Population Growth across Chinese Cities: Urban Land Expansion versus Housing Density Response to Demand Shocks

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Abstract

We study the extent to which responses in urban redevelopment density to demand shocks contribute to population growth, taking into account urban land expansions independent of demand shocks and changes in housing absorption by existing urban households. In the context of Chinese urban growth from 1998 to 2004 when liberalization in urban housing and labor markets significantly elevated labor mobility, we show that the variation in housing density response contributes as much to the cross-city population growth variance as do the variations in urban land expansion and per-capita housing demand growth. Furthermore, we find a number of institutional factors, such as local government efficiency, urban land reform and intra-city income inequality, to have significant influences on the housing-density supply elasticity.

Keywords: housing supply elasticity, urban redevelopment, demand shocks, urban growth

JEL classification: R11, R31, R58
Urban growth plays a central role in the economic growth of a nation. Productivity growth depends on the accumulation of tangible and intangible capital, which often occurs in cities (Lucas, Jr., 1988). Studies of urban growth inform us of the factors raising urban productivity (Glaeser, Scheinkman, and Shleifer, 1995; da Mata et al., 2007) and quality of living (Glaeser and Shapiro, 2003; Shapiro, 2006). Moreover, recent contributions to the endogenous growth literature show that the productivity growth of a nation depends on the way its cities grow to explore external economies in employment and in learning (e.g. Black and Henderson, 1999; Rossi-Hansberg and Wright, 2007). These studies, however, largely overlook the role of housing supply in determining 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. A number of recent studies, including Mayer and Somerville (2000), Quigley and Raphael (2005), Glaeser, Gyourko and Saks (2006) and Saks (2008), find that population and employment respond less positively to local employment demand shocks in cities with more restrictive land-use regulations.

Although the importance of housing supply for urban growth is widely recognized now, the literature appears inconclusive regarding the relative importance of the determinants of housing supply elasticity. In the U.S. context, there are different views regarding the role of land scarcity versus that of regulations in influencing housing supply elasticity. Glaeser (2004) maintains that, apart from regulatory constraints on construction, land scarcity would have little directly 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 local land scarcity—the fraction of area undevelopable within a radius of city center—and finds it to elevate housing price appreciation resulting from metropolitan population growth shocks, both directly and 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 critical role for economic development (Spence, Annez and Buckley, 2008), little about the determinants of housing supply elasticity has been documented.
Our objectives are twofold. First, we propose a simple model of housing-supply-based accounting of urban population growth for the purpose of studying cross-city variations in housing supply elasticity. Three important features distinguish this model from those employed in previous studies of housing-supply elasticity. One feature is its more complete accounting of the elements of net housing supply that accommodates population growth by including, in addition to the supply response to demand shocks, housing supply expansion independent of demand shocks and changes in housing absorption by existing urban population. These two additional elements can be especially important for a developing economy. The second feature is its identification of inter-city demand shocks built on population and employment mobility (e.g. Roback, 1982; Gabriel and Rosenthal, 2004). The mobility makes the demand for individual cities in a large economy highly elastic, at least in the long run, and removes their 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. Under perfectly elastic inter-city demand, as in Roback (1982), demand shocks would show up in inter-city land rent differentials whereas local housing-supply elasticity would determine urban size, not vice versa. Previous studies of long-run housing supply elasticity often seeks to identify the demand shifts through quantity changes (e.g., Harter-Dreiman, 2004; Saiz, 2008) and the supply constraints through their impact on local housing prices (e.g., Glaeser and Gyourko, 2003; Glaeser, Gyourko and Saks, 2005a and 2005b; Harter-Dreiman, 2004; Saiz, 2008). 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. We, therefore, seek to identify the demand shocks in terms of productivity growth and changing demand for local amenities through their impact on differential land rent (housing price) growth across cities. These two features enable us to evaluate housing supply elasticity more reliably across cities. The third feature is our control for changes in urban built-up area so as to distinguish housing supply elasticity due to redevelopment density increase, referred to as housing-density supply elasticity.
Our second objective is to extend the literature on urban growth and housing supply (Quigley and Raphael, 2005; Glaeser, Gyourko and Saks, 2006; and Saks, 2008) by identifying important institutional factors of housing supply pertaining to the stage of development of a developing economy. In the context of China’s remarkable urban development over the past decade, we examine such institutional factors as local government efficiency, urban land market reform, and intra-urban income inequality, in addition to physical conditions such as building age, per-capita road space, and population density in built-up area, in influencing the housing-density supply elasticity. 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.

