriptors. Quigley, Raphael, and Rosenthal (2002) describe a procedure for rendering the two counts generally comparable. Essentially, since African-American respondents are most likely to choose non-Hispanic White as a second racial identifier, everyone selecting African-American, even in conjunction with other races, is identified as African American. The same applies to Asian. All non-Hispanic Whites choosing only one racial category are identified as White. The Hispanic population is measured identically in 2000 and 1990.

[15] The 1990 values are collected from 100 percent data, but the 2000 data are estimated from a 5 percent sample. While the numbers are still generally comparable, the difference in collection methods introduces error into the regressions.

quartile, contract rents exhibited a higher growth rate than house values (see Table 1). Over the same period, median household incomes increased by 36 percent and the percentage of income devoted to gross rent (i.e. contract rent plus utilities and other housing expenditures) decreased slightly from 29.1 percent to 27.7 percent. Thus, while incomes boomed, rent burdens decreased only marginally, suggesting a relatively elastic relationship between income and price. For both values and rents, the third quartile shows the highest proportional and absolute growth – the third quartile value growth rate more than doubles that of the median value.

The census data advantageously provides broad coverage of every California municipality, which allows for a large sample size. The main disadvantage, however, is that the census levels off all values and rents to a fixed maximum threshold. In a cross sectional model, bias should be small. But where the regressions measure change in housing prices, given that the maximum threshold reported by the census essentially doubled between 1990 and 2000, the coefficients are biased towards infinity. Specifically, the rent threshold increased from 1001 to 2001 dollars, and the value threshold increased from 500,001 to 1,000,001 dollars. Thus, the most expensive cities would have shown a value increase of 500,000 dollars and a rent increase of 1000 dollars. Obviously, these dramatic changes greatly exceed both reasonable expectations and the changes typically observed in less expensive municipalities.

To address the situation, the sample replaces problematic observations with sample mean changes in the natural log of price. This maintains sample size but introduces error-in-variables bias into the regressions. The regressions discussed below employ this method. Unfortunately, the only other alternative is to drop the observations that exceed the maximum threshold. However, the resulting non-random sub-sample violates the least squares assumptions and likely biases coefficients toward zero. For comparison, Appendix A reports and discusses the truncated regressions. On the whole, both approaches, though biased, produce similar estimates.[16]

Instrumental Variables

The fourth and fifth sources of data act as exogenous instruments in the 2SLS procedure. Presumably, these variables only affect price through the housing supply outcomes reported by the permits indices. Quigley, Raphael, and Rosenthal (2002) construct two variables intended to characterize a wide range of locally enacted restrictions and to act as first stage estimators for the permits variables. These measures are drawn from two comprehensive city-level surveys of the local regulatory environment undertaken by Madelyn Glickfeld, Ned Levine, and the League of California Cities (LCC) in 1988 and 1992.[17] The questions pertain to zoning restrictions, growth controls, development incentives, density requirements, etc. – in other words, the survey generalizes and records specific enactments of growth management policy. Quigley et al. transform subsets of the raw data into two broadly descriptive indices. One characterizes the extent to which a municipality is “pro-growth;” the other, the extent to which a

[16] Comparing the regression results from truncated and non-truncated samples shows that, in fact, the non-truncated coefficients, despite including only a marginally greater number of observations, vastly exceed their truncated counterparts. Seemingly, if the dependent variables are left untreated, the results are biased toward infinity and completely unreliable.

[17] Glickfeld and Levine (1992) report the 1988 survey. Levine (1999) compares the 1988 and 1992 surveys.

municipality is “exclusionary”.[18] As defined, growth-hospitality includes encouraging growth through incentives or through the planning process. Exclusivity, however, describes the degree to which a municipality limits growth and installs policy that favors low-density and high-income housing. Quigley et al. predict that the degree of exclusivity should negatively correlate with both the Deviations Index and the Proportions Index. The pro-growth measure should vary positively with the Deviations Index and negatively with the Proportions Index.

While the “exclusivity” and “pro-growth” variables should reasonably reflect the regulatory conditions that influence local favorability towards single-family permits, these instruments alone prove relatively weak first stage estimators of the permits indices. That is, tests of the joint significance of the “exclusivity” and the “pro-growth” variables in first stage regressions typically yield low F-statistics.[19] The “exclusivity” index is particularly insignificant at conventional levels in each first stage estimate. The authors note that instrumental weakness challenges the reliability of their 2SLS results.

To increase overall instrumental relevance, the following 2SLS models add a fifth source of data: city-level voting outcomes for state propositions during the 1990s. Much of Fischel’s work, among others, has continuously promoted the theory that the median voter implements effective regulatory policy. For example, according to a summary of evidence in Fischel (1992), the median voter hypothesis applies where residents are most

[18] To be concise, this paper only summarizes the construction of these variables. For details on the specific components and distributions of each measure, see Quigley et al. (2002, pp. 16-17).

[19] According to a general rule of thumb, Stock and Watson’s *Introduction to Econometrics* (2003) categorizes a first-stage F-statistic less than ten as “weak” for models including one endogenous regressor, i.e. 2SLS estimates are biased and t-statistics are unreliable (p. 350).

likely to enshrine fiscal zoning as a land use priority. Along this line of reasoning, any assessment of popular opinion with regards to public policy has the potential to mirror bottoms-up political influence on regulatory policy. Thus, city-level voting results are employed as gauges of the local political climate.

If voters favor fiscal zoning, they may be more likely to pursue single-family detached development through available political processes. Assuming that fiscal conservatism towards land use generally correlates with overall political conservatism and that most residents consistently and exclusively favor either liberal or conservative policies, ballot-box responses should indirectly mirror land use priorities, even if the proposition itself does not directly concern land use policy. Moreover, almost any policy involving the use of public funds could indirectly skew development towards either lower or higher densities, so the relative support for state fiscal propositions may reflect the local preferences for development patterns.

Table 2 presents summary descriptions and simple correlations for selected propositions' percentage of "yes" votes from ballots between 1990 and 2000, the "pro-growth" and "exclusivity" measures, and the permits indices. As an example, consider Proposition 167, which shifts a larger share of the tax burden onto the wealthy, and Proposition 168, which reduces the local political barriers to public housing projects. Consistent with the assumption that these propositions install fiscally liberal policies, both outcomes correlate relatively strongly and negatively with the Proportions Index. In other words, on average, residents of single-family biased municipalities are more likely to vote more strongly in favor of preserving other fiscally conservative land use paradigms. Through this mechanism, the ballot-box outcomes bolster overall instrumental relevance and improve 2SLS reliability.[20]

V. Implementing a Simple Model of Housing Price Change:

A Preliminary Cross-Sectional Model for Housing Values and Rents

Recall the housing demand (4) and supply (5) equations presented earlier. In market equilibrium, substituting for quantity yields a reduced form equation for housing prices in a given city *i*:[21]

$$P_i = f(Y_i, pop_i, demo_i, reg_i, error_i) \qquad (6)$$

As this reduced form relationship is stochastic, the specification includes an error term ($error_i$).[22] The following independent variables, each drawn from census data, act as housing price determinants in the simple cross-sectional model:

1. *Natural Log of Population (100,000s).* All else equal, more populous municipalities should exhibit higher competition for land and, therefore, higher housing prices.
2. *Natural Log of Median Household Income ($10,000s).* Individuals with higher

[20] Obviously, these instruments are not observed exactly simultaneously with the permits indices. Their use as instruments presupposes that regulatory policy and popular political notions remain relatively fixed over the decade in question and exert a constant influence over permitted construction densities.
[21] This transformation follows Malpezzi (1996), among others.
[22] This analysis does not consider construction costs (*cost*) mentioned in the supply equation (4). Assuming that construction wages are higher in areas with higher income levels and that wage differences are the main source of inter-city variation in construction costs, the income and wealth variables serve as a proxy for construction costs.

incomes demand higher quality housing services. Also, median incomes are likely to be greater with access to an urbanized labor market. Thus, income should positively correlate with housing prices.

3. *Proportion of Population with a College Degree*. Human capital investments are probably more reliable than current income levels as indicators of lifetime income streams and, accordingly, housing consumption. All else equal, cities with a higher proportion of college graduates over the age of 25 should exhibit higher housing prices.
4. *Proportion of Population without a High School Degree*. Conversely, a higher proportion of high school dropouts should correlate with lower housing prices.
5. *MSA dummy*. To control for intrinsic differences between metropolitan and rural land and labor markets, a dummy variable equals one if the 2000 Census includes the observed municipality in a Metropolitan Statistical Area.
6. *Black, Hispanic, and Asian Proportions of the Population*. Quigley, Raphael, and Rosenthal (2002) suggest that the single-family permits bias is associated with racial and ethnic change. If so, how do these demographics interact with market prices?

For a given city *i*, the cross-sectional model may be expressed linearly:

$$\begin{aligned} \ln price_i = \; & \beta_1 + \beta_2(\ln pop_i) + \beta_3(\ln income_i) + \beta_4(college_i) + \beta_5(dropout_i) + \\ & \beta_6(black_i) + \beta_7(hispanic_i) + \beta_8(asian_i) + \beta_9(MSA_i) + error_i. \end{aligned} \quad (7)$$

Table 3 presents OLS estimates for the 1990 cross-sectional model. The dependant variables are the natural logs of first, second, and third quartile owner-occupied housing values and contract rents. In general, the equations, explaining about 80 percent of inter-jurisdictional variation, fit the sample well; and estimates have predictable interpretations.

