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The Long-Run Price Elasticity of Supply of New Residential Construction in the United States and the United Kingdom

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Most housing models, and most policy analysis, hinge on explicit or implicit estimates of the price elasticity of supply of housing: does the market respond to demand side shocks with more supply or higher prices? Building on a model originally developed by Steve Mayo, we estimate the price elasticity of supply of housing from new construction separately for the United States and for the United Kingdom. We examine the supply elasticity over a very long time frame—from the previous century. There is strong evidence of a "regime shift" in 1914–1947; over the entire period, prices rise in both countries, but not in a continuous manner. Post World War II, the United States is essentially flat, albeit with very large cycles. In the UK, relative housing prices generally rise postwar. According to our flow model, in the prewar United States our implied price elasticity is between 4 and 10, postwar it is between 6 and 13. In the prewar UK our implied price elasticity is between 1 and 4, postwar it is between 0 and 1. Stock adjustment models yield different price elasticities—surprisingly so, in our judgment. They range from 1 to 6 for the United States, and from 0 to 1 for the UK. We believe the stock adjustment models are particularly fruitful areas for additional work.

I. INTRODUCTION

Most housing models, and most policy analysis, hinge on explicit or implicit estimates of the price elasticity of supply of housing: does the market respond to demand side shocks with more supply or higher prices? Numerous surveys

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of the housing market literature (e.g., Olsen 1987, Smith *et al.* 1988, Maclennan 1982, Malpezzi 1996) have noted the relative lack of empirical estimates of supply elasticities, despite the central importance of these parameters for market analysis and policy.

Building on a model originally developed by Steve Mayo, we estimate the price elasticity of supply of housing from new construction separately for the United States and for the United Kingdom. The U.S. and the UK housing markets operate in very different regulatory and financial environments; these may be used to explain differences in estimated parameters. In addition to cross-country comparisons, we intend to extend the literature in three ways: (1) parametric use of demand information to aid identification; (2) application of recent time series methods to test maintained hypotheses prior to estimation; and (3) tests for stability of the key parameters over time. In particular, with respect to the latter point, we examine the supply elasticity over a very long time frame—from the previous century.

II. PREVIOUS RESEARCH

Our review of the supply literature is brief. DiPasquale (1999), Bartlett (1989), and Bramley *et al.* (1999) provide more discursive reviews, with the former focusing on the U.S. literature and the latter including a wider range of countries, including Britain.

The Price Elasticity of Supply in the United States

Muth (1960) is usually cited as the first econometric examination of the supply side of the U.S. housing market.[1] Muth regressed the real value of new construction against the relative price of housing and input prices, and also estimated an inverted model with housing prices on the left-hand side. In both cases Muth found no significant relationship between price and quantity, consistent with elastic supply. However, Muth limited his investigation to the interwar years 1919–1934.

Following Muth, Follain (1979) estimated a series of similar regressions, using postwar data (1947–1975) and examining issues of simultaneity and serial correlation. Follain found qualitatively similar results, consistent with elastic supply.

Olsen (1987) pointed out that both Muth's and Follain's studies included input prices on the right-hand side. Olsen (pp. 1017–1018) viewed this as a

[1]Among previous studies, some of which were cited by Muth, see Blank (1954), and Grebler *et al.* (1956).

misspecification, arguing that in general the relationship between supply and input prices should be independent of whether the supply curve is elastic or rising. He further noted that the inclusion of extraneous variables does not generally bias results, but may reduce efficiency. Since the hypothesis of interest is unfortunately formed as a null, this could be of some consequence. However Olsen was writing a review article, and did not attempt to reestimate these models.

DiPasquale and Wheaton (1994) estimated a stock adjustment model that incorporated a simple model of urban form. In the standard urban model of Alonso, Muth, and Mills, as the city grows, the value of the "average" location in the city rises, as the bid-rent curve shifts upward, even if the bid rent at the ever-expanding fringe is constant. In the event, using 1963–1990 data DiPasquale and Wheaton estimate some of the supply lowest elasticities extant, in the range 1.0 to 1.4.

More recently Blackley (1999) has estimated several models broadly similar to Follain's. She used a time span longer than that of most other papers reviewed here, namely 1950 to 1994. Like Follain, Blackley initially included input prices on the right-hand side of her specification, but she also tested Olsen's recommended sparser specification. Blackley's point estimates of supply elasticities range from 1.6 to 3.7.

Mayer and Somerville (2000) present results from an alternative stock adjustment model, in spirit related to DiPasquale and Wheaton's. While earlier work by Mayer and Somerville (1996) yielded estimates of between 1 and 4, this later effort found a stock elasticity of about 0.08 (i.e. a 10% increase in price yields a 0.8% increase in the total housing stock; a small number but a large magnititude, since in a given year the stock of housing is 50–100 times starts), but a flow elasticity of about 6 (starts increase 60% from a 10% price increase).

Several other papers that are not usually thought of as "supply elasticity papers" contain explicit or implicit estimates of such a parameter. Most of these papers have found or implied low elasticities. Kearl (1979) reported an elasticity of 1.6 for new construction, and Huang (1973) 2 for starts. Topel and Rosen's (1988) research on starts found a long run-elasticity of 3 using quarterly data from 1963 to 1983. Poterba (1991) also presented data that seemed to indicate a rising supply price. In general, this set of papers is characterized by models and data that constrain "long-run" adjustment to a few quarters or years. Also the particular years chosen were in at least the last two cases periods where real housing prices were rising. Had they extended their estimation forward or backward in time they would have included declining prices. Put another way, these estimates put lower bounds on the true long-run elasticity, but say nothing about how close to the bound the true parameter might be.

A number of other papers have appeared in the U.S. supply elasticity literature, but generally these have focused on supply from the existing stock (e.g., Ozanne and Struyk 1976), or have used a cross-section approach rather than time series (e.g., De Leeuw and Ekanem 1971, Rydell 1982, Stover 1986).[2] A good general survey of the literature on the existing stock is contained in Rydell's study.

The Price Elasticity of Supply in the United Kingdom

A number of papers and books have examined the UK housing market over time. Many of these (e.g., Habukkuk 1962) have been primarily historical and/or descriptive. While these have been invaluable, including providing the data for our present undertaking, only recently has a more analytic approach been taken. Nellis and Longbottom (1981) were among the first to estimate reduced form models of UK housing prices, but their focus was not particularly on behavioral elasticities. Drawing on their work, Buckley and Ermisch (1983) showed how results from Nellis and Longbottom, and from similar specifications, could be used to recover behavioral parameters, but they focused on the short run and on demand side parameters and said little about the long-run supply elasticity.

Muellbauer (1992) presents data on the asset price of housing over time in the UK, using data from 1960 to 1989. His analysis is primarily descriptive, and he makes several interesting comments and comparisons to German price history, to which we will refer below. Stern (1992) uses 1971–1989 data to estimate a two-stage least-squares model. Stern's model is distinguished by explicit consideration of the lag structure; he says nothing direct about the long-run supply elasticity, but he finds that prices adjust to increased supply only after a lag of several periods.

The only UK study we have found to date that explicitly estimates the price elasticity of supply using time series data is Whitehead (1974). Whitehead develops a series of related stock adjustment models, and estimates them using quarterly data from 1955 to 1972. She generally found inelastic supply, with various models yielding elasticities ranging from 0.5 to 2. Another study by Mayes (1979, cited by Bartlett 1989) also finds inelastic supply in the United Kingdom.

A number of other supply-oriented papers have also appeared in the UK literature, but as above, these have focused on supply from the existing stock (e.g., Maclennan 1978), or have used a cross-section approach rather than time series (e.g., Bramley 1993, Pryce 1999; these typically yield elasticities around 1 or less). Bartlett (1989) contains an excellent review, updated by Bramley *et al.* (1999).

