Journal of Urban Economics 48, 85–109 (2000)
doi:10.1006/juec.1999.2158, available online at http://www.idealibrary.com on IDEAL®

Residential Construction: Using the Urban Growth Model to Estimate Housing Supply[1]

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Received December 14, 1998; revised September 24, 1999

This article presents an empirical model of housing supply derived from urban growth theory. This approach describes new housing construction as a function of changes in house prices and costs rather than as a function of the levels of those variables, which previous studies have used. Empirical tests support this specification over the leading alternative models. Our estimates show that a 10% rise in real prices leads to an 0.8% increase in the housing stock, which is accomplished by a temporary 60% increase in the annual number of starts, spread over four quarters. © 2000 Academic Press

I. INTRODUCTION

Construction of new housing plays a critical role in the economy. Residential construction influences overall output directly, construction and manufacturing employment rises with housing starts, and indirectly through the multiplier effect, as new home buyers tend to purchase other consumer durables when they buy their house. Housing construction typically leads recessions and recoveries.[2] Changes in new housing supply affect the price of existing housing units, with implications for the wealth

[1]The authors thank Dennis Capozza, Denise DiPasquale, Jeff Fuhrer, Robert Helsley, Charlie Himmelberg, Glenn Hubbard, Stuart Rosenthal, Will Strange, and participants at the University of British Columbia Summer Seminar, Columbia Business School Macro Lunch, and AREUEA Annual Meeting in New Orleans for helpful comments and Margaret Enis for excellent research assistance. All errors are the responsibility of the authors.

[2]See Green [19] for an analysis of residential construction as a leading indicator.

0094-1190/00 $35.00

position of homeowners and housing affordability. Finally, the welfare and distribution effects of fundamental tax reform depend on the elasticity of housing supply.[3]

Most of the existing empirical work on housing treats residential construction like other types of investment. However, land makes housing different: land and thus housing prices must ensure a spatial equilibrium with a metropolitan area and land is inelastically supplied. In this paper we estimate a supply equation for new single-family residences that is based on the theoretical models of land development and urban growth. Our basic estimating equation treats single family starts as a function of changes in current and lagged house prices rather than of their level. This approach is also more consistent with the time series characteristics of the data than the more traditional levels specification. To compare our specification with existing empirical treatments of new housing supply we provide out-of-sample forecasts for the method developed here and the two others from the existing literature.

The nature of price equilibrium in housing markets suggests using changes in house prices to measure demand for new construction, rather than the level of house prices. House prices equilibrate the total quantity of housing, a stock variable, with the total demand for residential space. Housing starts are a flow variable, representing the change in the stock of housing, net of removals. Thus, starts should be a function of other flow variables, including the change in house prices. In balancing supply and demand, house prices ensure a spatial equilibrium for households within a housing market. As a result, house price levels must depend on variables that predict the size of a city, such as the opportunity cost of new land and expectations of future growth. In a steady state, some variables that explain the stock of housing may be uncorrelated with housing starts, which are the change in the stock of housing.

House prices reflect the price of structure (capital), which is elastically supplied in the long run, and land, which even in the long run is inelastically supplied. House prices and land prices typically move together while the price of structure tracks more closely with construction costs (see Rosenthal [30]). Land is not like other investment goods: the long-run cost curve for land is upward sloping. A one-time increase in demand that results in a larger city, and more construction to accommodate these additional households, also causes a permanent increase in land prices. This increase is necessary to ensure a spatial equilibrium for the now larger city. The literature on city size and urban growth (see Capozza and Helsley [5] among others) develops the dynamics of urban land and house prices.

[3]Capozza *et al.* [4] review the relationship between tax reform and real estate.

A simple example demonstrates the intuition of treating housing starts as a function of house price changes. Imagine a city composed of a stable number of homogeneous households. If the city is not growing and housing units do not deteriorate, when the housing market is in a long run equilibrium, house prices are constant and housing starts equal zero.[4] At the urban fringe house prices equal the value of land in agricultural use plus the cost of converting the land to residential use and building the structure.

Suppose that the city has an unexpected one-time influx of population. Demand for new residences increases, land and house prices rise, new construction occurs, and the city increases in size to accommodate the new residents. At the new equilibrium, the city is physically larger. House prices at the urban fringe are unchanged, but the fringe is now further from the city center. To ensure that households are indifferent between living in houses at the newer, more distant locations and existing units, the price of houses at developed locations must rise relative to their level prior to the demand shock. In the new spatial equilibrium, population is stable and there are no expectations of further growth, so starts are again equal to zero. Average house prices in the city are constant, but at their new higher level. Even in a city whose form does not follow the monocentric model, such as one with a number of suburban employment subcenters, our characterization of the new construction process still holds as long as the long-run supply curve for land is upward sloping. That is, as a city gets larger, existing locations become more valuable relative to new locations on the urban fringe.

In this example, starts occur only when the city makes the transition from one equilibrium to another, a period identified by the increase in the price level. A model where starts are a function of the price level would predict a permanent increase in the number of housing starts resulting from the one-time unexpected increase in population.[5] Yet, starts will increase only as needed to accommodate the new residents, a one-time event.

The methodology that we propose here has an easy analogy to the literature on investment and Tobin's *Q*. In the case of housing, new

[4] If households are heterogeneous and changing, then there can be new starts to meet the changing needs of the stable existing population. With depreciation, housing starts equal removals, in this case a constant percentage of the stock.

[5] If we allow for depreciation in the model, we can obtain a positive correlation between housing prices and starts. When the population of the city is higher, the city occupies a greater land area, so that housing prices are higher and the stock of housing is larger. With a constant removal rate, the larger city requires a greater number of housing starts to maintain its existing stock of units. Thus after the increase in population, starts would be higher than before, as would house prices.

