Housing Supply and Price Adjustment

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Abstract

We analyse two inter-related features of regional housing markets: determinants of new housing supply, and the impact of supply responsiveness on price dynamics. We demonstrate that a suitably specified q-theory model (including residential land values as well as construction costs) explains intended housing starts. Few prior studies have found significant land price effects, due either to their omission or (possibly) to incorrect data definition (use of agricultural rather than residential land values). We examine the interaction of supply responsiveness and price adjustment following demand shocks, using a new panel dataset covering 53 quarters across 73 regions of New Zealand. Regions with high supply responsiveness have relatively small price spikes following demand shocks, consistent with a rational response that limits house price jumps in regions with strong supply responses.

JEL classification
R21; R31; R38

Keywords
Housing supply; q-theory; house price dynamics

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1 Introduction

Regional house price dynamics in response to demand shocks are of central policy importance. Resources may be misallocated where short run prices diverge from equilibrium; prospective purchasers (sellers) suffer where prices are higher (lower) than equilibrium. Social policy concerns relating to housing affordability are magnified where prices jump well above long run equilibria.

Given these concerns, we analyse two inter-related features of regional housing markets: new housing supply determinants and the dynamics of price adjustment. We treat housing supply as fixed in the very short run, so demand shocks are reflected initially only in house prices rather than in quantities.[1] New house supply reacts to house prices and to costs of developing new houses. Rational agents anticipate these supply responses in their pricing decisions following a demand shock; thus the dynamics of price adjustment are related to anticipated local house supply responsiveness.

We explicitly model the determinants of new housing supply. Despite its theoretical and intuitive attractiveness, previous studies have struggled to apply a Tobin's "q" approach to explaining investment in housing.[2] We demonstrate that, with a suitably specified model, a q-theory specification does satisfactorily explain intended housing starts. We apply the theory using a new panel dataset covering 53 quarters across all 73 regions of mainland New Zealand (3,869 observations). Subsequently, we examine the impact that supply responsiveness has on price dynamics. Several recent international studies have examined the impact of regulation on prices. Our study supplements these analyses by explicitly examining the interaction of supply responsiveness with the speed and degree of price adjustment following demand shocks.

[1] Thus housing supply can be treated as exogenous in the house price equation.
[2] Harter-Dreiman (2003), for instance, notes that a broad consensus about the supply elasticity of housing does not exist.

Our results demonstrate, first, that application of q-theory to housing supply should include not only the house price and construction costs (which are standard in most prior specifications) but also the cost of land. We find strong, statistically significant, impacts of each relevant variable - including land prices - on new housing supply. We estimate the hypothesised relationship initially with coefficients restricted across regions and subsequently without cross-region restrictions. The cross-region restrictions are not accepted, implying that supply elasticities are region-specific, even after accounting for region-specific land prices. This finding is consistent with the existence of regulatory differences across local authorities. Second, our results demonstrate that regions with relatively high supply responsiveness have relatively small price spikes following demand shocks. This is consistent with a rational response that limits the jump in house prices in regions that have strong supply responses. In areas with weak supply response, prices jump further since new supply is not forthcoming in the near term to meet the higher demand.

To illustrate the supply results, Figure 1 presents a scatter plot of the relationship between new housing consents (expressed as a percentage of the existing housing stock) versus the logarithm of house prices relative to development costs.[3] The scatter plot covers all 3,869 observations using de-meaned data for each region. A clear positive supply relationship can be discerned. The relationship between supply responsiveness and the degree of price adjustment following a demand shock is summarised in a scatter plot for the relevant 73 regional supply and price adjustment parameters in Figure 3. The relationships underlying this plot are discussed further in the paper and are analysed in Table 3; for now, the posited negative relationship between the two can be seen.

Section 2 of the paper outlines our application of q-theory to housing investment and relates previous housing supply estimates to this framework. We present our estimation results together with robustness checks on the results. Section 3 outlines our model for house price dynamics and sets out our hypothesis

[3] Development costs are calculated as a logarithmically weighted average of construction costs (2/3) and land costs (1/3) in line with estimates subsequently in the paper.

linking supply responsiveness to price dynamics. We estimate the specified equations and conduct tests of the relationship between supply responsiveness and price dynamics. Section 4 has concluding remarks with pointers to future work. An appendix describes the data used in the study.

2 House Supply Determinants

We treat housing developers as profit-seeking agents. A developer seeks to build a new house where the expected house sale price exceeds the full costs of developing and building the house. We assume that a house built in period t is destined for sale in t+1. The expected sale price in t+1, given information at time t, is denoted PH^e_{t+1}. The developer's total costs (TC_t) comprise land costs borne in period t (PL_t), building costs (materials and labour) in period t (PB_t) and financing costs (determined by r_t, where r_t is the nominal interest rate, adjusted for a risk premium, between t and t+1).

We model the planned rate of change in housing supply between t and t+1 as being equal to the rate of new housing consents granted in period t (HC_t) relative to the existing housing stock (H_{t-1}).[4] We use housing consents as a measure of planned changes since a new house can be constructed legally only following the granting of a consent by the relevant territorial local authority (TLA), which is the unit of analysis in the study. Expressing the relationship in log-linear form, and allowing the coefficients (γ_{i0}, γ_{i1}) to be potentially region-specific,[5] we hypothesise that new housing supply for each region i is given by equation (1) (in which ε_{it} is an iid error term):

$$HC_{it}/H_{it-1} = \gamma_{i0} + \gamma_{i1}\ln\{PH^e_{it+1}/TC_{it}\} + \varepsilon_{it} \qquad (1)$$

We model expected house prices, PH^e_{it+1}, as a function of existing house prices in region i at period t, together with a region-specific growth factor

[4] We use H_{t-1} rather than H_t in the denominator since the former is known at time t. Since H is a stock variable, it makes virtually no difference which of the two we use. Also note that $HC_t/H_{t-1} \cong \Delta \ln H_t$ provided all consents in period t are converted into new houses in period t and provided there is no scrapping of existing houses. In practice, there is some scrapping and some consents are not actioned; these effects are minor and can be catered for in the time and region fixed effects in the estimated equations (where the effects are consistent over time and/or over regions) as well as in the error term (where the effects are random).
[5] Subsequently, we treat the term γ_{i0} as a nationwide constant, γ_0, plus a vector of area fixed effects.