The human capital variables perform as expected: the proportion of college graduates positively correlates with prices, and the proportion of residents without a high school diploma shows the opposite effect. By extension, a city with higher earnings potential, on average, reports higher housing prices. Therefore, lifetime income streams likely affect the quality of housing demanded in a given city. The largest negative effect of the dropout rate occurs in the third-quartile value and rent regressions (Columns 3 and 6, respectively). Thus, low levels of human capital depress prices greatest at the high end of the market.

Each regression rejects the null hypothesis for racial and ethnic proportions at the five percent level. Hispanic and Asian compositions have a greater positive effect on values than on rents. On the other hand, black proportions, though significant at the five percent level only in the first quartile value regression, consistently exhibit a negative coefficient. This may be a function of the ongoing migration of Hispanic and Asian populations to California. While Blacks have historically suffered disproportionate isolation in low-income areas, newer and more mobile migrating populations are more likely to follow job opportunities and economic growth. On average, newer immigrants probably tend to settle in growing (higher priced) local economies and avoid depressed (lower priced) areas. With respect to these strong hypothetical assumptions, the evidence provides reasonable validation.

Since the “pro-growth” and “exclusivity” indices characterize local regulatory environments circa 1990, these variables may be added to the cross-section model as political determinants of housing supply. The results are abbreviated in Table 4. Each regression rejects the null hypothesis for jointly excluding both regulatory variables at the one percent level of significance. In particular, the pro-growth index categorically shows a highly significant, negative effect. Assuming that growth-hospitable policies increase the stock of housing, promote lower densities, and provide employment for lower income residents, holding all else constant, a higher pro-growth index predictably implies a lower price. By extension, in so far as growth-favoring policy promotes housing construction as a response to demand, such municipalities likely demonstrate more elastic long run supply curves. On the other hand, the exclusivity measure offers less consistent results. Exclusive policies likely skew growth towards high-income residents and restrict housing construction in general, creating relatively inelastic supply curves and higher prices. Just this effect is significantly observed for median and third quartile values (columns 2 and 3), but the opposite sign occurs in the first quartile rent regression (column 4). Still, the effect is small for all models. On the whole, with respect to both measures of housing policy, the sensitivity of values exceeds that of rents.

The “Before and After” Transformation

As the census provides data for both 1990 and 2000, the cross-section specification may be transformed into a dynamic price change equation by merely subtracting the 1990 reduced form equation (7) from the 2000 version. Each variable, then, becomes a pooled difference between the 2000 and 1990 observations:

$$(\ln price^{2000} - \ln price^{1990})_i = \alpha_1 + \alpha_2(\ln pop^{2000} - \ln pop^{1990})_i + \ldots + \alpha_9 (MSA^{2000} - MSA^{1990})_i + \alpha_{10} (X^{2000} - X^{1990})_i + \ldots + error_i. \quad (8)$$

Equation (8) introduces *X*, a vector of unobserved variables that possibly influence housing prices. For an example of such an unmeasured price determinant, large universities probably have a profound influence on local housing markets, especially where a large proportion of the population is university-affiliated, as is the case in Berkeley or Davis. In another possibility, unobserved physical constraints like mountains or large lakes might act much like *de jure* growth controls to hinder the supply of developable land and cause price increases. Assuming such unobserved qualities of the local housing market are fixed over the decade, the "before and after" transformation holds constant any fixed factors that differ from one municipality to the next. Thus, the set of static unobserved variables $(X^{2000} - X^{1990})$, like $(MSA^{2000} - MSA^{1990})$, equals zero for every observation. Moreover, this transformation intuitively justifies including the permits variables, as the Deviations Index and the Proportions Index measure change in single-family permits weighted against the city's initial residential composition and its overall issuance of permits, respectively.

In the OLS results presented in Table 5, the change in natural log of price is regressed on a single permits variable with no other covariates. While each regression fails to reject the null hypothesis for the Proportions Index, the Deviations Index consistently shows a positive, significant effect. Thus, when controlled for no other determinants of price, higher than expected deviations in single-family detached permits are associated with more pronounced price increases over the decade.

Next, Table 6 depicts OLS regressions of housing prices on the rest of the "before and after" change variables and the Proportions Index. As in the single regression case (see Table 5), for each housing price measure, the Proportions Index estimates are not significant. The income coefficients, highly significant in all regressions, suggest that values have a more elastic relationship to income than rents. In all likelihood, nominal income and price, both of which increased in almost every sampled city, are jointly determined as a result of inflationary pressures and local growth. Still, the results corroborate the state-level observation that high-end values have responded most dramatically to California's growth in median incomes during the 1990s (see Table 1).

For all specified ethnic and racial categories, local growth negatively relates to price change. That is, after controlling for other variables, a net increase in the proportion of each major non-white racial and ethnic group is associated with slower growth in every value and rent quartile. The F-statistic rejects the null hypothesis for racial effects at the one percent level of significance in every regression model except third quartile rents (column 6). When comparing estimates across quartiles, Hispanic populations (the state's fastest growing demographic and the model's most consistently significant racial variable) exhibit the most dramatic effects at the high-end of the owner-occupied market and at the low-end of the rental market. To illustrate this effect, for a hypothetical town in which the Hispanic proportion increased one standard deviation

(0.056) over the course of the decade, first quartile rents would have increased roughly five percent less than in a similar town with no Hispanic growth, holding all else constant. The different inter-quartile estimates for rents and values suggest some possibilities about how the effect of ethnicity varies across housing submarkets. Assuming Hispanic populations are more likely to rent in low-end housing submarkets and White populations are more likely to buy at the high-end, the third quartile value's greater sensitivity to Hispanic growth may reflect a perceived disamenity associated with high-end homeowner's distaste for income and ethnic mixing.[23] Then, in so far as high-end owner-occupied housing prices incorporate the tastes of homebuyers, values would increase relatively slowly. Accordingly, if Hispanic residents tend to disproportionately consume low-quality rental housing, a high rate of Hispanic growth may slow the increase of first-quartile rents by skewing that submarket toward lower-qualities. In general, the variables measuring change in Asian proportional populations perform similarly across quartiles. Contrarily, the Black variables show the strongest and most significant effect on first quartile values.

When the regressions include the alternate permits measure, the Deviations Index, as a key explanatory variable, the "before and after" model performs similarly (see Table

[23] Quigley, Raphael, and Rosenthal's (2002) results somewhat substantiate this assumption. The proportional growth of Hispanic populations is negatively associated with the proportion of all new housing permits allocated to single-family detached units. Furthermore, Hispanic growth is more likely in cities with relatively high poverty rates, low proportions of college graduates, high proportions of high school dropouts, and larger overall populations. These data suggest that, on average, Hispanic populations tend toward cities that likely have high proportions of high-density housing and of rental units. For housing submarkets, these characterizations suggest that Hispanic populations disproportionately demand low-end rental housing. Appendix B further describes the association between racial demographic change and housing prices. Table B1 shows that a higher percentage of Hispanic households are renter-occupied.

7). Adding covariates reduces the estimated coefficient on the Deviations Index to the point where it only retains statistical significance at the five percent level in the case of third quartile rents. To interpret, a policy that exceeds historical proportions with respect to single-family permits is associated with a greater increase in third quartile rents. However, excluding a radical deviation, this effect is relatively small. As mentioned earlier, a higher Deviations Index may arise from a local policy skewed toward low-densities or from overall growth of the housing stock, not necessarily excluding higher-densities. Since the coefficient could reflect either a net supply increase or the alleged inflationary price influence of exclusive land use policy, it is not altogether surprising that the estimates in regressions (1) through (5) reveal ambiguous and null associations.

For the racial and ethnic change variables in Table 7, as in Table 6, only the third quartile rent regression (column 6) fails to reject the null hypothesis at the five percent level. A comparison of each racial and ethnic measure's performance across quartiles generally resembles the interpretation of the Proportions Index model versions: Asian, Hispanic, and Black population growth are likely to dampen housing prices, especially the owner-occupied values. Thus far, the regressions strongly imply that race, indeed, affects housing prices. Overall, the evidence demonstrates consistency with the conventional notion that housing prices and racial sorting are highly associated.

2SLS on the "Before and After" Model

As discussed earlier, the permits variables may be endogenously determined in the housing price models. Specifically, while exclusionary land use policies may restrict supply and increase housing prices, in the interest of maintaining high home values, landowners themselves may pursue such policies through the political process. In other words, causality runs both ways, and OLS estimators may be inconsistent. To mitigate the effects of simultaneous causality, the following extension of the "before and after" model uses a 2SLS procedure to draw out consistent estimates for the permits variables.