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 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).¹ 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.

¹ According to National Bureau of Statistics of China (NBSC), private sector provided 43.9 million jobs in Chinese cities between 2002 and 2006 (including 16.9 million in foreign-funded firms and 27 million in other forms of private firms). Meanwhile, employment in state-owned and collective-owned enterprises declined by 10.7 million.
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 inter-city demand shocks in China during the 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-density supply is assumed constant. We show that the variation in the housing-density supply elasticity contributes as much to cross-city variance in population growth as do the variations in urban land expansion and household housing demand growth. We further find that the institutional factors make more significant differences to urban redevelopment density in response to demand shocks than do physical conditions.

The remaining of the paper is organized as follows. Section I presents the model for housing-supply-based accounting of urban population growth and for measuring inter-city demand shocks. Section II describes the data and our measurements of urban growth, demand shocks, and the various explanatory variables. Section III reports the estimates of population growth accounting, with a focus on the determinants of housing-density supply elasticity. We conclude in Section IV.

I. 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

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2 According to the 2000 census, over 7.5% of the population moved within provinces between 1995 and 2000 and 2.7% moved across provinces. About 14% of the residents in Beijing, Shanghai, and Guangdong province in 2000, for example, were new arrivals after 1995. By comparison, about 3 percent of US population move across counties within a given state in any given year and an additional 3 percent move across states (Borjas, 1999, p10). 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.
cost rise with density. Let \( p \) be the housing price level. \( h \) is then competitively determined by the first-order condition:\(^3\)

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

(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-density supply.

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 \( Q \): \( A \cdot \left( Q^{\sigma} + C^{\sigma} \right)^{\sigma - 1} \), 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

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

(3)

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 \( \ln r \) around the base-period \( \ln r_0 \), we have\(^4\)

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

(4)

where

\[
\lambda_0 \equiv \frac{1 + \sigma r_0^{\sigma - 1}}{1 + r_0^{\sigma - 1}}
\]

is a positive constant representing the price elasticity of demand. We will assume identical growth in housing price and rent so that \( g_r = g_p \) and define

\[
S \equiv \lambda_0 g_r = \lambda_0 g_p
\]

as our measure of urban demand shocks. In the case where housing expenditure \( (rQ) \) is a quarter of household income \( (Y) \) and \( \sigma = 1/3 \), we have \( \lambda_0 = 1/2 \).

\(^3\) 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).

\(^4\) For more details, see Zheng, Fu and Liu (2009).
Let \( N \) denote the number of households in the city, where the housing market clears when

\[
N \cdot Q = h \cdot L. \tag{5}
\]

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:

\[
g_N = g_L + g_h - g_Q = g_L - g_Y + \left(1 + \frac{n}{\lambda_0}\right) S, \tag{6}
\]

Equation (6) shows that population growth in a city depends on four components: 1) growth in urban area, which expands the supply of developable land, 2) household income growth, which accounts for increased housing absorption by the existing population, 3) demand shocks to the city, summarized by the housing price growth (multiplied by the price elasticity of demand) \( S \), and 4) the housing-density supply elasticity \( \eta \). Note that the demand shock \( S \) serves both to reduce per-capita housing consumption \( Q \) (Equation (4)) and to motive additional supply of housing density \( h \) (Equation (2)).

