Structural Estimates of Housing Supply Elasticity across Chinese Cities

Yuming Fu,[a] Siqi Zheng,[b] Hongyu Liu [b]

[a] Corresponding author. Department of Real Estate, National University of Singapore, 4 Architecture Drive, Singapore 117655. Email: yuming.fu@nus.edu.sg, Tel: (65) 6516 4412, Fax: (65) 6774 8684.

[b] Institute of Real Estate Studies, Tsinghua University, Beijing 100084, China.
Email for Siqi Zheng: zhengsiqi@tsinghua.edu.cn, email for Hongyu Liu: liuhy@tsinghua.edu.cn.

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Abstract

We present a structural model of urban growth in a spatial equilibrium setting to aid the separation of the effects of demand shocks from those of the spatial variation in housing supply elasticity. The model is applied to an analysis of urban growth across Chines cities between 1998 and 2004, to evaluate the determinants of housing supply elasticity. The variation in supply, via urban expansion or price elasticity, is found to play a greater role in accounting for cross-city differences in population growth, but a lesser role for the differences in wage-rate and price growth, than do the demand shocks. The housing supply elasticity is lower in denser and more expensive cities but greater in cities of higher political status or with denser roads. Furthermore, our results show that local governments can raise the housing supply elasticity by lowering regulatory costs and improving land-use equity.

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

JEL classification: R11, R31, R58

1. Introduction

Cities grow in size as a result of demand shocks or eased housing supply constraints. Demand shocks arise from urban productivity growth (Black and Henderson, 2003; da Mata et al., 2007; Glaeser et al., 1995) or increased attractiveness of urban amenities (Glaeser and Shapiro, 2003; Shapiro, 2006). But the population response to these shocks varies considerably from city to city, manifesting varied housing supply elasticity (Gyourko et al., 2010). Explaining the spatial variation in housing supply elasticity has received much attention in recent economics literature (Glaeser and Gyourko, 2003; Glaeser et al., 2005a and 2005b; Glaeser et al., 2006; Gyourko, 2009), with a growing appreciation of the importance of local housing supply elasticity for spatial equilibrium and, consequently, for the impact of national economic policies (Gleaser and Gottlieb, 2009). Evidence that geography and land-use regulations affect housing supply elasticity is now widely documented for US cities (Green et al., 2005; Mayer and Somerville, 2000; Quigley and Raphael, 2005; Saiz, 2010; and Saks, 2008). But these studies generally adopt a "reduced-form" approach to identifying the supply elasticity, measuring it in terms of population change in response to price changes, or vice versa, as a result of exogenous demand shocks. Such an approach is not always reliable in view of spatial equilibrium conditions, which determine the population and housing price in each city simultaneously according to both the demand shocks and the supply elasticity. Separation of the effects of demand shocks from those of supply elasticity on urban growth performance requires a structural model.

The objectives of the present paper are twofold. First, we incorporate the spatial equilibrium conditions in an empirical urban growth model to aid the separation of the effects of demand shocks from those of housing supply elasticity. We adopt the spatial equilibrium model presented in Glaeser and Gottlieb (2009), which extents the classic cross-city equilibrium model of Rosen (1979) and Roback (1982) by including the production of a nontraded good (housing services) in addition to that of a traded good. It further incorporates nontraded capital as a factor input that shifts the supply and regulates the supply elasticity of each good. We allow the supply elasticity of housing services in the model to vary across cities, to more fully account for the cross-city variations in population growth, housing price growth and wage rate growth. The resulting structural model decomposes these urban growth measures into two parts—one part captures the effects of the demand shocks and supply shift at a baseline housing supply elasticity, and the other, the effects of the cross-city variation in housing supply elasticity.[1] It does so by delineating how the responses of each of the urban growth measures to demand shocks are affected by the supply elasticity.

Second, we apply the structural model to examining the differential housing supply elasticity across Chinese cities. Despite our growing knowledge about the determinants of housing supply elasticity across US metropolitan areas, few rigorous empirical studies of urban housing supply in developing economies are available (Arnott, 2009). In particular, little evidence has been documented regarding the influence of urban institutions on housing supply elasticity in developing countries, although overregulation of the formal housing sector in these countries is widely believed to inhibit the supply elasticity (De Soto, 2000). We seek to narrow this empirical knowledge gap about urban housing supply elasticity in developing-economy contexts, where urban density and land-use institutions differ considerably from developed economies. Against a backdrop of liberalizing land and labor markets and rapidly rising income in Chinese cities in the wake of China's integration into the global economy, our empirical analysis reveals important influences on housing supply elasticity by such institutional factors as urban income inequality, government efficiency, land cost, and the city's political status, in addition to the influences by physical conditions such as urban population density and road density. The shift from state-provided to market-based housing supply in Chinese cities in the 1990s was a notable example of market-oriented housing policy reform among the developing economies (Buckley and Kalarickal, 2005). Cities with elastic housing supply are instrumental to economic growth in developing economies (see Spence et al., 2009). The present study provides empirical evidence on the workings of the urban housing market in the post-reform Chinese cities, and our findings point to ways further institutional reforms could enable the market to be more responsive to urban demand shocks.

[1] Glaeser and Tobio (2008) present a structural accounting of changes in urban population, prices and wage rates across US regions. Their focus is on identifying the importance of the demand shocks as opposed to the supply shift in explaining the population growth of the Sunbelt region relative to other regions, assuming a constant supply elasticity.

Our sample includes 85 cities across China. We examine their population growth between 1998 and 2004, a period marked by substantial rise in labor mobility and urbanization. The urban share of population increased from 33.4 percent in 1998 to 41.8 percent by 2004. 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 the full liberalization of the private housing market in Chinese cities (Fu et al., 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 (more than one square meter per urban resident). In addition to increased marketability of existing urban homes, the fraction of new urban homes 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).[2] The liberalization of the urban housing market and labor market significantly lowered internal migration 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 et al., 2009).

We present the structural model in Section 2, incorporating the spatial equilibrium conditions to separate the effects of demand shocks from those of supply elasticity on urban growth measures. Section 3 describes the data and the measurements of urban performance (*i.e.* population, housing price, and urban wage rate), urban demand shocks, and housing supply shift. Section 4 reports the estimates of the structural coefficients and discusses the determinants of housing supply elasticity across Chinese cities. We conclude in Section 5.

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[2] According to National Bureau of Statistics of China (NBSC), private sector provided 43.9 million new 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.

2. The structural model

Following Glaeser and Gottlieb (2009), we describe the economy of a city, indexed by subscript j, in terms of perfectly mobile and homogeneous workers (household consumers) and two production technologies, one for a traded good Q_T and the other for a non-traded good Q_N. The non-traded good will be called housing. Equilibrium is defined by three endogenous urban performance variables: (1) a city-specific wage rate W_j, equal to the labor productivity in the city; (2) a city-specific housing price P_j, which compensates the urban productivity and amenity differentials to eliminate inter-city migration incentives; and (3) a city population size N_j, which clears both the labor and housing markets in the city.

As in Glaeser and Gottlieb (2009), the production technologies are of constant-return-to-scale with respect to three factors of production, namely (non-traded) local public capital Z (e.g. public infrastructure, utilities and institutions), (traded) private capital K, and labor L. Specifically:

$$Q_T = A_j Z_A^{\gamma} K_A^{1-\gamma} \left(\frac{L_A}{K_A} \right)^{1-\alpha}, \tag{1}$$

$$Q_N = H \cdot Z_H^{\eta_j} K_H^{1-\eta_j} \left(\frac{L_H}{K_H} \right)^{1-\mu}, \tag{2}$$

where subscript A and H indicate factor inputs for traded good Q_T and non-traded good Q_N respectively.[3] We assume that the technology for Q_T, in terms of public capital share γ and labor share $1-\alpha$, is identical across cities but the total factor productivity A_j is city-specific due to location heterogeneities. We further assume $A_j=a_j(N_j)^{\omega}$, where $\omega>0$ represents the strength of urbanization economy. For housing production, we assume the total factor productivity H to be identical across cities due to technology diffusion. We also assume a spatially invariant labor share $1-\mu$ in the non-traded sector. But the factor share for the local public capital η_j is assumed to be city-specific to allow heterogeneous housing supply elasticity across cities; the housing supply elasticity varies inversely with η_j. We assume $\gamma<\alpha<1$ and $\eta_j<\mu<1$ to ensure positive factor shares for the private capital K in both the traded and non-traded sectors. K is traded and, like Q_T, is the numeraire good.

[3] All the factor input and the output quantities are city specific in equilibrium; they are not indexed by j for notational simplicity.

The technologies described by Eq. (1) and Eq. (2) are compatible with competitive and constant-return-to-scale firms who take the local public capital Z as well as all prices as given in their hiring and output choices. Labor market clearing in the city entails $L_A+L_H=N_j$. For a given share of housing spending in household budget $1-\beta$, L_A and L_H are a fraction, $1-(1-\mu)(1-\beta)$ and $(1-\mu)(1-\beta)$ respectively, of the city population. The labor market clearing in city j, according to the marginal product of labor of the competitive firms, is given by:

$$W_j^{1-\alpha+\gamma} = \varphi_1 a_j Z_A^{\gamma} N_j^{\omega-\gamma}, \tag{3}$$

where φ_1 is a constant. Eq. (3) indicates that the labor productivity, hence the wage rate, increases with city-specific productivity shifter a_j and the public capital Z_A. Since the coefficient of urbanization economy ω is typically small in comparison with the factor share of the local public capital γ (some indicative evidence is offered in Glaeser and Gottlieb (2009)), the urban productivity is generally decreasing in urban population size holding the public capital Z_A constant.

We depart from Glaeser and Gottlieb (2009) by assuming a CES instead of a Cobb-Douglas utility function for households, so that the price elasticity of housing demand can be less than unity, to allow additional flexibility in the structural coefficients for given values of housing supply elasticity. Each worker (household) receives a wage rate W_j and spends it to obtain an identical utility level u from the consumption of the traded good G_T, the non-traded housing service G_N, and a local amenity public good indexed by θ_j:[4]

$$u = \theta_j \left(\phi \cdot G_T^{\frac{\sigma-1}{\sigma}} + (1-\phi) G_N^{\frac{\sigma-1}{\sigma}} \right)^{\frac{\sigma}{\sigma-1}} = \theta_j \left((1-\phi)^{\sigma} + \phi^{\sigma} P_j^{\sigma-1} \right)^{\frac{1}{\sigma-1}} \frac{W_j}{P_j}, \tag{4}$$

where $\sigma \geq 0$ is the elasticity of substitution and $0 \leq \phi \leq 1$ is a preference parameter. The utility u is spatially invariant as required by cross-city labor market clearing in equilibrium.

The housing demand function, according to Roy's identity, is

$$G_N = \frac{(1-\phi)^{\sigma}}{(1-\phi)^{\sigma} + \phi^{\sigma} P_j^{\sigma-1}} \frac{W_j}{P_j} \approx (1-\beta) P_0^{(\sigma-1)\beta} \frac{W_j}{P_j^{1-(1-\sigma)\beta}}, \tag{5}$$

[4] Implicitly factor incomes from capital Z and K become the saving in the economy.

where the approximation follows the linearization of $\ln\frac{(1-\phi)^{\sigma}}{(1-\phi)^{\sigma}+\phi^{\sigma}P_j^{\sigma-1}}$ with respect to P_j around a baseline price P_0 and $1-\beta\equiv\frac{(1-\phi)^{\sigma}}{(1-\phi)^{\sigma}+\phi^{\sigma}P_0^{\sigma-1}}$ is the housing expenditure share of income at the baseline price.[5] The magnitude of the price elasticity of housing demand is $\varepsilon_P\equiv 1-(1-\sigma)\beta$ which is less than unity for $1>\sigma>0$. The city-specific housing price that clears the inter-city labor market can be expressed by the following equation, obtained by applying logarithm to the indirect utility function in Eq. (4) and the linear approximation in footnote 5:

$$(1-\beta)\ln P_j=\ln W_j+\ln\theta_j+\varphi_2\,, \tag{6}$$

where φ_2 is a constant. Eq. (6) is the familiar Rosen-Roback expression of a compensating housing price that increases with city productivity and amenity.