Other Countries

The only other housing supply elasticity estimates extant, to our knowledge, are Malpezzi and Mayo's (1994) estimates for Malaysia (between 0 and 1), Korea

[2] Generally these studies find that the supply of housing from the existing stock is fairly inelastic. However, in the long run the elasticity of total supply is for all practical purposes determined by the construction elasticity, certainly if the construction elasticity exceeds the existing stock elasticity.

(between 0 and 1), and Thailand (statistically indistinguishable from infinity). Malpezzi and Mayo argue that the rank ordering is the same as the ordering of each country's regulatory environment, a point to which we return below.

Next we develop some simple models that will be the basis for our current tests and analysis.

III. THEORY

We consider four related tests of the price elasticity of supply, and present results in this paper from three of them.

(I) The first test is a simple one. If markets are elastic, then prices do not vary, at least in the long run. So our first tests are whether there is a trend in the relative price of new construction. More stringently, tests for stationarity are also applied.

(II) The second test follows Muth (1960) and Follain (1979). Muth argued that if the market is elastic, P and Q should be independent in reduced form. Muth and Follain each basically examined the coefficient of the quantity of housing services in a reduced form price equation. They interpreted a t-test for Q as a test of elastic supply. One problem with this procedure is that, by itself, it cannot differentiate between perfectly elastic and perfectly inelastic markets since, in both cases, the slope of the function is not estimable with any precision.[3] Thus, we will not present detailed results from this test in the paper, although we did estimate this model and will briefly indicate our findings.

(III) This model, first developed in Malpezzi and Mayo (1996) from an insight by Mayo, follows Muth and Follain, taking advantage of some later suggestions of Olsen (1987). Consider the following three-equation flow model of the housing market:

$$Q_{\mathrm{D}} = \alpha_0 + \alpha_1 P_{\mathrm{h}} + \alpha_2 Y + \alpha_3 D$$

$$Q_{\mathrm{S}} = \beta_0 + \beta_1 P_{\mathrm{h}}$$

$$Q_{\mathrm{D}} = Q_{\mathrm{s}}.$$

For convenience, let us work in the natural logarithms of all variables. Among other advantages, we can interpret coefficients (approximately) as elasticities. The variables are defined as Q_{D} is the log quantity of housing demanded, Q_{S} the log quantity of housing supplied, P_{h} the log of the relative price per unit of housing, Y the log of income, and D the log of population.

[3] In one case, the estimated coefficient may be close to 0 because the true slope is 0. In the other case, the coefficient may not be significantly different from 0 because there may be insufficient variation in the quantity supplied (fixed, or nearly so). Hence standard errors will be arbitrarily large.

Note in passing that the Muth-Follain test derives from the fact that if β_1 is infinite then $\alpha_1=0$. The reduced form of the system can be found by equating supply and demand and solving for the observable variable, P_h, the price of housing. This yields

$$P_h = \frac{\alpha_0 - \beta_0}{\beta_1 - \alpha_1} + \frac{\alpha_2}{\beta_1 - \alpha_1} Y + \frac{\alpha_3}{\beta_1 - \alpha_1} D.$$

Clearly, even neglecting the constant term, the coefficients of the right-hand side variables are not identified, but if we estimated the total coefficient of income, and already knew α_1 and α_2, clearly we could derive the parameter of interest.

Thus our procedure is to estimate the simple reduced form equations, and to identify the key underlying parameter of this investigation, the price elasticity of housing supply, based on parametric estimates of housing demand parameters from the literature. Recalling that all variables are in logs, and making the reduced form stochastic, we estimate that

$$P_h = \gamma_0 + \gamma_1 Y + \gamma_2 D + \varepsilon.$$

Thus we estimate of the price elasticity of housing supply as

$$\beta_1 = \frac{\alpha_2}{\gamma_1} + \alpha_1,$$

where γ_1 is the estimated elasticity of housing price with respect to income, and the parameters α_1 and α_2 are parametrically assumed. Since we don't know the elasticity with certainty, we will actually calculate β_1 for a range of assumptions, assuming the price elasticity of housing demand to lie in the interval -0.5 and -1, and assuming the long-run income elasticity of demand to be alternately 0.5 and 1.0. These values encompass most of the econometric estimates extant for our two countries (Mayo 1981, Wilkinson 1973). Furthermore, a fascinating study by Haines and Goodman (1992) demonstrates that, at least in the United States, demand parameters around the turn of the 19th century were remarkably similar to those a century later.

A brief analysis of the relationship among these parameters is instructive. We start with a well-founded prior: that the price elasticity of supply is between 0 and infinity. Now consider Fig. 1, which presents the transformation of regression coefficient γ_1 into price elasticity estimate β_1 (for given values of α_2 and α_1). Figure 1 shows that, if our estimate of coefficient α_1 is unconstrained, there are regions in which the transformation into price elasticity space is not "well behaved." Only in the upper right quadrant of the graph (positive γ_2) is the transformation sensible, becoming more elastic as γ_1 approaches zero from the left, and becoming inelastic as it becomes "more positive."

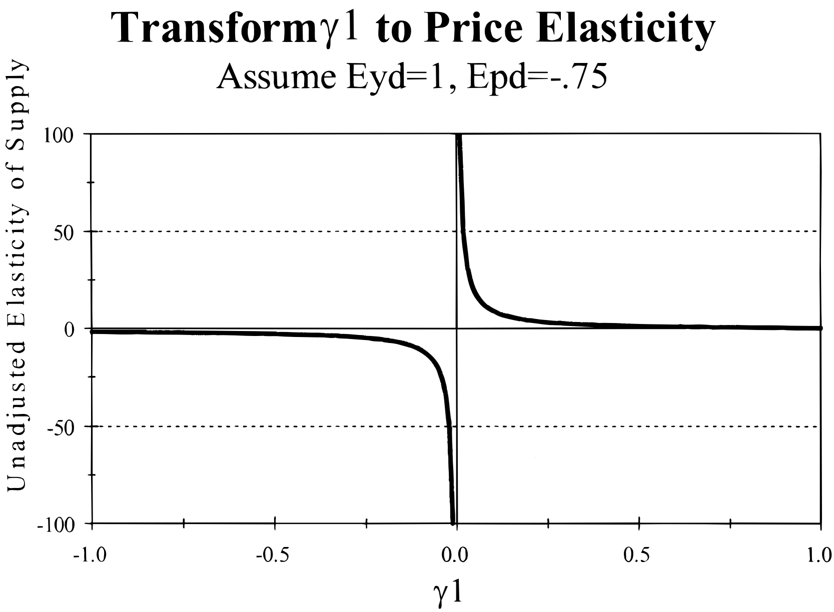

FIG. 1. Transform γ_1 to price elasticity. Assume Eyd $=1$, Epd $= -0.75$.

Thus, if the coefficient of income in the reduced form is negative, a naive interpretation of the results would be a downward-sloping supply curve for housing. Since coefficients are estimated, if we obtain negative coefficients we reject this interpretation on theoretical grounds. Rather, we interpret this as most consistent with perfectly elastic supply. In some other cases, our parameterized estimate of the supply elasticity could be, by naive arithmetic, "less inelastic than perfectly inelastic." On similar grounds, we would set such results equal to 0.

(IV) Our fourth procedure is an extension of the third. The flow model has a restrictive assumption embedded: that all adjustment takes place in a single year. A stock adjustment model may be preferred, given housing's durable nature, construction lags, and significant transactions costs. We can write

$$Q_D = \delta(K^* - K_{-1})$$

$$K^* = \alpha_0 + \alpha_1 P_h + \alpha_2 Y + \alpha_3 D$$

$$Q_S = \beta_0 + \beta_1 P_h$$

$$Q_D = Q_S,$$

where K_{-1} is the stock of housing in the preceding period, and K^* is the desired stock and δ is the adjustment per period.