(proxied by a vector of region-fixed effects, FE_i, with associated coefficient vector, λ_i) and a nationwide time-specific growth factor (proxied by time fixed effects, FE_t, with associated coefficient vector, λ_t). Thus for region i, we postulate:

$$\ln(PH^e_{it+1}) = \ln(PH_{it}) + \lambda_i FE_i + \lambda_t FE_t \qquad (2)$$

Total costs are modelled as a constant returns to scale Cobb-Douglas function of land and building costs; both land and building costs are relevant to the developer in deciding whether to purchase a plot of land plus construction materials and labour for a specific housing development. In addition, we include financial costs, reflecting the costs of borrowing to cover construction materials, labour and land from t until sale of the house in t+1:

$$TC_{it} = [e^{\alpha i} PL_{it}^{\beta i} PB_{it}^{1-\beta i}](1+r_t) \qquad (3)$$

Combining equations (1) to (3) we obtain the q-theory equation, (4), for housing supply:

$$HC_{it}/H_{it-1} = \lambda'_0 + \gamma_{i1}\ln\{PH_{it}/PB_{it}\} + \mu_i \ln\{PB_{it}/PL_{it}\} + \lambda'_i FE_i + \lambda'_t FE_t + \varepsilon_{it} \qquad (4)$$

where $\mu_i = \gamma_{i1}\beta_i$, the λ'_t incorporate the impact of r_t plus any time-specific risk premia plus other time fixed effects, the λ'_i incorporate all region-specific fixed effects from equations (1) to (3), and λ'_0 is the overall constant term excluding time and area fixed effects.

DiPasquale (1999) notes that in comparison with the large literature on housing demand, housing supply has been studied far less, and often with inconclusive results. Poterba (1984) and Topel and Rosen (1988) both take q-theory-related approaches to modelling housing supply. Poterba models net investment in housing structures as a function of real house prices, the price of non-residential construction, construction costs (real wages) and a measure for credit availability. The price of housing is the main determinant of new construction with estimated elasticities of new construction with respect to real house prices ranging from 0.5 to 2.3. Topel and Rosen estimate a supply function where housing starts are a function of real house prices and vector of cost shifters. They estimate a long-run supply elasticity of 3.0 and a short-run elasticity of about 1.0. Similar to Poterba, none of their cost measures is significant. Neither Topel and Rosen nor Poterba explicitly address the role of land, which is one of the most unique aspects of housing as an investment.[6]

More recent studies have sought to incorporate land as an input. DiPasquale and Wheaton (1994) estimate a model of housing construction that combines a stock adjustment process with a long run spatially-based definition of the equilibrium housing stock. Housing starts are a function of current housing prices, real interest rates, land costs (price of surrounding farmland), construction costs (material and labour) and the stock of housing in the previous period. Both their cost index and land price are statistically insignificant. We note here that the DiPasquale and Wheaton measure of land prices (farmland prices) differs from ours (residential values). In the presence of zoning restrictions that prevent residential subdivision of farmland,[7] the latter is likely to be the more relevant land measure; this may explain DiPasquale and Wheaton's finding that land prices have no significant effect on housing starts.

Recent work by Mayer and Somerville (1996, 2000a, 2000b) uses models of residential construction based on the theory of urban land development presented in Capozza and Helsey (1989). Mayer and Somerville construct a model where housing starts are a function of price and cost changes rather than levels, arguing that because housing starts are a flow variable, starts should be a function of other flow variables (i.e. price changes rather than levels). They contend that a model where starts are a function of the price level would predict a permanent increase in the number of housing starts resulting from a one-time increase in population or house prices. We consider, however, that this argument is incorrect.

[6] Poterba acknowledges the importance of land but omits it in his empirical work due to a lack of data. For related approaches, refer to Follain (1979) and Blackley (1999); see also Meen (2000) and Tsoukis and Westaway (1994).

[7] The importance of regulation in affecting prices and supply is discussed further in section 3.

The q-theory of investment, as derived here, links new investment to the levels of output and input prices. Both prices and quantities adjust to restore equilibrium following a shock, without the explosive increase in new investment posited by Mayer and Somerville. We therefore retain our explicitly derived housing supply equation, (4), as our modeling basis.

We estimate (4) using four separate approaches. Initially, we estimate the equation by pooled (ordinary) least squares (PLS), firstly with coefficients restricted to be identical across regions (other than the regional fixed effects) and secondly with unrestricted coefficients. We test the validity of the restrictions across regions. Subsequently, we estimate the equation using instrumental variables (IV). We do so because of the potential simultaneity between building consents in period t and the price and cost terms in period t. We estimate the IV equations in both restricted and unrestricted forms. The instruments comprise the set of variables appearing in the house price equation reported in section 3.[8] In all cases, we report standard errors using White period standard errors that are robust to arbitrary within cross-section residual autocorrelation. Results are reported in Table 1; all equations include both region and time fixed effects.

The key parameter determining responsiveness of new housing supply to demand shocks (which are reflected in house prices) is γ_{i1}. In the restricted PLS equation, γ_{i1} is significant at the 1% level [p=0.0000]. The size of γ_{i1} indicates that building consents rise by approximately 0.5% in response to a 1% increase in house prices (relative to building costs). The restricted IV equation indicates higher responsiveness; with building consents rising by approximately 1.1% in response to a 1% increase in house prices (relative to building costs). The difference between the two estimates suggests that prices and/or costs are themselves influenced by the supply response, consistent with our theoretical priors. We therefore treat the IV results as our preferred estimates.

[8] Specifically, after substituting the variables from the long run house price equation, (7), into the price adjustment equation, (9), the set of instruments is: $\ln XPROD_{it-1}$, $\ln XEMP_{it-1}$, $\ln DD_{it-1}$, UC_{it-1}, S_{it-2}. We include the full set of instruments to ensure theoretical consistency across the supply and price equations. We have also estimated the supply equation using four lags of each of the instruments to test robustness in case of serial correlation in the instruments. The IV estimates are virtually unchanged, as is the explanatory power of the equation. We therefore report only the results with the instrument set derived explicitly from (7) and (9).

The two restricted estimates find that land prices comprise less than half, but still a material portion, of total development costs; the implied β_i ($=\mu_i/\gamma_{i1}$) is estimated at 35% in the case of the IV estimate. In each case, the estimate of μ_i is significant at the 1% level; thus omission of land costs (as in a number of prior studies) will lead to omitted variables bias.

Columns 2 and 4 of Table 1 present the unrestricted PLS and IV estimates respectively. We do not present all 146 γ_{i1} and μ_i coefficients, but instead report their means (together with the mean of the implied β_i). An F-test for joint significance of each of γ_{i1} and μ_i is significant at the 1% level in each case. Thus when the specification is estimated at a disaggregated regional level, we still find that the q-theory specification holds and that land remains an important element of development costs.

More importantly for the purposes of this paper, a Wald test for the null hypothesis of equality of γ_{i1} $\forall i$ is decisively rejected [p = 0.0000]. This finding means that supply responsiveness differs across TLA regions. Without further information, we cannot determine whether the supply elasticities differ because of regulatory factors or because of geographical factors (e.g. mountainous land) or because of yet other factors.