The first stage estimates (see Table 8) include Quigley, Raphael, and Rosenthal's (2002) local land use characterizations and the statewide election results discussed earlier. Many of the available ballot-box results imply certain polarizing attitudes about local growth, exclusive land use policy, and more general tastes for political and fiscal conservatism; but such an in-depth political analysis exceeds the scope of this paper. Rather, for the narrow purpose of estimating the permits variables, maximization of instrumental relevance commands the set of first stage estimators. While these variables are sufficiently relevant to confidently estimate the Proportions Index, joint exclusion tests suggest that the available instruments provide slightly weaker estimates for the Deviations Index. As a result, the 2SLS estimator for the Deviations Index may suffer bias in the following regressions.[24]

Once again, each regression fails to reject the null hypothesis at acceptable levels of significance (see Table 9). However, the Deviations Index is significant at the ten percent level for first quartile values and rents when all covariates are included (columns 1 and 4). This implies, albeit weakly, that a higher measurement of the Deviations Index is associated with a lower inflation of housing prices over the course of the decade – an outcome consistent with the possibility that the Deviations Index signals a supply

[24] Instrumental weakness also plagues Quigley, Raphael, and Rosenthal (2002), especially for the Deviations Index. However, the addition of voting outcomes significantly increases instrumental relevance for both permits indices.

increase. Furthermore, marginal significance arises in low-end markets, suggesting that the lower submarket feels the greatest supply effects of the Deviations Index. On the other hand, the Proportions Index, a less ambiguously determined measure of single-family permits bias, lacks significance in each regression.

Thus, for both measures of bias toward single-family permits, the 2SLS regressions reveal no strong, consistent, causal effect on housing price changes. Rather, when considered as a whole, the "before and after" model lends greater importance to racial and ethnic change as determinants of 1990s housing prices in California. Quigley, Raphael, and Rosenthal (2002) suggest that these demographic changes respond to single-family permits variables. While difficulty marks the attempt to disentangle the relative price effects of these two factors (especially given the potential sources of bias), demographics seem to take precedence as a determinant of housing price. However, though the "before and after" model has potential for strong causal interpretations, the change variables are likely jointly determined with housing price in a dynamic local economy. This leads to additional sources of bias. In the next section, an alternate approach attempts to measure the association between the permits indices and housing prices by controlling for initial conditions in the local housing price market rather than for other changes in other price determinants.

The "Initial Conditions" Model

The "initial conditions" regressions measure 2000 housing prices as a function of the 1990 cross-section variables plus the permits indices,

$$(\ln price)_i^{2000} = f(\ln pop, \ln income, MSA, college, dropout, black, hispanic, asian)_i^{1990} + \gamma_9 (single\text{-}family)_i + error_i, \quad (9)$$

where *single-family* is either the Deviations Index or the Proportions Index. Subtracting l*n price*^{1990} from both sides of equation (9) enables an interpretative comparison to the "before and after" specifications:

$$(\ln price^{2000} - \ln price^{1990})_i = f(\ln pop, \ln income, MSA, college, dropout, black, hispanic, asian)_i^{1990} + \gamma_9 (single\text{-}family)_i - \gamma_{10} (\ln price^{1990})_i + error_i \quad (10)$$

Regression equation (10) measures the effect of the permits variables while controlling for local market conditions in 1990, including price. While this model does not account for changes in other local determinants of housing prices, the static variables control for inter-market differences in the determinants of housing prices and thereby isolate the association between housing prices and the permits variables. Thus, the "initial conditions" model, though it offers weaker causal interpretations, may provide more reliable estimates.

To improve instrumental relevance in the "initial conditions" model, the first stage regressions presented in Table 10 utilize slightly different sets of instruments from those employed in the "before and after" model. Once again, maximizing relevance guides the selection of available instruments. For both indices, joint tests indicate slightly less relevant instruments than those in the "before and after" 2SLS model and fail to confidently rule out a weak instruments bias.

In the Proportions Index models (see Table 11), OLS coefficients are significantly negative for median rent changes (column 5). Comparing OLS with 2SLS shows a more dramatic effect: the estimated sensitivity of rents increases in the 2SLS estimates (columns 4 through 6).[25] Moreover, the Proportions Index shows a significant effect for the measures of owner-occupied value changes in the 2SLS regressions. The coefficient varies between –0.408 (column 6) and –0.543 (column 2). To interpret with an example, when 1990 income and demographic conditions constant, a one standard deviation increase in the ratio of all new permits that are single family implies a 10 to 13 percent lower change in housing prices, depending on quartile and tenure type.[26] For each quartile, the effect on values exceeds the effect on rents. This result resonates with the assumption that single-family permits have a more direct interaction with the owner-occupied housing market. As multi-family units likely comprise a larger share of rental housing, single-family permits only indirectly increase rental supply. Filtering or some other mechanism through which housing submarkets interact apparently functions quite

[25] Incomplete data on the permits indices and instrumental variables decreases sample size in 2SLS regressions. Incidentally, all of the remaining observations are located within metropolitan statistical areas, so the variable *MSA dummy* is dropped from the 2SLS regressions. To ensure that the coefficients are comparable to the 2SLS estimates, *MSA dummy* is also dropped from the OLS regressions.

[26] For the sampled Proportions Index, the standard deviation is 0.238 and the mean is 0.758.

well in this model.[27] Still, these regressions only control for initial conditions and not for changes in other variables. Thus, these interpretations do not necessarily contradict the previous claim that price effects depend more on racial and ethnic changes than on the measurements of permits biases.

On the other hand, the Deviations Index performs poorly in the 2SLS regressions. Significant in the OLS estimates, a deviation from expectations with respect to single-family permits exerts a positive effect on prices. The estimates imply that a one standard deviation increase in the Deviations Index (i.e. approximately 1.82) correlates with an increase in housing price change of about 2 percent.[28] However, in the 2SLS results, significance disappears for each measure of housing price. Similarly, the sign on each coefficient switches from positive to negative. Thus, although the OLS coefficients suggest an associative relationship between housing prices and the degree of single-family deviation from historical proportions, insignificance in the more advanced 2SLS regressions refuses a strong interpretation. If nothing else, these results agree with the notion that the underlying determinants of the Deviations Index are intrinsically ambiguous. Some cities may produce a high value through a policy shift towards exclusive single-family housing. Others may issue single-family permits in response to a high rate of housing growth, multifamily units included. Accordingly, the net observed effect is not statistically different from zero in the 2SLS model. Unfortunately, however, all of the 2SLS "initial conditions" estimates may suffer from slightly weak

[27] Malpezzi and Green (1996) also provide evidence that filtering still works at the bottom of the market.

[28] Figure 1 presents a frequency distribution for the Deviations Index. The estimates predict a much greater effect for the multitude of cities showing a Deviations Index much higher than one standard deviation.

instruments.[29]

VI. Summary of Findings and Conclusion

When considered as a whole, the regressions presented above paint an interesting portrait of the static and dynamic determinants of housing prices in California. While biases plague many of the results, the several assertions are fairly reliable. The following characterizations summarize common themes in the regression interpretations.

First, income and wealth variables affect housing prices in the cross-sectional models. Median household income and education, a proxy for human capital, consistently reject the null hypothesis. Both present and potential future earnings tend to increase housing prices with a greater effect on values than on rents.

Second, racial and ethnic compositions have a clear association with both levels of and changes in housing prices. Again, values show a greater sensitivity to racial sorting. The levels of 1990 prices are higher, on average, in cities with a higher proportion of Hispanic and Asian residents. However, over the course of the 1990s, a net growth in the proportional composition of Black, Hispanic, or Asian residents likely dampens housing price inflation. As a whole, the outcomes offer strong support for the association between housing prices and racial sorting.

Third, the extent to which a municipality's land use regime is characterized as hospitable to growth varies negatively with price in the cross-section model. All else equal, cities with fewer political barriers to housing construction tend to have lower

[29] Specifically, joint significance tests on the instrumental variables yield low F-statistics.

housing prices. While the measure of local land use exclusivity performs less consistently, the results generally agree with the established body of literature and suggest that more restrictive policy increases housing price. Simply put, local regulatory policy, as characterized by the static policy survey variables, likely influences housing prices.

Fourth, the price effects of the single-family permits indices deliver fewer clear interpretations. With a few notable exceptions, OLS coefficients are generally insignificant. Even with the 2SLS method, the "before and after" regressions fail to reject the null hypothesis for the permits variables. On the other hand, the 2SLS "initial conditions" regressions, in which the control variables are more likely to be exogenously determined, suggest a significantly negative price effect of the proportion of new housing permits allocated to single-family construction. The evidence implies that the Proportions Index signals a net supply increase with the most pronounced effects on owner-occupied housing values.

Returning to the central research question, what does the data suggest about land use policy's impact on housing prices? In the 1990 cross section models, the policy survey variables corroborate much of the existing literature. Malpezzi (1996) suggests that a dynamic model accounting for changes as well as levels of prices is the next logical step (p. 237). With mixed results, such a dynamic model is constructed within this paper. Unfortunately, in this case, data limitations and the built-in ambiguities of the available change variables hinder the endeavor. In the dynamic models, the permits variables provide no evidence for the alleged price inflationary effect of a regulatory bias toward single-family permits.[30]

On the other hand, the "before and after" regressions offer stronger evidence for the association between racial and ethnic change and housing prices. To an extent, Quigley, Raphael, and Rosenthal (2002), in finding an association between the permits variables and racial outcomes, provide a basis to assume that density biases and demographic compositions are closely linked. If nothing else, the "before and after" models suggest that sorting by race has a more pronounced effect on prices than sorting by density.