A useful feature of Equation (6) is its allowance for population growth due to housing supply expansion independent of the demand shocks. Such supply expansion can be an important source of urban growth—Glaeser and Tobio (2008), for example, find the population growth in the US Sunbelt since 1980s to have resulted to a large extent from such housing supply expansion—but are often overlooked in extant cross-city studies of housing supply elasticity. In this study we will treat the growth of urban land use (built-up area) \( g_L \) as if independent of the demand shocks. This simplification is made for two reasons. First, to a large extent the conversion of rural land to urban uses is policy driven. All urban land in China is owned by the state and the urban governments monopolize the land-use conversion and the revenue from leasing out the converted land for private development. Incentivized by the fiscal decentralization reform in the 1990s (Jin and Zou, 2005) and the economic growth mandate, the urban governments used land-use conversion both to attract foreign direct investment and to raise additional revenue (Lichtenberg and Ding, 2009), often prompting tighter oversight by the central and provincial government to curb excessive urban land-use expansion. Second, the simplification allows us to focus on the contribution of housing-density supply elasticity to urban population growth.
The density of housing construction, both in the redevelopment of old built-up areas with predominantly low floor-to-area ratio FAR (or plot ratio) and in new developments, has a major influence on new housing supply in Chinese cities. Fu and Somerville (2001) show that the redevelopment FAR in Shanghai in the early 1990s (when it was the largest city in China) was influenced both by the land-use demand at the redevelopment site and by the road infrastructure availability in the vicinity. However, the high cost of re-housing displaced households, often born by the private developers, could delay and deter urban redevelopment in many central locations. The road infrastructure and the re-housing cost are some of the determinants of the housing-density elasticity $\eta$ that we will examine in this study. Our focus on the housing-density supply elasticity complements the extant studies of housing supply elasticity, which often focus on factors restricting the extensiveness of residential development. 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.

To measure the relative demand shocks across cities, we adopt the open-city assumption of Roback (1982). Although we do not believe the population mobility in China was unfettered in our study period, the internal migration rates documented by the 2000 Population Census (see footnote 2) show that the mobility is probably high enough for housing prices and urban wage rates to reveal the willingness to pay for urban amenities by the marginal migrants and the willingness to pay for urban productivity by the marginal firms. As shown in Gabriel and Rosenthal (2004), the residential rent premium equals the households’ willingness to pay for the local quality of living plus the urban wage-rate premium, which in turn equals firms’ willingness to pay for labor-augmenting urban productivity.

In our model, the indirect utility of a household in city $k$ and time period $t$ is given by:

$$v_t = A_t \left[ 1 + r_{st}^{-1} \right]^{1 \over \sigma-1} \frac{Y_{kt}}{r_{st}}. \quad (7)$$

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 $t=0$) as well as
city \( k \) (relative to the benchmark city \( k=0 \)), together with a first-order Tyler expansion of \( \ln r_{kt} \) around the base-period log rent in the benchmark city, \( \ln r_{0t} \), we obtain:\(^5\)

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

where \( S_k = \lambda_0 \ln \left( \frac{r_{kt}}{r_{0t}} \right) \) 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 (or the firms’ willingness to pay for urban productivity growth). Given that both the urban quality of living and productivity growth can be influenced by urban population growth, we will instrument \( S_k \) with the initial urban conditions that predict \( \Delta_t \ln A_{kt} \) and \( \Delta_t \ln Y_{kt} \).\(^6\)

II. 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, as well as several income-related city characteristics are derived. Additional city attributes are collected from Urban Statistics Yearbook to instrument \( S_k \) and 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 ***

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

\(^5\) Details of the derivation can be found in Zheng, Fu and Liu (2009).

\(^6\) 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.
genuine growth (Shen, 2005). There is no reliable way to adjust the population statistics for these reclassifications; 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 due to population reclassification. 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%.\(^7\)

*** Insert Figure 1 about here ***

**Measuring demand shocks \( S \)**

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. For a consistent measure of the demand shocks in a larger cross-section of cities and for a more representative housing stock in each city, we have to resort to indirect estimates. We adopt the estimates provided by Zheng, Fu and Liu (2009), who pool the 1998 and 2004 UHS data to estimate a household housing demand equation that accounts for household income and demographic attributes, identifying the

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\(^7\)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.
base-period (1998) city-specific land (housing) rent premium $R_k = \ln\left(\frac{r_{k0}}{r_0}\right)$ and rent growth $S_k = \lambda_0 \ln\left(\frac{r_{k0}}{r_{k0}}\right)$, respectively, as city fixed effects and these fixed effects multiplied by the year 2004 dummy. These $R_k$ and $S_k$ estimates, displayed in Figure 2, reflect the differential price incentives for housing consumption across cities as well as the change in the price incentive over the study period in individual cities (relative to the benchmark city Beijing).