We gain some insights into this issue by comparing the supply responsiveness of TLAs within New Zealand's major city, Auckland. The city comprises five 'core' TLAs[9] plus a further two TLAs that are semi-rural. We concentrate on the five core TLAs so as to compare 'like with like' as much as possible. McShane (1996) prepared an assessment of the impact of regulation on the 'housing and construction' components of the Consumers Price Index, using Auckland case studies. His interpretation of the local authorities' district plans rated Papakura local authority as the most development-friendly authority within the city.[10] He also found that Papakura had experienced lower increases in regulatory costs than experienced by the three other local authorities included in his study (North Shore, Waitakere, Auckland City). These insights into relative

[9] North Shore, Waitakere, Auckland City, Manukau City, Papakura.

[10] "The Papakura District Plan stands out on account of its relatively friendly attitude towards its people and the market-place" (McShane, 1996, p.5).

local authority housing supply responsiveness can be compared with our econometrically estimated findings regarding supply responsiveness.

We test the null hypothesis that γ_{i1} is equal for the five core TLAs. A Wald test for the null hypothesis of equality is rejected [p=0.0000]. This rejection is consistent with McShane's documentation of differences across local authorities. Further, these results are obtained after controlling for land prices, reflecting geographical constraints and potentially also other regulatory factors (e.g. zoning restrictions). To the extent that regulatory factors are responsible for the estimated supply responses, these factors must therefore be influencing the rate of construction of new dwellings rather than acting through land constraints.

We present the γ_{i1} (instrumental variables) estimates in Table 2 for the five core TLAs. Consistent with McShane's observations, Papakura has the highest supply responsiveness, albeit closely followed by Manukau (the latter was not covered in McShane's comparison). The supply responsiveness in these two TLAs is 50% higher than in Auckland City and is more than twice the responsiveness estimated in North Shore and Waitakere. The second column of Table 2 reports the estimated region fixed effect for each TLA, which may also reflect underlying regulatory differences. Consistent with the supply response parameters (and with McShane's documentation), Papakura has the highest regional fixed effect of the five core TLAs. Our estimates are therefore consistent with prior case study findings.

3 Dynamic House Price Responses

House prices are a 'jump' variable, equating short run housing demand with (fixed) short run supply. Quantities and prices of houses adjust over time to establish spatial equilibrium whereby the benefits and costs of living in an area are equated (Roback, 1982). Accordingly, long run trends in house prices can be explained as a function of economic and demographic factors. Pain and Westaway (1996) demonstrate that a standard consumer optimisation problem over current and future housing and non-housing consumption goods yields the inverted demand curve in period t:

$$\ln(p^h_t/p^c_t) = (1-\delta)\ln(\theta) - \delta\ln(h_t/pop_t) + \delta\ln(cx_t) - \ln(uc_t) \qquad (5)$$

where: p^h_t is the price of housing; p^c_t is the price of non-housing consumption goods; δ is the coefficient of relative risk aversion (within a representative constant relative risk aversion utility function); θ is the (constant) ratio of housing services to the housing stock, h_t; pop_t is population; cx_t is per capita non-housing consumption; and uc_t is the real user cost of capital.

Empirical application of this approach takes the stock of houses, h, as contemporaneously fixed (so can be treated as exogenous in the price equation). Prices adjust after a demand shock towards the equilibrium value determined by the observed stock of houses; however, the nature of price adjustment may differ depending on the nature of the short term supply response.[11]

The importance of housing supply responsiveness for price dynamics following a demand shock can be seen from the simple demand and supply graph in Figure 2. House prices and quantities (P and Q respectively) are depicted on the axes; the line marked D is the demand curve for houses; the line marked S^L is the

[11] See Grimes and Aitken (2004) for application of this approach to New Zealand; and Capozza et al (2002) for a related approach in the United States.

long run supply curve for houses. Equilibrium is initially at prices and quantities (P_0, Q_0). Housing demand then shifts permanently upwards (to the line marked D'), possibly due to a permanent positive employment increase for a given population within the region. In (5), this will be reflected through an increase in cx. In the long run, the price and quantity of housing settles at (P_L, Q_L).

In the short run, house supply does not respond fully. Consider the perfectly inelastic short run supply curve, S^S. Prices, in the short term, will jump to the short run equilibrium at P_1. Price P_1 can be derived directly from (5) using the new value for cx, holding all other variables constant, including h (since supply is assumed perfectly inelastic).

Now consider a case with more responsive short run housing supply, given by short run supply curve, S'^S. The new short run equilibrium price will be at P_2. Denote P_1-P_0 as ΔP_1 and P_2-P_0 as ΔP_2 . From Figure 2, we see that $\Delta P_2 = \phi\Delta P_1$, where $0<\phi<1$, and where $\partial\phi/\partial\gamma_{i1}<0$ (recalling that γ_{i1} is the short run elasticity of housing supply). If we were to calculate the new equilibrium price, P*, as P_1 (i.e. the short run equilibrium price derived on the assumption of a zero short run supply elasticity) and estimate the short run adjustment equation, (6),[12] we would expect the adjustment coefficient, ϕ, will be close to 1 (respectively 0) where short run supply is inelastic (elastic):

$$\Delta P = \phi(P^*_{t-1} - P_{t-1}) \qquad (6)$$

Glaeser and Gyourko (2005) emphasise the importance of house supply in determining house prices and mediating urban dynamics.[13] When supply expands quickly in response to demand pressures the housing stock and population can grow quickly with little pressure on house prices. Recent evidence

[12] "i" subscripts are added when the equation is estimated at the regional level. This specification assumes that short run adjustment responds to demand shocks in the previous quarter; other adjustment dynamics can be catered for through slightly different dynamic specifications.
[13] They note also that the durable nature of housing creates an asymmetry in growth; in declining areas, house prices are likely to fall before houses are demolished. We do not explore this aspect explicitly in the current paper, in part, because the 1991-2004 period has generally been one of expansion across much of New Zealand.

indicates that regulation plays an important role in affecting the elasticity of new housing supply; see, for example: Mayer and Somerville (2000a), Glaeser and Gyourko (2002, 2003), Glaeser et al (2005a, 2005b, 2005c), Green et al (2005) and Quigley and Raphael (2005). Glaeser et al (2005b) point out that new construction has fallen and housing prices have risen dramatically in a small, but increasing number of places. They argue that this is primarily due to increasing regulatory barriers to large-scale residential development. Green el al (2005) estimate supply elasticities for 44 U.S. metropolitan areas following a model based on Capozza and Helsey (1989). Using survey data on land regulation they estimate supply elasticities and find that areas that are heavily regulated exhibit lower elasticities.