Finally, the housing market clearing in the city entails equality between total expenditure on housing services $G_N N_j$ and total cost of housing service production $P_j Q_N$:

$$(1-\beta)\left(\frac{P_j}{P_0}\right)^{(1-\sigma)\beta}W_jN_j=P_jQ_N=\varphi_3H^{1/\eta_j}P_j^{1/\eta_j}W_j^{(\mu-1)/\eta_j}Z_H\,, \tag{7}$$

where φ_3 is a constant. Note that the housing supply Q_N is proportional to Z_H, increases with the housing-sector productivity H and price P_j but decreases with the wage rate in the city. The coefficient of price elasticity of housing supply is $\varepsilon_H\equiv 1/\eta_j-1$, which decreases in the factor share of the housing sector local public capital η_j: the capacity for the private sector to respond to housing demand shocks is more limited when land and regulatory costs (factor income due to local public capital) accounts for a larger share of the housing cost.

Eq. (3), Eq. (6) and Eq. (7) define the market-clearing wage rate W_j, housing price P_j and population N_j under spatial equilibrium conditions. As in Glaeser and Gottlieb (2009), we express these endogenous

[5] $\ln\frac{(1-\phi)^{\sigma}}{(1-\phi)^{\sigma}+\phi^{\sigma}P_j^{\sigma-1}}\approx\ln(1-\beta)+(1-\sigma)\beta\frac{P_j-P_0}{P_0}\approx\ln(1-\beta)+(1-\sigma)\beta\left(\ln P_j-\ln P_0\right)$.

urban performance measures in terms of their elasticity with respect to the three sources of exogenous shocks, namely, the housing supply shifter $HZ_H^{\eta_j}$, productivity shifter $a_j Z_A^{\gamma}$, and amenity shifter θ_j:

$$\begin{bmatrix} \ln N_j \\ \ln W_j \\ \ln P_j \end{bmatrix} = \begin{bmatrix} \kappa_N \\ \kappa_W \\ \kappa_P \end{bmatrix} + \boldsymbol{\Lambda} \begin{bmatrix} \ln\left(HZ_H^{\eta_j}\right) \\ \ln\left(a_j Z_A^{\gamma}\right) \\ \ln\theta_j \end{bmatrix}, \quad \boldsymbol{\Lambda} \equiv \begin{bmatrix} \lambda_{NH} & \lambda_{NA} & \lambda_{N\theta} \\ \lambda_{WH} & \lambda_{WA} & \lambda_{W\theta} \\ \lambda_{PH} & \lambda_{PA} & \lambda_{P\theta} \end{bmatrix}. \tag{8}$$

where κ_N, κ_W, and κ_P are constants and the coefficients of $\boldsymbol{\Lambda}$ (the structural coefficients) are derived in the appendix. Note that all these structural coefficients depend on the housing supply elasticity ε_H.

Applying intertemporal log differentiation to Eq. (3), Eq. (6) and Eq. (7) and using the notation $\dot{X} \equiv \ln\left(X_{t+1}\right) - \ln\left(X_t\right)$, we have the following dynamic market-clearing equations when the city grows:

$$b\dot{N}_j + \dot{W}_j = \left(\dot{a}_j + \gamma \dot{Z}_A\right) / \left(1 - \alpha + \gamma\right), \tag{9}$$

$$\left(1-\beta\right)\dot{P}_j - \dot{W}_j = \dot{\theta}_j - \dot{u}, \tag{10}$$

$$\dot{N}_j + \dot{W}_j + \left(1 - \varepsilon_P\right)\dot{P}_j = \left(\varepsilon_H + 1\right)\left(\dot{P}_j - \left(1-\mu\right)\dot{W}_j\right) + \left(\varepsilon_H + 1\right)\left(\dot{H} + \eta_j \dot{Z}_H\right), \tag{11}$$

where $b \equiv \left(\gamma - \omega\right)/\left(1 - \alpha + \gamma\right)$. Eq. (9) and Eq. (10) show how the productivity shock $\dot{a}_j + \gamma \dot{Z}_A$ and the amenity shock (adjusted by a uniform utility growth) $\dot{\theta}_j - \dot{u}$ can be measured by the observed urban growth measures, and thus be estimated, independently of the supply elasticity, although the individual growth measures are all dependent on the supply elasticity as Eq. (8) indicates.

Eq. (11) provides an accounting of housing demand growth (left-hand side) and housing supply adjustments (right-hand side) in a city. It highlights some of the limitations of the reduced form approach to identifying housing supply elasticity. Note, on the one hand, that the demand growth is accounted for not only by population growth but also by wage-rate growth and price growth (when the demand is not unitary elastic); on the other hand, the supply adjustment respond not only to the price growth but also to the wage-rate growth, which raises the cost. Moreover, the response of all the endogenous variables in Eq. (11) to the demand shocks depends on the supply elasticity; separating the effects of the demand shocks from those of the supply elasticity is not so straightforward. Applying demand shocks to Eq. (11), we obtain (see also Eq. (A.8) and Eq. (A.9) in the appendix):

$$\varepsilon_H = \frac{\lambda_{NA} + \lambda_{WA} + (1-\varepsilon_P)\lambda_{PA}}{\lambda_{PA} - (1-\mu)\lambda_{WA}} - 1 = \frac{\lambda_{NA}/\lambda_{PA} + (2-\mu)(1-\beta) - \varepsilon_P}{1-(1-\mu)(1-\beta)}, \tag{12}$$

$$\varepsilon_H = \frac{\lambda_{N\theta} + \lambda_{W\theta} + (1-\varepsilon_P)\lambda_{P\theta}}{\lambda_{P\theta} - (1-\mu)\lambda_{W\theta}} - 1 = \frac{1-(2-\mu)b - \varepsilon_P(\lambda_{P\theta}/\lambda_{N\theta})}{\lambda_{P\theta}/\lambda_{N\theta} + (1-\mu)b}, \tag{13}$$

where $\lambda_{NA}/\lambda_{PA}$ and $\lambda_{N\theta}/\lambda_{P\theta}$ are reduced-form estimators of housing supply elasticity, representing population response to price changes predicted by productivity and amenity shocks respectively. Solving for these reduced-form estimators in terms of ε_H, the housing supply elasticity in the structural model, we have

$$\lambda_{NA}/\lambda_{PA} = \left(1-(1-\mu)(1-\beta)\right)\varepsilon_H + \varepsilon_P - (2-\mu)(1-\beta), \tag{14}$$

$$\lambda_{N\theta}/\lambda_{P\theta} = \frac{\varepsilon_H + \varepsilon_P}{1-b-b(1-\mu)(1+\varepsilon_H)}. \tag{15}$$

Note that the two reduced-form estimators can have very different values and give biased measures of ε_H. In particular, $\lambda_{NA}/\lambda_{PA}$ increases with ε_H at a rate less than 1 and underestimates ε_H when ε_P is relatively small; in contrast $\lambda_{N\theta}/\lambda_{P\theta}$ increases faster than ε_H and always overestimates it since b is generally positive. Both $\lambda_{NA}/\lambda_{PA}$ and $\lambda_{N\theta}/\lambda_{P\theta}$ increase with the price elasticity of demand ε_P. Accurate identification of the supply elasticity, therefore, requires a clear distinction of the sources of the demand shocks.

To isolate the influence of the heterogeneous housing supply elasticity on the urban growth measures from those of the demand shocks and supply shifts, we linearize the structural coefficients in $\mathbf{\Lambda}$ in Eq. (8) around a baseline value $\bar{\varepsilon}_H$ (corresponding to $\bar{\eta}$). Applying intertemporal log differentiation to Eq. (8), using the first-order derivatives of the structural coefficients with respect to ε_H (given in Table A of the appendix), and taking into account the dynamic market-clearing equations, we obtain:

$$\begin{bmatrix} \dot{N}_j \\ \dot{W}_j \\ \dot{P}_j \end{bmatrix} \approx \bar{\mathbf{\Lambda}} \begin{bmatrix} \dot{H} + \bar{\eta}\dot{Z}_H \\ \dot{a}_j + \gamma \dot{Z}_A \\ \dot{\theta}_j - \dot{u} \end{bmatrix} + \begin{bmatrix} \bar{\lambda}_{NH} \\ \bar{\lambda}_{WH} \\ \bar{\lambda}_{PH} \end{bmatrix} S_j \bar{\eta}^2 \Delta\varepsilon_H, \tag{16}$$

where $\bar{\mathbf{\Lambda}}$ is the matrix $\mathbf{\Lambda}$ evaluated at $\bar{\eta}$, $\Delta\varepsilon_H \equiv \varepsilon_H - \bar{\varepsilon}_H$, and

$$\begin{aligned} S_j \equiv & \left(\bar{\lambda}_{NH} + \bar{\lambda}_{WH} + \bar{\lambda}_{PH}\left(1-\varepsilon_P\right)\right)\left(\dot{H} + \bar{\eta}\dot{Z}_H\right) \\ & + \left(\bar{\lambda}_{NA} + \bar{\lambda}_{WA} + \bar{\lambda}_{PA}\left(1-\varepsilon_P\right)\right)\left(\dot{a}_j + \gamma \dot{Z}_A\right) \\ & + \left(\bar{\lambda}_{N\theta} + \bar{\lambda}_{W\theta} + \bar{\lambda}_{P\theta}\left(1-\varepsilon_P\right)\right)\left(\dot{\theta}_H - \dot{u}\right) - \dot{Z}_H . \end{aligned}$$

Note that S_j is the growth in total housing expenditure net of the growth in housing-sector nontraded capital, $\dot{N}_j + \dot{W}_j + \left(1-\varepsilon_P\right)\dot{P}_j - \dot{Z}_H$, as a result of the demand and supply shocks evaluated at the baseline supply elasticity $\bar{\varepsilon}_H$. According to Eq. (11), $\bar{\eta}S_j$ equals the supply incentive response to the demand and supply shocks, $\dot{P}_j - \left(1-\mu\right)\dot{W} + \dot{H}$, evaluated at $\bar{\varepsilon}_H$. Note further that S_j always increases with the demand shocks (its coefficients with respect to $\dot{a}_j + \gamma \dot{Z}_A$ and $\dot{\theta}_j - \dot{u}$ are always positive); hence a greater supply elasticity ($\Delta\varepsilon_H > 0$) always make population growth more responsive (because $\bar{\lambda}_{NH} > 0$), but the wage-rate growth and price growth less responsive (because $\bar{\lambda}_{WH} < 0$ and $\bar{\lambda}_{PH} < 0$), to the demand shocks. Eq. (16) provides the framework for separately evaluating the effects of the demand and supply shocks on the urban growth measures with the supply elasticity held constant (first part on the right-hand side) and the effects of the spatial variation in the supply elasticity (the second part). Controlling for the variation in housing supply elasticity in the second part is necessary for the evaluation of urban growth responses at a constant supply elasticity in the first part, which in turn enables the computation of S_j necessary for the evaluation of the effects of supply elasticity variation. Eq. (16) also shows that the influence of the housing supply elasticity differences on the urban growth measures is equivalent to augmenting the supply shift by $S_j \bar{\eta}^2 \Delta\varepsilon_H$.