*** 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 ***

**Computing predicted demand shocks $\hat{S}$**

To instrument the demand shocks, we use city initial conditions in 1998 to predict the demand shock measure $S_k$ obtained from Zheng, Fu and Liu (2009). These initial conditions are divided into two groups, one for variables reflecting the quality of living and the other for variables relating to urban productivity growth. The variables in the first group predict the growth in the willingness to pay for urban quality of living (i.e. $\Delta, \ln A_i$ in Equation (8)) and include: city temperature index $\text{TEMP}$, 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; $SO2$ emission and $GREEN$ space availability in the city, to reflect the environmental quality; the average years of schooling of the adult population in the city, $EDU$, to reflect the quality of social-interaction environment (Shapiro, 2006); and the number of doctors ($DOCT$) and hospital beds ($H\_BED$) per 100 people, to reflect the city’s healthcare quality.

Variables in the second group predict the urban wage (productivity) growth $g_W$ (i.e. $\Delta, \ln Y_i$ 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 (\( P_{\text{CAPITAL}} \)) and coastal geographical position (\( \text{COASTAL} \)). 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.

Table 2 reports the OLS estimates of the determinants of the demand shocks \( S \). 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. \( r_0Q = Y/4 \) and \( \sigma = 1/3 \)), we have:

\[
\Delta_i \ln A_i = S_k/2 - g_W + \Delta_i \ln v, \tag{9}
\]

where \( \Delta_i \ln v \) is a constant representing the growth in household real income (utility). Using \( S_k/2 - g_W \) 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, \( DOCT \times H\_BED \)) over time. Finally, holding constant these amenity conditions, we find some degree of mean reversion in land rent differential across the cities.

*** Insert Table 2 about here ***

Column 2 of Table 2 reports the reduced-form estimates of both the amenity demand shocks \( \Delta_i \ln A_i \) 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. 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, respectively, economic recovery and exceptionally high local demand for labor. 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 \( \hat{S}_{\delta} \) using the estimates reported in Column 2 of Table 2. \( \hat{S}_{\delta} \) explains about half of the variance of \( S \).

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

The growth in the supply of developable land \( g_L \) is produced by two variables: the growth in urban built-up area \( g_{BuiltUp} \) 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 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.\(^8\) 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 supply elasticity \( \eta \) fall into two groups, one relating to redevelopment costs and the other to land-use pattern, which affects the allocation of land

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\(^8\) 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.
resources in the city. Several variables are included in the first group. The initial road space per capita of non-agricultural population, $\text{ROADPC}$, 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,\(^9\) 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 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 90\(^{th}\) 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 90\(^{th}\) percentile level of

\(^9\) 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.
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 variables accounting for cross-city differences in land-use pattern include (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; and (iii) 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 land-lease revenue; 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.

### III. Accounting population growth across cities

Between 1998 and 2004, by our conservative measure of $g_N$, cities in our sample on average grew nearly 20% in terms of non-agricultural population, which is quite remarkable growth. 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, therefore, would have been absorbed by increased household housing consumption $Q$ as a result of income growth. Shown in Table 1, the growth in $Q$ due to household income growth ($g_Y$) during the period averages 1.581 (assuming a unitary income elasticity), much greater than the average reduction in $Q$ due to price growth ($S=\lambda_{0}g_{r}$) of
0.365, indicating a substantial growth in household housing consumption \( g_Q = g_Y - S = 1.216 \) according to Equation (4) as well as a substantial increase in household real income according to Equation (9).\(^{10}\) 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 \( ROADPC \), with the initial urban quality indicated by \( R \), and with the ratio \( FDI/EMP_{SOE} \), but is negatively correlated with the household income growth \( g_Y \). Both the measured and the predicted demand shocks, \( S \) and \( \hat{S} \), are weakly correlated with \( g_N \) and \( \hat{S} \) has little correlation with those housing-supply related variables. In particular, \( \hat{S} \) is very weakly correlated with \( g_{BuiltUp} \) and \( g_{Road} \), suggesting that much of the urban land use expansion during the period was probably independent of the demand shocks.