Substituting the expression for K^* into the expression for Q_D, again equating supply and demand, and solving for P_h, yields

$$P_h = \frac{\delta\alpha_0 - \beta_0}{\beta_1 - \delta\alpha_1} + \frac{\delta\alpha_2}{\beta_1 - \delta\alpha_1} Y + \frac{\delta\alpha_3}{\beta_1 - \delta\alpha_1} D - \frac{\delta}{\beta_1 - \delta\alpha_1} K_{-1}.$$

Once again we estimate a reduced form

$$P_h = \gamma_0 + \gamma_1 Y + \gamma_2 D + \gamma_3 K_{-1} + \varepsilon,$$

but from the coefficient of income the price elasticity of housing supply is now

$$\beta_1 = \frac{\delta\alpha_2}{\gamma_1} + \delta\alpha_1,$$

where parameters α_1 and α_2 are parametrically assumed as before. The adjustment parameter δ can be varied parametrically. Our baseline estimate of δ is 0.3, following Muth (1960).[4] Note that if $=1$, model (IV) reduces to model (III).[5]

III. THE HOUSING MARKETS OF THE UNITED STATES AND THE UK

Figure 2 presents GNP *per capita* for the United States and the UK respectively. Despite the common notion that the postwar growth was "unprecedented," U.S. real output *per capita* grew faster in the prewar/pre-Depression period. Omitting the Depression and World War II years, the simple average of annual growth rates was 2.1% in 1890–1929, and 1.7% in 1947–1993. However, the decrease

[4]Few explicit estimates of the stock adjustment parameter are extant. A particularly careful development of a stock adjustment model can be found in Hanushek and Quigley (1979), along with estimates of the parameter δ somewhat higher than Muth's. However, Hanushek and Quigley derive their estimates from household level data rather than aggregate as used here, and as they point out, it is not clear that such parameters "aggregate up." In any event it is prudent to use the lower estimate of to see how much difference explicit stock adjustment could reasonably make in supply elasticity estimation.

[5]Of course, when actually estimating models III and IV in turn, the estimated coefficient of income will change, so the stock adjustment model with $=1$ will generally yield numerical results different from those of the flow model.

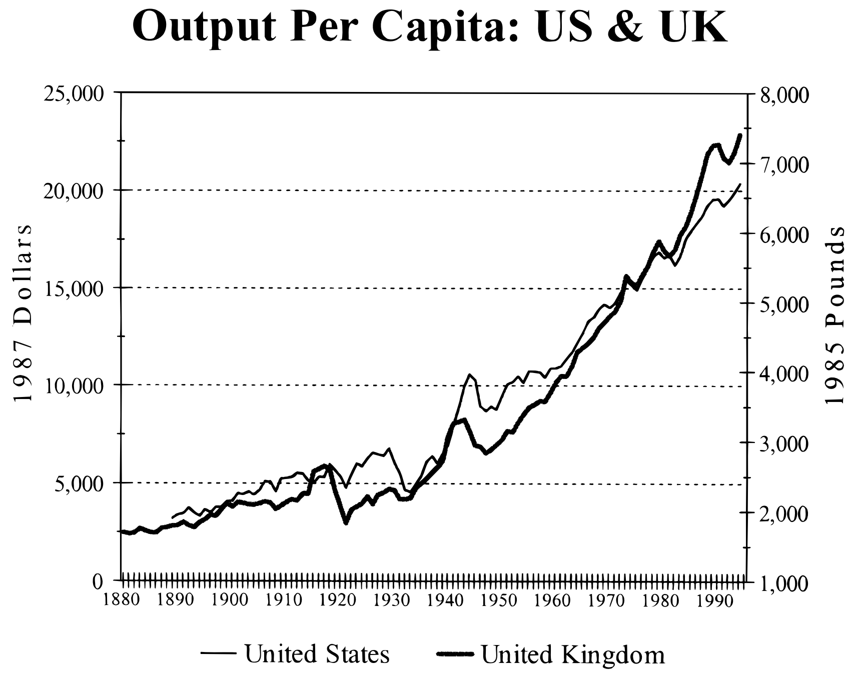

FIG. 2. Output per capita: United States (thin line) and UK (thick line).

in the variance of growth rates was dramatic. The standard deviation of annual growth rates was 6.3% in the prewar period, and only 2.6% in the postwar period.

The simple average of annual prewar UK real growth rates was much less, 1.3% *per annum*. Postwar rates averaged a healthier 1.9%. However, like the United States, the UK experienced a decline in the *variance* of growth rates; the standard deviations of annual rates were 3.2% prewar and 2.4% postwar.

It should be noted that these increases in *per capita* output are measured in each country's local currency. While significant fluctuations in both directions occurred, generally the pound fell relative to the dollar over the period under study. In 1880, for example, the pound was worth \$4.87, while as of 1994 it was worth \$1.53. Also, total GNP grew relatively faster in the United States over the period, as population growth was significantly faster. The UK's population grew from about 35 million in 1880 to 58 million in 1994, an annual rate of just under one-half percent *per annum*. U.S. population grew from 50 million in 1880 to 260 million in 1994, which is an annual rate of 1.5%.

Of course not all GNP is immediately available as household purchasing power in a given period. For the postwar period we have the advantage of data on disposable personal income *per capita* for both the United States and the UK.

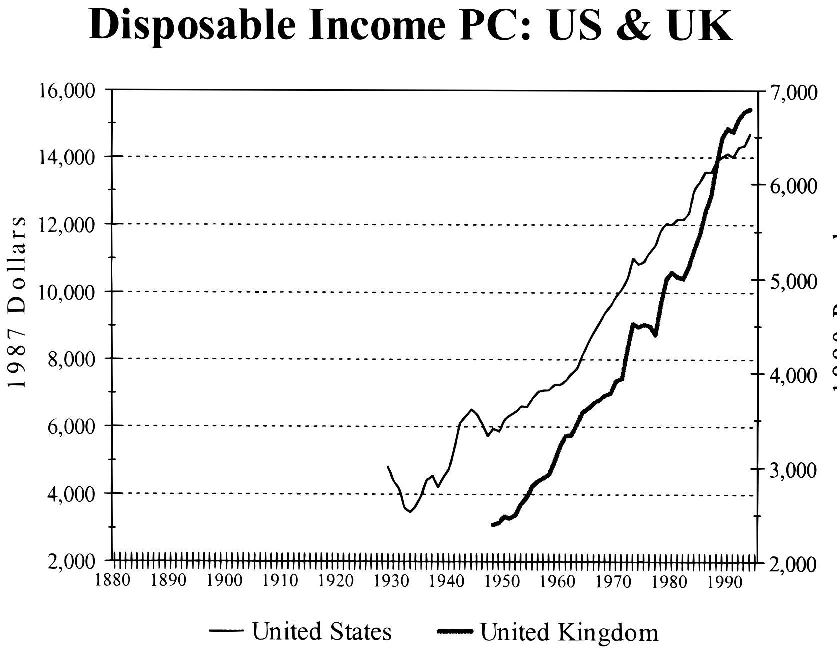

FIG. 3. Disposable income per capita: United States (thin line) and UK (thick line).

Figure 3 presents these data. The correlation between these data and their respective national GNP *per capita* figures are over 0.99 for both countries.

Housing starts (United States) and completions (UK) are shown in Fig. 4.[6] In both, there has been a substantial shift upward in this century. The initial shift was earlier in the UK, in the 1920s, followed by a precipitous decline. The United States also had a (proportionately smaller) shift in the 1920s, a more moderate decline and recovery, followed by the postwar plateau.

Another measure of housing is the real value of residential investment. This is shown in Fig. 5. This measure is more inclusive, as it comprises major additions and renovations as well as new units. Patterns are roughly similar, in that there is a shift around the great wars/Depression. Note the divergence between the countries for both measures of housing output during the 1930s, when the UK embarked on a massive program of slum clearance and housebuilding, including council housing.

In the United States, starts and residential investment are highly correlated for most years, but in the UK completions and real residential investment more often

[6]Starts data were readily available for a long time series for the United States, and completions for the UK. For each country, where starts and completions overlap they are highly correlated, both contemporaneously and with a one-year lag in starts.