Land-use regulations raise the land share of housing cost (η_j) and hence reduce housing supply elasticity: $\Delta\varepsilon_H < 0$. Equation (16) predicts that both wage rate and price are elevated as a result (since both $\bar{\lambda}_{WH}$ and $\bar{\lambda}_{PH}$ are negative), consistent with the findings in many extant studies of the price impact of land-use regulations (*e.g.*, Glaeser and Gyourko, 2003; Glaeser, Gyourko and Saks, 2005a, 2005b).[6] However, the structural model helps to clarify the causal channel of the regulatory effects on prices and wage rates. Equation (16) shows that these effects, which have the sign of $-(\gamma-\omega)$, are derived from the urban scale effect on productivity under fixed local public capital Z_A. In other words, it is a smaller population size as a result of the land-use regulations that raises urban productivity W_j and price P_j, rather than the higher price as a result of the land-use regulation that deters urban population.

3. Chinese urban growth: data and measurements

To apply the structural model to estimating housing supply elasticity across Chinese cities, several assumptions underlying the spatial equilibrium and the housing market need to be validated in the context of China's economic development. The spatial equilibrium requires free labor mobility across cities. Although barriers to labor mobility still existed, especially for low-skill migrant workers (see Fu and Gabriel, 2012), labor mobility had been substantially elevated during our sample period, from 1998 to 2004, as a result of the liberalization of urban labor market in the 1990s and urban housing market in 1998, as mentioned in the introduction.[7] Fu and Gabriel (2012) show that province-to-province labor migration during the early 1990s was highly responsive to inter-provincial economic incentives and Zheng et al. (2009) find the adjustments in housing prices across Chinese cities between 1998 and 2004 to be strongly influenced by the differential productivity and urban quality of life across the cities.

The spatial equilibrium model described in the last section assumes competitive supply of housing services. Since mid 1990s, urban homes previously owned by the state and work units have been

[6] The empirical evidence reported in extant studies, however, has been mixed, as noted in Harter-Dreiman (2004) and Quigley and Rosenthal (2005), possibly reflecting unreliable separation of the effect of demand shocks from that of supply elasticity (Ihlanfeldt, 2007).

[7] 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.

privatized (Fu et al., 2000). Urban homes are now generally privately owned and can be freely rented or sold at market prices.[8] Housing development is competitive. There were over 60 thousand residential developers in 2008, the majority of which are privately owned. The largest 10 developers are either publicly listed or privately owned and their market share was about 4.2% in 2004 (China Real Estate Top 10 Research Group, March 2009).

With the above qualified validation of the spatial equilibrium assumptions, we employ the structural model to examining the urban growth and housing supply between 1998 and 2004 in a sample of 85 cities across Chinese provinces. As documented in Zheng, Fu and Liu (2009), these cities are among those sampled by the Urban Household Survey (UHS), which is conducted by the National Bureau of Statistic of China (NBSC).[9] The household-level data from UHS are employed to estimate housing (service) prices P_j and urban wage rates W_j and their growth between 1998 and 2004. Urban population and other attributes that predict the demand shocks related to urban productivity and amenity, the supply shifts due to urban expansion, and housing supply elasticity are collected from Urban Statistical Yearbooks. These variables, which are listed in Table 1 together with their sample statistics and data sources, are explained below.

*** Insert Table 1 about here ***

3.1 Measuring urban endogenous variables

Urban population N_j and population growth $\dot{N}_j$

Chinese cities are delineated by administrative boundaries and typically include substantial rural areas and agricultural population. Our urban population data are based on the city non-agricultural population statistics from Urban Statistics Yearbook. The historical consistency of these urban population statistics

[8] Urban homeownership rate increased from below 20% in the early 1990s to over 50% by 1998. By 2004 the homeownership rate exceeded 80%. See Chamon and Prasad (2010).

[9] UHS 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, districts and neighborhoods are sorted and sampled at fixed distance, followed by the sampling of households within the selected neighborhoods. Cities missing necessary data Urban Statistics Yearbook are dropped.

is compromised, however, by administrative boundary changes occurred in a number of 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 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 associated with changes in city administrative boundaries. Third, we replace the abrupt changes with average population growth rate in the adjacent years surrounding each abrupt change. 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, $\dot{N}_j$, are plotted in Figure 1. The average population size of the 85 cities in 1998 is about 1.2 million; half of the cities have a population above 700,000. On average these cities grew by 18.5% during the six-year period (see Table 1), which probably understates the actual urban population growth due to failure to account for urban workers with rural Hukou (residential registration) statues.[10] But this measure of $\dot{N}_j$ 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%.[11]

*** Insert Figure 1 about here ***

[10] Chinese citizens are registered either as rural Hukou or urban Hukou according to the place of birth their mothers. Although population mobility rose substantially in the 1990s as a result of liberalized urban housing and job markets, conversion of rural Hukou to urban Hukou by rural migrant workers in cities was still tightly regulated.

[11] Glaeser and Gyourko (2005) show a strong asymmetry in housing supply elasticity for cities experiencing negative demand shocks due to housing durability. The cities in our sample generally experienced positive demand shocks during our study period.

Cross-city housing price differential p_j and price growth $\dot{P}_j$

Available housing price indexes in China over our sample period, such as *Zhong Fang* property price index and *Guo Fang Jing Qi* property index, are based on average per-square-meter sale prices of new homes, which are often of better quality than average homes in a city. They cover only a relatively small number of large cities and are not quality-adjusted. We need consistent measures of base-period price differentials across cities p_j and sample-period price growth $\dot{P}_j$ that reflect the cost of housing consumption for representative households in a relatively large number of cities. We thus resort to indirect estimates based on housing consumption incentives revealed by the housing choices made by households in the two time periods observed in the UHS data. We adopt the estimates provided by Zheng et al. (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 in addition to city fixed effects and the city fixed effects for 2004. These fixed effects capture the differential price incentives for housing consumption (or the user cost of housing) across the cities and the change in these incentives between 1998 and 2004; the former identify the base-period housing price differential $p_j \equiv \ln\left(P_j/P_0\right)$ multiplied by $1-\varepsilon_P$ (P_0 denotes the price level in Beijing) and the latter, log price growth $\dot{P}_j$ multiplied by the price elasticity ε_P.[12] These p_j and $\dot{P}_j$ estimates are displayed in Figure 2. Over our sample period, the average log price growth across the 85 cities is 0.73, with a standard deviation of about 27%. Figure 3 plots the population growth measure $\dot{N}_j$ against the price growth measure $\dot{P}_j$ across the 85 cities; as in Glaeser et al. (2006), these growth rates appear quite dispersed, suggesting possible wide variations in housing supply elasticity.

*** Insert Figure 2 about here ***

[12] The UHS sample size for individual cities ranges from 50 to over 1000 households depending on city size. The dependent variable of the housing demand equation is the log housing consumption quantity, computed according to the observed hedonic home attributes observed in each period and the 2004 market implicit prices of these attributes as weights. These implicit prices are estimated based on appraised market value of homes reported only in 2004. Assuming a constant elasticity of substitution σ=1/4 and a baseline housing expenditure share of income ($1-\beta$)=1/3 (hence a price elasticity ε_P=0.5), p_j equals 2 times the base-period city fixed effects and $\dot{P}_j$, 2 times the city fixed effects for 2004. General price inflation during our sample period was negligible.

*** Insert Figure 3 about here ***

Cross-city wage-rate differential w_j and wage-rate growth $\dot{W}_j$

Estimates of cross-city wage-rate differential and wage-rate growth are derived from UHS data based on Mincerian wage regression. These estimates are reported in Zheng et al. (2009), where the wage regression is applied to the pooled sample of households across the 85 cities in two time periods. The dependent variable is log household employment income and the independent variables include the household size, the gender, age, employment sector and education attainment of the household head (main income earner), as well as city fixed effects and the city fixed effects for year 2004. Similar to the household housing demand estimates discussed above, the city fixed effects in the wage regression identify the cross-city wage-rate differential (relative to the wage rate in Beijing) w_j and the city fixed effects for 2004 identify the wage-rate growth in individual cities between 1998 and 2004, $\dot{W}_j$. The average log wage-rate growth across is 1.45, with a standard deviation of about 14%.

Note that the standard deviation of the population growth statistics $\dot{N}_j$ is notably smaller than that of the wage-rate growth $\dot{W}_j$, which in turn is notably smaller than the standard deviation of the price growth measure $\dot{P}_j$. The excess volatility of $\dot{P}_j$ over that of $\dot{W}_j$ can be accounted for by the variance of $\dot{\theta}_j$ and the relatively small elasticity of the utility with respect to $\dot{P}_j$, $(1-\beta)<1$, according to Eq. (10). The relatively low standard deviation of $\dot{N}_j$, however, may be in part due to measurement errors, as $\dot{N}_j$ is measured by the change in non-agricultural population, which does not account for new migrant workers with rural Hukou.

3.2 Measuring urban demand shocks and urban expansion

The urban demand shocks, arising from $\dot{a}_j + \gamma \dot{Z}_A$ and $\dot{\theta}_j - \dot{u}$, can be measured by the endogenous variables independently of the housing supply elasticity, as Eq. (9) and Eq. (10) show; the supply shift

$\dot{H}+\bar{\eta}\dot{Z}_H$, however, cannot be so measured, as Eq. (11) shows. To distinguish the housing supply shift from the housing supply elasticity differential in the structural model, Eq. (16), we assume that the shift is $\dot{Z}_H$ is proportional to urban expansion rate UER_j, measured by observed growth in urban built-up area and in road space: $UER_j = \dot{Z}_H / \rho$, where $1/\rho$ is a constant elasticity. In China, urban expansion is largely determined by local public capital formation. We predict the demand shock measures, $b\dot{N}_j + \dot{W}_j = \left(\dot{a}_j + \gamma \dot{Z}_A\right) / (1-\alpha+\gamma)$ and $(1-\beta)\dot{P}_j - \dot{W}_j = \dot{\theta}_j - \dot{u}$, and the supply shift measure UER_j, by predetermined vectors of urban attributes, $\mathbf{x}_1$, $\mathbf{x}_2$ and $\mathbf{x}_3$ respectively. The predicted values are represented by the variables PDS_j, AMS_j and ZHS_j respectively; these variables are estimated from the following equations:

$$b\dot{N}_j + \dot{W}_j = PDS_j\left(\mathbf{x}_1\right) + v_{1j}, \tag{17}$$

$$\left(1-\beta\right)\dot{P}_j - \dot{W}_j = AMS_j\left(\mathbf{x}_2\right) + v_{2j}, \tag{18}$$

$$UER_j = ZHS_j\left(\mathbf{x}_3\right) + v_{3j}, \tag{19}$$

where v_1 through v_3 are random errors. A key determinant of PDS_j is the capital investment in the city over the sample period, measured by the per capita fixed investment between 1998 and 2004, $FK9804_j/N_j$. We instrument this variable using another vector of predetermined variables $\mathbf{x}_4$; the predicted value of ln($FK9804_j/N_j$), denoted by $CFK_j\left(\mathbf{x}_4\right)$, is estimated by the equation below, where v_4 is a random error:

$$\ln\left(FK9804_j / N_j\right) = CKF_j\left(\mathbf{x}_4\right) + v_{4j}. \tag{20}$$

Determinants of the productivity shock measure $PDS_j\left(\mathbf{x}_1\right)$

We postulate that the productivity growth in a city, manifested by $b\dot{N}_j + \dot{W}_j$, is influenced by the following variables: (i) 1998 land rent premium p_j (reflecting location productivity); (ii) per capita new fixed capital formation, $CFK_j\left(\mathbf{x}_4\right)$; (iii) 1998 state-owned-enterprise (SOE) share of urban employment $SOE\%_j$ (reflecting adverse incentive for productivity improvement); and (iv) the urban amenity shock

$AMS_j(\mathbf{x}_2)$ (reflecting demand on urban resources for non-productive uses). We further control for urban wage-rate premium w_j and unemployment rate UMP_j, to account for conditional convergence in urban productivity and cyclical productivity fluctuations.

