



Home-purchase restriction, property tax and housing price in China: A counterfactual analysis



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ABSTRACT

In this paper, we respectively evaluate the effects of home-purchase restrictions and the trial property taxes on housing prices in China using a counterfactual analysis. We modify the method of Hsiao, Ching and Wan (2012) by using the leave- n_v -out cross-validation criterion for the optimal choice of the control cities. Some Monte Carlo simulations illustrate the better performance of our method. We then construct the counterfactual growth rates of housing prices in Beijing, Shanghai and Chongqing using the selected control cities. We find that: (i) purchase restrictions reduced the annual growth rate of housing prices in Beijing by 7.69 percent; (ii) the trial property tax of Chongqing reduced the annual growth rate of housing prices by 2.52 percent; (iii) the trial property tax of Shanghai had no significant effect on housing prices.

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1. Introduction

To curb the soaring housing prices, the Chinese government has adopted a series of policies in recent years, among which, home-purchase restriction and property tax are two of the most important and influential ones. Home-purchase restriction was first started in Beijing in May 2010 and then progressively implemented in most major cities in China. It prohibits resident households from buying more than two homes and non-resident households from buying more than one home. The pilot programs of property tax were implemented in Shanghai and Chongqing in January 2011. The property tax in Shanghai targets the second and above homes purchased after January 28, 2011, while the property tax in Chongqing is mainly levied on high-end homes.

The purchase restriction and the property tax policies have aroused nationwide heated debates on: (i) whether purchase restrictions should be replaced by property taxes; (ii) which trial tax scheme is more effective in curbing housing prices, Shanghai scheme or Chongqing scheme.

Answering these questions involves estimating the effects of these policies separately, which, however, faces several difficulties. First, conventional policy evaluation methods like difference-in-difference require detailed micro-level data, which are hard to

obtain in general. Second, it is not an easy task to model how and why housing prices in a city have changed over time, as well as what factors have affected the changes. Third, there are many confounding policies going on, for example, the differentiated credit policy and the policy of indemnificatory housing etc.

Similar to Hsiao et al. (2012, hereafter HCW), we overcome the above issues by exploiting the dependence of housing prices among different cities and constructing the counterfactuals using data from cities without purchase restrictions and property tax. As housing prices of different cities are often driven by some common factors, information on other cities can help to predict housing prices in a particular city. Our real data analysis below (cf. Figs. 3, 5, and 7) confirms this point. Therefore, using housing prices of cities not subject to purchase restrictions or property tax, we estimate what would have happened to the growth rates of housing prices in Beijing, Shanghai and Chongqing if there had been no purchase restrictions or property tax. We then estimate the effects of purchase restrictions and property tax by the difference between the actual housing prices and the counterfactuals.

A subtle problem is the optimal choice of the control cities to construct the counterfactuals. As most major cities in China are subject to purchase restrictions, our candidate control cities for Beijing are mainly medium-sized ones. Since Chongqing is not subject to purchase restrictions,¹ we again use the medium-sized

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¹ There is a common misunderstanding that takes property tax in Chongqing for home-purchase restriction. We verified both from the municipal government of Chongqing and other sources, Chongqing has so far only adopted property tax.

cities to construct the counterfactuals for Chongqing's property tax so that the effects of the property tax are not contaminated by purchase restrictions. As Shanghai is subject to both policies, to separate the effect of property tax from that of the purchase restrictions we use cities with purchase restrictions implemented about the same time as Shanghai for the control group. A common problem for choosing the control cities is: on the one hand, we would like to select some cities that can fit the housing prices of Beijing, Chongqing and Shanghai well; on the other hand, too many control cities will also lower the model predictive ability. Our derivation below (cf. (11)) confirms these two points.

We propose to use the leave- n_v -out cross-validation ($CV(n_v)$) of Shao (1993) for the optimal choice of the control cities. We show that the $CV(n_v)$ objective function mimics the mean squared prediction error, and hence it is consistent in the sense that the probability of selecting the optimal model converges to 1 as $T \rightarrow \infty$. Moreover, the probability of selecting the model with the best predictive ability also converges to 1. Some Monte Carlo simulations show that in terms of mean squared prediction error $CV(n_v)$ performs better than AICC and AIC methods used by HCW, which supports our theoretical findings.