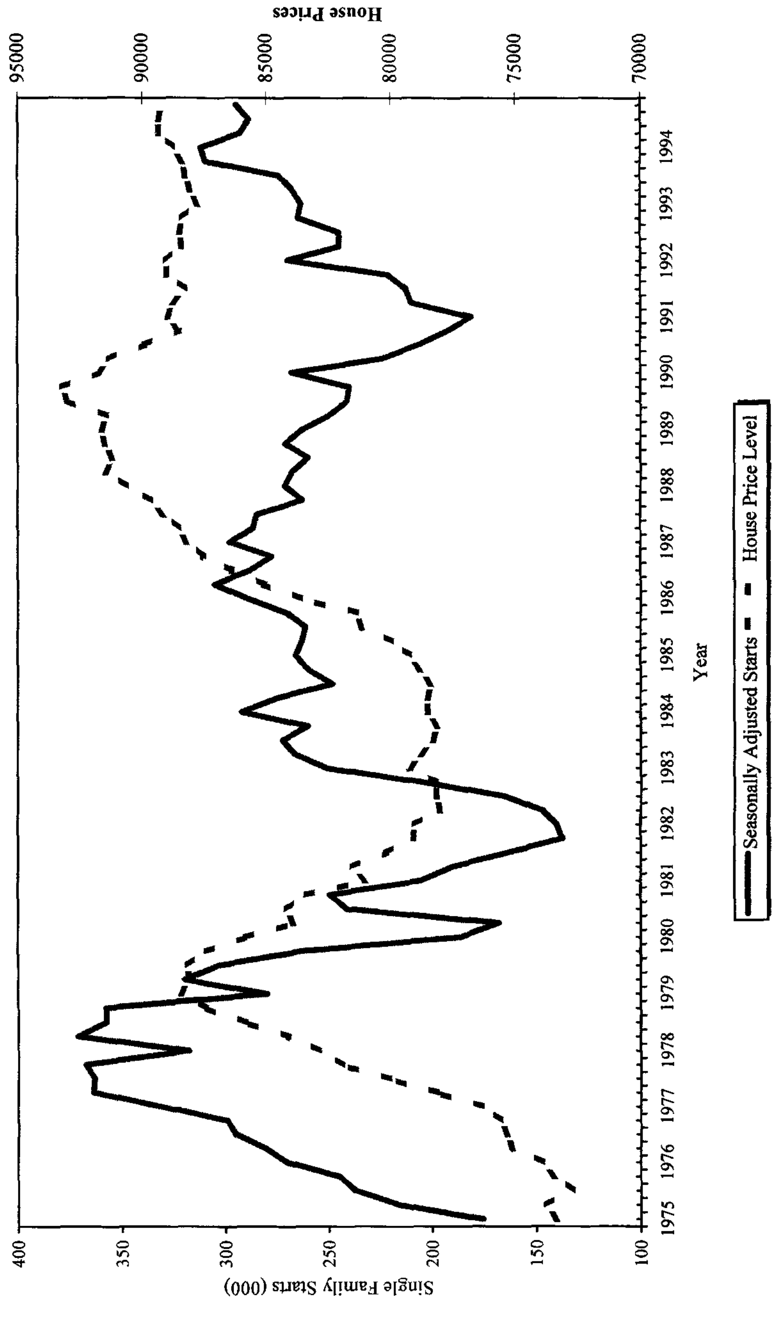

FIG. 1. Single-family starts vs prices: Seasonally adjusted starts and constant quality house prices.

investment (starts) should be positive as long as Q (the ratio of the market price of new housing divided by its construction cost) is greater than 1. In this case, construction costs include interest rates, the cost of materials and labor, and also land. While capital, materials, and labor have perfectly elastic supply curves in the long run, land within a given distance to the city is in limited supply. An increase in demand for housing raises the price of land and thus the price of housing. At a new equilibrium with a higher price of housing and land, $Q = 1$ and housing starts return to their steady state level—replacing depreciation of the existing stock.

The distinction between starts and prices is readily apparent in U.S. housing data as shown in Fig. 1. (An explanation of the data is provided in Section IV). Between 1987 and 1994, house prices remained above the level of earlier periods, yet starts during this period were consistently below the number of starts recorded in the late 1970s.[6] This figure suggests the limitations of using price levels to explain housing starts. By contrast, Fig. 2 shows that the relationship between housing starts and the change in house prices is much more consistent.

The empirical model in this paper also generates a stable measure of the true supply elasticity, the percentage change in the housing stock from a percentage change in house prices. This estimate will be quite small because housing starts are a small percentage of the stock; annual starts average 2.2% of the stock. A one-time increase in house prices leads to a one-time increase in the stock of housing, accomplished by a temporary increase in new construction, ignoring the replacement of units removed from the stock. Estimating starts as a function of house price levels can yield unusual predictions because a change in the level of house prices results in a permanent increase in new construction.

Treating starts as a function of house price changes is also consistent with the time series properties of housing stock and prices. Previous research (see Holland [20], Meese and Wallace [24], and Rosenthal [30]) finds that the real price of existing housing time series is not stationary, but that first differences in the price series form a stationary series. The stock of housing is also a nonstationary series, although starts themselves are stationary. Problems exist in estimating relationships between a stationary variable (starts) and a nonstationary variable (prices). Although over short time periods or in small samples these variables may be correlated, in the long run this correlation will disappear. Furthermore, regressions using multiple nonstationary series can lead to spurious correlations (Granger and Newbold [18]). If the stock of housing and real house

[6] Mayer and Somerville [22] demonstrate the effects of bank failure, credit crunch, and FIRREA on starts over this period only hold for the Northeast in the early 1990s.