Determinants of the amenity shock measure $AMS_j(\mathbf{x}_2)$

We use several urban attributes observed in 1998 to predict the amenity shock as manifested by $(1-\beta)\dot{P}_j - \dot{W}_j$, including: (i) city temperature index $TEMP_j$, 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;[13] (ii) SO_{2j} emission per dollar of GDP output, to reflect the environmental quality; (iii) the average years of schooling of the adult population in the city, EDU_j, to reflect the social environment quality (e.g., Shapiro, 2006). A dummy variable $COASTAL_j$, indicating whether or not a city is in the coastal region, and the 1998 relative land rent p_j are also include in $\mathbf{x}_2$ to further control for location heterogeneity.

Determinants of the housing supply shift measure $ZHS_j(\mathbf{x}_3)$

Urban land use and road space expansions underlie the growth of public capital Z_H for housing production. We observe each city's growth in urban built-up area *g_built* (between 1998 and 2003) and in road space *g_road* (between 1998 and 2004); we define urban expansion rate $UER_j \equiv$ (*g_built*×6/5+*g_road*)/2. Although urban governments control both land conversion and transport infrastructure investment, their choices can be influenced by urban land scarcity and demand conditions.[14] In particular, we postulate that $\mathbf{x}_3$ includes the base-period population density in urban area

[13] 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.

[14] All urban land in China is state owned and urban governments monopolize the right to convert non-urban land to urban uses as well as the granting of urban sites for redevelopment via long-term leases (see (see Fu and Somerville, 2001; Wu et al., 2010). Incentivized by the fiscal decentralization reform in the 1990s (Jin and Zou, 2005) and the political mandate to promote local economic growth, city 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.

DENSITY (reflecting the pressure for urban expansion), population density in city's rural area *DENSITY_RU* (reflecting the cost of land conversion, which increases with the displacement of rural residents), base-period housing price differential p_j and wage-rate differential w_j (both indicative of location productivity), and the based-period location amenity quality indicated by $AMS_j(\mathbf{x}_2)$ (additional demand for urban expansion). We also include in $\mathbf{x}_3$ base-period urban population N_j and the coastal location dummy to further control for urban expansion constraints.

Determinants of urban fixed investment $CFK_j(\mathbf{x}_4)$

The fixed investment in the city, measured by $FK9804_j/N_j$, is affected by the city's location quality and access to resources. More productive locations, indicated by higher base-period land-rent premium p_j and wage-rate premium w_j, are expected to attract more capital investment, which reinforces location productivity; such circular causation is well documented in the new economic geography literature (*e.g.* Krugman, 1991). We also expect cities that in 1998 have a more dominant presence of SOEs (indicated by *SOE*%), a significant presence of foreign direct investment (indicated by FDI share of fixed investment, *FDI*%), and a higher level of fixed investment per capita ($FK98_j/N_j$), as well as provincial capital or provincial level cities ($CAPTL_j$=1), to have access to more resources for investment over our sample period. In addition, we control for city size N_j and labor force quality (indicated by 1998 college graduates share of adult population in the city, *SCH16%*).

Estimates of the demand and supply shocks

We estimate Equation (17) through Equation (20) jointly using a GMM estimator. The results are reported in Table 2. We need to choose parameters $1-\beta$ and $b=(\gamma-\omega)/(1-\alpha+\gamma)$ in Equation (17) and (18). We choose a baseline housing-related share of consumer expenditure $1-\beta$ =1/3. The value of b is set to 0.3 in accordance to the assumptions of $1-\alpha$=40% (wage share of GDP),[15] γ=20%, and ω=2% (or any combination of ω and $\gamma=\left((1-\alpha)b+\omega\right)/(1-b)$). We validate the assumption of b=0.3 when we estimate the structural parameters.

[15] Workers' pay was 36.7% of GDP in 2005, according to All-China Federation of Trade Unions.

*** Insert Table 2 about here ***

The first column of Table 2 shows conditional convergence in urban productivity, which grows faster in lower-wage locations and where land rent is more expensive (reflecting location advantages), SOE share of employment is lower (more vibrant private enterprises), and per capita fixed investment is higher. The key instruments for the per capita fixed investment during our sample period include 1998 per capita fixed investment, FDI share of fixed investment and the college graduate share of the population in the city (as shown in the last column of Table 2). Interestingly, cities facing strong demand for their amenities (a more positive *AMS* value) see their productivity growth severely compromised, indicating diversion of resources to non-productive uses. Finally the productivity growth is affected by the cyclical condition of the city economy as indicated by the urban unemployment rate. In particular, the productivity growth is expected to be slowest, other things being equal, when the unemployment rate is near 5%; an unemployment rate below of above this rate (sort of natural unemployment rate) would predict above-trend productivity growth due to a tight labor market condition or due to expected cyclical recovery.

The second column shows conditional convergence in housing prices across cities: the price grows faster than the wage rate, reflecting positive amenity demand shocks, in coastal cities and in cities where climate (*TEMP*) is more moderate, air pollution level (SO_2) is lower, or average education level of the residents (*EDU*) is higher.

The third column reports the estimates of urban expansion rate, showing a faster expansion in denser cities (due to population pressure), in more productive cities (as reflected by higher land-rent premium p_j and urban wage-rate premium w_j), and in cities facing growing demand for their amenity. We also find that cities with larger population and in coastal regions face more severe constraints in land-use expansion. The effect of rural population density is more ambiguous; a higher rural density, on the one hand, raises the demand for urbanization, but on the other hand it raises the opportunity cost of land for urban expansion.

The last column in Table 2 shows the predictability of the fixed investment in individual cities over our sample period. As expected, the investment is driven by location productivity as reflected by higher p_j and w_j. Cities with a more dominant SOE sector, with a high level of per capita fixed investment in the base year (1998), or being a provincial capital or provincial-level city, seem to have greater access to financial resources to finance fixed investment over the sample period. A stronger presence of FDI in the city is indicative of the city's openness to global trade in 1998 and its location advantage and, hence, predicts stronger investment in subsequent years. A negative effect of population size on per capita fixed investment is suggestive of economies of scale: the requirement for per capita capital stock would be lower in larger cities for the same productivity level. Lastly, we note that the fixed investment is relatively lower in cities with a more educated workforce (as indicated by *SCH16*%); this finding reflects the fact that urban expansion during our sample period was not so much driven by innovations, which would rely more on education, but was mostly driven by investment in technologies that can take advantage of low-cost labor.

4. Estimating the structural coefficients and the cross-city determinants of housing supply elasticity

4.1 The empirical specification

To specify the housing supply elasticity differential $\Delta\varepsilon_H$ in Eq. (16), we examine several possible determinants of housing supply elasticity in Chinese cites. During the sample period, developers generally acquired residential sites from the local government at negotiated prices. These prices have to cover the cost of resettling displaced residents and various government charges. The structural model indicates that the housing supply elasticity ε_H decreases with the nontraded capital share (or land cost share) of housing price η_j. That share would be higher where developable sites are more expensive to acquire and regulations for housing development are more costly to comply. We hypothesize that a higher built-up area population density (*DENSITY*) and a higher base-period land rent premium p_j raise the cost of resettlement compensation and hence the land cost. In addition, the government in capital cities (*CAPTL*=1), with more resources at disposal, may be more inclined to promote urban redevelopment by refraining from levying excessive land-use and development charges. We also include the base-period urban road space per capita (*ROAD*) as an determinant of $\Delta\varepsilon_H$; Fu and Somerville (2001) find that development sites in Shanghai where adjacent roads are less congested are often granted a higher allowed floor-to-area ratio (FAR), which provides the developers greater flexibility in raising building density.

We further examine the impact of two institutional variables on housing supply elasticity: (i) income inequality within the city, measured by the ratio of 75 percentile household income to 25% percentile household income in 1998, denoted by *QR98*; and (ii) urban government efficiency, denoted by *G_EFF*. Between 1998 and 2004, urban governments produced few homes for low-income households and largely left the choice of new residential developments to the private market so as to maximize the revenue from land disposal. More extreme income inequality in a city is expected to produce more rampant low-density luxury residential development, making developable sites more scarce and hence the housing supply elasticity lower.[16]

G_EFF is derived from a World Bank study of investment environment across 120 Chinese cities (World Bank, 2006, Table B5). The World Bank study reports two scores, denoted *G_EFF_TFP* and *G_EFF_FDI* respectively, which summarize 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 90th percentile level; the *G_EFF_FDI* score is the expected gain in local foreign direct investment level (indicating investment

[16] In response to the lack of high-density new home supply, the State Council issued an executive guideline in 2006 to regulate the dwelling-size mix in new housing projects in cities (关于调整住房供应结构稳定住房价格的意见), requiring at least 70% of the dwelling units in new development projects to be no larger than 90 square meters in size.

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 -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 $G\_EFF \equiv -(G\_EFF\_TFP/\sigma_{TFP} + G\_EFF\_FDI/\sigma_{FDI})/2$, where σ_{TFP} and σ_{FDI} are the standard deviation of the *G_EFF_TFP* and *G_EFF_FDI* scores respectively. A higher *G_EFF* score indicates lower regulatory costs, which would raise the productivity and the investment profitability for private enterprises as well as for residential development.

Hence $\Delta\varepsilon_H = \varepsilon_H - \bar{\varepsilon}_H$ is specified by the following equation:

$$\begin{aligned}\Delta\varepsilon_H &= k_1 p_j + k_2 \ln\left(DENSITY_j\right) + k_3 CAPTL_j + k_4 \ln\left(ROAD_j\right) \\ &+ k_5\left(QR98_j > 1.842\right) + k_6 G\_EFF_j - k_0 + v_j,\end{aligned} \tag{21}$$

where v_j is a zero-mean random error and the constant k_0, which would increase with $\bar{\varepsilon}_H$, is chosen such that $\Delta\varepsilon_H$ has a zero mean value. Our hypotheses are that $k_1<0$, $k_2<0$, $k_3>0$, $k_4>0$, $k_5<0$ and $k_6>0$. Eq. (21) will be estimated jointly the structural coefficients $\bar{\Lambda}$ in Eq. (16), which now can be rewritten to incorporate the estimated demand and supply shocks:

$$\begin{bmatrix} \dot{N}_j \\ \dot{W}_j \\ \dot{P}_j \end{bmatrix} \approx \begin{bmatrix} \bar{\eta}\rho\bar{\lambda}_{NH} \\ \bar{\eta}\rho\bar{\lambda}_{WH} \\ \bar{\eta}\rho\bar{\lambda}_{PH} \end{bmatrix} \frac{\dot{H}}{\bar{\eta}\rho} + \bar{\Lambda} \begin{bmatrix} \bar{\eta}\rho ZHS_j \\ (1-\alpha+\gamma) PDS_j \\ AMS_j \end{bmatrix} + \begin{bmatrix} \bar{\eta}\rho\bar{\lambda}_{NH} \\ \bar{\eta}\rho\bar{\lambda}_{WH} \\ \bar{\eta}\rho\bar{\lambda}_{PH} \end{bmatrix} \frac{S_j}{\rho} \bar{\eta}\Delta\varepsilon_H. \tag{22}$$

Furthermore, S_j can be computed according to the estimated demand and supply shocks:

$$\begin{aligned} S_j &= \left(\bar{\lambda}_{NH} + \bar{\lambda}_{WH} + \bar{\lambda}_{PH}\left(1-\varepsilon_P\right)\right)\bar{\eta}\rho\left(\dot{H}/\bar{\eta}\rho + ZHS_j\right) \\ &+ \left(\bar{\lambda}_{NA} + \bar{\lambda}_{WA} + \bar{\lambda}_{PA}\left(1-\varepsilon_P\right)\right)\left(1-\alpha+\gamma\right) PDS_j \\ &+ \left(\bar{\lambda}_{N\theta} + \bar{\lambda}_{W\theta} + \bar{\lambda}_{P\theta}\left(1-\varepsilon_P\right)\right) AMS_j - \rho ZHS_j. \end{aligned} \tag{23}$$

4.2 Estimates of the structural model

Table 3 reports the correlation coefficients among the endogenous urban performance measures (namely $\dot{N}_j$, $\dot{W}_j$ and $\dot{P}_j$), the urban demand and supply shock measures (PDS_j, AMS_j and ZHS_j), and the determinants of the housing supply elasticity. We find, not surprisingly, that population growth is highly correlated with the urban expansion measure ZHS_j. In addition, the population growth, as well as ZHS_j, is positively correlated with the amenity shock AMS_j but somewhat negatively correlated with the productivity shock PDS_j. We also find cities with lower regulatory cost grew faster but those with very high income inequality grew relatively slowly over our sample period.