Building on these insights, we seek to determine whether local authorities that have relatively high short run supply responsiveness have less volatile price dynamics following demand shocks. In contrast to most of the cited studies, we have data that enable us both to estimate supply elasticities explicitly, and to estimate the dynamics of price adjustment. We have already estimated supply responsiveness (γ_{i1}) in section 2. In order to estimate price dynamics, we estimate a standard cointegrating regression for house prices (to determine P_i*), and then estimate a log change version of (6) to estimate ϕ_i. Our hypothesis is that areas with high values of γ_{i1} will have low values of ϕ_i; thus the relationship between γ_{i1} and ϕ_i will be significantly negative.

We base our specification of P* on the prior work of Grimes and Aitken (2004). That work established that the variables of interest are non-stationary and a cointegration approach to modelling prices is appropriate. The log of real house prices is regressed against log of dwelling density (lnDD; i.e. $\ln(h_t/pop_t)$ in (5)), the real user cost of capital (UC; i.e. uc_t in (5)),[14] plus two variables proxying determinants of per capita non-housing consumption, cx_t in (5), being: lnXPROD (the log of per capita regional production) and lnXEMP (the log of employment as a ratio of population of working age). We restrict the long run coefficients on

[14] We enter uc, multiplied by a freely estimated coefficient, rather than log(uc) since uc is negative in some quarters for some areas. As in Grimes and Aitken (2004) we proxy real capital gains expectations within the uc term as the last three years' annual rate of capital gain within the area.

these variables to be identical across areas reflecting shared underlying preferences. Area fixed effects are included; time fixed effects are replaced by inclusion of an area-specific time trend for each TLA to account for long run (deterministic) trends in tastes towards different regions (e.g. towards climate or coastal proximity). The resulting equation, estimated using pooled least squares,[15] is presented as (7):

$$\ln\{PH_{it}/PC_t\} = \underset{[0.0000]}{\underset{(0.0710)}{0.4514}}\ln XPROD_{it} + \underset{[0.0000]}{\underset{(0.1780)}{0.9484}}\ln XEMP_{it} - \underset{[0.0000]}{\underset{(0.2616)}{2.9811}}\ln DD_{it} - \underset{[0.0000]}{\underset{(0.0007)}{0.0138}}UC_{it} \quad (7)$$

Adj.R^2 = 0.972; s.e. = 0.0738; n=3,942 (1991q1-2004q2)
Constant, area FEs and area-specific time trends included but not reported; White period standard errors in round brackets; p-values in square brackets; mean of dependent variable = 4.6355 (std dev = 0.4416).

For the purposes of our study, it is the dynamic response of house prices to a demand shock - i.e. to the explanatory variables in (7) - that is of major interest. We estimate an adjustment equation, standard in the cointegration approach[16] and consistent with (6), as in (8):

$$\Delta\ln\{PH_{it}/PC_t\} = \eta_0 + \eta_{i1}[\ln\{PH_{it-1}/PC_{t-1}\}^* - \ln\{PH_{it-1}/PC_{t-1}\}] + \eta_2 S_{it-2} + \zeta_{it} \quad (8)$$

where ζ_{it} is an iid error term, and S_{it-2} is the ratio of house sales to housing stock in region i in period t-2. We include this variable since earlier research (Grimes and Aitken, 2004) indicates that prior sales strongly influence

[15] PLS is used since all variables are non-stationary (Grimes and Aitken, 2004), making PLS estimates super-consistent. The residual from (7) is stationary: the null of a unit root is rejected at p=0.0000 for each of the Levin, Lin and Chu t-test, the Breitung t-test, the Im, Pesaran and Shin W-statistic, the ADF-Fisher Chi-square statistic, and the PP-Fisher Chi-square statistic; the null of no unit root is not rejected by the Hadri Z-statistic (p=0.9385). Equation (7) therefore represents a cointegrating vector.

[16] Engle and Granger (1987). We have reversed the order of the variables within the adjustment term compared with the more usual cointegration adjustment specification to aid interpretation without any change to methodology.

price dynamics. The coefficient, η_{i1}, is the responsiveness of house prices in period t to a demand shock in period t-1 in region i. Initially we estimate (8) with η_{i1} restricted to be identical across regions and then estimate the panel with η_{i1} unrestricted.[17] Subsequently, we test whether the unrestricted estimates of η_{i1} are related systematically to estimated supply responsiveness in each area. Results of estimating (8), with η_{i1} restricted, using PLS[18] are shown as (9):

$$\Delta\ln\{PH_{it}/PC_t\} = \underset{\substack{(0.0034)\\ [0.0000]}}{-0.0170} + \underset{\substack{(0.0309)\\ [0.0000]}}{0.5816}[\ln\{PH_{it-1}/PC_{t-1}\}^* - \ln\{PH_{it-1}/PC_{t-1}\}] + \underset{\substack{(0.2202)\\ [0.0000]}}{1.6807}S_{it-2} \quad (9)$$

Adj.R^2 = 0.287; s.e. = 0.0671; n=3,869 (1991q2-2004q2)
White period standard errors in round brackets; p-values in square brackets.

When re-estimated with η_{i1} unrestricted, the estimates of η_{i1} vary across regions with a mean of 0.50 and standard deviation of 0.22. The standard deviation indicates that substantial variation in adjustment dynamics across regions is apparent. Figure 3 presents a scatter plot of the estimated price adjustment parameters (η_{i1}) against the estimated (IV) supply adjustment parameters (γ_{i1}). While the match is far from perfect, a negative relationship between the two, as predicted by theory, is apparent.

The significance of this relationship is explored further in Table 3. We hypothesise that η_{i1} will be smaller the larger is the estimate of γ_{i1}. Columns 1 and 2 report a cross-section regression of η_{i1} on a constant plus the PLS and IV estimates of γ_{i1} respectively. In each case, the coefficient on γ_{i1} is negative and

[17] The equation does not include area fixed effects; when added, they are jointly insignificant and other coefficients remain virtually unchanged.
[18] PLS is appropriate since all explanatory variables are lagged. There is little evidence of autocorrelation (DW=2.08) but since the Durbin-Watson statistic is not appropriate in the presence of a lagged dependent variable, we report (robust) White period standard errors. Equation estimates are unweighted; estimates are virtually unchanged using GLS with period weights; when cross-section weights are used the estimate of η_1 and η_2 fall a little to 0.4710 and 1.5027 respectively. Population-weighted estimates are similar, as are estimates weighted by housing stock and also by house sales.

significant at the 1% level.[19] This result is consistent with local authorities having low supply elasticities tending to have high price responses to demand shocks. Since land prices (which, *inter alia*, reflect geographical conditions) are already included in the supply equation, the difference in supply elasticity is most likely related to regulatory differences across local authorities.