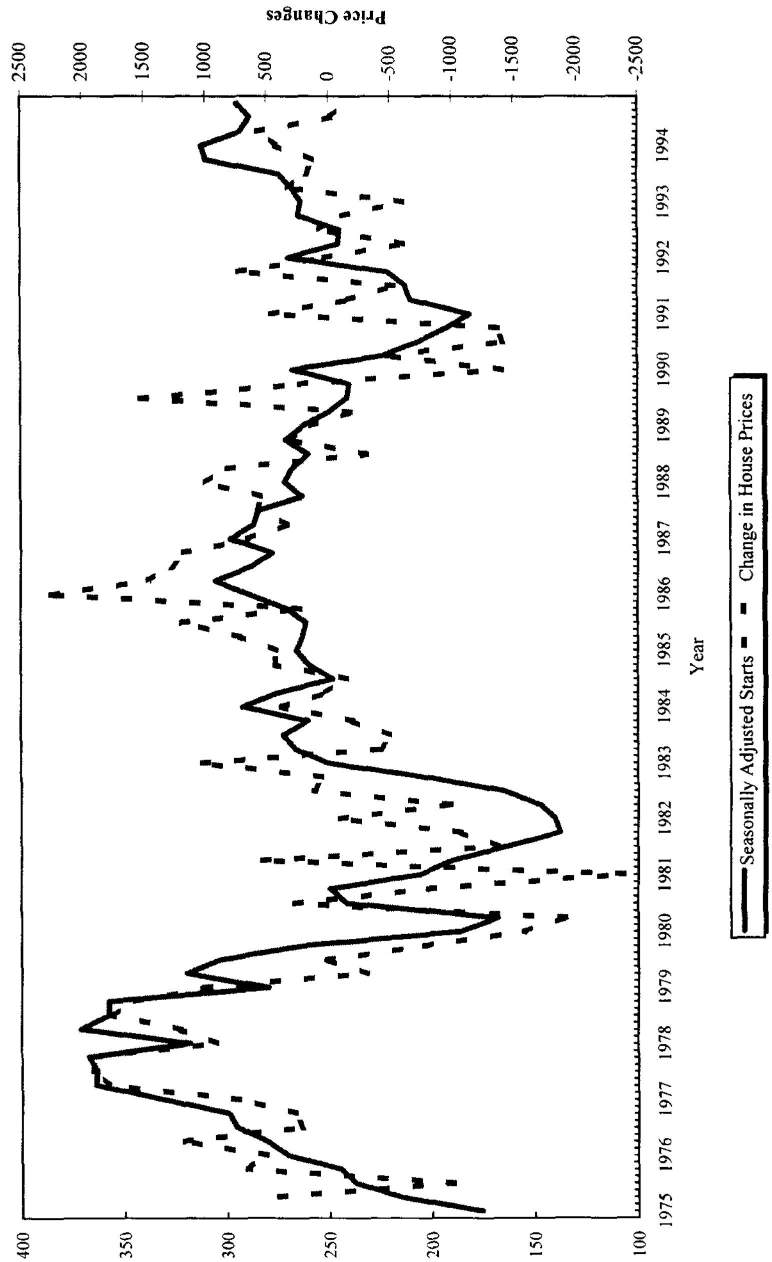

FIG. 2. Single-family starts vs change in prices: Seasonally adjusted starts and constant quality house prices.

prices are both stationary in their first differences, the proper econometric specification is to regress starts on price changes.

In the next section of the paper we review existing empirical treatments of housing supply that model starts as a function of house prices. We discuss the empirical model in Section III (a more formal treatment is provided in the Appendix). In Section IV, we use national time series data from 1975 to 1994 to estimate new single family construction. We also use out-of-sample forecasts to show that our model fits the data better than the specifications used in DiPasquale and Wheaton [12] and Topel and Rosen [34]. Section V concludes the paper with an agenda for future research.

II. REVIEW OF EMPIRICAL HOUSING SUPPLY RESEARCH

Existing empirical studies of housing supply use two approaches to estimate the relationship between starts and house prices.[7] In the first, housing supply and demand functions are combined into a single reduced form equation. The price elasticity of starts is derived from the coefficients on supply and demand shifters in the reduced form regression. Authors such as Muth [26], Follain [16], Stover [33], and Malpezzi and Maclennan [21] use this approach, a variant of the reduced-form equation, and none finds a statistically significant relationship between price levels and demand measures. From this they conclude that the supply curve for new housing is perfectly elastic.[8] This characterization is also evident in VAR time series models of housing prices such as Dreiman and Follain [13]. The second approach directly estimates the aggregate supply curve for new residences, modeling starts as a function of the level of house prices and various cost shifters. Research using this specification includes Poterba [28, 29]; Topel and Rosen [34], and DiPasquale and Wheaton [12], which yields much lower estimates of the price elasticity of starts, from 1.0 to 3.0. Blackley [2] compares these different approaches using a single long-run data set.

In their widely cited paper, Topel and Rosen [34] use an investment model to estimate single-family starts in the presence of dynamic marginal costs. They postulate that marginal costs rise with both the level of and changes in new construction activity. Consequently, following a positive demand shock, builders lower costs by smoothing their increase in output

[7]See DiPasquale [10] for a brief review of the literature on housing supply.

[8]Olsen [27] notes that if improperly specified, this approach can yield inconsistent coefficient estimates when both input prices and output are included as independent variables in a price equation. Follain includes a variety of specifications, only one of which is directly affected by this criticism. Malpezzi and Maclennan structure their tests to avoid the problem altogether.

over a number of periods rather than building all new units in one or two periods. In their empirical work, Topel and Rosen find evidence of this relationship: both lagged and future starts are correlated with current-period output. They find that the single-quarter price elasticity of starts is about one-third of the long-run value of 3.0, but convergence occurs quickly, in approximately one year.

Their model places new construction in a dynamic framework and the premise of the model, that builders smooth production in response to demand shocks, has much intuitive appeal. However, research on inventory investment has struggled to find empirical support for this approach. For example, Blinder and Maccini [3] demonstrate that actual production is typically more volatile than sales in the manufacturing sector, implying that little smoothing occurs. A similar result occurs in housing markets, where the coefficient of variation for starts is greater than that for sales of new houses (0.26 versus 0.19). This raises the question of whether the smoothing model is truly appropriate for housing investment.

The formal presentation in Topel and Rosen does not explicitly address the role of land. They explain smoothing by focusing on the effects on construction costs of the movement of resources across sectors. It is possible to apply the smoothing approach to land. Given delays in bringing land from agricultural to urban use and obtaining building permits, it may be optimal to have a smooth supply of permitted, developed sites ready for starts.[9]

DiPasquale and Wheaton [12] estimate a stock-adjustment model in which current starts are a function of the difference between desired stock and the stock in the previous period adjusted for removals. They use the current price level as a proxy for the desired stock and include an estimate of the lagged stock in their regressions. DiPasquale and Wheaton obtain results consistent with their model; the coefficient on prices is positive and the coefficient on lagged stock is negative. Their approach has the advantage that it recognizes the difference between the stock and flow of housing units, a condition ignored in some of the earlier empirical research. However, the stock of housing is notoriously difficult to measure in non-Census years; physical depreciation and removals are unobserved, removal is an endogenous action, and not all starts are completed with the same lag. Most researchers apply an estimated decennial removal rate, but this causes the variation in a stock variable to mimic the variation in starts within any single decade. Though DiPaquale and Wheaton's elasticity estimate is similar to those in Poterba and in Topel and Rosen, their

[9] Mayer and Somerville [22] find that production lags vary by region, with greater lags in the Northeast and the West, where land approval processes are slower, and fewer lags in the South and the Midwest.

specification suggests that the stock of housing adjusts very slowly to shocks: 2% of the gap between actual and desired stock is closed in any one year. This rate yields an adjustment period of 35 years, which seems too long. Curiously this rate is equal to the ratio of mean annual starts to the stock.

III. EMPIRICAL MODEL OF RESIDENTIAL CONSTRUCTION

New housing starts move the housing stock from one equilibrium to another following a positive demand shock. In the monocentric models of urban growth (Arnott and Lewis [1], Wheaton [35], and Capozza and Helsley [5]) this type of shock generates a permanent increase in land and house prices at all interior locations of the city. The size of a price change at any single interior location will measure the magnitude of the demand shock. The growth in population is accommodated by an expansion of the urban area because all new construction occurs at the fringe. In reality, some single-family construction occurs at in-fill sites or as the redevelopment of existing units, but the bulk of this construction does occur in suburban areas.[10] With stable construction costs, growth rates, and capital costs, house prices at the fringe are unchanged—though the fringe itself is increasingly distant from CBD. We present a formal derivation of this relationship between housing starts and changes in house prices in the Appendix.

