Evaluating Housing Supply Elasticity in Inter-city Spatial Equilibrium: Evidence from Chinese Urban Growth

Yuming FU*
Institute of Real Estate Studies and Department of Real Estate
National University of Singapore
4 Architecture Drive, Singapore 117655
Email rstfuym@nus.edu.sg

Siqi ZHENG
Institute of Real Estate Studies
Tsinghua University
Beijing 100084, China
Email Zhengsiqi@tsinghua.edu.cn

Hongyu LIU
Institute of Real Estate Studies
Tsinghua University
Beijing 100084, China
Email liuhy@tsinghua.edu.cn

Last version: September 2009; this version: January 2010

Funding: National University of Singapore Academic Research Fund (R297000079112).

Acknowledgements: We thank anonymous referees and research workshop participants at Southwestern University of Finance and Economics in China for helpful comments. Ren Rongrong provided excellent research assistance.

* Corresponding author
Evaluating Housing Supply Elasticity in Inter-city Spatial Equilibrium: Evidence from Chinese Urban Growth

Abstract

We examine the influences of urban redevelopment costs and land-use allocation efficiency and equity on the elasticity of housing development density with respect to urban demand shocks in the context of China’s urban growth from 1998 to 2004. We do so in an inter-city spatial equilibrium framework, where demand shocks are capitalized in urban land rent growth, which influences urban population growth through housing supply elasticity. Importantly, this approach distinguishes the supply elasticity revealed by the population growth effect of land rent shocks from the price effect of urban population shocks, which reveals the effect of agglomeration economies, a distinction not widely recognized in the extant housing supply studies. We further control for supply shifts due to urban land-use growth and increased housing absorption by existing urban residents due to income growth, in accounting the contribution of the demand shocks and housing density elasticity to urban growth, and find a number of institutional factors, including local government efficiency, privatization of state-owned enterprises, and income equality, to elevate the density elasticity, as predicted by their impact on development cost and land-use allocation efficiency and equity. Moreover, the variation in the density elasticity appears as important in explaining cross-city population growth as that in urban land-use growth.

Keywords: housing supply elasticity, urban growth, spatial equilibrium, Chinese economy

JEL classification: R11, R31, R58
1. Introduction

Cities play a central role in promoting productivity growth by augmenting private investment in tangible and intangible capital (Lucas, Jr., 1988). Studies of urban growth therefore inform us of the factors contributing to productivity growth (Glaeser, Scheinkman, and Shleifer, 1995; da Mata et al., 2007) and to the quality of living (Glaeser and Shapiro, 2003; Shapiro, 2006). Moreover, how cities grow shapes the endogenous growth of a nation (e.g. Black and Henderson, 1999; Rossi-Hansberg and Wright, 2007). These studies largely abstract from the influence of housing supply on urban growth. Gyourko, Mayer and Sinai (2006) show a wide variation in urban growth in response to demand shocks across US metropolitan areas, suggesting an important role for housing supply.

The influence of housing supply on urban growth is now widely recognized; Mayer and Somerville (2000), Quigley and Raphael (2005), Glaeser, Gyourko and Saks (2006) and Saks (2008), for example, find that local population and employment respond less positively to demand shocks in places with more restrictive land-use regulations. But the literature appears inconclusive regarding the importance of various determinants of housing supply elasticity, in particular the importance of land scarcity versus land-use regulations. Glaeser (2004) maintains that land scarcity by itself, apart from regulatory constraints on construction, would have little direct impact on housing supply elasticity. Evenson (2003) and Green, Malpezzi and Mayo (2005) find urban population density, in addition to land-use regulatory constraints, to have a negative effect on the supply elasticity, suggesting a possible influence of land scarcity. Using detailed topographic data, Saiz (2008) provides a more reliable exogenous measure of urban land scarcity—the share of undevelopable land area within a certain radius of city center—and finds it to elevate housing price appreciation in response to metropolitan population growth shocks; the topographic constraints reduce the supply elasticity directly as well as indirectly by causing more restrictive land-use regulations. In contrast, using panel time-series analysis, Harter-Dreiman (2004) finds no evidence of either land-use restrictions or population density having any significant influence on metropolitan area housing price response to housing demand shocks, whether in the short run or long run. In the context of developing economies, where urban growth plays a central role for economic development (Spence, Annez and Buckley, 2008), little about the determinants of housing supply elasticity has been documented.
This study contributes to the growing literature on urban growth and housing supply in two respects: we seek (i) to improve the identification of housing supply elasticity by incorporating inter-city spatial equilibrium in our analysis and (ii) to evaluate how urban redevelopment costs and land-use allocation efficiency and equity affect housing supply elasticity in the context of a developing economy. In the first respect, inter-city spatial equilibrium under perfect labor mobility entails compensating land rent growth in response to demand shocks arising from urban productivity growth and changes in amenity premium (Roback, 1982; Gabriel and Rosenthal, 2004; Glaeser and Gottlieb, 2009); the land rent growth, in turn, influences urban growth through housing supply elasticity. We therefore seek to evaluate housing supply elasticity in a framework of accounting urban population growth by the sources of net new housing supply, namely, supply shifts, changes in housing absorption by existing urban households, and housing construction in response to urban demand shocks. The accounting of the first two components is important especially in a developing-economy context, where changes in land-use policy and household housing demand growth are often very significant. Importantly, our approach distinguishes the housing supply elasticity revealed by population growth effect of land rent shocks from the price effect of population shocks, which reveals the marginal effect of agglomeration economies in a spatial equilibrium (Glaeser and Gottlieb, 2009). Such distinction seems not widely recognized in the extant housing supply studies (see, for example, Glaeser and Gyourko, 2003; Glaeser, Gyourko and Saks, 2005a and 2005b; Harter-Dreiman, 2004; Saiz, 2008), hence the conflicting findings of the impact of land scarcity and land-use regulations on supply elasticity.\(^1\) High labor mobility makes the demand for individual cities in a large economy highly elastic, at least in the long run, removing the cities’ market power to influence the local land rent by restricting housing supply. Indeed, Aura and Davidoff (2008) show that even if households have heterogeneous tastes for local amenities, so that labor mobility is not unfettered even in the long run, local housing quantity would have only a small effect on local housing prices for US cities under realistic parameter assumptions. Proper identification of housing supply elasticity, therefore, should be based on population response to demand shocks, holding constant the supply shifts and new absorption by existing resident households.

\(^1\) Quigley and Rosenthal (2005) note that the literature offers mix evidence regarding the regulatory impact on housing prices—Harter-Dreiman (2004) even finds somewhat negative effects—and Ihlanfeldt (2007) notes that the regulatory impact could be in part due to the local amenity consequences of the regulations.
In the second respect, to evaluate how urban redevelopment costs and land-use allocation affect housing supply elasticity in the context of a developing economy, we focus on the recent remarkable urban growth experience in China. In particular, we examine such institutional factors as local government efficiency, urban land-use reform, and income inequality, which often have significant influences on redevelopment costs and land-use allocation efficiency and equity in the context of a developing economy, in addition to the influences of physical conditions such as building age, per-capita road space, and urban population density. Studies of housing policies in developing economies have long emphasized the need for market enabling policies to improve housing supply elasticity but few have sought to empirically assess the determinants of housing supply elasticity that can be influenced by policies (see Buckley and Kalarickal, 2005, for a review). Our evaluation of the institutional factors would inform policy makers of effective market enabling instruments. Furthermore, in contrast to the extant studies of housing supply elasticity, which focus on regulatory and physical constraints mostly affecting the extensive margin of urban land uses (e.g. urban growth controls and topography), our analysis focuses on housing supply affected by the intensive margin of land use: how land-use density responds to urban demand shocks. Improving the responsiveness of housing development density to urban demand shocks is very important for market enabling policies in densely populated developing-economy cities.