Our results show that from May 2010 to November 2011 the average annual growth rate of housing prices in Beijing would had been 16.97% if there had been no purchase restriction, while the actual annual growth rate during this period is 9.28%. Therefore home-purchase restrictions reduced the annual growth rate of housing prices in Beijing by 7.69 percent. We find that the trial property tax in Chongqing reduced the annual growth rate of housing prices by 2.52 percent from February 2011 to November 2012, while the trial property tax in Shanghai had no significant effect on the housing prices.

As a robustness check, we examine the micro-level data by China Household Finance Survey (CHFS) and find that 31% of households living in Beijing are restricted from home purchasing, while only 0.6 percent of households living in Shanghai are likely to be subject to the trial property tax. Based on the data released by Chongqing Municipal Finance Bureau, 10% of households living in Chongqing are subject to the trial property tax in 2011.

Therefore, based on our analysis, our answers to the heated debates are: (i) home-purchase restrictions are more effective than property taxes, and it cannot be replaced by the property tax at least in the short-run, if the government is still willing to curb the housing prices; (ii) the property tax scheme of Chongqing is more effective than the one of Shanghai.

Finally, we relate our studies to some existing literature. Studying the effects of home-purchase restrictions on housing prices has not been established in the existing literature because home-purchase restrictions in such scale are not very common.² It is indirectly related to the literature of rationing. Howard (1977) provides a theoretical model to study how rationing affects consumption. Mackay and Whitney (1980) decompose the total effect of quantity constraint into substitution effect and income effect. Empirical studies on credit rationing include Berger and Udell (1992) and Steijvers and Voordeckers (2009) etc. It is important to note that credit rationing comes from the supply of credit, while home-purchase restrictions are imposed on the demand side.

There are many studies on the impact of property tax on housing prices. On the one hand, imposing property tax, holding public expenditures constant, might be capitalized into house values. The prices of houses fall by the present value because of the projected increase of holding costs, see e.g. Simon (1943),

Netzer (1966), van den Noord (2005) and Cebula (2009). On the other hand, increases in public expenditures, holding property tax constant, might have offsetting effects on housing prices because of the improved local public services quality, see e.g. Oates (1969), Hamilton (1976), Rosen and Fullerton (1977), Fische (1992) and Lang and Jian (2004). It is important to note that the trial property taxes in Chongqing and Shanghai are less likely to affect the quality of local public services because they have not been an important source of local tax revenue³ and the property tax revenue has been mainly used to finance the supply of cheap rental housing for low-income urban households. Recently, Bai et al. (2014) establish the consistency result of the HCW method for nonstationary data, and hence make the method a valuable tool for studying macroeconomic policies. They then apply the method to property tax policies without considering the home-purchase restrictions.

The rest of the paper is organized as follows. Section 2 introduces the institutional background of Chinese housing market and a simple theoretical model to study the effects of home-purchase restrictions and property tax on housing prices. Section 3 presents the econometric method and some Monte Carlo simulation results. Section 4 presents the empirical analysis. In Section 5 we conclude. The proofs are given in the Appendix.

2. Institutional background and a theoretical model

This section first introduces the institutional background of Chinese housing market. Then, we build a simple theoretical model to analyze the effects of purchase restrictions and property tax on housing prices, respectively.

2.1. Institutional background

Chinese housing market experienced four stages of development: nonexistence, early development, expansion, and policy interventions to control the soaring housing prices.

Between the founding of the People's Republic of China in 1949 and 1988, Chinese housing market was non-existent. In the urban areas, housing, as part of social welfare, was provided by government institutions or state-owned enterprises (*work units*). In the rural areas, households live in houses that were either built by themselves on the land assigned or provided directly by the village. During this period, there was no market for households to buy or sell a home.

A constitutional amendment in 1988 provided a legal foundation for private housing development. The government still retains the ownership of urban lands. However, home buyers acquire the right of land use for 70 years. Though residents were encouraged to purchase their residential homes from their *work units* at discounted prices, housing was still allocated as a welfare good rather than a commodity traded on the market. As a result, the secondary market for housing was not very liquid in the first decade after the reform.

The issue of the 23rd Decree by the State Council in 1998 marked the beginning of the modern Chinese housing market. *Work units* were not allowed to build or buy residential units for their employees. Instead, *work units* had to integrate housing benefits into their employees' salary to help them buy homes on the market. As a strategic move following the 1997 Asian financial crisis, the reform was designed to promote the commercial

² A few countries such as Australia, Brazil, and Argentina recently imposed home-purchase restrictions on foreign individuals and companies, but the scale is much smaller than the one in China.