The model in the Appendix also makes clear that starts should also be measured as a function of the change in construction costs as well as prices. The intuition is relatively straightforward; city size depends on the price and cost of housing. In this specification, costs include all construction-related expenses, such as materials, labor, and interest rates (the cost of carrying land and materials).

In the theoretical literature, development occurs instantly to meet the growth in the city's population. In reality there are delays in developing land from nonurban uses and then in constructing residential units on these lots. With these delays, developers must forecast demand several periods in advance of their expected completion dates. Also, finished lots

[10]Among the 44 MSAs surveyed in the 1989–1991 AHS metro files, only 16% of units less than 5 years of age are in the MSA's principal city. These units tend to be located in those MSAs, such as Ft. Worth, Oklahoma City, Phoenix, San Antonio, and San Diego, where the central city includes undeveloped land within its municipal boundary. In 18 of the 44 MSAs, fewer than 2% of newer units are located in the MSA's principal city.

must be available before builders can start to construct new housing.[11] With lags, the number of finished lots, ld, in any period is a function of forecasts of city growth made in earlier periods. With a lag length of one period, the supply of finished lots is a function of market information in the previous period, i.e., price and cost changes (Δp and Δc) in period $t-1$,

$$ld_t = f(E_{t-1}(\Delta p_t, \Delta c_t)) = g(\Delta p_{t-1}, \Delta c_{t-1}). \tag{1}$$

In this case, we assume that builder's expectations about future price and cost changes are a function of lagged changes in the same variables. The supply of developed lots acts as a constraint on the number of houses that can be constructed, thus forming an upper bound on actual starts s_t. The number of actual starts depends on the optimal unconstrained level given demand for new housing starts, s^*, and the supply of developed lots ld available at time t,

$$s_t = \min[s_t^*, ld_t]. \tag{2}$$

Combining (1) and (2), actual starts are a function of current and lagged price and cost changes,

$$s_t = \min[s_t^*(\Delta p_t, \Delta c_t), ld_t(\Delta p_{t-1}, \Delta c_{t-1})] \approx g(\Delta p_t, \Delta c_t, \Delta p_{t-1}, \Delta c_{t-1}). \tag{3}$$

The actual pattern of lags, and thus lags of price changes, that is appropriate is an empirical question we address in the estimation of Eq. (3) in the next section.

At first glance, our estimating Eq. (3) differs from the stock-flow specification in DiPasquale and Wheaton [12]. However, the two specifications do share a number of similarities. In their paper they use lagged stock in conjunction with current house prices to capture the relationship between demand for new units and the existing urban form. In the stock-flow model, the equilibrium stock in the current period is a function of the current price level, while lagged stock is a function of the previous period's price level. This suggests that in an estimating equation, a change in the stock can be described by Δp_t.

[11] We assume that the lag in converting new land to finished lots is longer than the time needed to construct structures on the lots because land development is the slowest part of the supply process. Negotiating the approvals and subdivision permitting process can take up to several years in some jurisdictions. In contrast, once permits are obtained houses can be constructed in less than 90 days.

IV. EMPIRICAL ANALYSIS

The specification presented above describes a supply function for new housing that is consistent with the land development process. While the model is most appropriate for a single city, the empirical work that follows uses national data. Doing so imposes two strong assumptions: first, that an urban form framework is applicable to national data; second, that there is a single national housing market, an assumption found in all other housing supply research using national data. Aggregation is problematic because it hides interesting variation in the timing of real cycles across regions and shrouds inter-metropolitan area movements in population (Goodman [17]). However, national housing starts are an important policy variable. Most existing studies of the supply of new housing use national data. By using national series we can compare the results of our approach to the existing literature on housing starts, a variable of interest in monitoring macroeconomic conditions.

We measure house price movements with the Freddie Mac repeat sales price index. By measuring price changes with repeat observations of the same housing units, the repeat sales methodology used in constructing the Freddie Mac series controls for location. This provides us with a price series consistent with the theoretical model, where we measure prices at a fixed location in the interior of the city. There is an extensive debate on the merits of various approaches to estimating quality controlled house price series. Meese and Wallace [25] indicate some of the problems with repeat sales indexes, which clearly affect the price series we use for analysis. To date, most researchers have used a hedonic new house price series (Census Series C-27) as the price variable in a supply equation. However, this series is subject to downward bias because of changes in the location of new housing; i.e., it does not fully adjust for the increase in location rents (prices) inside the city as growth occurs. Despite the problems with repeat sales indexes, the Freddie Mac series is conceptually more compatible with our treatment of supply.

Table 1 gives descriptive statistics for all of the variables used in the empirical work. The real house price series is calculated by taking the Freddie Mac index, which identifies quarterly changes in house prices for the years 1975 to 1994, and converting these changes to house price levels using the 1991 national hedonic house price estimated in DiPasquale and Somerville [11].[12] Real house prices increase by an average of \$224 per quarter (0.3% of the mean price level), with declines as large as −\$2,508 and increases as high as \$2,127. Most of these gains occurred in the late 1970s as real house prices leveled off in the 1980s. Housing starts also vary

[12] All dollar values are in third-quarter 1994 dollars, deflated using CPI-UX less shelter.

TABLE 1

Descriptive Statistics

Variable	Mean	Standard deviation	Minimum	Maximum
Stock (000)				
Level	59,851	4,443	51,481	67,166
Changes—Analogous to starts				
Starts (000)—Not seasonally adjusted				
Level	263.4	69.5	113.6	449.1
Starts (000)—Seasonally adjusted				
Level	259.9	52.5	137.4	371.2
Real house price ($)				
Level	84,154	5,714	72,706	93,305
Changes	224	943	−2,508	2,127
Real prime rate (%)				
Level	4.66	2.93	−1.67	11.02
Changes	0.08	1.32	−4.82	5.61
Estimated user cost				
Level	9.87	1.55	7.35	13.69
Changes	−0.01	0.34	−1.00	0.84
Median months to sale—New homes				
Level	6.68	1.66	3.70	11.60
Changes	−0.11	2.15	−6.90	4.70
Real material price index				
Level	0.97	0.06	0.87	1.08
Changes	−0.002	0.011	−0.023	0.026
Employment excluding construction (000)				
Level	105,467	10,489	87,127	122,567
Changes	421	612	−1,644	1,820
Married couples (000,000)				
Level	50.07	2.07	46.83	53.15
Changes	0.08	0.83	−0.05	0.37
Real energy price index				
Level	0.830	0.114	0.692	1.067
Changes	0.000	0.031	−0.090	0.081