*** Insert Table 3 about here **

We estimate the structural model represented by Eq. (21) and Eq. (22) in several steps, with different specifications of S_j and coefficient restrictions for $\bar{\mathbf{\Lambda}}$. We first assume S_j to be a constant; the resulting estimates of $\bar{\mathbf{\Lambda}}$ then are used to compute S_j according to Eq. (23). We show that the structural estimates of $\bar{\mathbf{\Lambda}}$ and Eq. (21) are robust whether S_j is treated as a constant or is determined by Eq. (23). The coefficient restrictions implied by equations (A1) through (A9) in the Appendix will be used to aid the identification of $\bar{\mathbf{\Lambda}}$. In particular, we will estimate the baseline housing supply elasticity $\bar{\varepsilon}_H$, which, together with parameters b, $(1-\beta)$, μ and ρ, determines all the coefficients in $\bar{\mathbf{\Lambda}}$.

It is useful to take a look first at the statistics of Eq. (11) to have a rough estimation of $\bar{\varepsilon}_H$ and $\dot{H}$. On the one hand, the mean and standard deviation of $\dot{N}_j + \dot{W}_j + (1-\varepsilon_P)\dot{P}_j - \dot{Z}_H$, whose predicted value at $\bar{\varepsilon}_H$ determines S_j, are 1.88 and 0.22 respectively; on the other hand, the mean and standard deviation of $\dot{P}_j - (1-\mu)\dot{W}$ are 0.138 and 0.28 respectively.[17] The difference in these mean statistics, according to Eq. (11), implies a very large $\dot{H}$ or a large $\bar{\varepsilon}_H$ or both. A $\bar{\varepsilon}_H$ value of 1, for example, would imply $\dot{H}$=0.8,

[17] $\dot{Z}_H$ is measured by ρUER_j, where ρ=0.2 as the structural coefficient estimates will show. We assume $1-\mu$=0.4 (the same value as $1-\alpha$) and $\varepsilon_P = 0.5$ (same as that used to compute $\dot{P}_j$ and p_j as explained in footnote 12).

which does not seem implausible in view of the strong total factor productivity growth in the economy during the period.[18] The relative magnitude of the standard deviation statistics, however, implies a very low $\bar{\varepsilon}_H$ (given that $\dot{H}$ is spatially invariant). The relatively small standard deviation of $\dot{N}_j + \dot{W}_j + (1-\varepsilon_P)\dot{P}_j - \dot{Z}_H$ is likely due to the suppressed variance of the population growth statistics (recall that $\dot{N}$ fails to take into account the new migrants with rural Hukou, who contribute importantly to differential population growth across Chinese cities). Eq. (22) indicates that an underestimated $\bar{\varepsilon}_H$ will bias k_0 and hence the estimate of $\dot{H}/\bar{\eta}\rho$, towards negative values. The estimate of $\dot{H}/\bar{\eta}\rho$ turns out to be very small and imprecise, ranging from slightly negative to slightly positive, across alternative specifications. For efficiency, we report in Table 4 the GMM estimates of the structural model with $\dot{H}/\bar{\eta}\rho$ set to zero.[19]

*** Insert Table 4 about here ***

The first four columns in Table 4 reports the structural estimates with a constant S_j, denoted by $\bar{S}$. The estimates in column 1 are based on the equation for population growth only. All the structural coefficients are of expected sign and are statistically significant. Furthermore, the estimates of the determinants of the supply elasticity differential are of expected sign, although only four of the six variables are statistically significant at least at 5% confidence level: capital cities and cities having higher base-period road density and low regulatory costs show higher housing supply elasticity (holding constant the urban built-up area), whereas those with higher base-period population density, high land rent premium, or high income inequality, show lower supply elasticity.

Columns 2 to 4 in Table 4 reports estimates including additional equations for the endogenous variables, with additional structural coefficients determined by parameters b and $1-\beta$ according to Equations (A1) through (A6). We set the baseline housing share of household expenditure at $1-\beta = 1/3$ but let the value

[18] Work Bank (2008) reports an overall TFP growth of almost 5% per year for China between 1999 and 2005.

[19] The estimation results are qualitatively similar when $\dot{H}/\bar{\eta}\rho$ is set to alternative values around zero.

of b determined by the estimation. The estimates of the structural coefficients for population growth remain positive and their statistical significance improves with the incorporation of the additional endogenous growth measures in the system of equations. The estimates for b is slightly greater than 0.3, the value used to estimate PDS_j in Table 2. The estimate for b, however, could be somewhat inflated by artificially low variability in our population growth measure, which depresses the magnitudes of the structural coefficients for $\dot{N}_j$ relative to those for $\dot{W}_j$ and $\dot{P}_j$. The estimates of the determinants of the supply elasticity differential remain qualitatively similar but show improved statistical significance when the wage-rate growth is included in the system of equations.

*** Insert Table 5 about here ***

Table 5 shows the calculation of all the structural coefficients according to the coefficient restrictions used in estimating the structural model. We also show the implied values of the reduced-form estimators of the housing supply elasticity and the implied value for parameter ρ, the reciprocal of the elasticity of urban expansion measure *ZHS* with respect to the growth in nontraded capital for the housing sector $\dot{Z}_H$. The values of the reduced-form estimators, $\bar{\lambda}_{NA}/\bar{\lambda}_{PA}$ and $\bar{\lambda}_{N\theta}/\bar{\lambda}_{P\theta}$, are less than 0.15, indicating a very small estimate for the baseline housing supply elasticity according to Eq. (14) and Eq. (15). These values, again, could be downwardly biased due to the suppressed volatility of the $\dot{N}_j$ statistics. The implied value of ρ is about 0.2.

In column 5, we incorporate Eq. (14) and Eq. (15), which are derived from Eq. (A8) and Eq. (A9), as additional coefficient restrictions, to determine $\bar{\lambda}_{NA}$ and $\bar{\lambda}_{N\theta}$ by the baseline housing supply elasticity $\bar{\varepsilon}_H$. To make the estimation more robust, we linearize Eq. (15) with respect to b and $\bar{\varepsilon}_H$ at b = 0.3 $\bar{\varepsilon}_H = 1$ respectively, to obtain $\bar{\lambda}_{N\theta}/\bar{\lambda}_{P\theta} = \left(3.26 + 12.76(b-0.3) + 3.02(\bar{\varepsilon}_H - 1)\right)$. We further assume $1-\mu$=0.4 (the same value as $1-\alpha$) and $\varepsilon_P = 0.5$ (same as that used to compute $\dot{P}_j$ and p_j as explained in footnote 12). The $\bar{\varepsilon}_H$ estimate is statistically significant but is very small, about 0.12, in line with the value of the reduced-form estimators. Such a small estimate of $\bar{\varepsilon}_H$ is unrealistic; it implies $\bar{\eta}$=0.9. In fact η has to be less that μ (which is set at 0.6) for the traded capital to play any role in housing service production. Although we believe the true value of $\overline{\varepsilon}_H$ to be relatively small, given that it represents the supply adjustment at the intensive margin (via construction density) and that the Chinese cities in our sample were generally of very high density already in the base period (the sample mean of *DENSITY* is over 11 thousand people per square kilometer in urban built-up area), an estimate of $\overline{\varepsilon}_H$ close to 1 (implying $\overline{\eta}$ close to 0.5) is probably more realistic. Saiz (2010) reports a mean supply elasticity (at both intensive and extensive margins) of about 2 across US metropolitan areas.

We use the structural coefficient estimates reported in column 4 of Table 4 (as well as those shown in the corresponding column of Table 5) to compute S_j according to Eq. (23), which gives the value of $\dot{N}_j + \dot{W}_j + (1-\varepsilon_P)\dot{P}_j - \dot{Z}_H$ predicted by the demand shocks and supply shifts at the baseline supply elasticity $\overline{\varepsilon}_H$. We use the resulting S_j, together a ρ value of 0.2 (see Table 5), to re-estimate the structural model and the determinants of $\overline{\eta}\Delta\varepsilon_H$. The results are reported in column 6 of Table 4 and they are similar to those reported in column 5. Note that the effects of land assembly costs (indicated by the base-period land rent premium and population density) on housing supply elasticity appear statistically less significant but those of city political status and base-period road density become statistically more significant.

As the last two rows in Table 3 show, the estimates $\overline{S}\overline{\eta}\Delta\varepsilon_H$ and $S_j\overline{\eta}\Delta\varepsilon_H$ are highly correlated (with a correlation coefficient of 0.978). They are also similarly correlated with the endogenous growth and exogenous shock measures. Thus the estimates of the structural coefficients as well as those determinants of the supply elasticity appear robust with respect to the measurement of the price incentive for supply adjustment S_j, whether it is assumed invariant or computed according to Eq. (23). The effects of the spatial variation in housing supply elasticity are mainly due to the variation in $\Delta\varepsilon_H$, not in S_j, although the standard deviation of S_j is much greater than that of $\Delta\varepsilon_H$. According to the estimates reported in the last column of Table 4, the standard deviation of $\overline{\eta}\Delta\varepsilon_H$ is 3.0%, which implies that $\Delta\varepsilon_H$ has a standard deviation of about 3.3% given $\overline{\varepsilon}_H = 1/\overline{\eta} - 1 = 0.117$. Further experiments show that $\bar{S}\bar{\eta}\Delta\varepsilon_H$ and $S_j\bar{\eta}\Delta\varepsilon_H$ are not correlated with the exogenous shocks, measured by $b\dot{N}_j + \dot{W}_j = \left(\dot{a}_j + \gamma \dot{Z}_A\right) / (1 - \alpha + \gamma)$, $(1-\beta)\dot{P}_j - \dot{W}_j = \dot{\theta}_j - \dot{u}$, and $UER_j = \dot{Z}_H / \rho$, once PDS_j, AMS_j and ZHS_j are respectively controlled for.