³ Based on data released by Chongqing Municipal Finance Bureau, the total trial property tax revenue in 2011 is 100 million RMB, which is only 0.12 percent of the local tax revenue. Based on China Real Estate Index System (CREIS) data, the estimated trial property tax revenue in Shanghai is about 300 million RMB in 2011, which is only 0.1 percent of the local tax revenue.

development of housing market and housing industry was chosen as the new growth engine. Chinese housing market was developed at an unprecedented pace in the first decade after the reform. Based on the Chinese Statistical Yearbook from National Bureau of Statistics (NBS), residential property investment in GDP has tripled from 2 percent in 2000 to 6 percent in 2011. In terms of the quantity of commercial housing sold, the annual amount increased almost 9-fold, from about 108 million square meters in 1998 to 934 million square meters in 2010.

The expansion of Chinese housing market is associated with soaring housing prices and falling housing affordability. Housing Price Index of China from NBS shows that housing prices increased by 78 percent from 1998 to 2010. To have a general picture on urban housing prices, we use Chinese Household Finance Survey data, conducted by Southwestern University of Finance and Economics and People's Bank of China. This survey data covers 8438 households randomly selected from 320 communities among 2585 counties. The data reveals that urban housing prices, based on the initial purchasing costs and home owners' estimates of current home values, are tripled from 1998 to 2010. The rising cost of buying a home becomes a major concern both for the public and the government. According to Wei and Allen (2010), the ratio of median housing prices to median monthly household income is 89 in the urban areas in 2010, while it is 50 in the US even before the 2008–2009 financial crisis. The ratio of the housing price to the monthly rental of a comparable house in coastal cities in China is 500 in 2010, while in major US cities it is on the order of 250 in 2010.

To curb the soaring housing prices, the government adopted a series of policies,⁴ among which home-purchase restrictions, first implemented in Beijing in 2010 and then progressively in more than 40 cities, are considered as the most strict housing policy in China. Despite of some minor differences of purchase restrictions across cities, in general purchase restrictions prohibit resident households (i.e. with registered permanent residence or *Hukou*) from buying more than two units of residential premises and non-resident households (i.e. without *Hukou*) from buying more than one unit of residential premises with proof of local tax receipts or social security records of 1 year. As a result, a fraction of households is restricted from home purchasing.

As home-purchase restrictions cannot be a long-run regulation, the government started the trial property taxes in Shanghai and Chongqing in 2011, which are expected to replace the purchase restriction order. The trial property taxes in Shanghai and Chongqing represent two different schemes of taxation. In Shanghai, property taxes are levied on the second and above units of residential premises (including existing second-hand residential premises and newly-built commercial residential premises) newly purchased by Shanghai-resident families, and the first and above units of residential premises newly purchased by non-resident families. For taxable households, 60 square meters per family member are exempt from taxes. The annual tax rate is 0.6 percent, and the tax rate is 0.4 percent for those homes prices lower than twice the average prices of the previous year's newly-built commercial residential premises.

In Chongqing, property taxes are levied on both old and newly bought independent houses, newly bought high-end residential premises, and the second and above units of ordinary homes (a home less than 140 square meters, around the average value of local comparable homes) newly purchased by non-resident families. For taxable households, 180 square meters are exempt

from taxes for independent houses and 100 square meters are exempt from taxes for high-end residential premises. The annual tax rate is 0.5 percent for those high-end homes with prices lower than 15,000 RMB per square meter and for the second or more unit of ordinary homes bought by non-resident households. The tax rate is 1 to 1.2 percent for those homes with prices higher than 15,000 RMB per square meter. As a result, the trial property taxes in both Shanghai and Chongqing are only applicable to a fraction of residential premises.

2.2. A theoretical model

This subsection presents a model to analyze the effects of purchase restrictions and property tax on housing prices, respectively. Considering the fact that the trial property taxes do not affect the quality of public goods in Shanghai and Chongqing, we focus our attention on the partial equilibrium holding public expenditures constant.

Consider a continuum of households indexed in $[0, 1]$. A household faces a stochastic income stream and chooses consumption, financial assets, and house holdings to maximize expected lifetime utility subject to a budget constraint. The problem of the household is

$$\max_{\{C_t, H_t, B_t\}} E_0 \sum_{t=0}^{\infty} \beta^t [\ln(C_t) + \eta \ln(H_t)]$$

Subject to $B_t = r_t B_{t-1} + Y_t - C_t - q_t(H_t - H_{t-1})$.