Note. Stock is constructed from the decennial census. Quarterly figures are calculated using housing starts and an implied decennial removal rate. Consequently changes in the stock are identical to starts with constant decennial trends. The real house price series is constructed using the Freddie Mac repeat sales price index and the 1991 estimated national hedonic price level from DiPasquale and Somerville [11]. The user costs series is calculated with DiPasquale and Wheaton's [12] methodology: the marginal tax rate is for the typical first time home buyer and property taxes are estimated to be 1.8%

significantly over the cycle. Quarterly starts range from 113,600 to 449,100 (0.2 to 0.8% of the total stock).[13]

We use real the real prime rate to measure the cost of financial inputs to builders. Most construction loans are financed at adjustable rates based on the prime rate. Any demand-side effects of changes in interest rates should be captured by the user cost, which is included as an instrument for price changes.[14] The user cost is calculated with the methodology of DiPasquale and Wheaton [12].[15]

Evidence of positive serial correlation for real house prices in the short run (Case and Shiller [8]) suggests that prices do not fully adjust to clear the housing market. In fact, time-to-sale also varies significantly at different parts of the cycle, with a low of 3.7 months to a high of 11.6 months in our survey period, and changes in this measure can precede real price movements. For example, at the beginning of a downturn, time on the market will typically rise several quarters before observed transaction prices begin to fall. We follow the existing literature and include the lagged value of the median number of months recently sold new homes were on the market in some of our regressions to capture demand factors that do not show up in price changes because of the stickiness in house prices.

By estimating starts as a function of changes in house prices, we address the econometric problems that occur because house prices are nonstationary (Holland [20], Meese and Wallace [24], and Rosenthal [30]). Augmented Dickey–Fuller tests for stationarity presented in Table 2 confirm that both starts and price changes are stationary, although the power of these tests can be quite low when samples are small (Faust [15]). These results show that the time series nature of the data is consistent with the

[13]The stock series is estimated using the starts series, the 1970, 1980, and 1990 Census counts of the stock, and the 1993 American Housing Survey estimate of the number of year-round single-family residences. The interdecennial removal rates are estimated so that starts minus total removals equals the stock in the next census year. The estimated annual rates for 1970–1980, 1980–1990, and 1990–1993 are 0.2, 0.5, and 0.47%, respectively. The 1990–1993 rate is assumed to hold for 1993–1994.

[14]The user cost of capital can differ from the prime rate because it is a long-term rate and demand for mortgages is affected by marginal tax rates. Interest rate spreads and marginal tax rates both vary significantly over time, so that the two series are only weakly correlated. (See Poterba [29].)

[15]The user cost equals $(1 - t_y - t_p - \pi)^*i$ where i is the effective interest rate for 30-year fixed mortgages, t_y is the marginal tax rate for the typical first-time home buyer, t_p is the property tax rate set at 1.8%, and π is the average of the current and the previous periods' inflation rate.

TABLE 2
Augmented Dickey Fuller Tests

Variable	Estimated alpha	ADF T-statistic	Reject unit root	ADF probability	ADF lags
Stock					
Level	1.001	0.749		0.991	5
Changes—Analogous to starts					
Starts					
Level	0.735	−3.312	**	0.013	5
Real house price					
Level	0.970	−1.417		0.574	4
Changes	0.571	−3.139	**	0.024	3
Real prime rate					
Level	0.832	−2.131		0.232	2
Changes	−0.349	−5.087	***	0.001	2
Estimated user cost					
Level	0.951	1.470		0.470	3
Changes	0.477	2.977	**	0.037	3
Median months to sale—New homes					
Level	0.715	−2.506		0.114	3
Changes	−0.571	−4.481	***	0.001	3
Real material price index					
Level	0.957	−2.245		0.146	3
Changes	0.170	−3.445	***	0.010	2
Employment excluding construction					
Level	1.000	−0.024		0.956	4
Changes	0.704	−2.833	*	0.054	3
Married couples					
Level	0.990	−2.128		0.233	4
Changes	0.508	−4.264	***	0.010	3
Real energy price index					
Level	0.959	−1.239		0.656	3
Changes	0.325	−3.465	***	0.009	2

Note. Reject the null hypothesis of a unit root (that alpha = 1) at the following levels of significance: ***, 1% level; **, 5% level; *, 10% level. All tests are two-sided ADF tests with seasonal dummies included in the regression. The Freddie Mac repeat sales index begins in 1975; to allow for consistent lags across variables, unit root tests are imposed for the years 1977–1994. With the exception of stock, including a trend does not alter the results. Changes in the detrended stock are nearly identical to the level of starts. By construction, the stock series equals the sum of starts minus a decennial trend removal rate.

model presented earlier.[16] Figures 1 and 2 in the Introduction also demonstrate this point.

[16] The empirical tests of the housing starts model presented are intended to explain short-term cyclical variations in housing investment (starts), rather than the long-run relationship between the housing stock and the price level. Other tests (not presented here) reject a cointegrating relationship between the stock and the real price of housing.

The model in Eq. (11) requires lagged price and cost changes, but the appropriate number of lags depends on the length of time required to obtain developed land, acquire housing permits, and builders' expectations about changes in future house prices. We use the following equation, with starts as a function of current and lagged changes in house prices, real interest rates, and construction costs,

$$s_t = g\left[\Delta p_t, \ldots, \Delta p_{t-j}, \Delta r_t, \Delta r_{t-1}, \Delta c_t, \Delta c_{t-1}\right]. \tag{4}$$

Because of the possible endogeneity between starts and both current period house prices and construction costs, we estimate (4) using an instrumental variables approach.[17]

Table 3 presents estimated coefficients from (4), along with several alternative specifications all with quarterly dummies and a time trend. We correct for serial correlation using an AR1 process. The Q-statistics indicate that the 95% chi-squared critical values are met for all regressions. The first regression is a direct estimate of Eq. (4). The regression yields plausible parameters; the coefficients on the current and first two lags of changes in prices and current interest rate changes are statistically different from zero at the 5% level.

Changes in house prices have the strongest effect on housing starts. In regression (1), a one-standard-deviation increase in real house prices ($943) increases aggregate starts by 18,300 units in the quarter of the increase and by 53,800 units over the course of a year. These figures are approximately 7.0 and 20.7% of mean quarterly starts, respectively. The alternative specifications in regressions (2) to (4) yield similar results though the aggregate effect on starts tends to be smaller; for instance, a total increase over a year of 45,200 units in regression (3).