The simpler, less ambiguous "initial conditions" model regressions, for which the 1990 variables provide convincingly exogenous controls for housing price changes, indicate that the Proportions Index signifies a supply increase for each quartile and tenure type. Where public policy favors low-density development (i.e. high levels of the Proportions Index) homeowners report lower values. Evidently, these lower values filter through the rental market. In comparison with the "before and after" model, the "initial conditions" model arguably provides a structurally weaker causal interpretation but avoids potential sources of additional endogeneity bias and, therefore, provides less ambiguous associative interpretations.

All told, the primary advances of this exercise are twofold. First, the "before and after" model introduces a simple, generalized, dynamic housing price model that accounts for changes in land use policy and housing prices. In a similar fashion, future research on the causal effect of land use policy should attempt to construct a more

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[30] At least, the effects of single-family bias are insignificant or negative in the strongest 2SLS models.

sophisticated, less ambiguous dynamic model. Second, the first stage regressions extend Quigley, Raphael, and Rosenthal's (2002) attempts to address the endogeneity of land use policy as depicted by the proportional bias toward single-family permits. The voting results, in concert with the policy survey measures, strengthen the set of potential instruments and thereby enhance 2SLS reliability. Since endogenously determined regulatory variables likely disturb many models that address land use and housing prices, like-minded researchers should consider similar methods. Aside from these developments, the findings concur with much of the conventional literature on land use regulation.

APPENDIX A: TRUNCATED REGRESSIONS

As mentioned earlier, each census employs a threshold above which all values or rents are reduced to the maximum value or rent.[1] This maximum threshold changes between 1990 and 2000. Thus, to address the problems arising when comparing housing prices across censuses, the results presented in the body of the paper rely on samples that replace missing price changes with sample averages. As an alternate approach, the following regressions simply drop observations for which the 1990 value or rent is above the maximum threshold recorded by the census. The previous results suffer Error-in-variables bias; the following suffer truncation bias. However, significance levels are generally similar in both sets of results.

The Cross-Section Regressions

The OLS regressions for the 1990 cross-sectional model in Table A1 rely on a truncated sample. Though the highest priced cities are dropped, the results generally resemble the previous versions (Table 1) and have similar interpretations. Only the median rent estimates fail to reject the null hypothesis for racial and ethnic proportions. Racial compositions have a greater effect on values than on rents. Black proportions, though insignificant at conventional levels, consistently exhibit a negative effect on price. Since the "pro-growth" and "exclusivity" indices characterize the local regulatory environment circa 1990, these variables may be added to the cross-section model as supply constraints. In Table A2 (like in Table 4), the "pro-growth" and "exclusivity"

[1] In the 1990 census, the maximum threshold value is 500,001 dollars; the maximum threshold rent is 1001 dollars. In the 2000 census, the maximum threshold value is 1,000,001 dollars; the maximum threshold rent is 2001 dollars

variables are added to the regressions. The pro-growth index categorically shows a highly significant effect while the exclusivity index, with the exception of first quartile rents, fails to reject the null. These results agree with the assumption that “pro-growth” municipalities provide fewer barriers to housing construction regardless of income and quality levels. Conversely, the exclusivity measure, though only significantly for first quartile rents, contradicts expectations. Specifically, exclusive policies that likely skew growth towards high-income residents and restrict housing construction in general should increase prices (all else equal), but the opposite effect is observed. Still, as in Table 4, the effect is small and insignificant in most of the regressions.

OLS “Before and After” Regressions

In the collection of OLS results presented in Table A3, the change in natural log of price is regressed on a single permits variable with no other covariates. While each regression fails to reject the null hypothesis for the Proportions Index, the Deviations Index is consistently positive at the five percent level of significance. These coefficients generally resemble the results based on the other sample (Table 5).

The next set of tables depict OLS regressions of housing prices on the “before and after” equation and the permits indices. Table A4 shows the Proportions Index case, and Table A5 shows the Deviations Index case. As in the single regression case, for each housing price measure, the Proportions Index estimates are not significant. The Deviations Index rejects the null hypothesis only for third quartile rents. Additionally, the income, human capital, and racial variables have coefficients and interpretations similar to those of the previous sample presented in Table 6.[2]

2SLS "Before and After" Regressions

Table A6 presents first stage estimates for the samples that drop observations for which the 1990 third quartile values exceed the maximum reported threshold.[3] For the 2SLS models presented in Table A7, both indices fails to reject the null hypothesis in each regression.[4] In the multiple regression models, the Deviations Index is significantly negative at the five percent level in the first quartile value regression and at the ten percent level in the first quartile rent regression. Thus, with respect to housing permits, a deviation from historical proportions of single-family units, holding all covariates equal, tends to slow value increases at the low-end of the market.

OLS and 2SLS "Initial Conditions" Regressions

Table A8 presents first stage estimates for the "initial conditions" model. For both permits index models, the joint test F-statistics fail to confidently rule out a weak instruments bias. In the Proportions Index models (Table A9), OLS and 2SLS coefficients are significantly negative for each quartile of rent change. Only significant in the OLS estimates, the effect of a single-family permits deviation from expectations reflects a positive effect on price growth. When applied to 2SLS, significance disappears

[2] These interpretations are discussed on pages 24-26.
[3] For each of the five other dependent variables, sample sizes and F-statistics are larger. The third quartile value first stages offer the weakest performance and, therefore, present a limiting case in Table A6.
[4] The single-variable regressions, however, reject the null hypothesis on the Proportions Index at the ten percent level for first quartile and median rents. The negative coefficients are generally consistent with the other models.

for each measure of housing price.

Comparison of Results

On the whole, the truncated sample produces regression results that closely resemble those produced by the sample that replaces missing prices with sample averages. While both approaches introduce bias, their consistency bolsters confidence in the conclusions presented in the body of this paper.

APPENDIX B: EXPANSION OF RACIAL AND ETHNIC TRENDS

In the "before and after" model regressions, as racial and ethnic composition changes, so does housing price. To extend this assertion, the single-variable OLS models presented in Appendix B utilize another set of variables developed by Quigley, Raphael, and Rosenthal (2002). Specifically, these variables are defined analogously to the Proportions Index and Deviations Index. The authors use these racial and ethnic indices to test whether demographic change in California municipalities responds to a single-family policy bias as measured by the corresponding permits indices. With regards to housing prices, these constructions allow a test of the association between local price change and the degree to which growth in a given municipality is skewed towards a certain racial or ethnic group.[1] The overly simplified single-variable regressions, though non-representative of any causal framework, further describe the correlations between housing price change and racial change. Moreover, as summarized in Table B1, different races and ethnicities, on average, demand different quantities of renter- and owner-occupied housing. This suggests that different races might show different effects for each of the six measures of housing price.

Proportion of Total Population Growth Attributable to a Given Race or Ethnicity

The proportional growth of a given racial/ethnic category *j* in a given city *i* equals the change in category *j* as a ratio of total city population growth:[2]

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[1] To address the data issues arising from the census' maximum threshold of recorded values, the following regressions use samples that estimate price changes for municipalities that exceed the threshold (as employed in the body of this paper) and not the truncation method (as pursued in Appendix A).

[2] For more detail on the construction of this variable, see Quigley, Raphael, and Rosenthal (2002).

$Proportional\ Change_j = \Delta population_{ji} / \Delta population_i$.

Table B2 presents single-variable OLS regressions of changes in the logs of first, second, and third quartile values and rents on the change in White, Black, Hispanic, and Asian compositions, respectively, as a proportion of total population growth. The data are taken from the 1990 and 2000 censuses.

The results show that owner occupied values increase with White population growth. In contrast, proportional growth in Black and Hispanic population dampens value change. Unlike the other racial variables, Asian proportional growth fails to reject the null hypothesis on values. Also, with the notable exception of the Hispanic variable, the racial variables typically produce insignificant effects on rent change.

Like the "before and after" model, these regressions provide evidence of the association between racial change and housing price change, especially with respect to owner-occupied housing values. Hispanic and Black population growth have a negative association with housing price. However, White population growth, which the "before and after" model does not consider, seems to have the opposite effect.

Proportional Deviation from Expectations for a Given Race or Ethnicity

The next racial/ethnic variable follows the construction of the Deviations Index.[3] It measures the proportion by which actual racial/ethnic population growth deviates from

[3] Again, Quigley, Raphael, and Rosenthal (2002) offer greater depth on this variable's construction.

expectations based on the initial inter-city racial distribution in 1990 and the total growth in this population between 1990 and 2000. The expected population change for category *j* in city *i* is defined:

$$expected\ \Delta population_{ji} = \Delta population_{j} * population_{ji} / population_{j},$$

where $population_j$ equals the total state population of racial/ethnic category *j*, and $population_{ji}$ equals the population of group *j* in city *i*. Accordingly, the proportionate deviation from expectations of population growth equals:

$$deviation_{ji} = (\Delta population_{ji} - expected\ \Delta population_{ji}) / expected\ \Delta population_{ji}.$$

Table B3 presents single-variable OLS regressions of changes in the logs of first, second, and third quartile values and rents on the proportionate deviations from expectations of White, Black, Hispanic, and Asian growth, respectively.