*** Insert Table 4 about here ***

The main results of this study, those related to the determinants of the housing-density supply elasticity, are reported in Table 4. The first column shows the OLS regression of population growth measure \( g_N \) on the growth in developable land \( g_L \equiv g_{BuiltUp} \times 0.44 + g_{Road} \times 0.56 \), the growth in household income \( g_Y \), and the predicted demand shock \( \hat{S} \), with the housing-density supply elasticity \( \eta \) assumed constant across cities. We assume a Cobb-Douglas technology for producing developable land by combining built-up area and road space at a constant return to scale with respective elasticity of 0.44 and 0.56 (chosen to achieve best OLS fit).\(^{11}\) The estimates have the expected signs suggested by Equation (6) but, with the assumption of a constant \( \eta \), the demand shocks contribute very little to the accounted cross-city variance in population growth.

\(^{10}\) 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.

\(^{11}\) We assume residential land use to expand at the same rate as the overall urban land use.
In column 2 of Table 4, the housing density supply is allowed to vary according to city attributes. Overall, the variation in the density elasticity accounts for additional 29% of the cross-city variance in population growth—the $R^2$ squared increases from 0.396 in column 1 to 0.685 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 $H_AGE$, 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 ($H_AGE > 13.5$). We further find provincial capital and provincial level cities ($P\_CAPITAL$) to have higher housing density elasticity, possibly due to more favorable compensation policies for displaced households.

In terms of land-use pattern, a significant presence of foreign direct investment relative to the size of the SOE employment share ($FDI/EMP\_SOE$) increases the housing density elasticity, possibly by accelerating the reform of urban land use rights and creating a more dynamic urban land market. We find cities with high income inequality among households in 1998 (and failing to mitigate it during the period), $INEQ$, to have 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. In addition, large income inequality in a province ($Y_{90/75\_Province} > 1.37$) reduces the housing density elasticity of its capital city.¹² Provincial capital cities often attract rich people in the province to buy home either for speculative investment or for consumption by their family members; hence province-wide household income inequality would contribute additional demand for luxury housing development in the provincial capital cities.

Finally 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 $G\_EFF$

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¹² Two key provincial level cities, Beijing and Shanghai, are assumed to be have luxury housing demand influenced by the national income inequality. The 5 provincial capital (provincial level) cities having $Y_{90/75\_Province} > 1.37$ are Beijing, Shanghai, and the capital cities of Guangdong, Xinjiang and Heilongjiang provinces.
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 negative exponential of $G\_EFF$ has a better explanatory power (the $R$ squared is improved from 0.773 to 0.788 when the negative exponential is used in place of the linear form), indicating that 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. The variation in housing-density supply elasticity accounts for as much cross-city population growth variance as do the variations in urban land expansion and household income growth (the $R$ squared obtained when $\eta$ is assumed constant, whether based on the full sample or the smaller sample of 68 cities, is only about half the $R$ squared value reported in column 3).

To evaluate the magnitude of the housing-density supply 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 initial 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 $g_r=20\%$ assuming $\lambda_0=0.5$) in this base-scenario city would generate 3.17 percent population growth, with urban land 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 urban land 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 developable land would add to total housing supply or 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). 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 $\eta/\lambda_0 = (\beta_3 - \beta_2)/\beta_1 = 1.63$, or $\eta = 0.814$ for $\lambda_0 = 0.5$.\footnote{In the accounting relationship, increased effects of $g_L$ ($\beta_1$) and $S$ ($\beta_3$) 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.}

If we further use the median value of the $\lambda_0 g_L/\tilde{s}$ ratio, which is 0.733 across the 85 cities, as an indication of the urban land expansion elasticity with respect to $g_r$, our base-scenario housing supply elasticity (from both urban land expansion 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 supply elasticity due to redevelopment, $\eta$, would be somewhat greater than the median elasticity of housing supply due to land supply expansion.