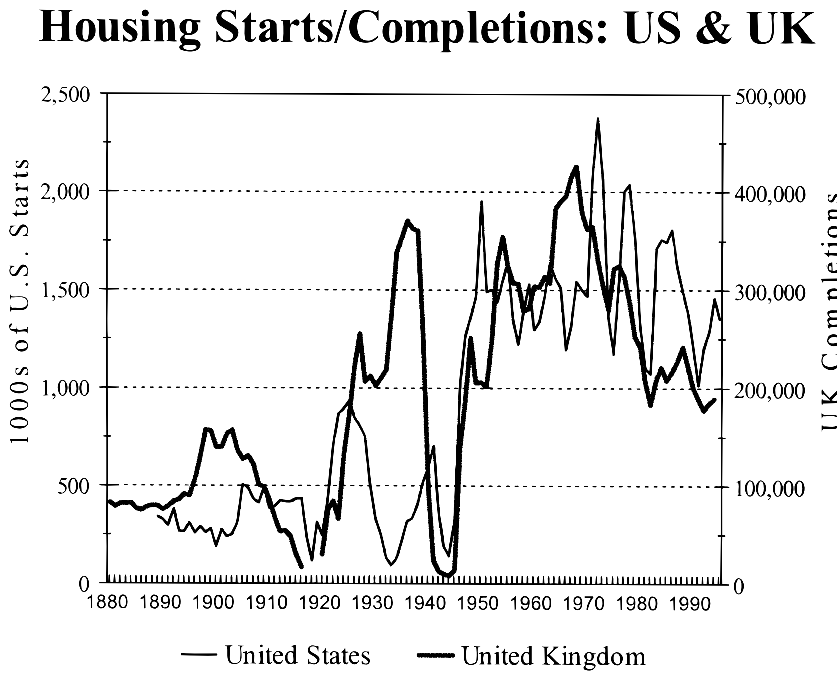

FIG. 4. Housing starts/completions: United States (thin line) and UK (thick line).

diverge. In particular, in the postwar period both series tracked closely in the UK up until about 1965. After 1965 real residential gross fixed capital formation remained near its peak while completions fell substantially. Further work is needed to find out to what extent this reflects a shift to improvements *in situ*.

Finally, we present our bellwether, the relative price of housing. The first, in Fig. 6, is based on national incomes accounts (NIA) deflators. Among studies that have used such cost-based indexes are Follain (1979). Conceptually, for many purposes we prefer transactions-based indexes to such cost-based indexes, but only cost-based indicators are available over such a long run. We use median house prices sold. Among studies that use similar transactions-based indicators are Topel and Rosen (1988) and Poterba (1991).

Of course there are many types of transactions-based housing price indexes in turn. Major categories include simple medians (as here), Laspayres or chain price indexes (such as the familiar rental consumer price indexes), indexes based on hedonic equations (such as Malpezzi *et al.* (1980)), and repeat sales indexes (such as those of Case and Shiller (1987)) Green and Malpezzi (2001) provide a review. Here we merely note that in practice most of these indices are highly correlated, although if data are available, repeat sales and hedonics have advantages over medians. The most important advantage is that, done properly, these

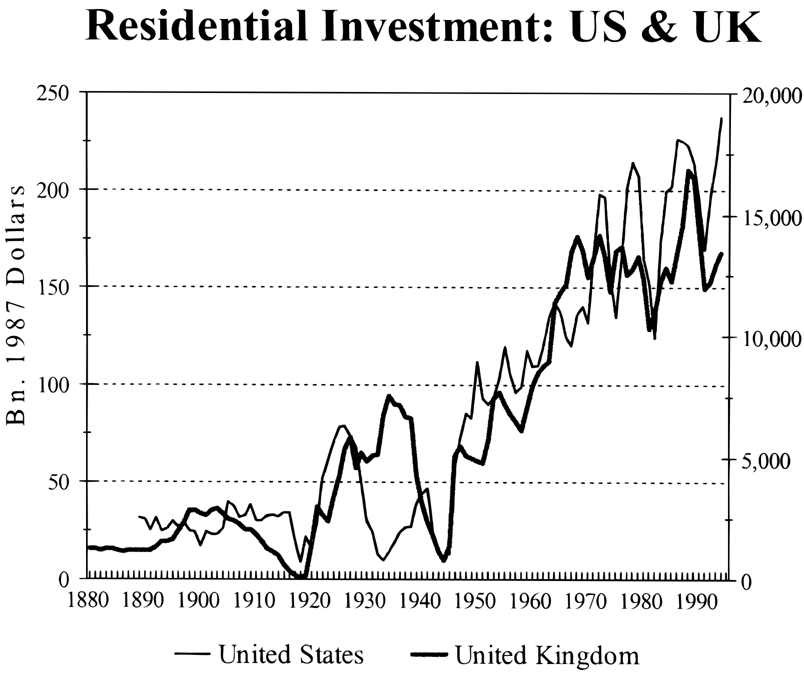

FIG. 5. Residential investment: United States (thin line) and UK (thick line).

latter methods adjust for changes in quality over time, while the simple medians do not. Hendershott and Thibodeau (1990) studied this problem and concluded that in the United States median house prices overstated true quality-adjusted price increases by about 2% *per annum*. We will make use of this adjustment below.

Turning now to our data, for overlapping years the two kinds of indexes are correlated, but more strongly in the United States than in the UK. Figures 7 and 8 shows real transactions-based price indexes for new housing for each country.

Another way in which these indexes differ is that the NIA cost-based indexes are driven by construction costs (hence "new construction" in our title, and in Follain's). Land development costs are included in such indexes, but not the cost of the raw land. The transactions-based indexes, on the other hand, comprise both construction costs and total land costs.

Given these differences, it is interesting that in the years for which we have matching data, the cost-based and transactions-based indexes are still strongly correlated both in changes and in levels. The turning points also match rather well. The most pronounced pattern is that the transactions-based indexes are more volatile, as we would expect. Thus, the cost-based indexes are less appropriate for studying short-run phenomena than long run. Transactions-based indexes would be particularly preferred whenever second moments are particularly critical, for

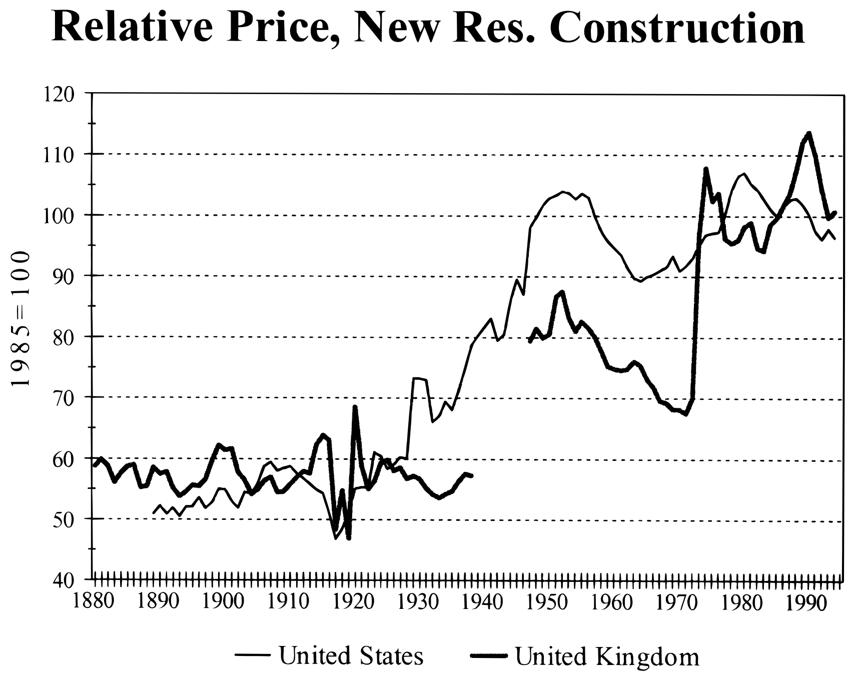

FIG. 6. Relative price and new residential construction: United States (thin line) and UK (thick line).

example, when examining property returns in portfolio models, but of course the present study focuses on long-run relationships.