In each regression, the White and Asian variables reject the null hypothesis at the one percent level of significance. For these races, a deviation from expectations is associated with a greater increase in housing values and rents in each quartile. Hispanic and Black deviations, however, tend to show a smaller, typically insignificant effect. Still, where the coefficients show significance, they suggest that a positive deviation from expected growth in Hispanic or Black populations is associated with a lower housing price change.[4]

[4] Specifically, the Black median (column 5) and third quartile (7) and the Hispanic first quartile (4) regressions produce a significantly negative coefficient.

The “before and after” regressions suggest that increasing ratios of Black, Hispanic, and Asian populations tends to decrease city housing prices, all else equal. Relying on simple single-variable regressions, this appendix extends and supplements the assertion that housing price change is not race-neutral. A high proportion of total population growth attributable to Hispanic or Black residents corresponds with lower housing prices. On the other hand, a positive deviation from expected growth in Asian populations seems to suggest more dramatic price inflation. White population growth, both as a proportion of total growth and as a proportionate deviation, is associated with higher housing price change. Obviously, these estimates do not control for other price determinants and certainly contain omitted variables biases. While the results provide no strong causal interpretation, the regressions clearly reinforce the claim that housing prices bear associations with the geographic sorting of demographic change.

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Table 1: Summary of California-Level Housing Prices, 1990 and 2000

	Owner-occupied House Values			*Contract Rents*		
	1st Quartile	Median	3rd Quartile	1st quartile	Median	3rd Quartile
1990 Census	127,100	195,500	294,800	414	561	728
2000 Census	140,900	211,500	346,000	503	677	917
Change (%)	11%	8%	17%	21%	20%	26%

Source: 1990 U.S. Census Summary File 1 and 2000 U.S. Census Summary File 3, state-level tabulations.

Table 2: Short Descriptions and Correlation Coefficients for Selected 1990s California State Proposition Outcomes

Prop . #	*Summary*	*Average Ratio of "Yes" votes*	**Correlation Coefficients** (342 observations)										
			Prop. 155	*Prop. 156*	*Prop. 160*	*Prop. 164*	*Prop. 167*	*Prop. 168*	*Prop. 199*	*Pro-growth*	*Exclusivity*	*Dev. Index*	*Prop. Index*
155	Bond for construction or improvement of public schools.	0.5194	1.000							-0.1194	0.0175	-0.1951	-0.2822
156	Bond to provide funds for inter- and intra-city rail transit.	0.4674	0.832	1.000						-0.2209	0.0407	-0.3105	-0.2973
160	Property tax exemption for unmarried widows of military deaths.	0.5164	0.315	0.259	1.000					0.1495	-0.0604	-0.1810	-0.3246
164	Establishes congressional Term limits	0.6378	-0.696	-0.473	-0.200	1.000							
167	Sales tax decreases, high-end income tax increases, and renter tax credits.	0.3884	0.520	0.416	0.359	-0.490	1.000			0.0772	0.1456	-0.0594	-0.3017
168	Facilitates bypassing voter approval for low-rent housing projects.	0.3959	0.829	0.748	0.350	-0.653	0.606	1.000		-0.1138	0.0146	-0.2614	-0.3756
199	Limits mobile home rent control, provides limited private sector, low-income rent subsidy.	0.3999	-0.135	-0.105	-0.191	0.258	-0.635	-0.252	1.000	-0.1393	-0.1823	0.0189	0.0775

Source: 1992, 1993, and 1996 voter pamphlets.

Table 3: OLS Regressions of the Simple Cross-Sectional Reduced Form Housing Price Models, 1990

	Dependent Variables					
	Natural Log of Value, 1990			Natural Log of Rent, 1990		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Natural Log of Population, 1990	0.023 (0.013)	0.023 (0.015)	0.030 (0.016)	0.056 (0.013) **	0.031 (0.010) **	0.026 (0.007) **
Natural Log of Median Income, 1990	0.663 (0.064) **	0.533 (0.072) **	0.407 (0.079) **	0.595 (0.058) **	0.543 (0.046) **	0.424 (0.043) **
Proportion College Graduate, 1990	1.373 (0.156) **	1.376 (0.165) **	1.215 (0.172) **	0.199 (0.120)	0.292 (0.095) **	0.164 (0.091)
Proportion High School Dropout, 1990	-1.216 (0.324) **	-1.546 (0.365) **	-1.859 (0.411) **	-0.621 (0.286) *	-0.558 (0.236) *	-0.928 (0.181) **
MSA Dummy	0.253 (0.052) **	0.288 (0.053) **	0.302 (0.056) **	0.186 (0.042) **	0.163 (0.034) **	0.179 (0.032) **
Proportion Black, 1990	-0.488 (0.245) *	-0.380 (0.234)	-0.386 (0.242)	-0.154 (0.164)	-0.034 (0.111)	-0.007 (0.102)
Proportion Hispanic, 1990	0.796 (0.179) **	0.911 (0.202) **	0.961 (0.226) **	0.234 (0.157)	0.269 (0.130) *	0.422 (0.098) **
Proportion Asian, 1990	0.440 (0.169) **	0.426 (0.172) *	0.400 (0.172)	0.306 (0.106) **	0.235 (0.085) **	0.222 (0.081) **
F-stat for jointly excluding racial proportions (p-value)	10.69 (0.000)	9.82 (0.000)	8.30 (0.000)	3.33 (0.020)	3.57 (0.014)	10.58 (0.000)
R^2	0.83	0.81	0.78	0.81	0.83	0.80

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
Estimates are based on a sample of 456 California municipalities.
*p<0.05; **p<0.01

Table 4: OLS Regressions of the Simple Cross-Sectional Reduced Form Housing Price Models, 1990

	Dependent Variables					
	Natural Log of Value, 1990			Natural Log of Rent, 1990		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Exclusivity	0.005 (0.003)	0.006 (0.002) *	0.007 (0.003) *	-0.005 (0.002) **	-0.002 (0.001)	-0.001 (0.001)
Pro-Growth Index	-0.035 (0.008) **	-0.036 (0.008) **	-0.035 (0.008) **	-0.015 (0.005) **	-0.015 (0.004) **	-0.016 (0.004) **
F-stat for jointly excluding regulatory variables (p-value)	9.37 (0.000)	10.28 (0.000)	9.94 (0.000)	12.95 (0.000)	10.77 (0.000)	7.86 (0.001)
R^2	0.81	0.79	0.76	0.79	0.82	0.77

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term and the covariates listed in Table 3. Estimates are based on a sample of 351 California cities.
*p<0.05; **p<0.01

Table 5: OLS Single Regressions of Price Change on the Permits Variables

	Dependent Variables					
	Natural Log of Value, 1990			Natural Log of Rent, 1990		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index	0.039	0.043	0.044	0.006	-0.022	-0.024
	(0.035)	(0.036)	(0.037)	(0.032)	(0.028)	(0.027)
R_2	0.003	0.003	0.003	0.000	0.003	0.003
Deviations Index	0.013	0.011	0.010	0.008	0.009	0.011
	(0.004)	(0.004)	(0.004)	(0.004)	(0.003)	(0.003)
	**	**	*	*	**	**
R^2	0.016	0.013	0.010	0.008	0.016	0.030

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
Estimates are based on samples of 454 and 450 California municipalities, respectively.
*p<0.05; **p<0.01

Table 6: OLS Regressions of the Proportions Index on the "Before and After" Reduced Form Housing Price Models

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index	-0.002 (0.029)	0.010 (0.030)	0.017 (0.031)	-0.010 (0.029)	-0.029 (0.022)	-0.036 (0.021)
Δ Natural Log of Population	-0.026 (0.031)	-0.043 (0.040)	-0.043 (0.050)	0.006 (0.034)	-0.009 (0.022)	-0.005 (0.021)
Δ Natural Log of Median Income	0.788 (0.086) **	0.806 (0.094) **	0.775 (0.128) **	0.569 (0.118) **	0.482 (0.072) **	0.445 (0.057) **
Δ Proportion College Graduate	-0.172 (0.196)	0.118 (0.202)	0.100 (0.245)	-0.250 (0.199)	0.213 (0.155)	0.264 (0.159)
Δ Proportion High School Dropout	-0.105 (0.204)	0.098 (0.203)	0.280 (0.282)	0.283 (0.200)	0.182 (0.142)	0.067 (0.130)
Δ Proportion Black	-1.496 (0.310) **	-1.082 (0.344) **	-0.892 (0.377) *	-0.422 (0.266)	-0.356 (0.129)	-0.291 (0.223)
Δ Proportion Hispanic	-0.540 (0.175) **	-0.669 (0.188) **	-0.755 (0.247) **	-0.866 (0.187) **	-0.503 (0.129) **	-0.256 (0.120) *
Δ Proportion Asian	-1.019 (0.186) **	-1.020 (0.190) **	-0.983 (0.207) **	-0.321 (0.150) *	-0.168 (0.137)	-0.130 (0.120)
F-stat for jointly excluding racial proportions (p-value)	19.24 (0.000)	14.39 (0.000)	9.90 (0.000)	7.15 (0.000)	5.88 (0.001)	2.27 (0.080)
R^2	0.37	0.37	0.31	0.25	0.30	0.27

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term. Estimates are based on a sample of 454 California municipalities.
*p<0.05, **p<0.01