Our sample includes 85 cities across China. We examine their population growth between 1998 and 2004, a period featuring substantial liberalization on labor mobility and rapid urbanization. The urban share of population increased from 33.4 percent in 1998 to 41.8 percent by 2004, at about twice the pace of urbanization in the decade to 1998. The strong demand for urban growth resulted from several factors. By 1998 the gradual reform of the urban housing welfare system had led to the end of the work-unit based state provision of urban housing and full liberalization of the private housing market in Chinese cities (Fu, Tse and Zhou, 2000). Private housing construction took off; between 1998 and 2004, housing construction accounted for about 6 percent of the national GDP and annual housing completion in cities was about 550 million square meters. In addition to increased

---

2 Green, Malpezzi, and Mayo (2005), for example, derive their specification of housing supply elasticity from the urban growth model of Capozza and Helsley (1989), where new construction occurs only at the urban boundary and the supply of developable land is physically limited by the urban radius.
marketability of existing urban homes, the fraction of new supply in cities sold freely at market prices doubled to over 60 percent. At the same time, urban labor market was liberalized as private sector jobs grew significantly; the non-farm private sector’s share of GDP increased from 43% in 1998 to over 57% in 2003 (OECD, 2005). The liberalization both in urban housing and labor markets significantly lowered labor mobility barriers and afforded Chinese people unprecedented freedom in choosing urban locations to live and work. The elevated labor mobility in this period, together with strong income growth that raised the demand for urban amenities, contributed to widening land rent differentials across Chinese cities (Zheng, Fu and Liu, 2009).

We find the relative land rent growth in Chinese cities during our study period significantly correlated with a number of initial urban conditions affecting the quality of living and urban productivity growth. The demand shocks predicted by these urban conditions, however, explain little of the cross-city variance in population growth when the elasticity of housing supply is assumed constant. We show that the variation in housing supply elasticity via density response to the demand shocks contributes as much to cross-city variance in population growth as do the variations in urban land-use growth and new housing absorption by existing urban households. We further find that the institutional factors make more significant differences to housing density elasticity than do physical conditions.

We present our framework of estimating housing supply elasticity, derived from inter-city spatial equilibrium and a housing-supply-based urban population growth accounting, in Section 2. Section 3 describes the data and our measurements of urban population growth, demand shocks, and the various explanatory variables. Section 4 reports the estimates of the

---

3 According to National Bureau of Statistics of China (NBSC), private sector provided 43.9 million urban jobs between 2002 and 2006 (16.9 million in foreign-funded firms and 27 million in other forms of private enterprises). Employment in state-owned and collective-owned enterprises declined by 10.7 million in the same period.

4 According to the 2000 census, over 7.5% of the population moved within provinces between 1995 and 2000 and 2.7% moved across provinces. Although this rate of migration is relatively low in comparison with economies like US, where about 3 percent of population move across states in any given year (Borjas, 1999, p10), it is substantially elevated in comparison to the 1990-1995 period, during which about 1% moved across provinces. But, given the huge national population size and a relatively low urbanization level, even a modest rate of migration, when destinations are concentrated in cities, can have a great impact on population change in cities; about 14% of the residents in Beijing, Shanghai, and Guangdong province in 2000, for example, are new arrivals after 1995. World Bank (2009, Box 5.3) provides a brief account of the evolution of labor mobility regulations in China and the recent surge in cross-region labor migration flows.
urban growth accounting, with a focus on the determinants of housing density elasticity. We conclude in Section 5.

2. A model of housing-supply-based accounting of urban population growth

To evaluate the influences of housing supply on urban growth, we follow Glaeser and Tobio (2008) in assuming that housing is produced competitively in a city from structure density \( h \) and land \( L \). We ignore the spatial structure of the city, so that \( h \) represents the average housing density in built-up area. The total quantity of housing supplied in the city is \( hL \). The cost of producing \( h \) units of structure per unit of land is \( c_0 h^\delta \), where \( \delta > 1 \) makes the average cost rise with density. Let \( p \) be the housing price level. \( h \) is then competitively determined by the first-order condition:\(^5\)

\[
p = \delta c_0 h^{\delta-1}.
\] (1)

The housing structure can be redeveloped over time as the housing price rises. The growth in housing density can be derived by applying log difference with respect to time to both sides of Equation (1) and re-arranging:

\[
g_h = \frac{g_p}{\delta - 1} = \eta \cdot g_p,
\] (2)

where \( g_x = \frac{d}{dt} \log(x) \) denotes the percentage growth in \( x \) and \( \eta = \frac{1}{\delta - 1} \), the price elasticity of housing supply via density choices.

Households in the city earn income \( Y \) and derive CES utility from tradable goods \( C \) (sold at a fixed price of unity) and non-tradable housing \( Q \): \( A \cdot (Q^\sigma + C^{\frac{\sigma-1}{\sigma}})^{\frac{1}{\sigma-1}} \), where \( A \) is city-specific quality of living and \( 0<\sigma<1 \) is the elasticity of substitution. The household consumption of housing \( Q \), at rent \( r \), is

\[
Q = \frac{1}{1 + r^{\sigma-1}} \frac{Y}{r}.
\] (3)

\(^5\) The perfect competition assumption is not unreasonable for Chinese urban residential market. There were over 60 thousand residential developers in 2008, the majority of which are privately owned. The largest 10 developers are public listed or all privately owned and their market share was about 4.2% in 2004 and 7.8% in 2008 (China Real Estate Top 10 Research Group, March 2009).
Applying log difference with respect to time to both sides of Equation (3) together with a first-order Taylor expansion with respect to time-\(t\) log rent \(\ln r_t\) around the base-period \(\ln r_0\), we have\(^6\)

\[
g_Q = g_Y - \lambda_0 g_r,
\]

where \(\lambda_0 = \frac{1 + \sigma r_0^{-1}}{1 + r_0^{-1}}\) represents the magnitude of the price elasticity of demand. In the case where housing expenditure \((rQ)\) is a quarter of household income \((Y)\) and \(\sigma=1/3\), we have \(\lambda_0=1/2\). We will assume a constant \(\lambda_0\) across cities and identical growth in housing price and rent so that \(g_r = g_p\). Following the inter-city spatial equilibrium literature, we define \(S = \lambda_0 g_r = \lambda_0 g_p\) as our measure of urban demand shocks; as noted in footnote 4, the large inflow of population into Chinese cities during our study period would support our application of the spatial equilibrium assumption to this study.

Let \(N\) denote the number of households in the city, where the housing market clears when

\[
N \cdot Q = h \cdot L.
\]

We assume a fixed household size so that the growth in population equals that in household number. Applying log difference with respect to time to both sides of Equation (5) and using Equation (2) and (4), we obtain the housing-supply-based accounting of population growth:

\[
g_N = g_L + g_h - g_Q = g_L - g_Y + \left(1 + \frac{\eta}{\lambda_0}\right) S,
\]

Equation (6) shows that population growth in a city depends on three components: 1) growth in urban land use; 2) growth in housing absorption by the existing population, driven by household income growth; and 3) demand shocks to the city, summarized by the housing price growth (multiplied by the price elasticity of demand) \(S\), which influences new housing supply via price elasticity \(\lambda_0\) and the elasticity of construction density \(\eta\).

Equation (6) is silent about how land-use growth responds to urban demand shocks. Supply of new land for urban uses often depends on public policies and regulations rather than being competitively determined. Nevertheless it is important to account for shifts in land supply for residential uses in order to correctly assess the response of housing construction to urban

\(^6\) For more details, see Zheng, Fu and Liu (2009).
Population Growth and Housing Supply across Chinese Cities

demand shocks. Glaeser and Tobio (2008), for example, find the population growth in the US Sunbelt since 1980s to have resulted to a large extent from the increase in land supply for housing construction independent of demand shocks; such supply shifts are often overlooked in extant cross-city studies of housing supply elasticity. Our analysis will control for the growth in urban land use (built-up area) $g_L$ in order to focus on the response of housing development density to urban demand shocks. The behavior of $g_L$ is in itself an interesting topic but is beyond the scope of the current study. In general, the new supply of land for urban uses is policy driven, as all urban land in China is owned by the state and the urban governments monopolize the conversion of rural land to urban uses.\footnote{Incentivized by the fiscal decentralization reform in the 1990s (Jin and Zou, 2005) and the political mandate to promote local economic growth, Chinese urban governments have used land-use conversion both to attract foreign direct investment and to raise local revenue (Lichtenberg and Ding, 2009), often prompting tighter oversight by the central government to curb excessive urban land-use expansion.}

The density of housing construction, both in the redevelopment of old built-up areas with predominantly low floor-to-area ratio (FAR) and in new development areas, would have a major influence on new housing supply in Chinese cities. Fu and Somerville (2001), for example, show that the redevelopment FAR in Shanghai in the early 1990s (when it was the largest city in China) was influenced both by the site-specific land-use demand and by the road infrastructure availability in the vicinity. In addition, the high cost of re-housing displaced households often delay and deter urban redevelopment projects. The road infrastructure condition and re-housing costs are among the determinants of the housing density elasticity $\eta$ that we will examine in this study, which also include institutional factors affecting urban land-use efficiency and equity as well as housing development costs.