Changes in real interest rates have a statistically significant effect on housing starts, but the effect is smaller in magnitude than that of changes in house prices. In regression (1), a one-time, one-standard-deviation (1.3 percentage points) increase in the real prime rate lowers total starts by 12,000 units, less than 5% of the average number of starts in a quarter. The effect is smaller, 8,000 units, in regression (3). The effect of changes in interest rates on demand for housing is captured in the price change variable because changes in the user cost are an instrument for price changes. This small direct effect of real interest rates on housing starts suggests that much of the effect of interest rates on the housing market occurs through demand rather than supply.

[17] Instruments include lagged changes in construction costs and current and lagged changes in the number of married couples, the user cost of capital, nonconstruction employment, and real energy prices, as well as current and lagged exogenous variables. These instruments are similar to those used by Topel and Rosen. Using microdata, Somerville [32] finds that construction costs do move with the volume of construction.

TABLE 3
Regression Results (1975–1994)

Variable	Regr. (1)	Regr. (2)	Regr. (3)	Regr. (4)
Change in price	0.0194	0.0176	0.0177	0.0189
	(0.0091)	(0.0089)	(0.0092)	(0.0087)
Change in price (−1)	0.0196	0.0192	0.0156	0.0159
	(0.0047)	(0.0047)	(0.0049)	(0.0048)
Change in price (−2)	0.0134	0.0132	0.0129	0.0129
	(0.0043)	(0.0043)	(0.0040)	(0.0040)
Change in price (−3)	0.0047	0.0045	0.0017	0.0017
	(0.0051)	(0.0051)	(0.0048)	(0.0048)
Change in real prime rate	−4.85	−4.88	−3.67	−3.49
	(2.44)	(2.37)	(2.32)	(2.33)
Change in real prime rate (−1)	−4.16	−4.33	−2.38	−2.24
	(2.50)	(2.47)	(2.40)	(2.41)
Stock (−1)				0.0012
				(0.0019)
Median months on market			−9.79	−9.33
Until Sold—New homes (−1)			(4.72)	(4.57)
Change in real building	−98.5		−19.4	14.7
Material cost index	(377.7)		(373.8)	(372.2)
Time trend	−0.053	−0.066	−0.228	
	(0.438)	(0.437)	(0.368)	
Constant	208.2	209.2	309.2	356.8
	(45.8)	(45.6)	(63.9)	(133.9)
Number of observations	76	76	76	76
Regression type	AR-IV	AR-IV	AR-IV	AR-IV
Log-liklihood	−352.5	−351.9	−348.3	−349.1
Estimated AR1 rho	0.67	0.67	0.60	0.60
Q-statistic(4)	6.34	7.47	5.71	5.31

Note. Standard errors in parentheses. All regressions use seasonally unadjusted series and include quarterly dummies. Instruments used for current changes in real house prices and building material costs. Instruments for current change in real house prices are current and lagged values of changes in nonconstruction employment, real energy prices, mortgage rates, and the number of married couples. We use lagged changes in the real building materials price index as instruments for changes in real materials prices. We also include lagged values of all exogenous variables as instruments.

As in other empirical housing supply studies, the coefficient on materials prices is not statistically different from zero. Lacking appropriate instruments at the national level that are uncorrelated with housing demand with which to correct for the endogeneity between starts and materials prices, we use lagged material prices. As demonstrated in regression (2), removing the materials cost variable from the basic equation has little effect on other coefficients.

Regression (3) in Table 3 augments the basic equation with the lagged values of median time-to-sale for new homes, a nonprice measure of market conditions. Consistent with the findings of both Topel and Rosen and DiPasquale and Wheaton, this variable is negative and significantly different from zero, suggesting that builders pay attention to sales rates as well as prices in deciding whether to start new homes because of the well-known stickiness of house prices. The inclusion of lagged time-to-sale slightly reduces the size of the coefficients on current and lagged price and interest rate changes. The coefficient on median time-to-sale in column (3) suggests that deviations in this variable have quite a large effect on new construction. A one-standard-deviation increase in the median time-to-sale lowers aggregate starts by 16,300 units, approximately 36% of the total effect of a one-standard-deviation change in price. Clearly, builders respond to nonprice signals of market conditions.

We include lagged stock in regression (4) to control for the role of depreciation in explaining new construction. With a constant depreciation rate, starts should increase with the stock as more units depreciate and need to be replaced. In the other specifications we assume that this depreciation is captured by the time trend and the constant. In DiPasquale and Wheaton, lagged stock is important because it describes aspects of the land market and urban growth not fully revealed in price levels. Once we control for price changes, the coefficient on lagged stock is not statistically different from zero and including it has little effect on the results. This suggests that our approach of using price changes is successful in capturing the process of urban growth in the housing starts equation.

Like DiPasquale and Wheaton, we differentiate between the elasticity of housing supply and that of housing starts. The former describes the percentage change in the entire stock of housing, while the latter describes the change in flow of new construction. Regression (1) in Table 3 generates an estimated supply elasticity of 0.08, so that a doubling of house prices would increase the entire stock by 8%. Our estimate is substantially lower than DiPasquale and Wheaton's estimate of 1.0–1.2. In their model a price increase leads to a permanent increase in starts. Also, in a stock adjustment model, the long-run price elasticities of starts and the stock must be equal. Since our model allows for transient increases in starts, this equality need not hold. The estimated coefficients from our model suggest that the complete response of the stock to a demand shock occurs within one year.

Comparing our elasticity of housing starts to existing estimates is difficult because our starts' elasticity is sensitive to the length of time over which we calculate the elasticity of starts. The longer the period, the lower the starts elasticity. According to our estimates, a one-time increase in prices increases starts in the current quarter and in each of the next three quarters. Afterwards, starts return to their previous level. From regression (1), a 1% increase in prices in a given quarter causes a 6.3% increase in starts in the same quarter. This is approximately six times the estimate from DiPasquale and Wheaton and Topel and Rosen's short-run elasticity. Over a year, a one-time 1% increase in prices causes annual starts to rise 3.7%. This is only slightly higher than Topel and Rosen's long-run estimate. The total transient increase in starts over 4 quarters is quite high when measured against mean quarterly starts, approximately 18.5% of the starts in a single quarter. Elasticity estimates for changes in the real prime rate and time-to-sale are much lower; both are consistently below -0.3.

In addition to comparing our regression coefficients and elasticities with those in other papers, we also present out-of-sample forecasts for our model, using regression (3), and those of DiPasquale and Wheaton [12] and Topel and Rosen [34]. Table 4 presents the estimates from each of the models using data from 1976 to 1987.[18] The coefficients from the DiPasquale and Wheaton specification are similar to those presented in their article. However, the estimated coefficients from the Topel and Rosen model appear to be less stable. In fact, the estimated coefficient on the current price level is not statistically different from zero, suggesting that the short time series limits the efficiency of their estimates.