\section*{IV. Conclusions}

We have sought to contribute to the literature on urban growth and housing supply in two respects. First, we provide a model of housing-supply-based accounting of urban population growth that treats cities as price takers in an economy of mobile workers and that takes into account housing supply expansions independent of demand shocks and changes in housing absorption by existing households, so as to improve the identification of urban housing supply response to demand shocks. Second, applying this model in the context of a developing economy, \textit{i.e.} China during the 1998~2004 period, when liberalization in urban housing and labor markets led to a surge in labor mobility and urban growth, we offer an empirical assessment of the importance of urban land expansion versus housing-density supply elasticity to urban growth. We find the cross-city variation in the density response to demand shocks to contribute as much to the cross-city population growth variance as do the variations in urban land expansion and per-capita housing demand growth. Furthermore, we find a number of institutional factors, such as local government efficiency, land-use property rights reform and intra-city income inequality, to have significant influences on the housing-

\footnote{$\beta_1$ 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, manufacturing land use often drive urban expansion 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 $\eta = 1.02$.}
density supply elasticity, in addition to physical conditions, such as average building age of homes, per-capita road space and urban population density, whose influences appear statistically relatively weak.

Although our empirical results need be viewed with some caution given the approximate measures of urban population growth and demand shocks that we have to rely on for our analysis, a few preliminary policy implications can be drawn. First, the potential for increased housing density in urban built-up area to meet the urban population growth demand is significant, even in the high-population-density cities like those in China. Second, government policies can enable the private housing market to produce more homes in urban areas by lowering the tax and regulatory burdens on private enterprises, enhancing the marketability of land use rights, channeling more land resources from luxury housing development to higher-density housing projects, and investing more in urban transport infrastructure.

15 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 r_0^{k-1}}{1 + \sigma r_0^{k-1}} \ln \frac{r_0}{r_k} ). 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 (( \ln \frac{Y}{Y_0} )), 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 (°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>( SO_2 )</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 \left( \frac{Y_k}{Y_0} \right) ), 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 \left( \frac{Y}{Y_0} \right) ), 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_{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_{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>
<tr>
<td>Variable</td>
<td>Description</td>
<td>1998</td>
<td>2004</td>
<td>2010</td>
<td>2020</td>
<td>2023</td>
<td>2024</td>
</tr>
<tr>
<td>---------------</td>
<td>-----------------------------------------------------------------------------</td>
<td>------</td>
<td>------</td>
<td>------</td>
<td>------</td>
<td>------</td>
<td>------</td>
</tr>
<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>24.00</td>
<td>10.40</td>
<td>2.673</td>
<td>85</td>
</tr>
<tr>
<td><strong>DENSITY</strong></td>
<td>City built-up area over non-agricultural population, 10,000 people /sqkm, 1998. [3]</td>
<td>1.131</td>
<td>1.107</td>
<td>2.169</td>
<td>0.520</td>
<td>0.326</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>0.076</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>0.162</td>
<td>68</td>
</tr>
<tr>
<td><strong>G_EFF</strong></td>
<td>City government efficiency score = 0.5 × (G_EFF_TFP / stdev(G_EFF_TFP) + G_EFF_FDI / stdev(G_EFF_FDI))</td>
<td>2.100</td>
<td>2.109</td>
<td>3.745</td>
<td>-0.134</td>
<td>0.894</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>0.026</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>0.059</td>
<td>85</td>
</tr>
<tr>
<td><strong>Y_{75/25}</strong></td>
<td>City 75^{th} percentile household income over 25^{th} percentile household income, 1998. [1]</td>
<td>1.715</td>
<td>1.679</td>
<td>2.380</td>
<td>1.400</td>
<td>0.182</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>0.350</td>
<td>85</td>
</tr>
<tr>
<td><strong>Y_{90/75_P}</strong></td>
<td>Provincial 90^{th} percentile household income over 75^{th} 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>0.062</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 $S, gW, R,$ 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 $S/2-g_w$</th>
<th>Column 2 $S$</th>
</tr>
</thead>
<tbody>
<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>0.114</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>$\hat{S}$</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>$g_N$</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>$g_{\text{BuiltUp}}$</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>$g_{\text{Road}}$</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>$\ln(\text{ROADPC})$</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>$H_{\text{AGE}}$</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>$\text{FDI}/\text{EMP}_{\text{SOE}}$</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>$\text{INEQ}$</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>$P_{\text{CAPITAL}}$</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>$\ln(\text{DENSITY})$</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 determinants