3. **Data and measurements**

Our sample of 85 cities across Chinese provinces is documented in Zheng, Fu and Liu (2009). These cities are covered by the national Urban Household Survey (UHS) conducted by the National Bureau of Statistic of China (NBSC) in both year 1998 and 2004, from which our demand shock estimates $S_k$, city land rent premium $R_k$, as well as several income-related city characteristics are derived. UHS is conducted annually and consistently at the national level. Additional city attributes are collected from Urban Statistics Yearbook to instrument $S_k$ and

\footnote{Incentivized by the fiscal decentralization reform in the 1990s (Jin and Zou, 2005) and the political mandate to promote local economic growth, Chinese urban governments have used land-use conversion both to attract foreign direct investment and to raise local revenue (Lichtenberg and Ding, 2009), often prompting tighter oversight by the central government to curb excessive urban land-use expansion.}
explain the cross-city variation of \( \eta \) in Equation (6). These variables are described in Table 1, which also reports their sample statistics and data sources.

*** Insert Table 1 about here ***

Note that Chinese cities are geographically determined by administrative boundaries and typically include substantial rural areas and agricultural population. Urban growth involves conversion of rural land within the city jurisdiction into urban uses and conversion of the displaced farmers into urban residents and workers. In addition, urban population grew by receiving rural migrants.

**Measuring urban population growth** \( g_N \)

Our urban population data are based on the city non-agricultural population statistics from Urban Statistics Yearbook. The reliability of these urban population statistics are compromised, however, by administrative boundary changes of many cities during our sample period, which resulted in population changes due to reclassification rather than genuine growth (Shen, 2005). There is no reliable way to adjust the population statistics for these boundary changes; nevertheless we make a best-effort attempt using the following eyeballing procedure suggested by NBSC. First, we calculate the year-to-year growth in city non-agricultural population between 1990 and 2004. Second, we eyeball the growth pattern for individual cities and identify abrupt changes in the growth rate; we assume these abrupt changes to be associated with changes in city administrative boundaries. Third, we replace the abrupt changes with average population growth rate in the adjacent years. Finally we use these “smoothed” population growth figures to compute the city non-agriculture population in each year from the 1990 base non-agricultural population. The resulting adjusted urban (non-agricultural) population level in 1998 and the non-agricultural population growth between 1998 and 2004, \( g_N \), are plotted in Figure 1. The average population size of the 85 cities in 1998 is about 1.2 million; half of the cities are above 700,000 in population size. On average these cities grew by 18.5% during the six-year period (see Table 1), which probably understates the actual urban population growth for failure to account for new migrants without official urban household registration. But this measure of \( g_N \) offers reasonable consistency across cities, which is important for the purpose of our study. The (non-
agricultural) population growth ranges from zero to 40% in our sample of 85 cities, with a standard deviation of about 9%. 8

*** Insert Figure 1 about here ***

Measuring cross-city land rent differential $R$ and land rent growth $S$

Urban demand shocks are indicated by land rent growth $S$. But available housing price indexes in China, such as Zhong Fang property price index and Guo Fang Jing Qi property index, are based on average sale prices per square meter rather than constant quality indexes. They cover only a relatively small number of large cities and focus on the price changes of new homes. To obtain a consistent measure of base-period land rent premium $R$ and sample-period land rent growth $S$ for a relatively large cross-section of cities with respect to a more representative housing stock in each city, we have to resort to indirect estimates based on urban household survey data. Essentially these estimates represent the variations in price incentives across cities and over time as revealed by household housing consumption choices. We adopt the estimates provided by Zheng, Fu and Liu (2009), who pool the 1998 (base period) and 2004 UHS data to estimate a household housing demand equation that accounts for household income and demographic attributes, identifying the base-period (indexed by subscript $t=0$) land rent premium for city $k$ (relative to the base-period land rent in benchmark city, Beijing, $r_0$) $R_k = \ln(r_{k0}/r_0)$ and rent growth $S_k = \lambda_0 \ln(r_{kt}/r_{k0})$, respectively, as city fixed effects and these fixed effects multiplied by the year 2004 dummy. These $R_k$ and $S_k$ estimates are displayed in Figure 2.

*** Insert Figure 2 about here ***

Figure 3 plots the population growth measure $g_N$ against the demand shock measure $S$ across the 85 cities; the growth rates appear quite dispersed, suggesting possible wide variations in housing supply elasticity.

*** Insert Figure 3 about here ***

Glaeser and Gyourko (2005) show a strong asymmetry in housing supply elasticity for cities experiencing negative demand shocks due to housing durability. None of the cities in our sample seem to have been subject to negative demand shocks during our study period.
Computing predicted demand shocks $\hat{S}$

Land rent growth may be influenced by housing supply due to short-run overshooting in land rent adjustment to demand shocks or due to the sensitivity of urban productivity and amenity to urban size. In order to mitigate this potential endogeneity problem, we seek to predict the cross-city land rent growth $S_k$ based on base-period urban conditions that would influence the growth in urban productivity and amenity premium in a spatial equilibrium framework (Roback, 1982; Glaeser and Gottlieb, 2009). Specifically, consider a representative household described in Section 2, whose indirect utility in city $k$ and time period $t$ is given by:

$$v_t = A_{kt} \left[1 + r_k^{-\sigma}\right]^{-1} \frac{Y_{kt}}{r_{kt}}.$$ (7)

The independence of $v_t$ with respect to $k$ is a result of population mobility. Applying log difference to Equation (7) with respect to time $t$ (relative to the base-period $t=0$) and city index $k$ (relative to the benchmark city $k=0$), together with a first-order Taylor expansion of $\ln r_{kt}$ around the base-period log rent in the benchmark city, $\ln r_{00}$, we obtain: 9

$$S_k = (1 + \sigma r_0^{-\sigma})(\Delta_t \ln A_{kt} + \Delta_t \ln Y_{kt} - \Delta_t \ln v_t),$$ (8)

where $S_k = \lambda_0 \ln (r_k / r_0)$ represents the demand shock experienced by city $k$ over our study period from 1998 to 2004, $\Delta_t \ln A_{kt}$ is the growth in the households’ willingness to pay for the quality of living (amenities) in city $k$, and $\Delta_t \ln Y_{kt}$ is growth in urban wage premium (the firms’ willingness to pay for urban productivity differential). Equation (8) suggests two groups of predictors for observed urban land rent growth $S_k$. The variables in the first group predict the growth in the willingness to pay for urban quality of living; they include: city temperature index $TEMP$, defined as the distance of the city’s summer and winter temperature combination to the combination of the minimum summer temperature and maximum winter temperature across the cities, to reflect the severity of local climate; 10 $SO2$ emission and $GREEN$ space availability in the city, to reflect the environmental quality; the average years

---

9 Details of the derivation can be found in Zheng, Fu and Liu (2009).

10 A low $TEMP$ index indicates a relatively more temperate climate, whereas a high $TEMP$ index indicates a climate with either hot summer or harsh winter. Humidity would also be an important factor of climate amenity; however, the temperature zones are most important indicator of the overall climate amenity in China.
Variables in the second group predict the urban wage (productivity) growth $g_w$ (i.e. $\Delta \ln Y_t$, in Equation (8)) and include: employment structure measured by the tertiary sector share of employment $EMP\_TER$; enterprise competitiveness measured by the employment share of state-owned enterprises $EMP\_SOE$; and the 1998 unemployment rate $UMP$. In addition, we will control for the initial cross-city land rent and wage rate differentials, $R$ and $W$ respectively, as well as the cities’ administrative rank as a provincial capital or a provincial level city ($P\_CAPITAL$) and coastal geographical position ($COASTAL$). The $W$ and $g_w$ estimates are provided in Zheng, Fu and Liu (2009), where they are derived from a Mincerian wage regression with city and city-year fixed effects using household data from UHS.