Here E_0 is the expectation operator based on period 0 information, $\beta < 1$ represents discount factor, C_t is consumption, H_t is consumption of the service flow provided by houses, and η is utility parameter related to housing service. In the budget constraint, B_t is financial asset at the end of period t , r_t is the gross real interest rate, Y_t is labor income during period t , and q_t is the price of houses. The prices of consumption goods is unit 1 so that q_t is also the relative price of houses in terms of consumption goods.

The first-order conditions for the problem are

$$\frac{1}{C_t} = \beta E_t \frac{r_{t+1}}{C_{t+1}} \quad (1)$$

$$\frac{\eta}{H_t} = q_t \left[\frac{1}{C_t} - \beta E_t \frac{\pi_{t+1}}{C_{t+1}} \right] \quad (2)$$

$$B_t = r_t B_{t-1} + Y_t - C_t - q_t(H_t - H_{t-1}). \quad (3)$$

Here, the first efficiency condition indicates that the marginal utility cost of consumption in period t must equal the present value of the expected marginal utility benefit of consumption. The second condition specifies the equality of the rates of intratemporal/intertemporal substitution between consumption and houses and their relative prices. In Eq. (2), $\pi_{t+1} = q_{t+1}/q_t$ is the gross inflation rate of the relative price of houses. The third condition is the binding budget constraint.

In order to derive the implications of home purchase restriction and property tax without specifying a general equilibrium model, we assume that r_{t+1} is known at time t and the supply of house is a constant $\bar{H}$. In addition, in equilibrium $B_t = r_t B_{t-1}$. On the steady state, the aggregate demand for houses is

$$H = \frac{\eta Y}{q(1 - \beta)}. \quad (4)$$

In equilibrium, the housing price is

$$q = \frac{\eta Y}{\bar{H}(1 - \beta)}. \quad (5)$$

⁴ See Deng et al. (2011) and Wu et al. (2012) for other policies like expanding the supply of cheap rental housing for low-income urban households, differentiated credit and raising down-payment.

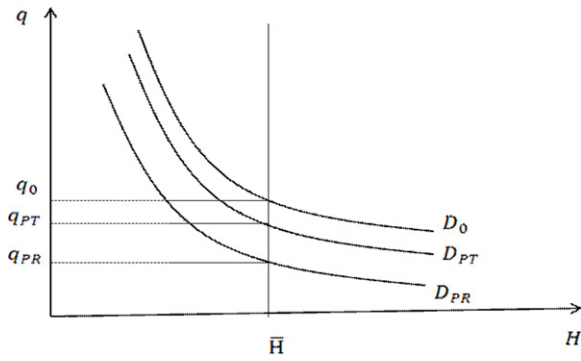


Fig. 1. The effects of home purchase restriction and property tax.

We can now analyze the effects of home-purchase restrictions and property tax on housing prices, respectively, using Eq. (5) as a benchmark.

Suppose that a fraction α of households is restricted to purchase more than 2 units of homes. The aggregate demand becomes $(1 - \alpha)H$. In comparison to the initial equilibrium, purchase restrictions reduce the aggregate demand by shifting the demand curve to the left as in Fig. 1.

At the new equilibrium, the housing price is q_{PR} , which is lower than the initial price q_0 . Therefore, purchase restrictions lower housing prices by reducing the aggregate demand for houses. This result is also valid when the supply curve slopes upward with lower elasticity than the demand.

As for property tax, only a fraction γ of houses is taxed with a tax rate τ . The additional holding cost is τqH , and the impact on the aggregate demand is $\gamma \tau qH$. This is reflected by shifting slightly the initial aggregate demand curve D_0 to the left to D_{PT} . At the new equilibrium, the housing price q_{PT} is lower than the initial prices. Therefore, property tax lowers housing prices by reducing the aggregate demand. However, the magnitude of the effect depends on the proportion of taxable houses.

3. Empirical methodology and Monte Carlo simulations

In this section, we modify the method of HCW by using the leave- n_v -out cross-validation ($CV(n_v)$) for the optimal choice of the control cities. We then conduct some Monte Carlo simulations to study the performance of the modified procedure.

3.1. Empirical methodology

We illustrate our empirical method to evaluate the policy effects using the home-purchase restriction order in Beijing. The effects of property tax can be analyzed similarly.