Table 7: OLS Regressions of the Deviations Index on the "Before and After" Reduced Form Housing Price Models

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Deviations Index	0.005 (0.005)	0.003 (0.004)	-0.002 (0.005)	0.004 (0.003)	0.006 (0.003)	0.008 (0.003) *
Δ Natural Log of Population	-0.039 (0.038)	0.010 (0.053)	0.051 (0.071)	-0.028 (0.033)	-0.000 (0.039)	0.000 (0.038)
Δ Natural Log of Median Income	0.727 (0.088) **	0.625 (0.158) **	0.514 (0.223) *	0.533 (0.105) **	0.362 (0.107) **	0.305 (0.107) **
Δ Proportion College Graduate	-0.200 (0.199)	0.026 (0.214)	0.019 (0.268)	-0.233 (0.199)	0.194 (0.161)	0.220 (0.169)
Δ Proportion High School Dropout	-0.158 (0.201)	0.000 (0.209)	0.158 (0.296)	0.219 (0.194)	0.134 (0.146)	0.036 (0.137)
Δ Proportion Black	-1.525 (0.323) **	-1.347 (0.390) **	-1.315 (0.457) **	-0.395 (0.262)	-0.506 (0.276)	-0.447 (0.285)
Δ Proportion Hispanic	-0.479 (0.167) **	-0.749 (0.200) **	-0.907 (0.279) **	-0.708 (0.175) **	-0.506 (0.137) **	-0.318 (0.135) *
Δ Proportion Asian	-0.919 (0.182) **	-0.922 (0.196) **	-0.886 (0.211) **	-0.215 (0.149)	-0.071 (0.143)	-0.010 (0.124)
F-stat for jointly excluding racial proportions (p-value)	18.65 (0.000)	15.55 (0.000)	11.53 (0.000)	5.62 (0.001)	5.05 (0.002)	2.41 (0.066)
R^2	0.37	0.32	0.24	0.24	0.25	0.22

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term. Estimates are based on a sample of 450 California municipalities.
*p<0.05, **p<0.01

Table 8: First Stage Relationships between Growth Control Measures, Ballot-Box Outcomes, and the Permits Indices ("Before and After" Model)

	Endogenous Explanatory Variable			
	Proportions Index		Deviations Index	
	No covariates	All other covariates	No covariates	All other covariates
Pro-Growth	-0.018 (0.006) **	-0.019 (0.006) **	0.197 (0.051) **	0.065 (0.033) *
Exclusivity	-0.003 (0.002)	-0.004 (0.002) *	-	-
Proposition 156	-0.168 (0.175)	-0.052 (0.186)	-	-
Proposition 160	-0.837 (0.233) **	-0.779 (0.257) **	-6.613 (1.966) **	-3.350 (1.165) **
Proposition 164	-	-	3.637 (1.427) *	3.140 (0.938) **
Proposition 168	-0.745 (0.201) **	-0.885 (0.204) **	-	-
Proposition 199	-0.223 (0.154)	-0.200 (0.148)	-	-
F-stat for jointly excluding instruments (p-value)	15.88 (0.000)	11.90 (0.000)	10.14 (0.000)	8.82 (0.000)
Sample size	347	347	342	342

Standard errors are in parenthesis. All models include a constant term and the other included exogenous covariates listed in Table 6.
*p<0.05, **p<0.01

Table 9: 2SLS Regressions of Housing Prices on the Permits Indices ("Before and After" Model)

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index - No covariates	-0.125 (0.096)	-0.100 (0.098)	-0.045 (0.096)	-0.118 (0.079)	-0.104 (0.079)	-0.059 (0.073)
Proportions Index - All other covariates	-0.079 (0.083)	-0.038 (0.084)	-0.013 (0.082)	-0.084 (0.078)	-0.068 (0.071)	-0.043 (0.062)
Deviations Index - No covariates	-0.031 (0.019)	-0.018 (0.018)	-0.000 (0.017)	-0.024 (0.015)	-0.013 (0.013)	-0.000 (0.012)
Deviations Index - All other covariates	-0.056 (0.030)	-0.033 (0.029)	-0.000 (0.028)	-0.044 (0.026)	-0.026 (0.023)	0.009 (0.021)

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term and the covariates listed in Tables 6 and 7. The Proportions Index models are estimated on a sample of 347 cities, and the Deviations Index models are estimated on a sample of 342 cities.
*p<0.05, **p<0.01

Table 10: Selected First Stage Relationships between Growth Control Measures, Ballot-Box Outcomes, and the Permits Indices ("Initial Conditions" Model)

	Endogenous Explanatory Variable	
	Proportions Index	Deviations Index
Pro-Growth	-0.017 (0.006) **	0.113 (0.047) *
Proposition 155	-	9.519 (2.297) **
Proposition 160	-	-8.785 (2.351) **
Proposition 164	-	6.921 (2.196) **
Proposition 168	-0.765 (0.182) **	-
Proposition 199	-0.472 (0.175) **	-
F-stat for jointly excluding instruments (p-value)	9.29 (0.000)	7.23 (0.000)
Sample size	347	343

Standard errors are in parenthesis. All models include a constant term. The regressions appearing here are selected because F-statistics for jointly excluding the instruments are lowest, but all other regressions perform similarly. The presented Proportions Index results correspond to the third quartile rent model, while the Deviations Index results correspond to the first quartile value model.
*p<0.05, **p<0.01

Table 11: Abbreviated OLS and 2SLS Regressions of Housing Prices on the Permits Indices ("Initial Conditions" Model)

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index - OLS	-0.038 (0.046)	-0.040 (0.048)	-0.039 (0.047)	-0.058 (0.037)	-0.074 (0.037) *	-0.067 (0.035)
R^2	0.28	0.28	0.25	0.18	0.14	0.19
Proportions Index - 2SLS	-0.485 (0.199) *	-0.543 (0.216) *	-0.497 (0.211) *	-0.459 (0.156) **	-0.463 (0.148) **	-0.408 (0.140) **
F-stat for jointly excluding instruments (p-value)	9.93 (0.000)	9.58 (0.000)	9.46 (0.000)	11.17 (0.000)	10.08 (0.000)	9.29 (0.000)
Deviations Index - OLS	0.013 (0.004) **	0.013 (0.004) **	0.012 (0.005) **	0.010 (0.004) **	0.012 (0.003) **	0.014 (0.003) **
R^2	0.30	0.30	0.25	0.20	0.15	0.14
Deviations Index - 2SLS	0.007 (0.023)	-0.001 (0.022)	-0.004 (0.020)	-0.024 (0.022)	-0.019 (0.018)	-0.009 (0.017)
F-stat for jointly excluding instruments (p-value)	7.23 (0.000)	7.60 (0.000)	8.11 (0.000)	8.13 (0.000)	7.84 (0.000)	7.45 (0.000)

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term, the covariates listed in Table 3 (except *MSA dummy*), and the 1990 natural log of housing price.
*p<0.05, **p<0.01

Appendix Table A1: OLS Regressions of the Simple Cross-Sectional Reduced Form Housing Price Models, 1990

	Dependent Variables					
	Natural Log of Value, 1990			Natural Log of Rent, 1990		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Natural Log of Population, 1990	0.018 (0.013)	0.001 (0.015)	0.014 (0.018)	0.057 (0.014) **	0.026 (0.010) **	0.012 (0.007)
Natural Log of Median Income, 1990	0.715 (0.071) **	0.683 (0.083) **	0.407 (0.079) **	0.585 (0.066) **	0.618 (0.051) **	0.641 (0.046) **
Proportion College Graduate, 1990	1.423 (.168) **	1.576 (0.197) **	1.656 (0.252) **	0.187 (0.121)	0.390 (0.104) **	0.507 (0.123) **
Proportion High School Dropout, 1990	-1.092 (0.344) **	-1.117 (0.418) **	-1.135 (0.504) *	-0.645 (0.292) *	-0.342 (0.241)	-0.308 (0.190)
MSA Dummy	0.243 (0.053) **	0.263 (0.055) **	0.259 (0.059) **	0.187 (0.042) **	0.150 (0.035) **	0.179 (0.032) **
Proportion Black, 1990	-0.454 (0.245)	-0.308 (0.234)	-0.272 (0.234)	-0.158 (0.165)	-0.003 (0.108)	-0.066 (0.094)
Proportion Hispanic, 1990	0.762 (0.186) **	0.775 (0.222) **	0.742 (0.263) **	0.241 (0.159)	0.198 (0.208) *	0.231 (0.097) *
Proportion Asian, 1990	0.436 (0.177) *	0.406 (0.186) *	0.292 (0.198)	0.307 (0.108) **	0.208 (0.087) *	0.230 (0.094) *
F-stat for jointly excluding racial proportions (p-value)	9.08 (0.000)	6.03 (0.001)	3.50 (0.012)	3.29 (0.021)	2.40 (0.067)	4.56 (0.004)
R^2	0.80	0.81	0.75	0.82	0.83	0.79
Sample size	441	429	411	436	456	396

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05; **p<0.01

Appendix Table A2: OLS Regressions of the Simple Cross-Sectional Reduced Form Housing Price Models, 1990