*** Insert Table 2 about here ***

Table 2 reports the OLS estimates of the determinants of the demand shock $S$. Column 1 focus on the determinants of the growth in willingness to pay for urban quality of living. According to Equation (8) and assuming $\lambda_0=1/2$ (i.e. $r_0Q=\frac{Y}{4}$ and $\sigma=1/3$), we have:

$$\Delta \ln A_t = S_t/2 - g_w + \Delta \ln v_t,$$

where $\Delta \ln v_t$ is a constant representing the growth in household real income during our sample period. Using $S_t/2 - g_w$ as the dependent variable, we find a significant negative effect of $TEMP$, which suggests an increasing aversion to severe summer or winter temperatures as people get richer. Similarly, people are increasingly willing to pay to living in cleaner cities (lower SO2 emission per unit of GDP). Whereas the availability of green space contributes positively to the city’s demand shock, the social interaction quality ($EDU$) has an even stronger positive influence. People also increasingly value the quality of city’s healthcare services (produced by the combination of doctors and hospital beds, $DOCT\times H\_BED$). Finally, holding constant these amenity conditions, we find some degree of mean reversion in land rent differential across the cities.

Column 2 of Table 2 reports the reduced-form estimates of both the amenity demand shocks $\Delta \ln A_t$ and the productivity shocks $g_w$. We use $S$ as the dependent variable and include
additional independent variables to account for productivity growth. We find higher productivity growth in cities with initial location advantages (higher $R$). The employment structure matters but its effect is nonlinear. Cities with a tertiary sector employing either a very large share of local workforce or a very small share, probably specializing in either service export or manufacturing export, seem to experience a relatively bigger productivity growth between 1998 and 2004; a city where the tertiary sector accounts for around 47% of urban employment appears to be least competitive. Cities also benefit somewhat from the presence of state-owned enterprises (SOEs), which were typically capital rich compared to the mostly small domestic private or collectively-owned enterprises during the time, as long as the SOEs’ share of employment is not excessive (i.e. $EMP_{SOE} > 0.53$) so that the cost of inefficient incentive structures outweighs the benefit of the capital endowment. In addition, cities with either a very high unemployment rate (above 10%) or a very low unemployment rate (below 3%) would experience a higher productivity growth than those with relatively normal unemployment rates, as a result of economic recovery and exceptionally high local demand for labor, respectively. Holding these productivity shocks constant, we find significant wage-rate (productivity) convergence across the cities as the negative coefficient of $W$ suggests. Finally, we control for the cities’ administrative rank ($P_{CAPITAL}$) and geographic positions ($COASTAL$) and find neither to have any additional influence on amenity demand or productivity shocks. We compute the predicted demand shock $\hat{S}_k$ using the estimates reported in Column 2 of Table 2. $\hat{S}_k$, which is predicted by both productivity shocks and urban amenity premium shocks, explains about half of the variance of $S$.

Measuring urban land-use expansion, per-capita housing demand growth, and the determinants of housing density elasticity

The growth in urban land use within a city, $g_L$, is produced by two variables: the growth in urban built-up area $g_{BuiltUp}$ from land-use conversion and the growth in urban road space $g_{Road}$; their mean values are 0.347 and 0.708, respectively, as shown in Table 1. We use household income growth to measure the growth in household housing demand $Q$. We define the city

---

11 Extant studies often use productivity shocks alone to instrument the local housing price growth. Malpezzi and Maclennan (2001), for example, use income change and Mayer and Somerville (2000) and Quigley and Raphael (2005) use predicted local employment growth.
average household income growth $g_Y$ as the mean of the growth in the 25th percentile household income and that in the 75th percentile household income. The mean $g_Y$ across the 85 cities is 1.58. Note that $g_Y$ is different from $g_W$ in that the latter reflects the wage growth for a constant bundle of human capital but the former reflects also income growth due to increased household human capital and other sources of income.

The determinants of the housing density elasticity $\eta$ fall into two groups, one relating to redevelopment costs and the other to the allocation of land resources in the city in terms of both equity and efficiency. Several variables are included in the first group. The initial road space per capita of non-agricultural population, $ROADPC$, is expected to support a greater $\eta$ by reducing the marginal social cost of housing density and hence relaxing FAR regulations (see Fu and Somerville, 2001). The average age of the housing stock in the city, $H_{AGE}$, varying from 10 to 24 years in our city sample, is expected to raise the housing replacement demand and hence the cost of net new construction. Land assembly cost, in terms of compensation to the households displaced by urban redevelopment, is expected to increase with both the age of the housing stock (the likelihood of household displacement) and the initial land rent premium $R$ (the cost of compensation). The land assembly cost would further increase with the population density in the built-up area $DENSITY$, which would increase the number of households in any redevelopment site. The redevelopment cost, however, could be lower in $P_{CAPITAL}$ cities, if the governments of these high-ranking cities were more generous in subsidizing urban redevelopment or more effective in relocating households affected by urban redevelopment.

We further include a measure of local government efficiency as an additional determinant of the redevelopment cost, using two scores reported in a World Bank study of investment environment across 120 Chinese cities (World Bank 2006, Table B5) for ranking the local government efficiency. These two scores, denoted $G_{EFF_TFP}$ and $G_{EFF_FDI}$, are estimated based on local firms’ feedback regarding i) tax burden (local taxes as a fraction of value added), ii) government service quality (the number of days necessary for import and

---

12 The median household income statistics are not available to us. Also, since UHS does not use a fixed household penal, the household income growth statistics could be influenced by the urban population growth, although it is not obvious whether the influence would be positive or negative.

13 This variable is based on the report by home owners in the 2004 UHS and hence is not totally exogenous with respect to the population growth between 1998 and 2004. Unfortunately, building age was not reported in the 1998 UHS.
export to clear custom), iii) red tap (the amount of time spent dealing with government regulators), and iv) corruption cost (the amount of entertainment expenses per unit of sales). The $G_{EFF\_TFP}$ score is the expected gain in total factor productivity (TFP) by the local firms had the city’s conditions in those four aspects of government efficiency were improved to the 90th percentile level; the $G_{EFF\_FDI}$ score is the expected gain in local foreign direct investment level (indicating investment profitability) had the same improvements were achieved. Table 1 shows that $G_{EFF\_TFP}$ and $G_{EFF\_FDI}$ range from -1% to 7.6% and from -22% to 16%, respectively, among the 68 cities in our sample for which the scores are available. A negative score indicates that the city exceeded the 90th percentile level of efficiency; whereas a large positive score means that the city had much to improve. We define our measure of local government efficiency $G_{EFF}$ to be the average of the standardized $G_{EFF\_TFP}$ and $G_{EFF\_FDI}$ scores and we expect a lower $G_{EFF}$ score to correlate with lower redevelopment costs; the conditions that raise the productivity and the investment profitability for private enterprises generally should do the same for residential development.

The second group of the determinants of $\eta$ relate to land-use efficiency and equity in individual cities. The land-use efficiency at the city level is affected by land-use reform (World Bank, 1993), as urban land was allocated administratively during the planned economy and SOEs had little incentive to economize on their land uses. We use the ratio of two variables to indicate the land-use reform and land-use efficiency in a city: (i) the presence of $FDI$, defined as the cumulative foreign direct investment as a fraction of city’s fixed investment between 1990 and 1998, which often boosted urban land reform and redevelopment; (ii) the size of SOE sector, measured by $EMP\_SOE$, to account for the amount of non-marketable state land occupied by SOEs and their employee housing projects, which often impeded the urban redevelopment in response to demand shocks. To measure the equity of land use allocation in a city, we include two variables: the extent of income inequality in the city $Y_{75/25}$, defined as the ratio of the city 75th percentile household income to the 25th percentile, and that in the province $Y_{90/75\_Province}$, defined as the ratio of the provincial 90th percentile household income over the 75th percentile. Between 1998 and 2004, urban governments produced few homes for low-income households and largely left the choice of new residential developments to private developers so as to maximize the revenue from land-

---

14 For Beijing and Shanghai, $Y_{90/75\_Province}$ is defined as national 90th percentile household income over the 75th percentile household income, as these cities tend to attract housing demand of the wealthy across the country.
use conversion; hence more extreme income inequality would be expected to produce more excessive supply of low-density luxury homes at the expense of the overall housing density in the city.15

4. Estimating urban population growth and housing density elasticity

Between 1998 and 2004, by our conservative measure of $g_N$, the 85 cities in our sample on average grew nearly 20% in terms of non-agricultural population (based on fixed city administrative boundaries), a quite remarkable rate of growth for cities of an average size of 1.2 million people in 1998. This rate of growth, however, is considerably smaller in comparison with the average growth in urban built-up area $g_{BuiltUp}$ and urban road space $g_{Road}$, about 35% and 70% respectively (see Table 1) during the same period. A considerable amount of increased housing supply during that period was absorbed by increased household housing consumption $Q$ as a result of household income growth. Shown in Table 1, household income growth ($g_Y$) during the period averages 1.581. If we assume a unitary income elasticity, the resulting growth in housing absorption by existing urban households would be much greater than the average reduction in $Q$ due to price growth ($S = \lambda_0 g_r$) of 0.365. The net growth in housing absorption is $g_Q = g_Y - S = 1.216$ according to Equation (4), which, together with non-negative change in urban amenity premium, imply a substantial increase in household real income $v_t$ according to Equation (9).16 Thus taking into account $g_Q$ in Equation (6) is essential for identifying the influence of housing supply on urban population growth.