*** Insert Table 6 about here ***

Despite the small estimate of $\bar{\varepsilon}_H$, the model estimates indicate a significant contribution of the differential housing supply elasticity to the urban growth performance across Chinese cities. Table 6 shows the contribution of the exogenous shocks and the supply elasticity differential to the cross-city variation in the endogenous growth measures. Among the exogenous shocks, the amenity shock appears to vary most across the cities, with a standard deviation of about 10%; the variations of the productivity shocks, the supply shift and the supply adjustment due to price elasticity $\bar{\eta}\bar{S}\Delta\varepsilon_H$ appear similar, each of which has a standard deviation of about 5%. As the structural coefficient values in Panel A of Table 6 show, the population growth is most responsive to the supply shift at the baseline value of the supply elasticity but least responsive to the amenity shock; wage rate is most responsive to the productivity shock but least to the amenity shock; price is most responsive to the productivity shock as well but least to the supply shift. In addition, the structural coefficient estimates with respect to $\bar{\eta}\bar{S}\Delta\varepsilon_H$ indicate, according to the first-order derivatives of the structural coefficients (see Table A in the Appendix), that the population response to the demand shocks is most affected by the supply elasticity but the wage-rate response is least affected.

Panel B of Table 6 shows the standard deviation of the component factors of the endogenous growth measures. With respect to the cross-city variation in population growth, the variation in supply shift, as well as that in supply elasticity, is more important than the variations in productivity shock and amenity shock as a contributing factor. The variation in population growth due to the demand shocks (the productivity shock and the amenity shock combined) is about one quarter of that due to the variation in supply (the supply shift and the supply elasticity differential combined). In contrast, the cross-city variations in wage-rate growth and price growth are mostly due to the differences in demand shocks;

the standard deviation due to the demand shocks is more than 4 and 3 times that due to the supply variation for wage-rate growth and price growth, respectively. The dampened response of population growth to the demand shocks, as well as the elevated responses of wage-rate growth and price growth, is consistent with a low baseline housing supply elasticity among the Chinese cities.

5. Conclusions

By incorporating the spatial equilibrium conditions in the analysis of cross-city growth performance, we are able to determine, via the structural coefficients of the model, how the responses of each of the endogenous variables (urban population, wage rate, and housing price) to the demand and supply shocks are affected by the supply elasticity. The structural model thus aids not only the estimation of the demand shocks independently of the supply elasticity but also the identification of the spatial variation in the supply elasticity from the combinations of the endogenous growth measures. Across Chinese cities post the housing market liberalization, over the period from 1998 to 2004, the model estimates show, on the one hand, relatively small population response to the demand shocks but large wage-rate and price responses, consistent with a relatively low baseline housing supply elasticity among Chinese cities. On the other hand, the cross-city variation in housing supply accounts for most of the differences in urban population growth but relatively little of the differences in wage-rate and price growth.

The spatial variation in the supply elasticity is predictable by the observed base-period urban attributes. In particular, a higher population density in the built-up area and a higher land rent premium reduce the supply elasticity by raising the cost of land assembly for redevelopment. The local government in provincial capital or provincial-level cities or in cities with more abundant road space per capita, appears more inclined to encourage urban redevelopment by lowering redevelopment charges or allowing higher building density, thus elevating the housing supply elasticity. Moreover, we find that cities with more extreme income inequality among residents tend to have lower housing supply elasticity, possibly due to the excessive use of urban land for low-density luxury housing developments, which aggravates land scarcity. In addition, cities with a more efficient local government, which imposes lower regulatory costs on business, are found to have higher housing supply elasticity. These results are useful for informing housing policies, showing that local governments can play an important role in raising housing supply elasticity, even in high-density built-up areas, by lowering regulatory costs, improving the equity of land-use allocation, and investing in public transport infrastructure.

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Appendix. The structural coefficients in $\mathbf{\Lambda}$ and their first order derivatives with respect to the supply elasticity $\varepsilon_H \equiv 1/\eta_j - 1$

To determine the coefficients of $\mathbf{\Lambda}$, we differentiate Equation (3), (6) and (7), respectively, with respect to the exogenous variables $HZ_H^{\eta_j}$, $a_j Z_A^{\gamma}$ and θ_j and obtain:

$$\lambda_{WH} = -b\lambda_{NH}\,,\quad \lambda_{WA} = 1/(1-\alpha+\gamma) - b\lambda_{NA}\,,\ \lambda_{W\theta} = -b\lambda_{N\theta}\,, \tag{A1-3}$$

$$(1-\beta)\lambda_{PH} = \lambda_{WH}\,,\ (1-\beta)\lambda_{PA} = \lambda_{WA}\,,\ (1-\beta)\lambda_{P\theta} = \lambda_{W\theta} + 1\,, \tag{A4-6}$$

$$\lambda_{NH} + \lambda_{WH} + (1-\varepsilon_P)\lambda_{PH} = \left(\lambda_{PH} - (1-\mu)\lambda_{WH} + 1\right)/\eta_j\,, \tag{A7}$$

$$\lambda_{NA} + \lambda_{WA} + (1-\varepsilon_P)\lambda_{PA} = \left(\lambda_{PA} - (1-\mu)\lambda_{WA}\right)/\eta_j\,, \tag{A8}$$

$$\lambda_{N\theta} + \lambda_{W\theta} + (1-\varepsilon_P)\lambda_{P\theta} = \left(\lambda_{P\theta} - (1-\mu)\lambda_{W\theta}\right)/\eta_j\,, \tag{A9}$$

where $b \equiv (\gamma-\omega)/(1-\alpha+\gamma)$. Solving the above 9 linear equations for the λ coefficients yields their expression in Table A. A smaller coefficient of price elasticity of housing demand ε_P makes λ_{NA}, $\lambda_{N\theta}$ and $\lambda_{W\theta}$ smaller, but other structural coefficients greater, in magnitude.

Table A

Structural coefficients λ	$\frac{d\lambda}{d\varepsilon_H}$ evaluated at $\eta_j = \bar{\eta}$
$\lambda_{NH} = \frac{(1-\beta)(1-\alpha+\gamma)}{\Phi}$	$\bar{\lambda}_{NH}\left(\bar{\lambda}_{NH} + \bar{\lambda}_{WH} + \bar{\lambda}_{PH}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{NA} = \frac{(\mu+\beta-\mu\beta)-(1-\beta)\eta_j-(1-\varepsilon_P)\eta_j}{\Phi}$	$\bar{\lambda}_{NH}\left(\bar{\lambda}_{NA} + \bar{\lambda}_{WA} + \bar{\lambda}_{PA}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{N\theta} = \frac{(1-\alpha+\gamma)\left(1-(1-\varepsilon_P)\eta_j\right)}{\Phi}$	$\bar{\lambda}_{NH}\left(\bar{\lambda}_{N\theta} + \bar{\lambda}_{W\theta} + \bar{\lambda}_{P\theta}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{WH} = \frac{-(1-\beta)(\gamma-\omega)}{\Phi}$	$\bar{\lambda}_{WH}\left(\bar{\lambda}_{NH} + \bar{\lambda}_{WH} + \bar{\lambda}_{PH}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{WA} = \frac{(1-\beta)\eta_j}{\Phi}$	$\bar{\lambda}_{WH}\left(\bar{\lambda}_{NA} + \bar{\lambda}_{WA} + \bar{\lambda}_{PA}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{W\theta} = \frac{-(\gamma-\omega)\left(1-(1-\varepsilon_P)\eta_j\right)}{\Phi}$	$\bar{\lambda}_{WH}\left(\bar{\lambda}_{N\theta} + \bar{\lambda}_{W\theta} + \bar{\lambda}_{P\theta}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{PH} = \frac{-(\gamma-\omega)}{\Phi}$	$\bar{\lambda}_{PH}\left(\bar{\lambda}_{NH} + \bar{\lambda}_{WH} + \bar{\lambda}_{PH}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{PA} = \frac{\eta_j}{\Phi}$	$\bar{\lambda}_{PH}\left(\bar{\lambda}_{NA} + \bar{\lambda}_{WA} + \bar{\lambda}_{PA}(1-\varepsilon_P)\right)\bar{\eta}^2$
$\lambda_{P\theta} = \frac{(1-\alpha+\omega)\eta_j-(1-\mu)(\gamma-\omega)}{\Phi}$	$\bar{\lambda}_{PH}\left(\bar{\lambda}_{N\theta} + \bar{\lambda}_{W\theta} + \bar{\lambda}_{P\theta}(1-\varepsilon_P)\right)\bar{\eta}^2$

Note: $\Phi \equiv (1-\beta)(1-\alpha+\omega)\eta_j + (\mu+\beta-\mu\beta)(\gamma-\omega) - (\gamma-\omega)\eta_j(1-\varepsilon_P) > 0$.

Table 1. Variable definition, data source and sample statistics

Variables	Definition [Data source]	Mean	Median	Max	Min	Std. Dev.	No. of obs.
$\dot{N}_j$	City "smoothed" log non-agricultural population growth, 1998-2004. [3]	0.185	0.195	0.404	-0.006	0.092	85
$\dot{W}_j$	Urban wage-rate growth, 1998-2004; inferred from Mincerian wage regression. [2]	1.451	1.451	1.885	1.194	0.135	85
$\dot{P}_j$	Housing price growth, 1998-2004; inferred from household housing demand. [2]	0.731	0.710	1.401	0.080	0.273	85
N_j	City non-agricultural population, 1998 [3].	119.1	70.5	893.7	10.1	138.4	85
w_j	City wage-rate premium: $w_j = \ln\left(W_j/W_0\right)$, 1998, Beijing: j=0. [2]	-0.370	-0.408	0.846	-0.892	0.283	85
p_j	City housing-price premium $p_j = \ln\left(P_j/P_0\right)$, 1998, Beijing: j=0. [2]	-2.015	-2.013	0.000	-3.732	0.822	85
g_built	Growth in urban built-up area, 1998-2003. [3]	0.347	0.288	1.386	-0.145	0.309	85
g_road	Growth in road space, 1998-2004. [3]	0.708	0.635	1.887	0.000	0.368	85
UER_j	(*g_built*×6/5+*g_road*)/2	0.563	0.519	1.686	0.000	0.321	85
DENSITY	City non-agricultural population over built-up area, 10,000 people /sqkm, 1998. [3]	1.131	1.107	2.169	0.520	0.326	85
DENSITY_RU	City agricultural population over non-built-up area, 10,000 people /sqkm, 1998. [3]	0.049	0.044	0.185	0.002	0.032	85
TEMP	City temperature index = vector distance between the city's summer and winter temperature (°C/100) and the sample min. summer temperature and max. winter temperature. [3]	0.195	0.170	0.348	0.082	0.083	85
COASTAL	Dummy for cities in coastal provinces.	0.459	0.000	1.000	0.000	0.501	85
SO_2	City SO_2 emission over GDP (ton per million Yuan), 1998. [3]	3.127	1.665	22.05	0.099	4.120	85
EDU	City average years of schooling of adult population (years). [3]	11.62	11.61	12.87	9.92	0.541	85
$\ln(FK9804/N_j)$	Natural log of cumulative fixed capital formation between 1998 and 2004 per person of 1998 non-agricultural population.	11.78	11.75	13.54	10.20	0.715	85

ln($FK98/N_j$)	Natural log of fixed capital formation in 1998 per person of non-agricultural population	8.824	8.638	11.122	7.534	0.732	85
CAPTL	Dummy for provincial capital city or provincial level city.	0.341	0.000	1.000	0.000	0.477	85
FDI%	1998 FDI share of fixed investment	0.014	0.009	0.057	0.000	0.014	85
SOE%	State-owned-enterprise (SOE) share of city employment, 1998. [3]	0.686	0.702	0.944	0.369	0.121	85
SCH16%	College graduate (with at least 16 years of schooling) share of urban adult population. [3]	0.059	0.048	0.175	0.018	0.037	85
UMP	City urban-area unemployment rate. [3]	0.039	0.032	0.154	0.000	0.029	85
ROAD	Road space per urban non-agricultural population (sqm), 1998. [3]	7.813	6.921	27.04	2.195	4.084	85
QR98	City 75th percentile household income over 25th percentile household income, 1998. [1]	1.715	1.679	2.380	1.400	0.182	85
G_EFF_FTP	Expected gain in total factor productivity (TFP) by local firms had the city government efficiency indicators improved to the 90th percentile level. [4]	0.169	0.165	0.340	-0.010	0.076	68
G_EFF_FDI	Expected gain in local foreign direct investment level had the city government efficiency indicators improved to the 90th percentile level. [4]	0.319	0.330	0.660	-0.220	0.162	68
G_EFF	City government efficiency score = -0.5×(*G_EFF_TFP* /stdev(*G_EFF_TFP*) + *G_EFF_FDI* /stdev(*G_EFF_FDI*)).	-2.100	-2.109	0.134	-3.745	0.894	68
AMS_j	Estimated amenity shock, based on the estimates in Table 2.	-1.200	-1.198	-0.971	-1.497	0.099	85
PDS_j	Estimated productivity shock, based on the estimates in Table 2.	1.503	1.505	1.714	1.304	0.084	85
ZHS_j	Estimated housing supply shift, based on the estimates in Table 2.	0.578	0.551	1.240	0.008	0.237	85