First let us examine our longest time series, of cost-based new construction price measures, presented in Fig. 6. In both countries, the 19th century was flat. There was a sharp rise in the 1920s—but much sharper in the UK than in the United States. The 1920s rise was followed by a 1930s bust, and recent fluctuations around 100. Note the high variability of both series, but especially the UK data in the 1920s and 1930s.

Note that for the United States, there was no strong trend in the price data for the period studied by Follain (1947–1975), consistent with his findings of an elastic market. There was a trend in price for the period studied by Muth (1915–1935, with war years omitted), but it was uncorrelated with starts and investment, which were highly volatile during that period.[7] There was a fairly pronounced upward trend from 1963 to the early/mid-1980s, so it is not surprising that when studying this period Poterba (using 1963–1990 data) and Topel and Rosen (using 1963–1984 data) found evidence of a rising supply price. The fact that long-run

[7]It was this lack of correlation that Muth interpreted as evidence of perfect elasticity.

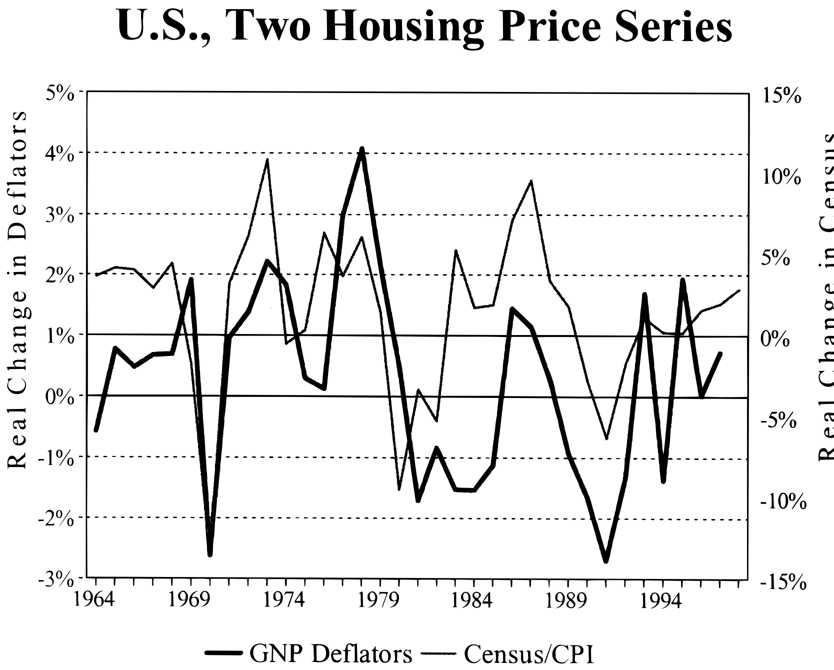

FIG. 7. Two housing price series, United States: GNP deflators (thick line) and census/CPI (thin line).

data reveal much more than the comparatively short series previously studied confirms the potential usefulness of our long-run approach. It suggests, in fact, that in the housing market the long run is likely to be long indeed.[8]

IV. ESTIMATION

The model was estimated using annual data, described in detail in the Appendix. United Kingdom data were available for 1850 to 1995, and U.S. data were available for 1889 to 1994.

Unless otherwise mentioned explicitly, following Follain, our residential price measure is the NIA deflator for residential construction, deflated by the GDP deflator. We also study price indexes based on medians, deflated and adjusted downwards by 2% *per annum* to correct for quality changes, following Hendershott and Thibodeau (1990).[9] We use two residential output measures: (1) the

[8]See Ball and Wood (1999) for a long-run look at output, in contrast to our examination of prices.

[9]Quality change could have, of course, been different in the UK, but we know of no similar estimates for that country.

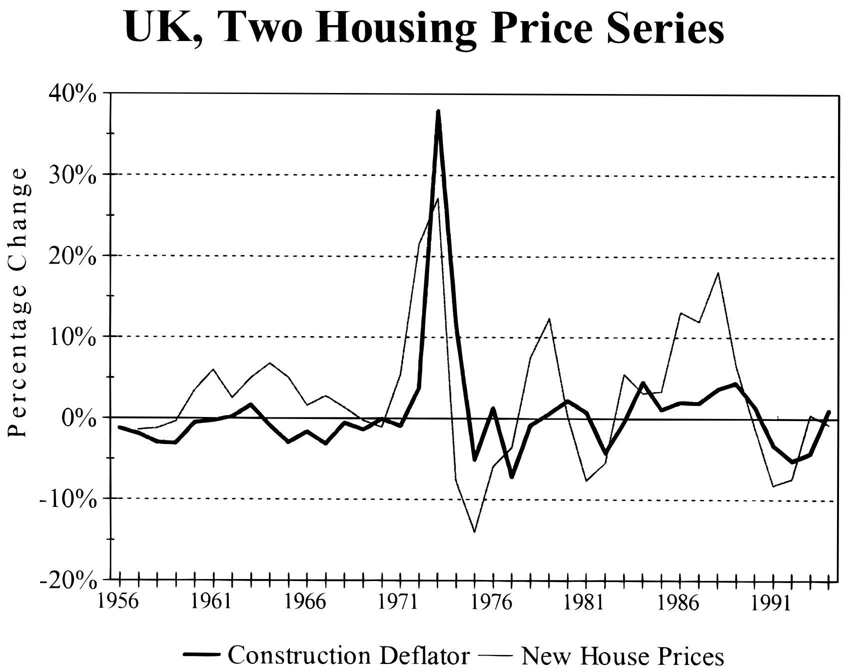

FIG. 8. Two housing price series, UK: Construction deflator (thick line) and new house prices (thin line).

real value of residential construction in each country and (2) either starts (United States) or completions (UK). Our primary income measure is real GDP (UK) or GNP (United States) per capita. For both countries, our alternative measure (available for fewer years) is real disposable income per capita. Demographics are proxied simply by population. Data are discussed further in the Appendix.

Pretests of the Data

Our statistical analysis maintains certain classical assumptions that are often violated with time series data. In this section we briefly discuss some of the tests undertaken prior to analysis.[10]

We began by plotting autocorrelation functions and computing augmented Dickey–Fuller tests for unit roots for each data series.[11] The unit root test is in effect a test of the hypothesis that data are a random walk. As a random walk

[10]Many discussions of these tests are available. See, for example, Granger and Newbold (1986) or Pindyck and Rubinfeld (1991).

[11]Critical values for the test are taken from MacKinnon (1991), from tables provided in Eviews econometric software.

does not have a finite variance, our usual significance tests will be incorrect. If we cannot reject that hypothesis, we might obtain spurious results.[12]

We can see from visual inspection of Fig. 6 that prices do not *look* stationary in either country; that is, it does not appear that the data have finite means and variances that are invariant with respect to time. However, neither are variances obviously exploding.[13]

Another issue to examine further is the possible break during the "war period" (1914 to 1947). One way to think about a stationary series is that it reverts to some mean, or perhaps a trend, after some shock has moved it away from said mean or trend. A number of papers have pointed to the World Wars and Great Depression as *persistent* shocks that have moved series away from means or trends permanently, but that other years experience smaller, "normal" shocks, which are transitory. Perron (1989) is perhaps the seminal work in this area; see Zelhorst and De Haan (1995) and Evans and Quigley (1995) as examples of other recent literature in this area. To us this suggests segmenting our estimates pre- and postwar, and omitting the war period.[14]

Another reason to proceed with formal tests is that if the data are reexpressed in logs (not shown), visual inspection suggests prices may be reverting to some trend value, so formal tests were conducted with both the untransformed variables and with logs. Since we will be working with logarithms in our models, here we focus on their stationarity tests.