*** Insert Table 3 about here ***

Examining the correlation coefficients in Table 3, we note that the population growth $g_N$ is positively correlated with the growth in urban built-up area $g_{BuiltUp}$ and in urban road space $g_{Road}$, with the initial road space availability $ROADPC$, the initial urban quality indicated by $R$, and the ratio $FDI/EMP\_SOE$, but is negatively correlated with the household income growth

---

15 In 2006, the State Council issued an executive guideline on the size mix of new housing supply in cities (关于调整住房供应结构稳定住房价格的意见), which requires 70% of the dwelling units in new development projects to be no larger than 90 square meters in size, so as to increase the supply of more affordable urban dwelling units.

16 According to the Urban Statistics Yearbook, the average housing living area per person across 219 (of the 287) cities at prefectural level or above increased by 55%, from 13.2 sqm in 1998 to 20.4 sqm in 2004. Of course, the increased living area per person is just part of the improvement in housing condition; much of the increased spending on housing would be in the form of improved building, community and location qualities.
Both the measured and the predicted demand shocks,  and ̂S, are weakly correlated with \( g_N \) and ̂S has little correlation with those housing-supply related variables. In particular, ̂S is very weakly correlated with ̂g\(_{\text{BuiltUp}}\) and ̂g\(_{\text{Road}}\), suggesting that much of the urban land use expansion resulting from land-use conversion and road construction during the period was probably independent of the demand shocks.

*** Insert Table 4 about here ***

The main results of this study, the OLS estimates of Equation (6), which accounts population growth in terms of housing supply, are reported in Table 4. We employ the predicted land rent growth ̂S as our measure of the demand shocks. The regression reported in first column assumes a constant density elasticity \( \eta \). The effects of ̂g\(_{\text{BuiltUp}}\) and ̂g\(_{\text{Road}}\) are separately positive and significant, indicating that both the land-use conversion and road construction within cities contribute to new housing supply. The growth in household income \( g_Y \) has a significant and negative effect on population growth, as the increased absorption of housing supply by the existing urban population reduces the new supply available to new comers. However, with the assumption of a constant \( \eta \), the demand shock ̂S contributes very little to the accounted cross-city variance in population growth.

In column 2 of Table 4, the housing density supply is allowed to vary according to city attributes. In addition, we combine the effects of ̂g\(_{\text{BuiltUp}}\) and ̂g\(_{\text{Road}}\), assuming the growth in residential land use to be produced by a constant-return-to-scale Cobb-Douglas function of land-use conversion and road construction: ̂gL \( \equiv \) ̂g\(_{\text{BuiltUp}}\)×0.44 + ̂g\(_{\text{Road}}\)×0.56; the linear coefficients are chosen proportional to the estimates of the separate effects if the two variables are independently included in column 2. Doing such combination does not affect the estimates of the other variables in column 2 but does somewhat improve their \( t \)-statistics due to increased degree of freedom, given the relative small number of observations in our sample. Overall, the variation in housing density elasticity accounts for an additional 24% of the cross-city variance in population growth—the adjusted \( R^2 \) increases from 0.399 in column 1 to 0.637 in column 2. As expected, the initial availability of road space ROADPC supports higher housing density in the city but the average age of the housing stock \( H_{\text{AGE}} \), the initial land-rent cost premium \( R \), and the built-up-area population density \( DENSITY \) raise the redevelopment cost (hence reducing \( \eta \)), although the last two variables are not statistically
significant. To mitigate the endogeneity problem, the $H_{AGE}$ effects are evaluated in two ways, through a fixed effect for the 12 cities with the most aging housing stock ($H_{AGE} > 18$) and through the effect of $R$ in the 59 cities with a relatively older housing stock ($H_{AGE} > 13.5$; the redevelopment cost in these cities would be more sensitive to $R$ because of the need to resettle the residents in the older buildings replaced by new development). We further find provincial capital and provincial level cities ($P_{CAPITAL}$) to have higher housing density elasticity, possibly due to more favorable support for urban redevelopment projects.

Land-use efficiency, indicated by the presence of foreign direct investment relative to the size of the SOE employment share ($FDI/EMP_{SOE}$), contributes to a higher housing density elasticity. In contrast, the inequity of land-use allocation, as indicated by the fixed effect $INEQ$ of cities with high income inequality in 1998, appears to lower housing density elasticity, consistent with our hypothesis that the intra-city income inequality encourages excessive supply of luxury housing developments at the expense of higher-density more-affordable housing developments. $INEQ$ selects 12 (of the 85) cities that have $Y_{75/25} > 1.81$ in base period and fail to reduce $Y_{75/25}$ during the sample period. The selection could be somewhat influenced by the urban growth experience; but higher growth would likely add more low-income population to the city and cause $Y_{75/25}$ to rise. Thus any endogeneity in $INEQ$ selection is likely to bias the estimate to understate the negative effect of $INEQ$ on urban growth. In addition, large income inequality in a province ($Y_{90/75\_Providence} > 1.37$) reduces the housing density elasticity of its capital city.\textsuperscript{17} Provincial capital cities often attract the wealthy in the province to buy luxury homes; hence province-wide household income inequality would contribute additional demand for luxury housing development in the provincial capital cities.

Note that some of the independent variables in the regression, including ln($ROADPC$), ln($DENSITY$), $R$, and $FDI/EMP_{SOE}$, are shifted by a fixed value, typically the sample median. We do so to facilitate the evaluation of the demand shock effect in a meaningfully defined base-scenario city. We will evaluate the magnitude of the housing density elasticity $\eta$ for such a base-scenario city using the estimates reported in column 3.

\textsuperscript{17} The 5 provincial capital (provincial level) cities having $Y_{90/75\_Providence} > 1.37$ are Beijing, Shanghai, and the capital cities of Guangdong, Xinjiang and Heilongjiang provinces.
In column 3 we include $G_{EFF}$ to further account for cross-city differences in redevelopment cost due to the variation in local government efficiency. We find the $G_{EFF}$ score to have a significant impact on the housing density response to demand shocks, suggesting an increased redevelopment cost where the local government imposes greater tax and regulatory burden on enterprises. Further, the $G_{EFF}$ effect is nonlinear; when the negative exponential function is used (as shown in Table 4) in place of the linear form (no shown), the $R^2$ squared increased from 0.773 to 0.788. In other words, the redevelopment cost would increase quickly as $G_{EFF}$ deteriorates from the best (low) score. The estimates for the other variables pertaining to redevelopment costs become more significant when the local government efficiency is taken into account. It is notable that the estimates for the bottom half of the determinants of $\eta$, relating to institutional factors, are generally more statistically significant, than those associated with the upper half of the determinants, relating to urban physical conditions. Moreover, the variation in housing density elasticity accounts for nearly as much cross-city population growth variance as do the variations in urban land-use growth and household income growth (the adjusted $R^2$ squared obtained when $\eta$ is assumed constant, whether based on the full sample or the smaller sample of 68 cities, is roughly about half the adjusted $R^2$ squared value reported in column 3).