The supply equation includes area fixed effects. It is possible that areas which have strong supply growth, irrespective of price elasticity of supply, may have different price dynamics compared with regions with slower supply growth. Our supply equations also include the effects of land prices. We hypothesise that areas with high land price growth may be subject to planning regulations that inhibit new residential land development.[20] If there is such an effect, we expect areas with high residential land price growth (possibly reflecting greater regulatory constraints) to exhibit stronger price dynamics and hence a larger η_{i1}.

We extend the equations in Table 2 to take account of these additional factors. Column 3 adds the area fixed effects from the (PLS) supply equation and also adds the rate of increase of residential land prices for each area over the sample period. Column 4 presents the corresponding results based on the IV supply equation. The coefficient on γ_{i1} stays negative and significant at the 5% level in each equation. The coefficient on residential land price increases is positive (as expected if regulatory constraints that affect land are present); they are significant at the 10% level using the PLS-based estimates and at 12% using the IV-based estimates. The area FEs are not significantly different from zero in either equation (p= 0.89 and 0.77 respectively) indicating that price dynamics are not related to the underlying housing supply growth within an area.

Columns 5 and 6 drop the area fixed effects from the equation. The resulting estimates of the effects of γ_{i1} and of land price increases are little

[19] We note that the regressor (γ_{i1}) is stochastic and so its p-value in this and subsequent regressions may be under-stated. The η_{i1} and γ_{i1} estimates are, however, obtained within a systems context since the (lagged) determinants of house prices are included as the instruments in the supply equation, so reducing the potential for the error term within γ_{i1} to be correlated with that for η_{i1}. Accordingly, the strong significance of γ_{i1} shown across all the equations in Table 3 indicates that the estimated negative relationship between the two parameters is likely to be robust.

[20] Such regulations may well be positively correlated with regulations inhibiting new construction (given the land supply).

changed, although the former are now significant at 1% in each case. The results indicate that local authorities with low supply elasticities and (more tentatively) high land price increases face more volatile price adjustment in response to demand shocks, possibly due to regulatory constraints.

The analysis above is primarily concerned with determining the process of dynamic supply and price adjustments to shocks. We note here that the data also indicate a (tentative) relationship between long run supply responsiveness and long run house price increases. Our previous estimates indicate that in the short run, higher house prices (relative to costs) induce higher new house supply; thus there is a positive short run relationship between house prices and new supply. Over the full sample period, however, there is a negative relationship between new supply (summed over the entire period and expressed relative to the initial housing stock in each region) and house price increases (relative to local land price increases). Despite the short run positive relationship, the cross-sectional (long run) correlation coefficient between the two is -0.17 (significant at the 15% level).

This long run result is consistent with the nature of long run price shifts indicated by Figure 2, in which a flatter (more responsive) supply schedule reduces the long run price increase consequent on an increase in demand (i.e. reduces the gap between P_L and P_0). While not the main focus of the paper, this consistency of long run outcomes with short run dynamics, in line with theoretical priors, is a useful robustness check on the dynamic findings which are the major focus of the paper.

4 Conclusions

Housing supply and house price dynamics are inextricably inter-related. We find that an increase in house prices (relative to total development costs) raises new house supply with an elasticity of between 0.5% and 1.1%, the upper estimate corresponding to our (preferred) instrumental variables estimate. Unlike most previous studies, we incorporate land prices as well as construction costs in our measure of total development costs; and unlike the few studies that have incorporated land prices, we find both land and construction costs to have a statistically significant effect on new house construction. One reason that may explain why we find land prices to be significant while others studies do not (apart from the different country setting) is that we use residential land values rather than values of surrounding agricultural land. The latter may not be the appropriate land price measure in circumstances where zoning restrictions provide the binding constraint as to whether land is used for residential or agricultural purposes.

We find that the estimated supply responsiveness coefficient is inversely related to the estimated coefficient determining price dynamics. Prices react more strongly to a demand change in local authorities in which housing supply responsiveness is low compared with those in which supply responsiveness is high. A 1% increase in equilibrium prices (calculated on the basis of the existing housing stock) leads to an immediate 0.56% jump in house prices in an authority with lower quartile supply responsiveness, compared with a 0.45% jump in an authority with upper quartile supply responsiveness.

Since responsiveness is faster in the latter case, the length of time that prices are raised above their long run equilibrium is also shorter in the more responsive than the less responsive authority. Thus the impact on resource (mis)allocation is more pronounced in the less responsive authority and lasts for longer than if it were more responsive.

Without specific regulatory data, we cannot attribute the differences in supply responsiveness to regulatory versus other factors. However our results for the local authorities within Auckland are consistent with the case study findings of McShane (1996) for the same city. In addition, the fact that we have controlled for land prices in calculating these estimates implies that geographical factors are most likely not the 'culprit' in determining supply responsiveness.

Land prices themselves have a strong impact on new house construction. A 1% increase in land prices is estimated to lift total development costs by 0.33% (using the restricted IV estimate), in turn reducing house supply by an estimated 0.37%. Thus regulations, such as zoning restrictions, that impact on the availability of residential land (forcing up residential land prices) induce lower house supply and raise house prices in affected authorities.

These results are consistent with, and extend, the findings of recent studies in the United States on links between housing supply and house price dynamics. The key contribution of this paper - apart from extending the evidence base to another country - is to highlight explicitly the role of residential land prices in determining supply responses and thence price dynamics. Similar use of residential, rather than agricultural, land prices may warrant investigation in countries, such as the United States, in which land prices have hitherto not featured prominently in estimated house supply relationships.

Appendix A: Data

We use a quarterly dataset of median house prices for New Zealand for the period 1991q1-2004q2 covering 73 Territorial Local Authorities (TLA). House price data are sourced from Quotable Value New Zealand (QVNZ), a state-owned organisation. The data include median sales prices, median capital values (i.e. official valuations used for property tax purposes) and the number of residential property sales at TLA level. QVNZ provides data for residential dwellings covering several categories. In this analysis we use residential dwellings defined as those dwellings of a fully detached or semi-detached style on their own clearly defined piece of land. We use median rather than mean data as this is less susceptible to being distorted by extremely low or high observations.