Note: The sources of data include: [1] Urban Household Survey (UHS); [2] Zheng, Fu and Liu (2009), which provide the estimates of $\dot{P}_j$, $\dot{W}_j$, p_j and w_j using household data from [1]; [3] Urban Statistics Yearbook; [4] World Bank (2006, Table B5).

Table 2. Estimates of urban demand shocks and housing supply shift

Equation (17) Dependent variable = $b\dot{N}_j + \dot{W}_j$ Predicted value = PDS_j		Equation (18) Dependent variable = $(1-\beta)\dot{P}_j - \dot{W}_j$ Predicted value = AMS_j		Equation (19) Dependent variable = UER_j Predicted value = ZHS_j		Equation (20) Dependent variable = ln($FK9804_j/N_j$) Predicted value = CFK_j	
w_j	-0.265 (8.5)***	p_j	-0.036 (3.2)***	w_j	0.525 (5.3)***	w_j	0.590 (5.3)***
p_j	0.036 (4.2)***	*TEMP*	-0.605 (5.3)***	p_j	0.196 (6.5)***	p_j	0.262 (5.8)***
SOE%	-0.223 (5.1)***	*COASTAL*	0.084 (3.6)***	ln(*DENSITY*)	0.470 (6.8)***	*SOE*%	0.846 (2.9)***
CFK	0.064 (4.6)***	SO_2	-0.006 (2.7)***	ln(*DENSITY_RU*)	0.017 (0.8)	ln(N_j)	-0.302 (7.7)***
AMS	-0.558 (9.6)***	*EDU*	0.141 (7.1)***	ln(N_j)	-0.140 (4.4)***	*CAPTL*	0.414 (6.1)***
UMP	-1.618 (3.6)***	Constant	-2.810 (11.1)***	*COASTAL*	-0.191 (3.9)***	ln($FK98/N_j$)	0.422 (8.2)***
UMP^2	16.270 (5.3)***			*AMS*	2.405 (10)***	*FDI*%	5.569 (2.4)**
Constant	0.230 (1.0)			Constant	0.437 (2.6)***	*SCH16*%	-6.048 (8.0)***
						Constant	9.659 (14)***
R squared	0.406	*R* squared	0.325	*R* squared	0.409	*R* squared	0.723

Note: Equation (17) through (20) are jointly estimated using GMM (cross-section White covariance) and common instruments that include all the predetermined variables in these equations. We choose $1-\beta = 1/3$ and b = 0.3. t-statistics are in parentheses; ***, **, and * denote, respectively, statistical significance at 1%, 5% and 10% levels. The number of observations is 85.

Table 3. Correlation matrix for urban growth measures and exogenous shocks (68 cities)

Variables	ID	1	2	3	4	5	6	7	8	9	10	11	12	13
$\dot{N}_j$	1													
$\dot{W}_j$	2	-0.259												
$\dot{P}_j$	3	0.112	0.121											
PDS_j	4	-0.144	0.655	0.091										
AMS_j	5	0.447	-0.519	0.248	-0.769									
ZHS_j	6	0.627	-0.240	0.145	-0.239	0.496								
p_j	7	0.404	0.063	0.128	0.168	0.117	0.722							
ln(*DENSITY*)	8	-0.074	0.238	0.032	0.003	0.022	0.201	0.082						
CAPTL	9	0.177	0.122	-0.033	0.001	0.050	0.264	0.304	0.518					
ln(*ROAD*)	10	0.499	-0.439	-0.080	-0.270	0.349	0.369	0.150	-0.417	-0.246				
Top 20% cities of highest income inequality: *QR98*>1.842	11	-0.224	-0.120	0.055	-0.156	0.091	-0.208	-0.069	-0.127	-0.087	-0.162			
G_EFF	12	-0.505	0.094	-0.137	-0.092	-0.172	-0.545	-0.651	-0.014	0.027	-0.263	0.022		
$\bar{\eta}\bar{S}\Delta\varepsilon_H$	13	0.531	-0.379	-0.124	-0.203	0.242	0.196	-0.045	-0.416	0.034	0.806	-0.383	-0.212	
$\bar{\eta}S_j\Delta\varepsilon_H$	14	0.592	-0.381	-0.113	-0.202	0.299	0.326	0.094	-0.331	0.117	0.850	-0.376	-0.260	0.978

Note: $\bar{\eta}\bar{S}\Delta\varepsilon_H$ and $\bar{\eta}S_j\Delta\varepsilon_H$ are computed, respectively, based on the estimates reported in columns 5 and 6 of Table 4.

Table 4. GMM estimates of the structural model and cross-city housing supply elasticity determinants

Column	1	2	3	4	5	6
Dependent variables	$\dot{N}_j$	$\dot{N}_j, \dot{W}_j$	$\dot{N}_j, \dot{P}_j$	$\dot{N}_j$, $\dot{W}_j$, $\dot{P}_j$	$\dot{N}_j$, $\dot{W}_j$, $\dot{P}_j$	$\dot{N}_j$, $\dot{W}_j$, $\dot{P}_j$
S_j estimate	$\overline{S}$	$\overline{S}$	$\overline{S}$	$\overline{S}$	$\overline{S}$	S_j
Structural coefficients in Eq. (22)						
$\overline{\eta}\rho\overline{\lambda}_{NH}$	0.169 (3.1)***	0.195 (3.8)***	0.142 (2.6)***	0.176 (3.6)***	0.180 (3.5)***	0.154 (4.5)***
$(1-\alpha+\gamma)\overline{\lambda}_{NA}$	0.345 (3.8)***	0.352 (4.1)***	0.386 (5.2)***	0.363 (5.1)***		
$\overline{\lambda}_{N\theta}$	0.360 (2.9)***	0.385 (3.3)***	0.393 (3.9)***	0.385 (4.1)***		
$b=(\gamma-\omega)/(1-\alpha+\gamma)$		0.337 (6.4)***	0.319 (8.5)***	0.325 (11)***	0.265 (51)***	0.264 (34)***
$\overline{\varepsilon}_H$					0.116 (3.5)***	0.117 (2.6)***
Supply elasticity differential, Eq. (21)	$(\overline{S}/\rho)\overline{\eta}\Delta\varepsilon_H$					$\overline{\eta}\Delta\varepsilon_H$
p_j	-0.255 (3.8)***	-0.280 (4.6)***	-0.228 (3.4)***	-0.218 (4.0)***	-0.163 (3.4)***	-0.015 (2.5)**
ln(*DENSITY*)	-0.450 (3.1)***	-0.509 (4.1)***	-0.344 (2.3)**	-0.367 (3.5)***	-0.309 (2.9)***	-0.032 (2.2)**
CAPTL	0.362 (2.2)**	0.354 (2.7)***	0.386 (1.9)*	0.294 (2.3)**	0.260 (2.3)**	0.035 (2.9)***
ln(*ROAD*)	0.277 (1.3)	0.226 (1.4)	0.448 (1.6)	0.323 (1.8)*	0.395 (2.0)**	0.063 (2.5)**
QR98>1.842	-0.143 (1.6)	-0.130 (2.2)**	-0.176 (1.5)	-0.177 (2.4)**	-0.171 (2.2)**	-0.019 (2.2)**
G_EFF	0.224 (2.8)***	0.218 (3.4)***	0.213 (2.3)**	0.168 (2.8)***	0.105 (2.2)**	0.011 (1.9)*
k_0	0.647	0.597	0.977	0.760	0.924	0.140
R squared	0.64	0.63, 0.46	0.64, 0.28	0.64, 0.45, 0.28	0.62, 0.46, 0.26	0.61, 0.46, 0.27

Note: The table reports estimates of Eq. (22) with $\Delta\varepsilon_H$ specified by Eq. (21). The GMM instruments include all the independent variables. Parameters $1-\beta$, $1-\mu$ and ε_P are set at 1/3, 0.4, and 0.5, respectively. In addition, $\dot{H}/\overline{\eta}\rho$ is set to zero and k_0 is chosen for mean $\Delta\varepsilon_H$ to be zero. S_j is the predicted value of $\dot{N}_j+\dot{W}_j+(1-\varepsilon_P)\dot{P}_j-\dot{Z}_H$ according to the structural estimates reported in column 5. The coefficient restrictions are given in Table 6. t-statistics based on cross-section White covariance are in parentheses; ***, **, and * denote, respectively, statistical significance at 1%, 5% and 10% levels. The number of observations is 68.

Table 5. Calculation of the structural coefficients

Structural coefficients in Eq. (22)	The coefficient restrictions used in estimating Eq. (22) in Table 4	Corresponding columns in Table 4	
		4	5
$b=(\gamma-\omega)/(1-\alpha+\gamma)$		**0.325**	**0.265**
$\bar{\eta}\rho\bar{\lambda}_{NH}$		**0.176**	**0.180**
$(1-\alpha+\gamma)\bar{\lambda}_{NA}$	$=\left((1-(1-\mu)(1-\beta))[\bar{\varepsilon}_H]+\varepsilon_P-(2-\mu)(1-\beta)\right)(1-\alpha+\lambda)\bar{\lambda}_{PA}$	**0.363**	0.363
$\bar{\lambda}_{N\theta}$	$=\left(3.26+12.76([b]-0.3)+3.02([\bar{\varepsilon}_H]-1)\right)\bar{\lambda}_{P\theta}$	**0.385**	0.385
$\bar{\eta}\rho\bar{\lambda}_{WH}$	$=-[b]\left[\bar{\eta}\rho\bar{\lambda}_{NH}\right]$	-0.057	-0.048
$(1-\alpha+\gamma)\bar{\lambda}_{WA}$	$=1-[b]\left[(1-\alpha+\gamma)\bar{\lambda}_{NA}\right]$	0.882	0.904
$\bar{\lambda}_{W\theta}$	$=-[b]\left[\bar{\lambda}_{N\theta}\right]$	-0.118	-0.096
$\bar{\eta}\rho\bar{\lambda}_{PH}$	$=-[b]\left[\bar{\eta}\rho\bar{\lambda}_{NH}\right]/(1-\beta)$	-0.172	-0.143
$(1-\alpha+\gamma)\bar{\lambda}_{PA}$	$=\left(1-[b]\left[(1-\alpha+\gamma)\bar{\lambda}_{NA}\right]\right)/(1-\beta)$	2.646	2.711
$\bar{\lambda}_{P\theta}$	$=\left(1-[b]\left[\bar{\lambda}_{N\theta}\right]\right)/(1-\beta)$	2.625	2.694
$\bar{\varepsilon}_H$			**0.116**
$\bar{\lambda}_{NA}/\bar{\lambda}_{PA}$		0.137	0.134
$\bar{\lambda}_{N\theta}/\bar{\lambda}_{P\theta}$		0.147	0.143
ρ	$=-\bar{\eta}\rho\bar{\lambda}_{WH}/\left(b(1-\alpha+\gamma)\bar{\lambda}_{WA}\right)=\left[\bar{\eta}\rho\bar{\lambda}_{NH}\right]/\left(1-[b]\left[(1-\alpha+\gamma)\bar{\lambda}_{NA}\right]\right)$	0.200	0.199

Note: The bold numbers are estimates reported in Table 4. The other numbers are calculated based on the bold numbers and the relevant coefficient restriction equations. These restrictions are based on Eq. (A1)~(A6), Eq. (14) and the linearized Eq. (15). The expressions in brackets represent input variables. The value of $1-\beta$, $1-\mu$ and ε_P are set at 1/3, 0.4, and 0.5, respectively. The expression for $\bar{\lambda}_{N\theta}/\bar{\lambda}_{P\theta}$ is derived from linearization of Eq. (15) around b = 0.3 and $\bar{\varepsilon}_H=1$.