<table>
<thead>
<tr>
<th>Column</th>
<th>1</th>
<th>2</th>
<th>3</th>
</tr>
</thead>
<tbody>
<tr>
<td>Independent \ Dependent variables</td>
<td>$g_N$</td>
<td>$g_N$</td>
<td>$g_N$</td>
</tr>
<tr>
<td>Constant</td>
<td>0.320 (4.2)**</td>
<td>0.225 (3.3)**</td>
<td>0.226 (3.1)**</td>
</tr>
<tr>
<td>$g_L \equiv g_{Built} \times 0.44 + g_{Road} \times 0.56$</td>
<td>0.162 (7.1)**</td>
<td>0.177 (7.7)**</td>
<td>0.136 (7.4)**</td>
</tr>
<tr>
<td>$g_Y$</td>
<td>-0.166 (3.8)**</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($ROADPC / 6.9$) $\times \hat{S}$</td>
<td>0.109 (2.3)**</td>
<td>0.115 (2.3)**</td>
<td>0.115 (2.3)**</td>
</tr>
<tr>
<td>$(H_{AGE} &gt; 18) \times \hat{S}$</td>
<td>-0.090 (1.9)*</td>
<td>-0.126 (2.5)**</td>
<td>-0.126 (2.5)**</td>
</tr>
<tr>
<td>$(H_{AGE} &gt; 13.5) \times (R - $ln$(0.5)) \times \hat{S}$</td>
<td>-0.083 (1.6)</td>
<td>-0.181 (2.9)**</td>
<td>-0.181 (2.9)**</td>
</tr>
<tr>
<td>ln($DENSITY / 1.1$) $\times \hat{S}$</td>
<td>-0.104 (1.4)</td>
<td>-0.132 (2.0)**</td>
<td>-0.132 (2.0)**</td>
</tr>
<tr>
<td>$P_{CAPITAL} \times \hat{S}$</td>
<td>0.146 (3.2)**</td>
<td>0.218 (4.7)**</td>
<td>0.218 (4.7)**</td>
</tr>
<tr>
<td>exp($-0.71 \times G_{EFF}$) $\times \hat{S}$</td>
<td>0.418 (4.7)**</td>
<td>0.418 (4.7)**</td>
<td>0.418 (4.7)**</td>
</tr>
<tr>
<td>$(FDI / EMP_{SOE} - 0.02) \times \hat{S}$</td>
<td>1.140 (3.2)**</td>
<td>1.010 (4.8)**</td>
<td>1.010 (4.8)**</td>
</tr>
<tr>
<td>$INEQ \times \hat{S}$</td>
<td>-0.171 (2.7)**</td>
<td>-0.228 (4.4)**</td>
<td>-0.228 (4.4)**</td>
</tr>
<tr>
<td>$P_{CAPITAL} \times (Y_{90/75 \text{ Province}} &gt; 1.37) \times \hat{S}$</td>
<td>-0.154 (3.2)**</td>
<td>-0.112 (2.0)**</td>
<td>-0.112 (2.0)**</td>
</tr>
<tr>
<td>R-squared</td>
<td>0.396</td>
<td>0.685</td>
<td>0.788</td>
</tr>
<tr>
<td>Number of observations</td>
<td>85</td>
<td>85</td>
<td>68</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.
Figure 1. 1998 population size and 1998-2004 population growth across 85 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