We mix-adjust the median data for each TLA, recognising that the types of property sold vary from year to year within a TLA. Our mix-adjustment procedure builds on the valuation-based approach of Bourassa et al (2004). We hypothesise that the observed house sales price, SP_{zt}, comprises three components: a (flexible) trend component (SPF_{zt}), a component due to the mix of houses sold in each year MIX_{zt}, and a random element, χ_{zt}, that is orthogonal to MIX_{zt}. We assume that SP_{zt} is proportional to SPF_{zt}; thus we maintain that the following relationship exists explaining the observed SP_{zt} data:

$$\ln(SP_{zt}) = \ln(SPF_{zt}) + c_{0z} + c_{1z}\ln(MIX_{zt}) + \chi_{zt} \qquad (10)$$

Our MIX_{zt} variable is obtained by taking the ratio of the median capital value for houses sold each year to the trend (HP filtered) median capital value of houses sold, where the latter is a proxy for the capital value of the "typical" house within a TLA. Our SPF_{zt} variable is formed as the HP filtered median sales price. We estimate (14) for each TLA and then derive the mix adjusted price (P_{zt}):[21]

[21] The standard deviation of the mix-adjusted sales price is lower than for the raw median sales price for every TLA. The smoothing is greater for TLAs with more volatile data; volatility, in turn, is related inversely to the number of TLA sales. Together, these results indicate that the adjustment is indeed compensating for mix differences within TLAs across years.

$$P_{zt} = \exp[\ln(SP_{zt}) - c_{1z}\ln(MIX_{zt})] \equiv \exp[\ln(SPF_{zt}) + c_{0z} + \chi_{zt}] \tag{11}$$

We obtain land prices relating to residential properties from QVNZ. QVNZ valuations (which are generally conducted on a three yearly cycle) split residential property values into structures and land components. We use these data to construct a flexible trend representing TLA land values over the full period.

The number of TLA residential sales per quarter is also sourced from QVNZ. These data are used to form S, the ratio of house sales to the housing stock in each region. Data on the housing stock in each TLA are available from the 1991, 1996 and 2001 censuses. They are interpolated to form quarterly observations.

Housing consents are available on a quarterly basis from Statistics New Zealand. These data include residential houses with a value greater than $4,999.[22] The data are seasonally adjusted using X-12 Arima.

Construction cost data are sourced from the trade publication, *New Zealand Building Economist*, and are available on a quarterly basis from 1992 to 2004 for six regions covering the entire country. We use the cost for standard dwellings[23] which represents average installed prices. The cost includes trade materials prices, labour rates, plus allowance (according to local conditions) for overheads, subcontractors, and subcontractors' profit where applicable.

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[22] Includes houses not attached to others, unit/flat/townhouse/studio attached and unattached horizontally, apartment blocks attached vertically, granny flat unattached, dwellings added to other buildings, communal accommodation and other residential dwellings not elsewhere included.

[23] Standard house specification: 2001 onwards: 94m2; 3 bedroom; level site; timber pile base; fibre cement base lining with plastic vents; timber steps; fibre cement weatherboards; R 2.2 batts to walls, R 2.4 batts to ceilings; truss gable roof with ceiling battens; Zincalume roofing and accessories; aluminium joinery; particle board floor; Gib board to walls and ceilings; shower over bath; separate wc; separate laundry with ss tub and cupboard under; 12 lights; 16 power outlets; average quality wallpaper; conventional four element stove. 1992 - 2000: 94m2; 3 bedroom; level site; concrete pile basement/fibre cement lined; concrete steps; weatherboards; all exterior walls and ceilings lined with 75mm batts; corrugated iron gable roof; timber joinery; particle board floor; gibraltar board walls; sloping ceiling with exposed rafters to dining room/lounge; flat ceiling to other areas; separate shower/bath/laundry; separate WC; 12 lights; 16 power points; average quality wallpapers; conventional four ring stove.

Per capita TLA production, XPROD, is formed by weighting quarterly GDP by industry[24] (one-digit ANZSIC[25]) by industry employment data from the 1991, 1996 and 2001 censuses.[26] This variable is therefore akin to a Bartik index (Bartik, 1991). While data unavailability necessitates the use of national industry production data (albeit weighted by TLA-specific weights), an advantage is that we thereby mitigate any endogeneity issues which could arise from use of TLA-specific production data (if such data were available).

Data for labour force participation (XEMP: ratio employed over total usually resident population aged over 15) was obtained from the 1986, 1991, 1996 and 2001 censuses and linearly interpolated to form quarterly observations.

UC, the real user cost of capital, was formulated exactly as in Grimes et al (2004). It comprises the real 90-day bank bill rate (i.e. the nominal rate less annual CPI inflation) minus the expected rate of real capital gain on housing within a TLA. The past three years' annual rate of real capital gain on houses at the TLA level is used as the measure of expected real capital gains.[27] The quarterly consumers price index is sourced from the Reserve Bank of New Zealand.

[24] The GDP data (constant prices) was seasonally adjusted using X-12 ARIMA.
[25] Australian New Zealand Standard Industrial Classification. There are 18 industry groups in the one-digit classification: Agriculture, forestry and fish (AGR), Mining (MIN), Manufacturing (MAN), Electricity, gas and water (EGW), Construction (CON), Wholesale trade (WHO), Retail trade (RET), Accommodation, cafes and restaurants (ACR), Transport and storage (TRN), Communication services (COM), Finance and insurance (FIN), Property and business services (PRP), Government administration and defence (GOV), Education (EDN), Health and community services (HEA), Cultural and recreational services (CUL), Personal and other services (PER) and Industry not specified (NSP).
[26] The Census employment data has been linearly interpolated to quarterly observations.
[27] Grimes et al (2004) tested a variety of expected real capital gains proxies, finding that extrapolative expectations based on the past three years region-specific developments performed best.

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Table 1: House Supply Responsiveness

	PLS Restricted	PLS Unrestricted	IV Restricted	IV Unrestricted
γ_{i1}	0.0048 (0.0010) [0.0000]	Mean=0.0048 [<0.01]	0.0112 (0.0023) [0.0000]	Mean=0.0070 [<0.01]
μ_i	0.0009 (0.0003) [0.0032]	Mean=0.0017 [<0.01]	0.0035 (0.0013) [0.0080]	Mean=0.0020 [<0.01]
β_i (Implied)	0.1795	0.3465	0.3094	0.2780
Adj.R^2	0.778	0.833	0.727	0.822
s.e.	0.0015	0.0013	0.0017	0.0014
n	3,869	3,869	3,869	3,869

Notes:
Estimated equation: $HC_{it}/H_{it-1} = \lambda'_0 + \gamma_{i1}\ln\{PH_{it}/PB_{it}\} + \mu_i\ln\{PB_{it}/PL_{it}\} + \lambda'_iFE_i + \lambda'_tFE_t + \varepsilon_{it}$
where implied $\beta_i=\mu_i/\gamma_{i1}$ (calculated at means of μ_i and γ_{i1} for unrestricted estimates).
White period standard errors in round brackets; p-values in square brackets.
PLS is pooled (ordinary) least squares estimation; IV is instrumental variables estimation.
All equations estimated over 1991q2 - 2004q2.
Mean (standard deviation) of dependent variable = 0.0038 (0.0032).
Constant, region fixed effects and time fixed effects included but not reported.
Adj.R^2 is the adjusted R^2; s.e. is the equation standard error.
n is number of observations (all equations cover 73 regions for 53 quarters).
Individual coefficients not reported in unrestricted equations (means of coefficients are reported);
p-values in unrestricted equations refer to F-test for significance of all 73 regional coefficients.
IV instruments: $\ln XPROD_{it-1}$, $\ln XEMP_{it-1}$, $\ln DD_{it-1}$, UC_{it-1}, S_{it-2}, $TIME_t$, constant); see section 3 and Appendix for descriptions.