Next, we use the coefficients from these estimates to develop forecasts for 1988 to 1994, which we present in Fig. 3. Overall, our model performs better than those that use price levels: the standard error of our forecast is 24.2, well below that of the other two models, 32.9 for DiPasquale and Wheaton and 36.4 for Topel and Rosen. These forecasts use estimated values for lag starts, i.e., dynamic updating, rather than actual starts. The problems with using unobservable future values to model current starts means that the forecast for the Topel and Rosen model in Fig. 3 is based in regression (4) on Table 4 rather than on their actual specification, which is regression (3). While we cannot draw definitive conclusions from these forecasts because of the relatively small sample size and the difficulties in generating forecasts from the Topel and Rosen specification, the above results indicate that the model we present here compares quite favorably with existing treatments of housing supply.

V. CONCLUSION

In this paper we develop an empirical model of new single-family housing supply that reflects the role of land in producing new housing and the theoretical treatments of urban growth. The conceptual approach is

[18] In their papers DiPasquale and Wheaton and Topel and Rosen use different price variables and interest rate measures than we do here. To facilitate direct comparisons, all of the forecast regressions use the Freddie Mac price index.

TABLE 4
Forecast Regressions (1976–1987)

Variable	Regr. (1)	Regr. (2)	Regr. (3)	Regr. (4)
Change in price	0.001			
	(0.011)			
Change in price (−1)	0.009			
	(0.007)			
Change in price (−2)	0.015			
	(0.006)			
Change in price (−3)	0.001			
	(0.006)			
Price level		0.0034	0.0005	0.0001
		(0.0013)	(0.0004)	(0.0006)
Starts (−1)				0.455
				(0.100)
Starts (−1) + .98* starts(1)			0.386	
			(0.041)	
Starts(1)*.98				0.253
				(0.161)
Stock (−1)		−0.0066		
		(0.0025)		
Change in real prime rate	−4.14			
	(3.36)			
Change in real prime rate (−1)	−4.43			
	(3.61)			
Real T-bill		−4.34		
		(2.26)		
Expected real T-bill (−1)			−3.16	−4.66
			(1.28)	(1.82)
Change in employment		0.042		
		(0.007)		
Expected inflation (−1)			−0.221	−1.552
			1.085	1.650
Median months on market until Sold for new homes sold (−1)	−20.5	−23.8	−13.0	−13.2
	(7.9)	(3.6)	(3.4)	(3.8)
Constant	376.7	501.9	70.9	140.6
	(69.5)	(93.1)	(46.3)	(93.7)
Model type	Price changes	DiPasquale and Wheaton	Topel and Rosen	Topel and Rosen
Number of observations	47	48	48	48
Regression type	IV-AR	OLS	IV	IV
Adjusted-R sq		0.93	0.92	0.92
Log likelihood	−221.3			
Durbin Watson	2.08	1.72	2.75	2.69

Note. Standard errors in parentheses. All regressions use seasonally unadjusted series and include quarterly dummies. Instruments used for current real house price variables and in the Topel and Rosen regressions for both lagged and future starts. The instruments are similar to those used in the regressions in Table 3. The regressions replicating Topel and Rosen's model use their instruments along with lagged median months on the market.

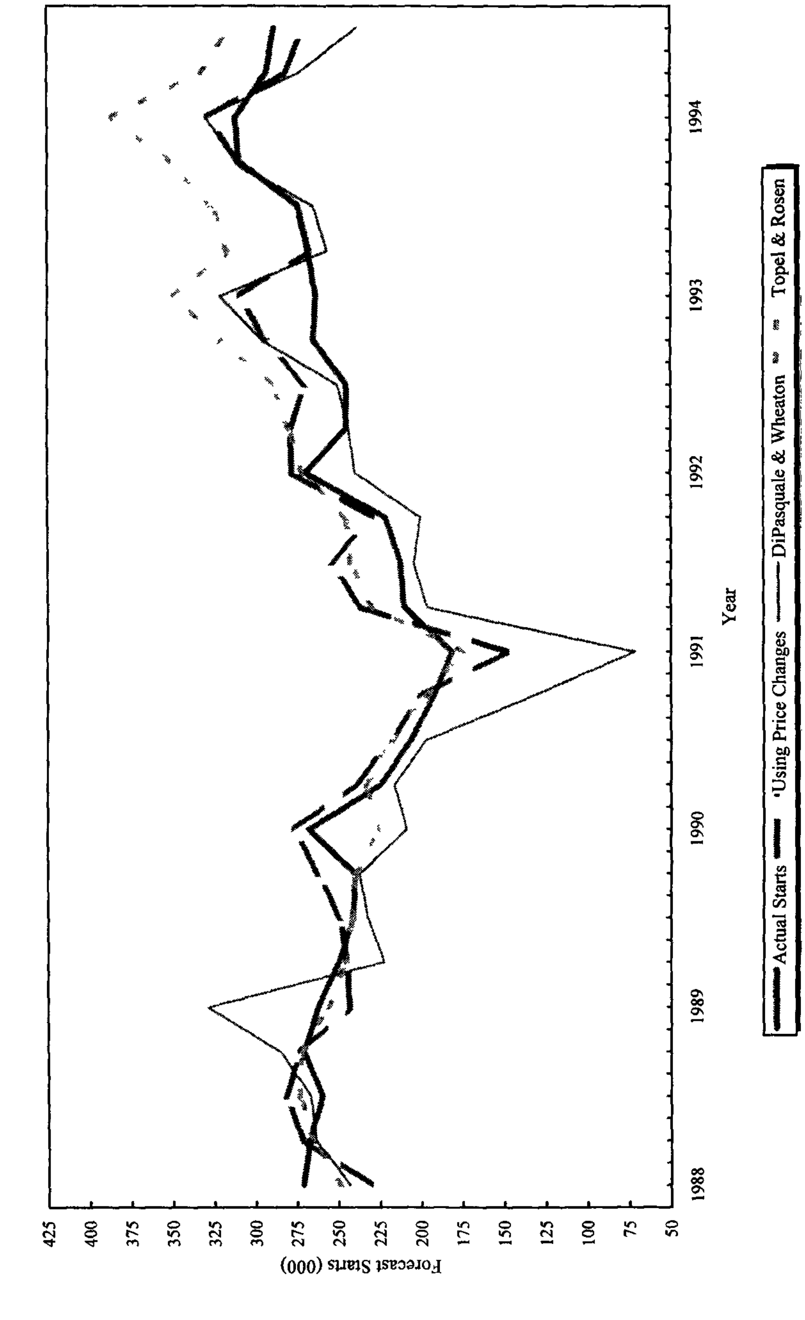

FIG. 3. Starts: Forecast vs actual, comparing out of sample starts forecasts.

also consistent with the time series properties of housing market data. Empirical estimates support our treatment of housing starts as a function of changes in existing and lagged house prices and costs. The significance of up to three lags of price changes and one lag of interest rate changes suggests that lags in the development process cause housing starts to take time to respond to demand shocks. In our estimates the stock adjusts to a demand shock within one year, much faster than the 35 years in the DiPasquale and Wheaton model and similar to the year adjustment period in Topel and Rosen. Consistent with other research, we find that the lagged value of expected time-to-sale has good explanatory power, which suggests that nonprice measures of demand are important in explaining new construction. Finally, our model performs quite favorably when its forecasts are compared with forecasts of two other widely cited models of new housing supply.