	Dependent Variables					
	Natural Log of Value, 1990			Natural Log of Rent, 1990		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Natural Log of Population, 1990	0.016 (0.016)	0.004 (0.019)	0.008 (0.022)	0.081 (0.016) **	0.040 (0.011) **	0.018 (0.008) *
Natural Log of Median Income, 1990	0.751 (0.079) **	0.723 (0.097) **	0.645 (0.116) **	0.576 (0.072) **	0.619 (0.054) **	0.642 (0.046) **
Proportion College Graduate, 1990	1.202 (0.203) **	1.333 (0.243) **	1.333 (0.315) **	0.066 (0.127)	0.304 (0.105) **	0.434 (0.120) **
Proportion High School Dropout, 1990	-1.141 (0.423) **	-1.185 (0.540) *	-1.312 (0.659) *	-0.787 (0.368) *	-0.360 (0.311)	-0.371 (0.235)
MSA Dummy	1.251 (0.035) **	1.107 (0.036) **	1.110 (0.038) **	0.887 (0.021) **	0.762 (0.017) **	0.734 (0.017) **
Proportion Black, 1990	-0.213 (0.227)	-0.053 (.219)	0.024 (0.352)	-0.098 (0.151)	0.046 (0.100)	0.137 (0.084)
Proportion Hispanic, 1990	0.841 (0.236) **	0.860 (0.294) **	0.869 (0.462) *	0.330 (0.208)	0.220 (0.174)	0.281 (0.131) *
Proportion Asian, 1990	0.561 (0.206) **	0.538 (0.215) *	0.462 (0.222) *	0.282 (0.119) *	0.216 (0.095) *	0.246 (0.097) *
Exclusivity	0.004 (0.003)	0.004 (0.003)	0.007 (0.004)	-0.005 (0.002) **	-0.002 (0.001)	-0.001 (0.002)
Pro-Growth Index	-0.035 (0.008) **	-0.034 (0.008) **	-0.035 (0.008) **	-0.016 (0.005) **	-0.015 (0.004) **	-0.013 (0.004) **
F-stat for jointly excluding regulatory variables (p-value)	9.03 (0.000)	8.16 (0.000)	8.91 (0.000)	13.14 (0.000)	10.26 (0.000)	6.97 (0.001)
R^2	0.78	0.75	0.71	0.78	0.80	0.76
Sample size	337	326	308	345	333	298

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05; **p<.01

Appendix Table A3: OLS Regressions of Price Change on the Permits Variables

	Dependant Variables					
	Natural Log of Value, 1990			Natural Log of Rent, 1990		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index	0.040 (0.036)	0.046 (0.038)	-0.050 (0.042)	0.006 (0.033)	-0.023 (0.029)	-0.028 (0.032)
R^2	0.003	0.003	0.004	0.000	0.002	.003
Sample size	439	427	409	447	434	394
Deviations Index	0.013 (0.004) **	0.011 (0.004) **	0.011 (0.005) *	0.008 (0.004) *	0.009 (0.003) **	0.012 (0.003) **
R^2	0.016	0.013	0.012	0.008	0.017	0.032
Sample size	435	423	405	443	430	390

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05; **p<0.01

Appendix Table A4: OLS Regressions of the Proportions Index on the "Before and After" Reduced Form Housing Price Models

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index	0.002 (0.030)	0.014 (0.031)	0.020 (0.034)	-0.011 (0.030)	-0.026 (0.023)	-0.029 (0.024)
Δ Natural Log of Population	-0.033 (0.030)	-0.060 (0.038)	-0.076 (0.047)	0.006 (0.034)	-0.015 (0.023)	-0.031 (0.021)
Δ Natural Log of Median Income	0.848 (0.089) **	0.923 (0.096) **	0.969 (0.138) **	0.577 (0.121) **	0.514 (0.078) **	0.576 (0.069) **
Δ Proportion College Graduate	-0.206 (0.200)	0.114 (0.206)	0.326 (0.246)	-0.244 (0.204)	0.237 (0.162)	0.369 (0.177) *
Δ Proportion High School Dropout	-0.030 (0.210)	0.239 (0.207)	0.532 (0.288)	0.297 (0.204)	0.226 (0.145)	0.210 (0.139)
Δ Proportion Black	-1.436 (0.305) **	-0.952 (0.336) **	-0.688 (0.374)	-0.411 (0.267)	-0.311 (0.228)	-0.141 (0.232)
Δ Proportion Hispanic	-0.596 (0.184) **	-0.763 (0.187) **	-0.899 (0.244) **	-0.873 (0.192) **	-0.524 (0.134) **	-0.310 (0.132) *
Δ Proportion Asian	-1.074 (0.184) **	-1.107 (0.194) **	-1.274 (0.215) **	-0.344 (0.153) *	-0.181 (0.144)	-0.163 (0.146)
F-stat for jointly excluding racial proportions (p-value)	19.97 (0.000)	15.28 (0.000)	12.95 (0.000)	6.93 (0.000)	5.58 (0.001)	2.07 (0.104)
R^2	0.38	0.40	0.38	0.25	0.31	0.33
Sample size	439	427	409	447	434	394

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05, **p<0.01

Appendix Table A5: OLS Regressions of the Deviations Index on the "Before and After" Reduced Form Housing Price Models

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Deviations Index	0.005 (0.005)	0.001 (0.005)	-0.002 (0.005)	0.004 (0.003)	0.005 0.003	0.007 (0.003) *
Δ Natural Log of Population	-0.041 (0.039)	0.007 (0.058)	0.052 (0.082)	-0.028 (0.033)	-0.001 0.041	-0.006 (0.048)
Δ Natural Log of Median Income	0.773 (0.096) **	0.699 (0.187) **	0.609 (0.287) *	0.538 (0.108) **	0.379 0.119 **	0.368 (0.152) *
Δ Proportion College Graduate	-0.239 (0.203)	0.016 (0.219)	0.141 (0.287)	-0.228 (0.203)	0.218 0.165	0.303 (0.193)
Δ Proportion High School Dropout	-0.101 (0.206)	0.098 (0.217)	0.323 (0.316)	0.231 (0.197)	0.169 0.149	0.129 (0.150)
Δ Proportion Black	-1.485 (0.182) **	-1.276 (0.406) **	-1.252 (0.509) *	-0.388 (0.263)	-0.484 0.285	-0.401 (0.332)
Δ Proportion Hispanic	-0.531 (0.172) **	-0.835 (0.202) **	-1.05 (0.285) **	-0.715 (0.153) **	-0.534 0.143 **	-0.382 (0.152) *
Δ Proportion Asian	-0.971 (0.182) **	-1.01 (0.204) **	-1.162 (0.241) **	-0.234 (0.153)	-0.087 0.151	-0.010 (0.168)
F-stat for jointly excluding racial proportions (p-value)	19.32 (0.000)	16.22 (0.000)	13.87 (0.000)	5.46 (0.001)	4.91 (0.002)	2.25 (0.082)
R^2	0.38	0.34	0.27	0.24	0.26	0.26
Sample size	435	423	405	443	430	390

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05, **p<0.01

Appendix Table A6: First Stage Relationships between Growth Control Measures, Ballot-Box Outcomes, and the Permits Indices ("Before and After" Model)

	Endogenous Explanatory Variable			
	Proportions Index		Deviations Index	
	No covariates	All other covariates	No covariates	All other covariates
Pro-Growth	-0.012 (0.006) *	-0.013 (0.006) *	0.167 (0.052) **	0.061 (0.033)
Exclusivity	-0.001 (0.002)	-0.001 (0.002)	-	-
Proposition 156	-0.511 (0.189) **	-0.300 (0.212)	-	-
Proposition 160	-0.462 (0.246)	-0.323 (0.269)	-8.430 (2.045) **	-2.964 (1.304) *
Proposition 164	-	-	3.384 (1.448) *	3.017 (0.974) **
Proposition 168	-0.745 (0.201) **	-0.835 (0.212)	-	-
Proposition 199	-0.635 (0.174) **	-0.602 (0.178)	-	-
F-stat for jointly excluding instruments (p-value)	16.71 (0.000)	11.90 (0.000)	11.68 (0.000)	6.75 (0.000)
Sample size	305	305	301	301

Standard errors are in parenthesis. All models include a constant term. The regressions are estimated on the sample of cities for which 1990 third quartile values are below the maximum reported value of 1,000,001 dollars.
*p<0.05, **p<0.01

Appendix Table A7: 2SLS Regressions of Housing Prices on the Permits Indices ("Before and After" Model)

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index - No covariates	-0.135 (0.102)	-0.140 (0.107)	-0.100 (0.109)	-0.135 (0.082)	-0.141 (0.082)	-0.116 (0.078)
F-stat for jointly excluding instruments	16.05	15.95	16.71	15.55	16.66	19.25
Proportions Index - All other covariates	-0.089 (0.091)	-0.075 (0.096)	-0.041 (0.101)	-0.102 (0.083)	-0.075 (0.077)	-0.065 (0.068)
F-stat for jointly excluding instruments	12.97	12.13	11.90	13.19	13.11	13.55
Sample Size	334	323	305	341	330	295
Deviations Index - No covariates	-0.026 (0.019)	-0.007 (0.018)	-0.017 (0.018)	-0.022 (0.015)	-0.009 (0.013)	0.008 (0.013)
F-stat for jointly excluding instruments	13.07	12.90	11.68	13.63	13.35	12.19
Deviations Index - All other covariates	-0.066 (0.031)*	-0.050 (0.031)	-0.026 (0.034)	-0.047 (0.027)	-0.032 (0.024)	-0.023 (0.025)
F-stat for jointly excluding instruments	9.10	8.40	6.75	9.20	8.64	6.81
Sample size	330	319	301	337	326	291