To evaluate the magnitude of the housing density elasticity $\eta$, we consider a base-scenario city: a non-capital city with median road space per capita ($ROADPC = 6.9$) in 1998, not very old housing stock ($H_{AGE}<18$), half the base-period land rent premium of Beijing ($R=\ln(0.5)$), a median population density in the built-up area ($DENSITY=1.1$), a median ratio of $FDI_{/EMP\_SOE}$, a moderate level of household income inequality ($INEQ=0$), and a government efficiency at the 90th percentile level ($G_{EFF}=0$). The estimates in column 3 shows that a 10 percent demand shock ($\hat{S}=10\%$, or $g_s=20\%$ assuming $\lambda_0=0.5$) in this base-scenario city would generate about 3.17 percent population growth via increased housing density (and reduced per-capita housing consumption) in the city, with urban land use and household income held constant ($g_L=g_Y=0$). This estimated effect of $\hat{S}$ ($\beta_3=0.317$) is certainly biased downward due to measurement errors in the various growth measures but is still reasonably strong in magnitude compared with the effect of the urban land-use growth $g_L \equiv g_{BuiltUp} \times 0.44 + g_{Road} \times 0.56$ ($\beta_1=0.136$) and the effect of household income growth $g_Y$ ($\beta_2=0.097$). Each percentage point increase in $\hat{S}$ generates in the base-scenario city, via reduced per-capita housing consumption and increased housing density, an additional housing
supply for population growth equivalent to 2~3 times what one percent increase in the urban land use would add to total housing supply or what one percent increase in household income would add to total housing consumption. The relative magnitude of these estimates could provide a rough indication of the size of \( \eta \) in Equation (6).\(^{18}\) Assuming an equal effect of \( g_Y \) and \( S \) on household housing consumption growth \( g_Q \) (Equation (4)) and an equal effect of \( g_L \) and \( g_h \) on housing output, we have \( \eta / \lambda_0 = (\beta_i - \beta_1) / \beta_1 = 1.63 \), or \( \eta = 0.814 \) for \( \lambda_0 = 0.5 \).\(^{19}\) If we further use the median value of the \( \lambda_0 g_L / S \) ratio, which is 0.733 across the 85 cities, as an indication of the urban land use elasticity with respect to \( g_h \), our base-scenario housing supply elasticity (from both urban land use growth and housing density increase) would be around 1.55, above the median housing supply elasticity of 1.34 for the 95 US metropolitan areas reported in Saiz (2008). Furthermore, for the base-scenario city, the housing density elasticity \( \eta \) would be somewhat greater than the median elasticity of urban land-use growth.

5. Conclusions

We have sought to improve the identification of housing supply elasticity by incorporating inter-city spatial equilibrium in our analysis of the determinants of housing density elasticity in the context of China’s recent urban growth. In the open-city spatial equilibrium framework, urban land rent growth is a sufficient indication of the demand shocks experienced by individual cities and the resulting population growth reveals three components of net new housing supply, namely, new home construction due to urban land-use growth, growth in housing absorption by existing urban households, and increase in housing density in response to the demand shocks. Identifying housing supply elasticity via quantity (urban population growth) response to urban land rent growth, rather than via reverse response, is important, for the response of land rent growth to population shocks would reveals marginal agglomeration economies rather than housing supply elasticity in the spatial equilibrium framework. Our analysis takes urban land use growth in individual cities as independent of the demand shocks in order to focus on the housing density elasticity. Such a focus is by choice, but the model

---

\(^{18}\) In the accounting framework, increased effects of \( g_L \) (\( \beta_1 \)) and \( S \) (\( \beta_5 \)) must be compensated by an increased effect of \( g_Y \) (\( \beta_2 \)), and vice versa, holding constant the amount of the variance of \( g_N \) explained by the regression.

\(^{19}\) \( \beta_i \) coefficient would be larger, hence the \( \eta \) value smaller, if the growth in residential land use is smaller than the overall urban land use growth \( g_L \). Although residential land use generally dominates urban expansion in developed economies, growth in manufacturing land use is often significant in Chinese cities. The \( \eta \) value would be larger if the elasticity of substitution \( \sigma \), hence \( \lambda_0 \), is greater; for \( \sigma = 1/2 \), for example, \( \lambda_0 = 5/8 \) and \( \eta = 1.02 \).
allows urban land-use growth to be responsive to urban demand shocks. In the context of China’s urban growth since early 1990s, on the one hand, converting rural land within city political boundaries to urban uses is the monopoly of the city government, whose political mandate centers on promoting local economic growth; on the other hand, housing development density is often driven by profit maximization motives in response to the demand shocks.

Our empirical analysis contributes to the knowledge of how redevelopment costs and land-use allocation efficiency and equity affect the response of housing development density to urban demand shocks in the context of a developing economy, complementing an existing literature that to a large extent focuses on housing supply affected by the extensive margin of land-use regulations and constraints. In particular, we find local government efficiency to be an important source of inter-city variation in redevelopment costs, urban economic reform indicated by the inflow of FDI and privatization of SOEs to raise urban land-use efficiency, and income inequality to produce land-use inequity that reduces housing density elasticity. These institutional factors appear to have more significant influence on the cross-city variation in the housing density elasticity than do urban physical conditions. It is also notable that, even in Chinese cities where population density is generally very high, responses in housing development density to urban demand shocks is as important as urban land-use growth with respect to new housing supply to accommodate urban growth.

Although our empirical results need be viewed with some caution given the approximate measures of urban population growth and land rent growth that we have to rely on for our analysis, a few preliminary implications can be drawn for housing market enabling policies. First, there appears a significant potential for expanding new housing supply in densely populated developing-economy cities by improving market responsiveness in development density choices to demand shocks. Second, in addition to investment in urban infrastructure, improving local government efficiency and enhancing urban land-use allocation efficiency and equity can have a significant influence in raising housing density elasticity. Further research to test these implications in the context of more developing economies would be a worthwhile effort.

The median population density in our sample of 85 Chinese cities is about 11,000 people/sqkm. By comparison, the density was about 4,000 people per sqkm in New York city according to the 1990 census.
References


Table 1. Variable definition, data source and sample statistics

<table>
<thead>
<tr>
<th>Variables</th>
<th>Definition [Data source]</th>
<th>Mean</th>
<th>Median</th>
<th>Max</th>
<th>Min</th>
<th>Std. Dev.</th>
<th>No. of obs.</th>
</tr>
</thead>
<tbody>
<tr>
<td>$g_N$</td>
<td>City “smoothed” log growth in non-agricultural population growth, 1998-2004. [3]</td>
<td>0.185</td>
<td>0.195</td>
<td>0.404</td>
<td>0.006</td>
<td>0.092</td>
<td>85</td>
</tr>
<tr>
<td>$S$</td>
<td>Urban demand shock, 1998-2004: $S_k = \frac{1 + \sigma^\alpha_k}{1 + \sigma^\alpha_0} \ln \frac{r_k}{r_0}$. Adopted from [2].</td>
<td>0.365</td>
<td>0.355</td>
<td>0.701</td>
<td>0.040</td>
<td>0.137</td>
<td>85</td>
</tr>
<tr>
<td>$\hat{S}$</td>
<td>Predicted $S$, based on the estimates in the last column of Table 2.</td>
<td>0.365</td>
<td>0.360</td>
<td>0.577</td>
<td>0.030</td>
<td>0.098</td>
<td>85</td>
</tr>
<tr>
<td>$g_W$</td>
<td>Urban wage (productivity) growth ($\Delta \ln Y_{uk}$), 1998-2004. Adopted from [2].</td>
<td>1.451</td>
<td>1.451</td>
<td>1.885</td>
<td>1.194</td>
<td>0.135</td>
<td>85</td>
</tr>
<tr>
<td>TEMP</td>
<td>City temperature index: the distance of the city’s summer and winter temperature ($^\circ$C/100) combination to the combination of the min. summer temperature and max. winter temperature across the cities. [3]</td>
<td>0.195</td>
<td>0.170</td>
<td>0.348</td>
<td>0.082</td>
<td>0.083</td>
<td>85</td>
</tr>
<tr>
<td>$SO2$</td>
<td>City SO2 emission over GDP (ton per million Yuan), 1998. [3]</td>
<td>3.127</td>
<td>1.665</td>
<td>22.05</td>
<td>0.099</td>
<td>4.120</td>
<td>85</td>
</tr>
<tr>
<td>GREEN</td>
<td>City green space per capita, (sqm), 1998. [3]</td>
<td>2.792</td>
<td>1.900</td>
<td>27.92</td>
<td>0.190</td>
<td>3.588</td>
<td>85</td>
</tr>
<tr>
<td>$EDU$</td>
<td>City average years of schooling of adult population (years). [3]</td>
<td>11.62</td>
<td>11.61</td>
<td>12.87</td>
<td>9.918</td>
<td>0.541</td>
<td>85</td>
</tr>
<tr>
<td>$DOCT$</td>
<td>Urban area number of doctors per 100 non-agricultural population, 1998. [3]</td>
<td>0.618</td>
<td>0.540</td>
<td>1.542</td>
<td>0.093</td>
<td>0.289</td>
<td>85</td>
</tr>
<tr>
<td>$H_BED$</td>
<td>Urban area number of hospital beds per 100 non-agricultural population, 1998. [3]</td>
<td>1.008</td>
<td>0.904</td>
<td>2.247</td>
<td>0.196</td>
<td>0.385</td>
<td>85</td>
</tr>
<tr>
<td>$R$</td>
<td>City land rent premium: $R_k = \ln(r_k/r_0)$, 1998, Beijing=0. Adopted from [2]</td>
<td>-1.008</td>
<td>-1.007</td>
<td>0.000</td>
<td>-1.866</td>
<td>0.411</td>
<td>85</td>
</tr>
<tr>
<td>$EMP_TER$</td>
<td>City tertiary-sector employment share, 1998. [3]</td>
<td>0.438</td>
<td>0.440</td>
<td>0.676</td>
<td>0.249</td>
<td>0.095</td>
<td>85</td>
</tr>
<tr>
<td>$EMP_SOE$</td>
<td>City state-owned-enterprise (SOE) employment share, 1998. [3]</td>
<td>0.686</td>
<td>0.702</td>
<td>0.944</td>
<td>0.369</td>
<td>0.121</td>
<td>85</td>
</tr>
<tr>
<td>$UMP$</td>
<td>City urban-area unemployment rate. [3]</td>
<td>0.039</td>
<td>0.032</td>
<td>0.154</td>
<td>0.000</td>
<td>0.029</td>
<td>85</td>
</tr>
<tr>
<td>$W$</td>
<td>City wage (productivity) premium: $\ln(Y_{uk}/Y_t)$, 1998, Beijing=0. [2]</td>
<td>-0.370</td>
<td>-0.408</td>
<td>0.846</td>
<td>-0.892</td>
<td>0.283</td>
<td>85</td>
</tr>
<tr>
<td>$P_CAPITAL$</td>
<td>Dummy for provincial capital city or provincial level city.</td>
<td>0.341</td>
<td>0.000</td>
<td>1</td>
<td>0</td>
<td>0.477</td>
<td>85</td>
</tr>
<tr>
<td>$COASTAL$</td>
<td>Dummy for cities in coastal provinces.</td>
<td>0.459</td>
<td>0.000</td>
<td>1</td>
<td>0</td>
<td>0.501</td>
<td>85</td>
</tr>
<tr>
<td>$g_{\text{BuiltUp}}$</td>
<td>Growth in urban built-up area, 1998-2004. [3]</td>
<td>0.347</td>
<td>0.288</td>
<td>1.386</td>
<td>-0.145</td>
<td>0.309</td>
<td>85</td>
</tr>
<tr>
<td>$g_{\text{Road}}$</td>
<td>Growth in road space, 1998-2004. [3]</td>
<td>0.708</td>
<td>0.635</td>
<td>1.887</td>
<td>0.000</td>
<td>0.368</td>
<td>85</td>
</tr>
<tr>
<td>$g_Y$</td>
<td>Average of the log growth in the 25th percentile household income and that in the 75th percentile household income, 1998-2004. [1]</td>
<td>1.581</td>
<td>1.596</td>
<td>2.059</td>
<td>1.206</td>
<td>0.150</td>
<td>85</td>
</tr>
</tbody>
</table>
### Descriptive Statistics