Table 6. Contributions of the exogenous shocks to the cross-city variation of the urban growth measures

Panel A							Panel B					
Exogenous shocks (and their empirical measures)			$\dot{a}_j + \gamma \dot{Z}_A = 0.6PDS$	$\dot{\theta}_j - \dot{u} = AMS$	$\dot{Z}_H = \rho ZHS$	$\bar{\eta}\bar{S}\Delta\varepsilon_H$	Standard deviation of the component factors of the endogenous growth measures					
Standard deviation			0.050	0.096	0.049	0.043	$\bar{\lambda}_{XA}\left(\dot{a}_j + \gamma \dot{Z}_A\right)$	$\bar{\lambda}_{X\theta}\left(\dot{\theta}_j - \dot{u}\right)$	$\bar{\lambda}_{XA}\left(\dot{a}_j + \gamma \dot{Z}_A\right) + \bar{\lambda}_{X\theta}\left(\dot{\theta}_j - \dot{u}\right)$	$\bar{\lambda}_{XH}\bar{\eta}\dot{Z}_H$	$\bar{\lambda}_{XH}\bar{\eta}^2\bar{S}\Delta\varepsilon_H$	$\bar{\lambda}_{XH}\bar{\eta}(\dot{Z}_H + \bar{\eta}\bar{S}\Delta\varepsilon_H)$
		Standard deviation	Structural coefficients $\bar{\lambda}$									
Endogenous growth measures (X)	Population $\dot{N}_j$	0.091	0.605	0.385	0.904		0.030	0.037	0.025	0.044	0.039	0.064
	Wage rate $\dot{W}_j$	0.141	1.506	-0.096	-0.239		0.075	0.009	0.083	0.012	0.010	0.017
	Price $\dot{P}_j$	0.283	4.519	2.694	-0.718		0.226	0.260	0.168	0.035	0.031	0.051

Note: Calculation is based on the same sample of 65 cities as in Table 4. The structural coefficients and $\Delta\varepsilon_H$ are computed based on the estimates reported in column 5 of Table 4. $(1-\alpha+\gamma) = 0.6$ and $\rho = 0.2$.

Figure 1. 1998 urban population size and 1998-2004 population growth across 85 Chinese cities

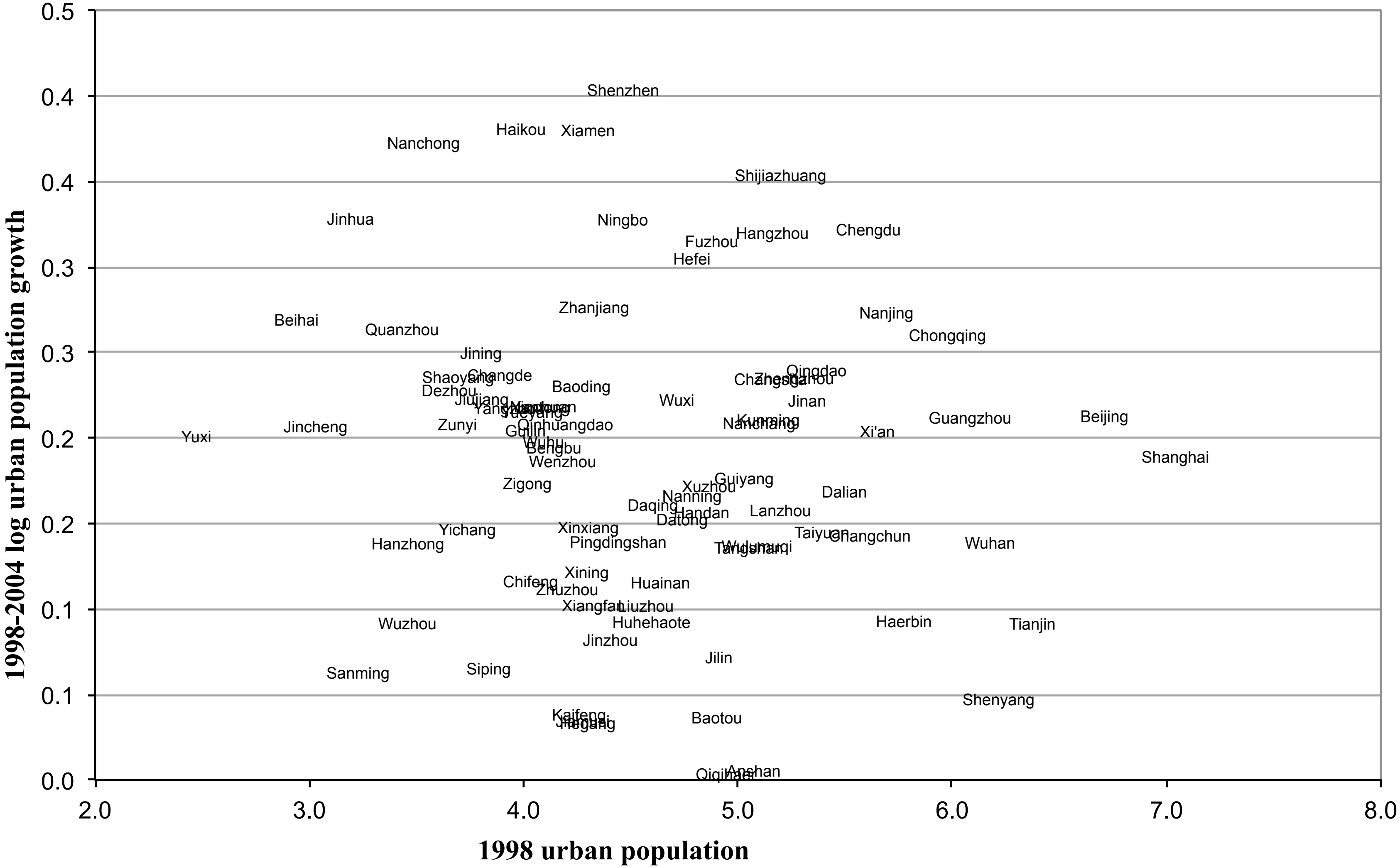

Figure 2. Housing price differential and price growth across 85 Chinese cities

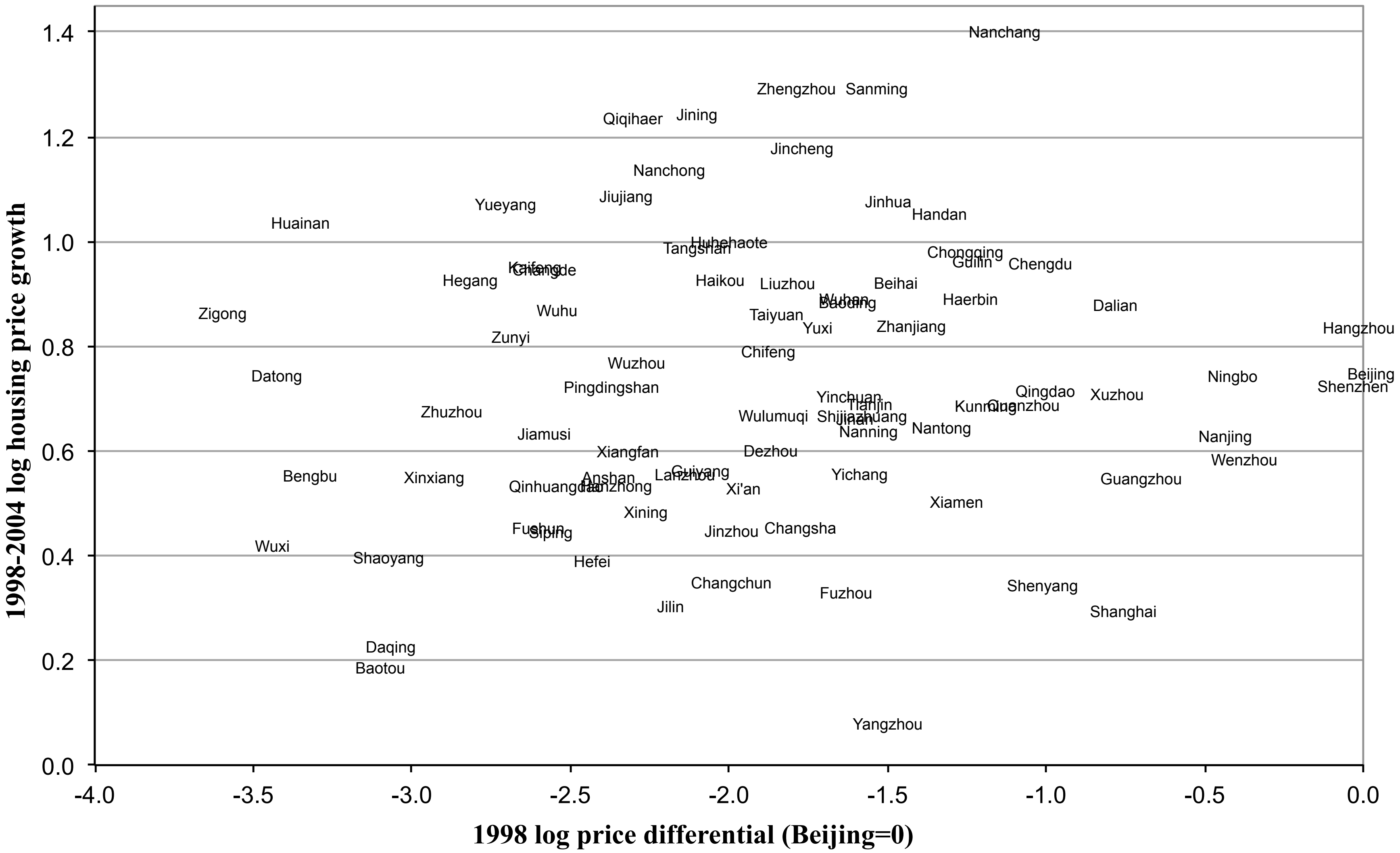

Figure 3. Population growth and housing price growth across 85 Chinese cities, 1998-2004