Cost-Based Price Indexes. For the United States 1889–1994, we could not reject the hypothesis of a unit root (i.e., could not reject the hypothesis of a random walk). For the full sample, the Dickey–Fuller test statistic was -1.46. Critical values are -3.15 at 10% Type I error, and -4.04 at 1%.[15] However, as we noted above the power of the Dickey–Fuller test is not generally regarded as large.

For the UK 1850–1995, the test statistic is -1.92. Critical values are as above, so we cannot reject the hypothesis of a unit root in UK prices over the entire period.

[12]NB, as in Pindyck and Rubinfeld (1991, p. 462): "The unit root test allows us to reject (or fail to reject) the hypothesis that a variable is *not* a random walk. A failure to reject is . . . only weak evidence in favor of the random walk hypothesis."

[13]Obviously the World War–Great Depression years 1914–1947 (hereafter the "war period") have a much higher variance, but postwar the second moments damped down.

[14]Dropping the war period is also consistent with the concerns raised above about the quality of data from these periods. Of course the tradeoff is that by dropping such data we may be closer to maintained hypotheses about error terms, and losing poor-quality data, we may also be losing valuable information. We did estimate models over the entire period for comparison but will not present them here.

[15]Tables in papers such as MacKinnon (1991), and in our software, give critical values at two or three conventional "significance levels." Little is known about the loss functions involved in these tests. Bayesian measures, such as the associated probability of a test statistic so large under the null, are not readily available.

Two or more series that are individually nonstationary may be cointegrated, in which case error terms from a regression equation—a linear combination of the variables—will be stationary.[16] It is the stationarity of the error term that is required for validity of our additional tests.

We compute the Johansen cointegration test statistic for the cointegrating vector of the log of real housing price and our two candidate-independent variables (population and real income per capita, plus a trend term). For the United States 1889–1994, the statistic is 27.42. The critical values are 29.68 at 5% and 35.65 at 1%, so we fail to reject no cointegration at the 5% level; it is possible these variables are "drifting apart."

For the UK 1850–1995, the test statistic is 30.17, compared to a 1% critical value of 35.65 and a 5% critical value of 29.68, so we can reject the hypothesis of "no cointegration," at conventional levels.

Next we turn to pretests performed separately on prewar and postwar data. For the United States, we cannot reject the unit root hypothesis in log price with prewar (pre-1914) data. The test statistic is −1.95, and the 10% critical value is −3.26. However we can reject the hypothesis that the vectors log price, log GDP per capita, and log population are not cointegrated. The test statistic is 29.77, and the 5% critical value is 29.68. For the U.S. postwar (post-1947), we can reject the unit root hypothesis in log price (test statistic is −3.71, and 5% critical value is −3.50). We also find a cointegrating vector, the Johansen test statistic, is 44.31 while the 1% critical value is 35.65.

For the UK, we cannot reject the unit root hypothesis in log price with prewar data. The test statistic is −2.31, and the 10% critical value is −3.17. However we can reject the hypothesis that the vectors log price, log GDP per capita, and log population are not cointegrated. The test statistic is 41.40, and the 1% critical value is 35.65. For the postwar UK, we can not reject the unit root hypothesis in log price. The test statistic is −2.64, and the 10% critical value is −3.18. Strictly speaking, in a classical sense, we cannot reject the hypothesis of no cointegrating vector in the postwar UK. The Johansen test statistic is 28.18 while the 5% critical value is 29.68.[17] Somewhat loosely, we would argue that the test statistic is "close to" the cutoff, such that if we estimate regression relationships in the face of such results there might be a roughly 1 in 20 chance that we might be faced with "uncointegrated" vectors, with possible attendant spurious correlations.

Transactions Based Price Indexes. These measures are only available postwar (post-1963 in the United States, and post-1956 in the UK). For this test, we also use the real disposable income per capita series, since it is arguably better than GDP per capita, and is available for the relevant period. We find that, for the

[16]More precisely, if the true error terms are stationary, and if we have the correct model, the estimated errors (the residuals) should exhibit stationary behavior.

[17]Available software and tables do not provide, say, cutoffs at 10% levels; and as noted above there are no associated probabilities readily available with these statistics.

United States, we can reject the unit root hypothesis in this log price series at the 10% level (test statistic −3.43, 5% cutoff −3.57, 10% cutoff −3.22). We can reject the “no cointegration” hypothesis (Johansen statistic 31.41, compared to a 5% cutoff of 29.68).

For the UK, we can reject the unit root hypothesis in this log price series at the 1% level (test statistic −4.71, 1% cutoff −4.22). We can also reject the “no cointegration” hypothesis (Johansen statistic 59.48, compared to a 1% cutoff of 35.65).

Given all these tests, we are left with several conclusions. If we estimate these models with the full span of the data, minimum variances and *t*-tests must be interpreted cautiously. However, we note again that the inability to reject random walks might be due to the low power of the tests. Generally, however, if we estimate our models as segmented by pre- and postwar periods (outside 1914–1947), our cointegration tests give us confidence in the results, for the latter reason, and because we have a strong prior that the interwar period is a shock, or a series of related shocks, persistent and unlike any other in the data, or likely to be encountered in the forseeable future. Note that this still leaves us with some long-time series for estimation, approaching 50 years postwar. None of the other papers surveyed on this topic approach such a span.

The econometric literature suggests that problems of spurious correlation from nonstationarity are more serious in short-time periods. Some studies find processes revert to means or trends only with long lags, a point to which we return below. See, for example, Frankel and Meese (1987), who find the U.S./UK real exchange rate reverts to trend, but only over a century.

Another issue, noted by Perron (1989) among others, is that a process might be generally stationary, i.e., shocks are usually temporary and the series reverts to means or trends; that there are shocks and there are *shocks*, occasional shocks that are so great as to be persistent. We argue that the war years 1914–1947 can be so characterized, and the data are consistent with such an interpretation.[18]

Model I (Time Trend) Results

The simplest model is a *t*-test of the hypothesis that price is constant over time, i.e., that a regression of real relative price of housing against a time trend yields a zero coefficient. Table I presents these results for a variety of periods for the both countries. For the United States, there is obviously a positive trend in real relative prices over the entire period, and the prewar (pre-1913) period. However, there is no trend in the postwar period. For the 1947–1994 sample the *t*-statistic for the time trend is 0.69; the probability of observing such a large

[18]Another possible explanation for the amazing apparent volatility and persistence of war year era shocks is that during such periods national income accounts are unreliable. See, for example, debates on data in Legerbott (1986), Romer (1986), Weir (1986), and Higgs (1992).

t under the null is 0.49. Thus the simplest test is consistent with an elastic *postwar* supply.

Table I also gives clues to some of the results from other literature. Recall from Fig. 6 that the cycles in relative housing prices are long, and our argument is that results will be sensitive to the time span studied. Holding the price measure constant (log relative price of new residential construction) we can vary the period to match the data from different studies. The trend from Follain's (1979) study years is actually *negative*; his conclusion that the market is elastic is therefore unsurprising. Using the same variable, but for the "Topel and Rosen years" (1963 to 1983) and the "Poterba years" (1963 to 1990), we find rising prices.

Of course results can vary due to choice of variable as well as time. The last two rows of Table I yield some insights. The penultimate row presents the time trend for the alternate price measure for each country, basically the real median price of new houses, undadjusted for quality. These data are available after 1963 in the United States and after 1947 in the UK. Note that the U.S. median house price results imply a 1.3% *per annum* increase over the span of the data; using the new construction series yields a lower estimate of about 0.4% *per annum*, but recall that the medians are unadjusted for quality. If we adopt the Hendershott–Thibodeau adjustment, discussed above, and subtract 2% *per annum*, we see that prices are not rising when this other series is used. If the adjustment is accurate, in fact, prices are falling.