<table>
<thead>
<tr>
<th>Variable</th>
<th>Description</th>
<th>Mean</th>
<th>Median</th>
<th>Min</th>
<th>Max</th>
<th>Sample Size</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>H_AGE</strong></td>
<td>City average building age of homes in 2004. [1]</td>
<td>15.07</td>
<td>14.80</td>
<td>10.40</td>
<td>24.00</td>
<td>85</td>
</tr>
<tr>
<td><strong>DENSITY</strong></td>
<td>City non-agricultural population over built-up area, 10,000 people /sqkm, 1998. [3]</td>
<td>1.131</td>
<td>1.107</td>
<td>0.520</td>
<td>2.169</td>
<td>85</td>
</tr>
<tr>
<td><strong>G_EFF_TFP</strong></td>
<td>The expected gain in total factor productivity (TFP) by local firms had the city government efficiency indicators were improved to the 90th percentile level. [4]</td>
<td>0.169</td>
<td>0.165</td>
<td>0.340</td>
<td>-0.010</td>
<td>68</td>
</tr>
<tr>
<td><strong>G_EFF_FDI</strong></td>
<td>The expected gain in local foreign direct investment level had the city government efficiency indicators were improved to the 90th percentile level. [4]</td>
<td>0.319</td>
<td>0.330</td>
<td>0.660</td>
<td>-0.220</td>
<td>68</td>
</tr>
<tr>
<td><strong>G_EFF</strong></td>
<td>City government efficiency score (= 0.5 \times \frac{(G_EFF_TFP / \text{stdev}(G_EFF_TFP)) + (G_EFF_FDI / \text{stdev}(G_EFF_FDI))}{2})</td>
<td>2.100</td>
<td>2.109</td>
<td>3.745</td>
<td>-0.134</td>
<td>68</td>
</tr>
<tr>
<td><strong>FDI</strong></td>
<td>Foreign direct investment share of city fixed investment from 1990 to 1998. [3]</td>
<td>0.022</td>
<td>0.014</td>
<td>0.136</td>
<td>0.001</td>
<td>85</td>
</tr>
<tr>
<td><strong>FDI/EMP_SOE</strong></td>
<td></td>
<td>0.039</td>
<td>0.020</td>
<td>0.368</td>
<td>0.001</td>
<td>85</td>
</tr>
<tr>
<td><strong>Y_{75/25}</strong></td>
<td>City 75th percentile household income over 25th percentile household income, 1998. [1]</td>
<td>1.715</td>
<td>1.679</td>
<td>2.380</td>
<td>1.400</td>
<td>85</td>
</tr>
<tr>
<td><strong>INEQ</strong></td>
<td>High-inequality city dummy: cities with (Y_{75/25}&gt;1.81) and non-decreasing. [1]</td>
<td>0.141</td>
<td>0.000</td>
<td>1</td>
<td>0</td>
<td>85</td>
</tr>
<tr>
<td><strong>Y_{90/75_Province}</strong></td>
<td>Provincial 90th percentile household income over 75th percentile household income, for provincial capital and provincial level cities (except Beijing and Shanghai, for which the value is based on the national percentile household income), 1998. [1]</td>
<td>1.321</td>
<td>1.309</td>
<td>1.508</td>
<td>1.243</td>
<td>29</td>
</tr>
</tbody>
</table>

Note: The sources of data include: [1] Urban Household Survey (UHS); [2] Zheng, Fu and Liu (2009), which compute \(S, g_w, R,\) and \(W\) using household data from [1]; [3] Urban Statistics Yearbook; [4] World Bank (2006, Table B5). Our sample of 85 cities are in both the 1998 and 2004 UHS. UHS is conducted by the National Bureau of Statistic of China (NBSC) and covers about 200 cities each year, representing all provinces and population-size groups. Cities are sorted by average wage within each group and sampled at fixed distances. In each city, streets and neighborhoods are sorted and sampled at fixed distance, followed by the sampling of households within the selected neighborhoods. Cities missing necessary data in [3] are dropped.
Table 2. OLS estimates of cross-city demand shock determinants

<table>
<thead>
<tr>
<th>Independent \ Dependent variables</th>
<th>Column 1</th>
<th>Column 2</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>( \frac{S}{2-g_w} )</td>
<td>( S )</td>
</tr>
<tr>
<td>Constant</td>
<td>-2.431 (6.6) ***</td>
<td>0.341 (1.0)</td>
</tr>
<tr>
<td>TEMP</td>
<td>-0.588 (3.7) ***</td>
<td>-0.539 (3.4) ***</td>
</tr>
<tr>
<td>SO2</td>
<td>-0.006 (2.3) **</td>
<td>-0.008 (3.1) ***</td>
</tr>
<tr>
<td>GREEN</td>
<td>0.004 (1.7)</td>
<td>0.005 (2.6) **</td>
</tr>
<tr>
<td>EDU</td>
<td>0.102 (3.3) ***</td>
<td>0.063 (3.0) ***</td>
</tr>
<tr>
<td>DOCT×H_BED</td>
<td>0.047 (3.1) ***</td>
<td>0.064 (3.3) ***</td>
</tr>
<tr>
<td>R</td>
<td>-0.064 (2.0) **</td>
<td>0.136 (3.2) ***</td>
</tr>
<tr>
<td>EMP_TER×(EMP_TER−0.94)</td>
<td>3.226 (3.9) ***</td>
<td></td>
</tr>
<tr>
<td>EMP_SOE×(EMP_SOE−1.06)</td>
<td>-0.447 (1.5)</td>
<td></td>
</tr>
<tr>
<td>UMP×(UMP−0.1)</td>
<td>26.75 (5.1) ***</td>
<td></td>
</tr>
<tr>
<td>W</td>
<td>-0.359 (6.0) ***</td>
<td></td>
</tr>
<tr>
<td>P_CAPITAL</td>
<td>-0.016 (0.6)</td>
<td></td>
</tr>
<tr>
<td>COASTAL</td>
<td>-0.027 (0.8)</td>
<td></td>
</tr>
<tr>
<td>R squared</td>
<td>0.385</td>
<td>0.515</td>
</tr>
</tbody>
</table>