Our estimates suggest a fairly moderate response of supply to house price changes. A 10% rise in real house prices leads to an 0.8% increase in the housing stock; this is accomplished by an immediate 63% increase in quarterly starts. Over a year, annual starts increase by a total of 37%. This analysis highlights the difference between a housing supply elasticity and the starts elasticity found in the existing literature. In our modeling framework, a one-time increase in housing prices leads to a temporary rather than permanent increase in new construction, yielding a finite increase in the stock of housing.

While our model performs well on national data, it is ideally suited for estimating housing supply functions for individual metropolitan areas. As noted above, the level of housing prices can vary across housing markets for reasons that have little to do with the demand for new housing, including differences in population, land availability, and the expected rate of growth. Consequently, applying the conventional specification of starts as a function of the current level of house prices may generate misleading results in cross-sectional or panel analysis. By using changes in house prices instead of house price levels, our model avoids this problem.

In future research we hope to use this model to estimate housing supply functions for different metropolitan area markets. Cross-sectional supply functions would allow us to study the effects of factors such as government building restrictions and the opportunity cost of vacant land on the price of housing and on the speed with which quantity supplied responds to demand shocks. Differences across markets in the elasticity of supply and the speed of adjustment may well explain why prices are higher and real estate cycles seem to be more pronounced in markets such as California and the Northeast, where development restrictions are tighter and buildable land within commuting distance of the downtown is scarce.

APPENDIX

This appendix formally derives the starts–house-price-change relationship from the Capozza–Helsley urban growth model [5]. Builders convert raw land to urban use by developing the land and constructing housing on the finished lots. With fixed lot and house sizes, and no uncertainty, developers maximize profits by selecting development time t^* to convert land at location d, given agricultural land rent r_a, house rents r_h, and structure cost c_h,

$$\max_{t^*} \pi(\gamma, d) = \int_T^{t^*} r_a e^{-i(\gamma - T)} d\gamma + \int_{t^*}^{\infty} r_h(d,t) e^{-i(\gamma - T)} d\gamma - c_h e^{-i(t^* - T)}. \tag{5}$$

The solution to this problem is the optimal development time in the absence of uncertainty,

$$r_h(d, t^*) = r_a + ic_h. \tag{6}$$

Conversion occurs when the price of housing at a currently undeveloped location exceeds the agricultural value and cost of conversion. At the fringe, land owners must be indifferent between leaving land in its existing agricultural use or developing it, so location rents must equal zero.

Equilibrium house rents in the monocentric city depend on city size b (the distance from the core to the city border), transport cost k, structure cost, and agricultural land rent.[19] At time T and distance d from the city center, the price of a house is given by the present discounted value of house rents,

$$p(d,T) = \int_T^{\infty} [r_a + ic_h + k(b_t - d)] e^{-i(t-T)} dt. \tag{7}$$

The terms on the right hand side are the rent for the land in its current use (the forgone agricultural rent), the rental value of the house's structure capital, and the location rent needed for a spatial equilibrium. The latter depends on the linear transportation cost k multiplied by the distance a house is located from the urban fringe.

[19] General results from a polycentric city will be similar, as long as all employment centers are in the interior of the urban area. The mathematics of these solutions are substantially more complicated. As a result, we use the simpler, though less realistic, monocentric model to convey the qualitative results of this approach. Also, we assume that each house embodies one unit of land.

Current prices depend on the city's expected growth rate g. Solving the integral in (7) for prices yields

$$p(d,T) = \frac{r_a}{i} + c_h + \frac{k(b_T - d)}{i} + \frac{kb_T g}{i(i-g)}. \tag{8}$$

The first three terms are the present value of the components of current rent. The last term is the present value of the expected increase in rent at location d. It is derived from future increases in location rents that will be needed to ensure a spatial equilibrium as the city grows at the expected rate g.

Instead of defining prices as a function of the distance from the border to the city center, we rearrange Eq. (4) to express the border as a function of house prices at a fixed interior location d ($d \leq b_T$),

$$b_T = (i-g)\left[\frac{p(d,T) - c_h}{k} - \frac{r_a}{ki} + \frac{d}{i}\right]. \tag{9}$$

The city border is a function of the population, which, given a fixed ratio of population to land, determines the total stock of housing.

At any point in time, the stock of units is the sum of all housing built, with adjustments for abandonment and demolitions. When there is no undeveloped land in the city, the stock can be used to define city size. In a circular city of radius b_T with θ radians of developed area and fixed lot size of 1, the total stock H_T can be described by the city's radius, the distance to the boundary,

$$H_T = \pi b_T^2 * \frac{2\pi}{\theta} = \frac{2\pi^2 b_T^2}{\theta}. \tag{10}$$

The stock changes with housing starts, ignoring removals. Given (10) we can describe starts between two periods as a function of the change in the city's area. In a simple case with no abandonment or demolition, and if all development occurs in a smooth process (no leapfrogging), starts s^* can be expressed as a function of the change price (and cost) levels from $T-1$ to T,

$$s_T^* = F(p(d,T), c_h(T)) - F(p(d,T-1), c_h(T-1)). \tag{11}$$

This equation highlights the connection between our treatment of starts and that found in DiPasquale and Wheaton's work [12]. They use a stock-flow model, so starts depend on optimal current and actual lagged stock. However, in our approach these two variables are themselves functions of current and lagged prices and costs.[20]

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[20] If we explicitly solve (10), starts in (11) are a quadratic function of $p(\bar{u}, T)$ and $p(\bar{u}, T-1)$. The quadratic specification is accurate when development is a smooth, continuous process around the entire ring of the city border. However, land assembly, topographical constraints, nonmarket uses of land, and differences in land use regulations across space all serve to prevent this pattern of development. They also make it difficult to identify the precise form of the underlying relationship between distance to the border and city population. If we are willing to assume a linear instead of a circular city, starts would depend explicitly on the differenced $p - c_h$.

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