TABLE I
Time Trend Coefficients Log Relative Price of New Construction

	United States	United Kingdom
All years	0.0082 (24.04)	0.0049 (18.14)
Pre-1913	0.0061 (8.23)	0.0017 (3.14)
Post-1947	0.0003 (0.69)	0.0077 (6.42)
"Follain years" (1947–1975)	−0.0043 (−5.03)	
"Topel and Rosen years" (1963–1983)	0.0096 (12.10)	
"Poterba years" (1963–1990)	0.0061 (8.23)	
Postwar, using log median real price of housing*	0.0133 (10.10)	0.0280 (17.40)
New construction, using same years as median real price of housing*	0.0040 (4.98)	0.0121 (8.04)

Note. t-statistics in parentheses. Years are 1947 to 1994 for UK, 1963 to 1995 for United States. Median prices are unadjusted for quality change, see text.

Turning now to the UK data, we see that prices are rising overall, in the prewar period, and (unlike the United States) in the postwar period. While the trend averages about 0.8% *per annum* postwar, Fig. 6 clearly showed there are substantial deviations from trend in most years.

These simple time trend results are informative, but by themselves do not yield numerical estimates of elasticities. We next turn to models III and IV, which do. First we will discuss the basic regression results, then derive some elasticity estimates.

Model III and IV Results: United States

Table II presents the regression estimates for models III and IV for the United States.[19] For all models, but especially for the flow models, inspection of residual plots indicated a high degree of serial correlation. We therefore estimated each model once using OLS, and once using a Cochrane–Orcutt correction for serial correlation (the AR(1) term in the tables).

TABLE II
Selected U.S. Regressions: Dependent Variable: Log Relative Price of New Residential Construction

Sample	Prewar	Prewar	Postwar	Postwar	Postwar	Postwar
Log GDP PC						
coef	0.088	0.092	0.682	0.076	− 0.035	0.101
se	0.127	0.110	0.166	0.130	0.146	0.138
prob > *t*	0.493	0.413	0.000	0.561	0.809	0.467
Log Population						
coef	0.215	0.193	−0.981	−0.336	−4.676	−2.851
se	0.161	0.158	0.243	0.282	0.514	0.656
prob > *t*	0.195	0.235	0.000	0.239	0.000	0.000
Log stock(−1)						
coef					1.780	0.996
se					0.235	0.267
prob > *t*					0.000	0.001
AR(1)						
coef		0.525		0.857		0.627
se		0.210		0.062		0.080
prob > *t*		0.021		0.000		0.000
Constant						
coef	−3.794	−3.578	5.425	3.365	43.556	26.064
se	0.827	1.110	1.424	2.720	5.123	6.214
prob > *t*	0.000	0.004	0.000	0.223	0.000	0.000
R^2	0.75	0.80	0.27	0.82	0.68	0.84
No. of observations	25	24	48	48	48	48
DW	0.96	1.79	0.19	0.78	0.33	0.60

[19] As noted above, model II is omitted.

Prewar (pre-1914) the simple flow models yielded fairly low coefficients on income, insignificantly different from 0, suggesting fairly elastic supply. The AR(1) term, while significant, did not make a qualitatively large difference in the estimation of this model. We were unable to estimate stock adjustment models for the prewar period, because of the paucity of good stock data.

Postwar (post-1947), the OLS estimate of the simple flow model yielded a much higher coefficient on log income, significantly different from 0. Interestingly, the autoregressive correction made a very large difference in the results, including the income coefficient, which drops down by a factor of over 8, and becomes insignificant. Given the large and regular cycles evident postwar in Fig. 6, above, perhaps the strong performance of the autoregressive term is unsurprising.

The postwar stock adjustment model also yields insignificant income coefficients for the United States. The point estimate of the OLS model is actually negative, though insignificantly different from 0. The point estimate of the Cochrane–Orcutt estimate is positive but again insignificant. Perhaps it is not surprising that the AR(1) term makes a much bigger difference in the flow model than in the stock adjustment model. Serial correlation may be due to omitted variables (among other causes), and it could be argued that the stock adjustment model, while simple, is “more complete.” Thus there would be less omitted variable effect for the AR(1) term to pick up.

Taken as a whole, with the exception of the OLS postwar flow estimates, these results are consistent with a fairly elastic long-run supply of housing.

Model III and IV Results: UK

Table III presents the corresponding regression estimates for the UK. Once again we estimated each model once using OLS and once using a Cochrane–Orcutt correction.

Prewar the simple flow models yielded somewhat higher coefficients on income than in the United States. The correction for serial correlation also makes more of a difference in the UK prewar; with the correction, the coefficient of income is not significantly different from 0, but without it the coefficient is larger and significant.

The postwar OLS estimate of the flow model yielded much higher coefficients of income in the UK than in the United States significantly different from 0, with or without the AR(1) correction. The autoregressive correction did reduce the size of the income coefficient, as in the United States, but by much less in proportionate terms.

In contrast to the United States, the postwar stock adjustment model for the UK yields positive and significant income coefficients. Once again, the Cochrane–Orcutt correction makes less difference for the stock adjustment models than for the flow models.

From these estimates, it appears that supply in the UK was more elastic prewar

TABLE III
Selected UK Regressions: Dependent Variable: Log Relative Price of New Residential Construction

Sample	Prewar	Prewar	Postwar	Postwar	Postwar	Postwar
Log GDP PC						
coef	0.558	0.209	1.436	1.012	0.799	0.728
se	0.215	0.275	0.196	0.314	0.249	0.354
prob > t	0.012	0.452	0.000	0.002	0.003	0.046
Log Population						
coef	−0.462	0.007	−6.453	−3.760	−9.066	−6.540
se	0.259	0.336	1.145	2.090	1.278	2.032
prob > t	0.081	0.983	0.000	0.079	0.000	0.003
Log Stock(−1)						
coef					0.850	0.628
se					0.236	0.273
prob > t					0.001	0.027
AR(1)						
coef		0.558		0.837		0.794
se		0.109		0.089		0.105
prob > t		0.000		0.000		0.000
Constant						
coef	4.715	2.386	62.687	36.891	91.894	66.145
se	1.212	1.723	10.888	20.613	12.953	20.272
prob > t	0.000	0.172	0.000	0.081	0.000	0.002
R^2	0.25	0.51	0.68	0.89	0.76	0.91
No. of observations	58	57	48	47	46	45
DW	0.93	2.10	0.38	1.20	0.44	1.18

than postwar. Taken as a whole, the postwar estimates suggest a less than perfectly elastic long-run supply of housing in the UK.

V. THE PRICE ELASTICITY OF SUPPLY IN THE UNITED STATES AND UK COMPARED

As discussed above, we can calculate the value of β_1 under a range of assumptions about α_1 and α_2 (and for the stock adjustment model). Table IV presents some representative results. All calculations in the table are from regression results including the Cochrane–Orcutt correction.

In the prewar United States our implied price elasticities from the flow models are between 4 and 10; postwar it is between 6 and 13. In the prewar UK our implied price elasticity from flow models is between 1 and 4; postwar it is between 0 and less than 1. Postwar stock adjustment elasticities are lower in both countries. They range from a little over 1 to 6 for the United States and from 0 to less than 1 for the UK.