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term and the covariates listed in Tables A4 and A5.
*p<0.05, **p<0.01

Appendix Table A8: First Stage Relationships between Growth Control Measures, Ballot-Box Outcomes, and the Permits Indices ("Initial Conditions" Model)

	Endogenous Explanatory Variable	
	Proportions Index	Deviations Index
	All other covariates	All other covariates
Pro-Growth	-0.017 (0.006) **	0.134 (0.052) **
Proposition 155	-	10.466 (2.600) **
Proposition 160	-	-10.69 (2.698) **
Proposition 164	-	9.633 (2.522) **
Proposition 168	-0.499 (0.180) **	-
Proposition 199	-0.499 (0.180) **	-
F-stat for jointly excluding instruments (p-value)	7.57 (0.000)	8.39 (0.000)
Sample size	305	301

Standard errors are in parenthesis. All models include a constant term. The regressions are estimated on the sample of cities for which third quartile housing values are below the maximum reported value of 1,000,001 dollars.
*p<0.05, **p<0.01

Appendix Table A9: Abbreviated OLS and 2SLS Regressions of Housing Prices on the Permits Indices ("Initial Conditions" Model)

	Dependent Variables					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
Proportions Index - OLS	-0.034 (0.046)	-0.029 (0.048)	-0.018 (0.050)	-0.071 (0.031) *	-0.080 (0.029) **	-0.075 (0.035) *
R^2	0.29	0.31	0.28	0.29	0.28	0.25
Sample Size	439	427	409	447	434	394
Proportions Index - 2SLS	-0.452 (0.196) *	-0.493 (0.215) *	-0.468 (0.235) *	-0.440 (0.155) **	-0.445 (0.142) **	-0.305 (0.111) **
F-stat for jointly excluding instruments	9.45	8.92	7.57	10.81	10.17	10.30
Sample Size	334	323	305	341	330	295
Deviations Index - OLS	0.013 (0.004) **	0.012 (0.004) **	0.010 (0.005) *	0.005 (0.003)	0.008 (0.003) *	0.010 (0.003) **
R^2	0.31	0.32	0.29	0.28	0.25	0.25
Sample Size	435	423	405	443	430	390
Deviations Index - 2SLS	0.000 (0.023)	-0.009 (0.021)	-0.009 (-0.024)	-0.024 (0.023)	-0.014 (0.016)	-0.014 (0.015)
F-stat for jointly excluding instruments	7.81	8.37	8.39	8.21	8.43	8.25
Sample size	330	319	301	337	326	291

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term, the covariates listed in Tables A1 and A2 (except *MSA dummy*), and the 1990 natural log of housing price.
*p<0.05, **p<0.01

Appendix Table B1: Summary of California-Level Tenure by Race, 2000

Race/Ethnicity	*Total households*	*Owner-occupied*	*% Owning*	*Renter-occupied*	*% Renting*
White	7,756,027	4,856,237	63%	2,899,599	37%
Black[a]	777,973	302,518	39%	475,455	61%
Hispanic	2,564,765	1,121,940	44%	1,442,825	56%
Asian[a]	1,110,698	613,743	55%	496,955	45%

Source: 1990 U.S. Census Summary File 1 and 2000 U.S. Census Summary File 3, state-level tabulations.
[a] Here, Black and Asian population are defined as Black Alone and Asian Alone, respectively. However, the ownership rates are probably still highly correlated with the definitions of 2000 Asian and Black populations in the rest of the analysis (see page 15, footnote).

Appendix Table B2: OLS Regressions of Housing Price Change on Racial Change as a Proportion of Total Population Growth

Independent Variable: Proportional Growth[a]	*Dependent Variables*					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
White (N=441)	0.015 (0.006) **	0.019 (0.005) **	0.019 (0.005) **	0.000 (0.005)	0.004 (0.003)	0.004 (0.003)
Black (N=451)	-0.035 (0.013) **	-0.035 (0.013) **	-0.025 (0.013)	-0.011 (0.007)	-0.011 (0.006)	-0.013 (0.005) *
Hispanic (N=448)	-0.017 (0.008) *	-0.023 (0.008) **	-0.025 (0.008) **	-0.021 (0.006) **	-0.017 (0.004) **	-0.012 (0.005) **
Asian (N=450)	-0.008 (0.008)	-0.007 (0.007)	-0.005 (0.007)	0.015 (0.007) *	0.008 (0.005)	0.004 (0.004)

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05, **p<0.01.
[a]For each city, the racial/ethnic proportional growth are defined analogously to the Proportions Index:
Proportionate growth = Δ race or ethnicity / Δ total population.

Appendix Table B3: OLS Regressions of Housing Price Change on Racial Deviations from Expected Population Growth

Independent Variable: Proportional Deviation[a]	*Dependent Variables*					
	Δ Natural Log of Value			Δ Natural Log of Rent		
	1st Quartile (1)	Median (2)	3rd Quartile (3)	1st Quartile (4)	Median (5)	3rd Quartile (6)
White (N=450)	0.014 (0.003) **	0.017 (0.003) **	0.017 (0.003) **	0.010 (0.002) **	0.009 (0.002) **	0.010 (0.001) **
Black (N=413)	-0.000 (0.002)	-0.001 (0.002)	-0.001 (0.002)	0.000 (0.002)	-0.003 (0.001) *	-0.002 (0.001) *
Hispanic (N=453)	0.001 (0.007)	0.000 (0.007)	0.005 (0.007)	-0.018 (0.007) *	-0.010 (0.005)	0.001 (0.005)
Asian (N=449)	0.010 (0.004) *	0.013 (0.004) **	0.015 (0.005) **	0.015 (0.004) **	0.012 (0.003) **	0.012 (0.003) **

Heteroskedasticity-robust standard errors are in parenthesis. All models include a constant term.
*p<0.05, **p<0.01.
[a]The racial/ethnic deviations from historical proportions are defined analogously to the Deviations Index:

Proportionate deviation from expectations = $(\Delta\ pop_{ji} - \Delta\ expected\ pop_{ji}) / \Delta\ expected\ pop_{ji}$
(expected $population_{ji} = \Delta pop_j * pop_{ji} / pop_i$, j = given race or ethnicity, and i = given city).

Figure 1

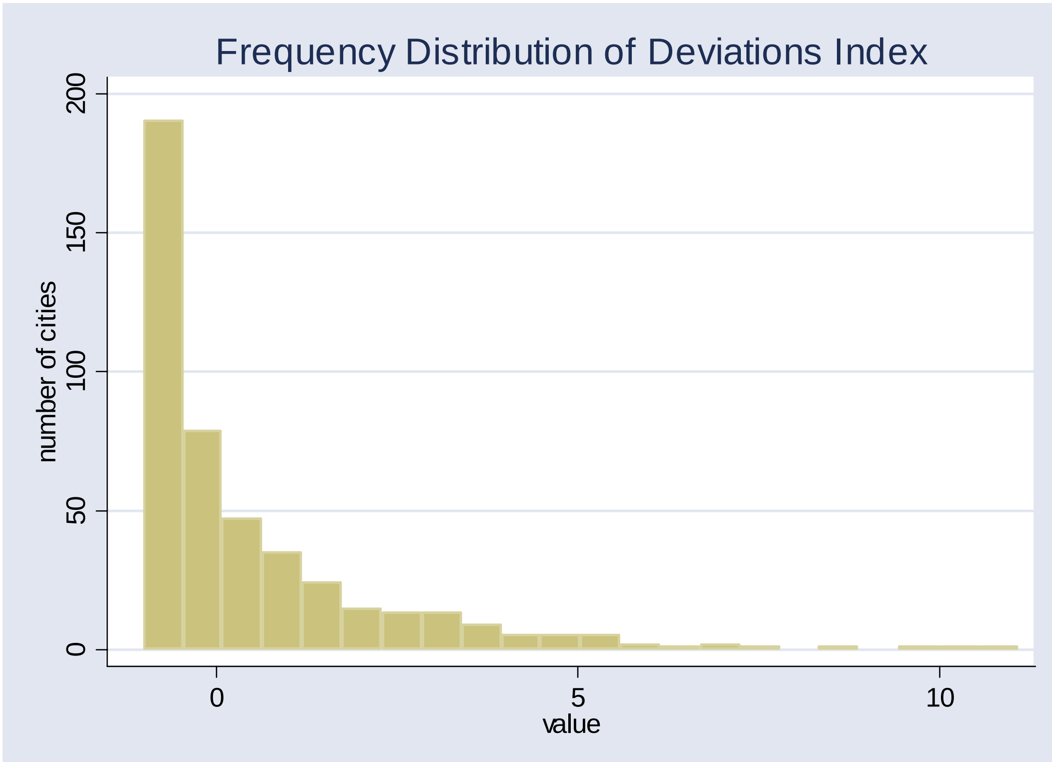

Figure 2