Note: t-statistics in parentheses are based on White Heteroskedasticity-Consistent standard errors & covariance. ***, **, and * denote, respectively, statistical significance at 1%, 5% and 10% levels. The number of observations is 85.
Table 3. Correlation matrix of urban demand shocks, population growth, and the housing-supply related variables (85 cities)

<table>
<thead>
<tr>
<th>Variables (ID)</th>
<th>ID</th>
<th>2</th>
<th>3</th>
<th>4</th>
<th>5</th>
<th>6</th>
<th>7</th>
<th>8</th>
<th>9</th>
<th>10</th>
<th>11</th>
<th>12</th>
</tr>
</thead>
<tbody>
<tr>
<td>S</td>
<td>1</td>
<td>0.718</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Ŝ</td>
<td>2</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>gN</td>
<td>3</td>
<td>0.112</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>g_BuiltUp</td>
<td>4</td>
<td>0.032</td>
<td>0.577</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>g_Road</td>
<td>5</td>
<td>0.020</td>
<td>0.434</td>
<td>0.508</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>g_Y</td>
<td>6</td>
<td>0.041</td>
<td>-0.332</td>
<td>-0.177</td>
<td>-0.072</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>ln(ROADPC)</td>
<td>7</td>
<td>0.002</td>
<td>0.479</td>
<td>0.275</td>
<td>-0.084</td>
<td>-0.261</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>H_AGE</td>
<td>8</td>
<td>-0.134</td>
<td>-0.095</td>
<td>-0.064</td>
<td>0.030</td>
<td>0.073</td>
<td>-0.207</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>R</td>
<td>9</td>
<td>0.134</td>
<td>0.410</td>
<td>0.434</td>
<td>0.401</td>
<td>-0.022</td>
<td>0.247</td>
<td>0.048</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>FDI/EMP_SOE</td>
<td>10</td>
<td>-0.070</td>
<td>0.475</td>
<td>0.329</td>
<td>0.073</td>
<td>-0.125</td>
<td>0.405</td>
<td>0.075</td>
<td>0.368</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>INEQ</td>
<td>11</td>
<td>-0.043</td>
<td>-0.148</td>
<td>0.079</td>
<td>0.131</td>
<td>-0.083</td>
<td>-0.132</td>
<td>-0.166</td>
<td>-0.019</td>
<td>-0.020</td>
<td></td>
<td></td>
</tr>
<tr>
<td>P_CAPITAL</td>
<td>12</td>
<td>-0.088</td>
<td>0.192</td>
<td>0.218</td>
<td>0.158</td>
<td>0.029</td>
<td>-0.106</td>
<td>0.130</td>
<td>0.375</td>
<td>0.010</td>
<td>0.136</td>
<td></td>
</tr>
<tr>
<td>ln(DENSITY)</td>
<td>13</td>
<td>-0.086</td>
<td>-0.114</td>
<td>0.181</td>
<td>0.077</td>
<td>0.106</td>
<td>-0.477</td>
<td>0.426</td>
<td>0.029</td>
<td>-0.089</td>
<td>0.008</td>
<td>0.476</td>
</tr>
</tbody>
</table>
Table 4. OLS estimates of cross-city population growth and housing density elasticity

<table>
<thead>
<tr>
<th>Independent \ Dependent variables</th>
<th>1</th>
<th>2</th>
<th>3</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>0.309 (3.9)**</td>
<td>0.225 (3.3)**</td>
<td>0.226 (3.1)**</td>
</tr>
<tr>
<td>$g_{\text{BuiltUp}}$</td>
<td>0.129 (4.8)**</td>
<td>0.177 (7.7)**</td>
<td>0.136 (7.4)**</td>
</tr>
<tr>
<td>$g_{\text{Road}}$</td>
<td>0.049 (2.1)**</td>
<td>0.136 (7.4)**</td>
<td>0.317 (2.8)**</td>
</tr>
<tr>
<td>$g_t = g_{\text{BuiltUp}} \times 0.44 + g_{\text{Road}} \times 0.56$</td>
<td>0.129 (4.8)**</td>
<td>0.177 (7.7)**</td>
<td>0.136 (7.4)**</td>
</tr>
<tr>
<td>$g_y$</td>
<td>-0.152 (3.5)**</td>
<td>-0.109 (2.4)**</td>
<td>-0.097 (2.1)**</td>
</tr>
<tr>
<td>$\hat{S}$</td>
<td>0.102 (1.4)</td>
<td>0.049 (0.7)</td>
<td>0.317 (2.8)**</td>
</tr>
<tr>
<td>$\ln(\text{ROADPC}/6.9) \times \hat{S}$</td>
<td>0.109 (2.3)**</td>
<td>0.115 (2.3)**</td>
<td></td>
</tr>
<tr>
<td>$(\text{H} _ \text{AGE}&gt;18) \times \hat{S}$</td>
<td>-0.090 (1.9)*</td>
<td>-0.126 (2.5)**</td>
<td></td>
</tr>
<tr>
<td>$(\text{H} _ \text{AGE}&gt;13.5) \times (\text{R} - \ln(0.5)) \times \hat{S}$</td>
<td>-0.083 (1.6)</td>
<td>-0.181 (2.9)**</td>
<td></td>
</tr>
<tr>
<td>$\ln(\text{DENSITY}/1.1) \times \hat{S}$</td>
<td>-0.104 (1.4)</td>
<td>-0.132 (2.0)**</td>
<td></td>
</tr>
<tr>
<td>$\text{P} _ \text{CAPITAL} \times \hat{S}$</td>
<td>0.146 (3.2)**</td>
<td>0.218 (4.7)**</td>
<td></td>
</tr>
<tr>
<td>$(\exp(-0.71\times G _ \text{EFF}) - 1) \times \hat{S}$</td>
<td>0.418 (4.7)**</td>
<td></td>
<td></td>
</tr>
<tr>
<td>$(\text{FDI}/\text{EMP _ SOE} - 0.02) \times \hat{S}$</td>
<td>1.140 (3.2)**</td>
<td>1.010 (4.8)**</td>
<td></td>
</tr>
<tr>
<td>$\text{INEQ} \times \hat{S}$</td>
<td>-0.171 (2.7)**</td>
<td>-0.228 (4.4)**</td>
<td></td>
</tr>
<tr>
<td>$\text{P} _ \text{CAPITAL} \times (Y_{90/75 _ \text{Province}} &gt; 1.37) \times \hat{S}$</td>
<td>-0.154 (3.2)**</td>
<td>-0.112 (2.0)**</td>
<td></td>
</tr>
<tr>
<td>$R$ squared</td>
<td>0.427</td>
<td>0.685</td>
<td>0.788</td>
</tr>
<tr>
<td>Adjusted $R$ squared</td>
<td>0.399</td>
<td>0.637</td>
<td>0.742</td>
</tr>
<tr>
<td>Number of observations</td>
<td>85</td>
<td>85</td>
<td>68</td>
</tr>
</tbody>
</table>

Note: The regression equation is given by Equation (6). *-statistics in parentheses are based on White Heteroskedasticity-Consistent standard errors & covariance. ***, **, and * denote, respectively, statistical significance at 1%, 5% and 10% levels.
Figure 1. 1998 population size and 1998-2004 population growth across 85 Chinese cities

1998 Non-agricultural Population, natural logarithm

Urban Population Growth (adjusted)

0.00 0.05 0.10 0.15 0.20 0.25 0.30 0.35 0.40 0.45

0.00 0.05 0.10 0.15 0.20 0.25 0.30 0.35 0.40 0.45

Population Growth and Housing Supply across Chinese Cities
Figure 2. Imputed land rent differential $R$ and rent growth $S$ across 85 Chinese cities, 1998-2004
Figure 3. Population growth and land rent growth (demand shocks) across 85 Chinese cities, 1998-2004