TABLE IV
Supply Price Elasticities from Selected Estimates

	United States	United Kingdom
Prewar, model III		
$\alpha_1 = -0.5,\ \alpha_2 = 1$	10.4	4.3
$\alpha_1 = -0.1,\ \alpha_2 = 1$	9.9	3.8
$\alpha_1 = -0.5,\ \alpha_2 = 0.5$	4.9	1.9
$\alpha_1 = -0.1,\ \alpha_2 = 0.5$	4.4	1.4
Postwar, model III		
$\alpha_1 = -0.5,\ \alpha_2 = 1$	12.7	0.5
$\alpha_1 = -0.1,\ \alpha_2 = 1$	12.2	0.0
$\alpha_1 = -0.5,\ \alpha_2 = 0.5$	6.1	0.0
$\alpha_1 = -0.1,\ \alpha_2 = 0.5$	5.6	0.0
Postwar, model IV		
$\alpha_1 = -0.5,\ \alpha_2 = 1,\ \delta = 0.3$	2.8	0.3
$\alpha_1 = -0.1,\ \alpha_2 = 1,\ \delta = 0.3$	2.7	0.1
$\alpha_1 = -0.5,\ \alpha_2 = 0.5,\ \delta = 0.3$	1.3	0.1
$\alpha_1 = -0.1,\ \alpha_2 = 0.5,\ \delta = 0.3$	1.2	0.0
$\alpha_1 = -0.5,\ \alpha_2 = 1,\ \delta = 0.6$	5.6	0.5
$\alpha_1 = -0.1,\ \alpha_2 = 1,\ \delta = 0.6$	5.4	0.2
$\alpha_1 = -0.5,\ \alpha_2 = 0.5,\ \delta = 0.6$	2.7	0.1
$\alpha_1 = -0.1,\ \alpha_2 = 0.5,\ \delta = 0.6$	2.4	0.0

One take on these results is that these intervals are very broad, and that in fact even they understate the imprecision of our knowledge of these elasticities, as these calculations held the coefficient estimate fixed, while they in fact have a distribution. Thus, unsurprisingly, better estimates of the key parameters α_1, α_2 and γ and of the coefficients will improve the precision of our estimates of the price elasticity of supply.

Another take is that, while the range of estimates presented is wide, there is a consistent pattern that the U.S. market is more elastic, for any given set of assumptions. It is also the case that we have deliberately picked a wide range of parameter values, whereas many readers will have there own, narrower priors. For these readers, “you pays your money and you takes your choice.”

Interpretation of Results

The United States and the UK have many similarities, including a broadly similar legal and institutional framework for real property. However, there are a number of important differences. Among others, the United States has directed its financial system more toward housing, has greater tax advantages, and has a less restrictive planning and regulatory environment.[20] For these reasons, we would not be surprised that the market for new residential construction is more responsive here.

In both countries, something is gained by considering a long time series of aggregate data, as we think our results show, but it is important to also note what is lost. In the UK the housing market in the southeast, centered on London, is often under greater demand pressure than, say, Newcastle or Wales. Tenure patterns and regulatory environments also vary by location, as documented in Maclennan *et al.* (1994) and Bramley (1993), among others. It is therefore hardly surprising that the time path of UK housing prices also varies significantly by region, as documented in Giussani and Hadjimatheou (1991), to give but one example. Broadly similar points can be made about the United States; see Goodman (1998) in general, and Abraham and Hendershott (1996) and Malpezzi (1999a) for analysis of local housing prices. This suggests large potential gains to estimating supply elasticities at the metropolitan or regional level, and analyzing the determinants of any differences, as begun in Green *et al.* (2000).

Perhaps the strongest result, and one of the most interesting, was the existence of a tremendous regime shift in both countries around the time of the World Wars and the Depression. While we can make conjectures about the undoubted economic dislocations of that long period and point to possible problems with data from such a volatile period, in the end this shift remains a tantalizing mystery ripe for future work.

The divergences between elasticities estimated from stock and flow models appear surprisingly high. More work on the relationship between the stock and flow models could prove particularly fruitful. More rigorous separation of short- and long-run elasticities would be an important component of this research program.

We also note that models like this one can and should be applied to other countries; once estimates are available for a wider range of places, more systematic analysis of the reasons for differences in market responsiveness becomes feasible. Finally, it is surprising that we know of no estimates of the supply elasticity for nonresidential real estate; these could also be developed with a model like ours.

VI. SUMMARY AND CONCLUSIONS

This paper has shown, *inter alia*, that:

- The United States and the UK have broadly similar patterns in residential new construction and prices, including a major regime shift shortly before the middle of this past century.

[20]For evidence and discussion of these propositions, see Bramley *et al.* (1999) and Evans (1999), among others.

- There is strong evidence of a “regime shift” in 1914–1947.
- Over the entire period, prices rise in both countries, but not in a continuous manner. Post-World War II, the United States is essentially flat, albeit with very large cycles. In the UK, relative housing prices generally rise postwar.
- According to our flow model, in the prewar United States our implied price elasticity is between 4 and 10; postwar it is between 6 and 13. In the prewar UK our implied price elasticity is between 1 and 4; postwar it is between 0 and 1.
- Stock adjustment models yield different price elasticities—surprisingly so, in our judgement. They range from 1 to 6 for the United States, and from 0 to 1 for the UK. We believe the stock adjustment models are particularly fruitful areas for additional work.

Limitations of the present study include the following. This study focuses on aggregate national data, when housing markets are local and diverse. We have not really investigated different lag structures, or explicit short-run/long-run multipliers. Thus, many issues remain for future research. Future research can, *inter alia*:

- Work more on identifying alternative lag structures, an explicit time-varying parameter model, and causality.
- Construct and analyze a long-run user cost measure, in contrast to the NIA measures.
- Estimate similar models for other countries, *inter alia* to permit better identification of the relationship between institutional features of the market and these outcomes.
- Estimate disaggregated models by region and/or metropolitan areas.
- Extend these models to nonresidential real estate.
- Estimate rates of quality change in the UK and other housing markets, following Hendershott and Thibodeau (1990).
- Study and estimate explicit stock adjustment parameters for both countries.

DATA APPENDIX

Main variables are printed below. Most have multiple sources—constant price variables were rebased to common years, and results were chained to obtain the full series. Simple transformations are not listed: a “D” prefix denotes an annual percentage change in a variable, and an “L” prefix denotes log.

United States

POP: Population, in 1000s. Source: *Historical Statistics of the U.S.*, updated with data from the *Economic Report of the President*, various issues.

GDPPCR: Real GDP Per Capita, in 1987 dollars. Source: *Economic Report of the President*, and *Historical Statistics*.

STARTS: Housing starts, in 1000s. Source: *Historical Statistics of the U.S.*, updated with data from the U.S. Department of Housing and Urban Development's *U.S. Housing Market Conditions*, various issues.

GPIRES87: Gross private domestic investment, residential, in 1987 dollars. Source: Grebler *et al.* (1956) and *Historical Statistics of the U.S.* (1889 to 1958); and *Economic Report of the President* (post-1959).

PDRES87: Price deflator for GPIRES87. Ratio of current to 1987 GPIRES. Source: same as GPIRES87.

PRNC87: Relative price of new construction: ratio of PDRES to the GNP Implicit Price Deflator. Source: Same as GPIRES87.

PNAREX: Median NAR existing single-family sales prices (nominal). Source: National Association of Realtors.

FRTWNRES: Net stock of fixed reproducible tangible wealth in residential capital, $ 1987 Billion. Source: BEA, *Fixed Reproducible Tangible Wealth in the United States*.

United Kingdom

POPUK: Population, in 1000s. Source: Mitchell (1988), and Central Statistical Office, *Annual Abstract of Statistics*, various issues.

GDPPC: Real GDP Per Capita, in 1985 pounds. Source: Mitchell (1988) and Central Statistical Office, *National Accounts Blue Book*, various issues.

COMPUK: Housing completions, in 1000s. Source: Source: Mitchell (1988), and HMSO, *Housing and Construction Statistics*, various issues.

DWGFC85: Gross private domestic investment, residential, in 1985 pounds. Source: Mitchell (1988) and Central Statistical Office, *National Accounts Blue Book*, various issues.

DWDFL85: Price deflator for DWGFC85. Ratio of current to 1985 DWGFC. Source: same as DWGFC85.

RPDW: Relative price of new construction: ratio of PDRES to the GNP Implicit Price Deflator.

NEWHP85: Average new house price, in 1985 pounds.

NSDEL75: Net stock of residential capital, at 1975 replacement cost.

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