Table 2: Auckland Core TLA Supply Parameter Estimates

TLA	Supply Responsiveness (γ_{i1})	Region Fixed Effect (λ'_i)
North Shore	0.63%	-0.16%
Waitakere	0.73%	0.78%
Auckland City	1.05%	0.81%
Manukau	1.56%	1.37%
Papakura	1.58%	1.91%

Notes:
Parameters are instrumental variables estimates consistent with Table 1, corresponding to equation (4); thus a 1% rise in house prices relative to costs increases quarterly housing consents relative to the housing stock by 0.63% in North Shore.

Table 3: Relationship of Price Adjustment to Supply Responsiveness

Supply Eq:	PLS	IV	PLS	IV	PLS	IV
Constant	0.5800 (0.0319) [0.0000]	0.5756 (0.0329) [0.0000]	0.5428 (0.0416) [0.0000]	0.5339 (0.0445) [0.0000]	0.5409 (0.0390) [0.0000]	0.5369 (0.0403) [0.0000]
γ_{i1}	-15.7699 (4.4138) [0.0006]	-10.1839 (3.1753) [0.0020]	-15.7958 (5.1194) [0.0029]	-9.4402 (4.0430) [0.0225]	-15.4430 (4.3599) [0.0007]	-9.8550 (3.1457) [0.0025]
AFE_i	-	-	0.3327 (2.4818) [0.8938]	-0.4006 (2.4264) [0.8693]	-	-
Land-Inc_i	-	-	0.0137 (0.0081) [0.0965]	0.0134 (0.0083) [0.1100]	0.0138 (0.0081) [0.0926]	0.0134 (0.0082) [0.1084]
Adj.R^2	0.140	0.114	0.151	0.122	0.163	0.134
s.e.	0.203	0.206	0.202	0.206	0.201	0.204
n	73	73	73	73	73	73

Notes:
Estimated equation: $\eta_{i1} = c_0 + c_1 * \gamma_{i1} + c_2 * AFE_i + c_3 * \text{Land-Inc}_i + \xi_i$
where η_{i1} is the estimated price adjustment parameter from (9); γ_{i1} is the estimated supply elasticity from the PLS and IV supply equation respectively (corresponding to the equation heading); AFE_i are the area fixed effects from the PLS and IV supply equations; Land-Inc_i is the rate of residential land price increase over the sample period; ξ_i is an iid error term;
c_0, c_1, c_2 and c_3 are estimated coefficients from a cross-section regression.
Standard errors in round brackets; p-values in square brackets.

Figure 1: New House Supply versus House Prices/Development Costs

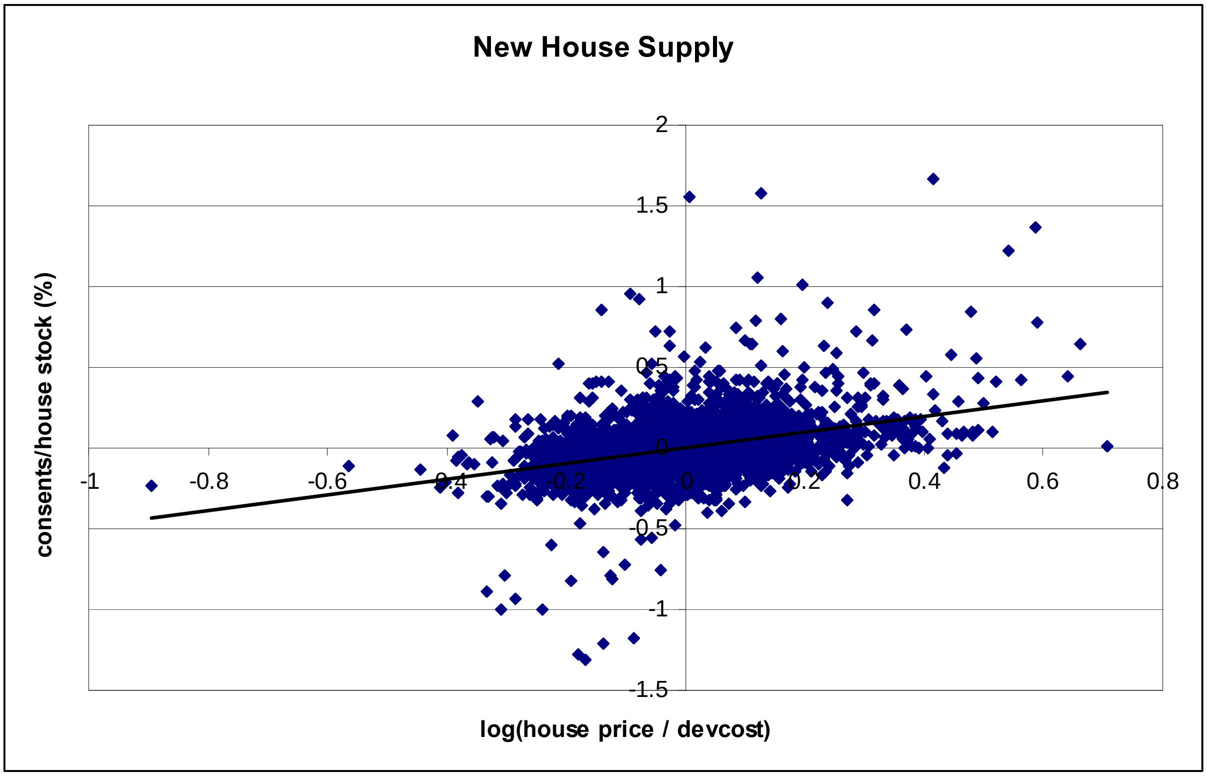


Figure 2: House Price Adjustment Dynamics

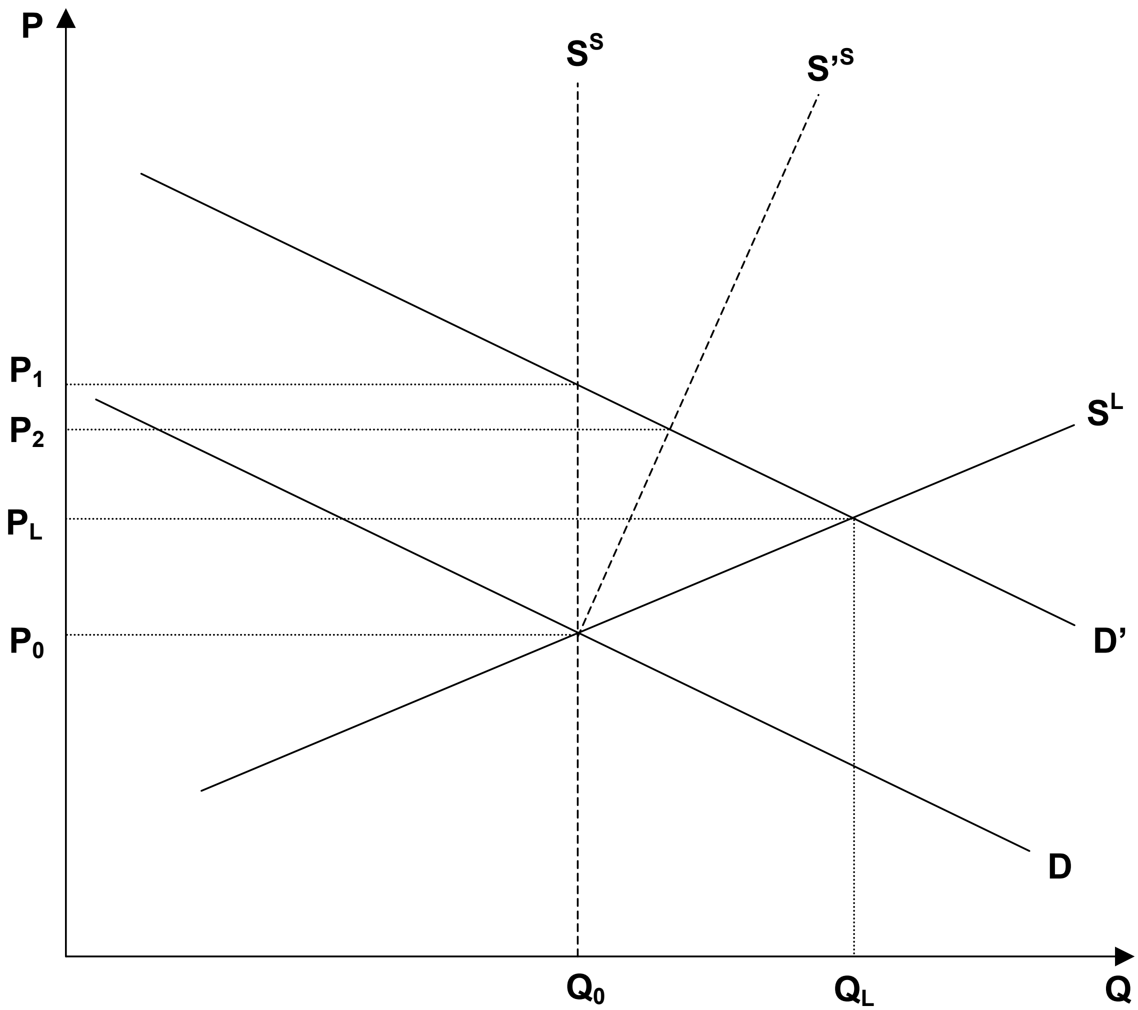
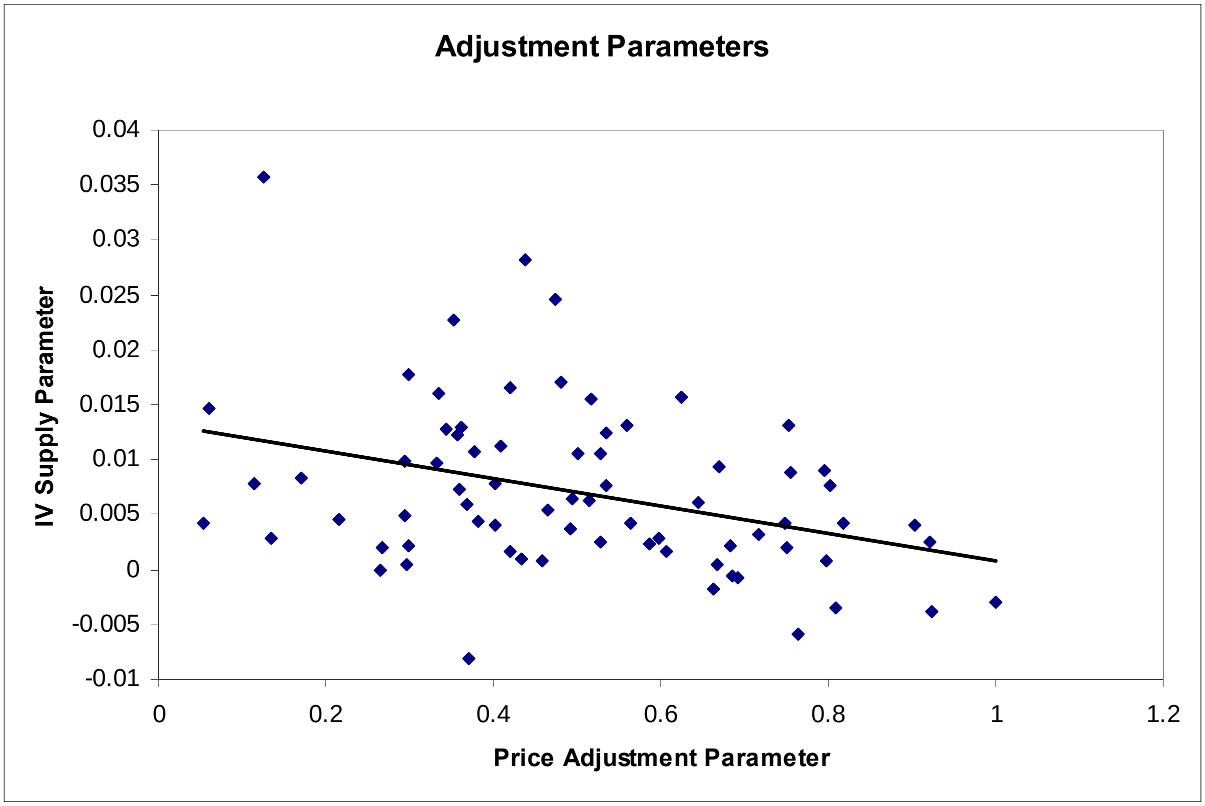


Figure 3: Estimated IV Price and Supply Adjustment Parameters (η_{i1} and γ_{i1})



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