Let g_{it}^0 denote the growth rate of housing prices of city i at time t without purchase restrictions. Without loss of generality, we let $i = 1$ correspond to Beijing. Following Hsiao et al. (2012), we assume that there exist K (unobservable) common factors $\mathbf{f}_t$ that drive home prices of all cities to change over time. These factors can be national GDP growth rate, macro policies, interest rate and changes in public expectations etc. Therefore we have the following factor model

$$g_{it}^0 = \mathbf{b}_i' \mathbf{f}_t + \alpha_i + \varepsilon_{it}, \quad i = 1, \dots, N, \quad t = 1, \dots, T, \quad (6)$$

where $\mathbf{b}_i$ denotes the $K \times 1$ vector of factor loading for city i , α_i is the fixed city-specific effect, and ε_{it} is the idiosyncratic error term with $E(\varepsilon_{it}) = 0$. Stacking $N \times 1$ g_{it}^0 into a vector yields

$$\mathbf{g}_t^0 = \mathbf{B} \mathbf{f}_t + \boldsymbol{\alpha} + \boldsymbol{\varepsilon}_t, \quad (7)$$

where $\mathbf{g}_t^0 = (g_{1t}^0, \dots, g_{Nt}^0)'$, $\boldsymbol{\alpha} = (\alpha_1, \dots, \alpha_N)'$, $\boldsymbol{\varepsilon}_t = (\varepsilon_{1t}, \dots, \varepsilon_{Nt})'$, and $\mathbf{B} = (\mathbf{b}_1, \dots, \mathbf{b}_N)'$ is the $N \times K$ factor loading matrix.

Let g_{it}^1 denote the growth rate of housing prices of city i at time t under purchase restrictions. Often we do not simultaneously observe g_{it}^0 and g_{it}^1 . The observed data

$$g_{it} = d_{it} g_{it}^1 + (1 - d_{it}) g_{it}^0,$$

where $d_{it} = 1$ if city i is under purchase restrictions at time t , and $d_{it} = 0$ otherwise.

At time $T_1 + 1$, here May 2010, the purchase restriction order took effect in Beijing. Therefore,

$$g_{1t} = g_{1t}^1 \quad \text{for } t = T_1 + 1, \dots, T.$$

For cities not subject to the purchase restriction we have

$$g_{it} = g_{it}^0 \quad \text{for } i = 2, \dots, N, \quad t = 1, \dots, T. \quad (8)$$

The effect of purchase restrictions in Beijing at time t will be

$$\Delta_{1t} = g_{1t}^1 - g_{1t}^0,$$

if we can predict g_{1t}^0 , $t = T_1 + 1, \dots, T$. If both N and T are large, we can use the procedure of Bai and Ng (2002) to identify $\mathbf{f}_t$ and $\mathbf{b}_1$ in Eq. (6), and hence predict g_{1t}^0 . However, neither N nor T is large in our data.

Instead, we adopt a procedure similar to HCW to construct the counterfactual g_{1t}^0 , for which we need the following two assumptions:

Assumption A1. $\{\boldsymbol{\varepsilon}_t\}$ is a stationary and ergodic process with $E[\boldsymbol{\varepsilon}_t] = 0$, $E[\boldsymbol{\varepsilon}_t \mathbf{f}_t'] = 0$ and $E[\boldsymbol{\varepsilon}_t \boldsymbol{\varepsilon}_t'] = \mathbf{V}$, where $\mathbf{V}$ is a diagonal constant matrix.

Assumption A2. $\text{Rank}(\mathbf{B}) = K$.

Assumption A1 is standard in the literature. Assumption A2 implies that the number of observable cross-sectional units, N , is greater than the number of common time-varying factors $\mathbf{f}_t$. With these assumptions and Eq. (7) one can derive that⁵

$$g_{1t}^0 = \bar{\alpha} + \mathbf{a}_{-1}' \mathbf{g}_{-1t} + \varepsilon_{1t}^*, \quad (9)$$

where henceforth

$$\mathbf{g}_{-1t} = (g_{2t}, \dots, g_{Nt})',$$

$\bar{\alpha}$ and $\mathbf{a}_{-1}$ are constants, and ε_{1t}^* is some error term uncorrelated with $\mathbf{g}_{-1t}$.

Eq. (9) suggests using $\mathbf{g}_{-1t}$ in lieu of $\mathbf{f}_t$ to predict g_{1t}^0 . One problem is the choice of the control cities $\mathbf{g}_{-1t}$. As discussed in HCW, using all the available cities is generally not the optimal choice. HCW propose a two-step procedure to select the optimal number of cities, which is essentially minimizing the Akaike information criterion (AIC, Akaike, 1973, 1974) or the corrected Akaike information criterion (AICC, Hurvich and Tsai, 1989). However, as shown in Shao (1993, 1996), AIC and AICC are 'asymptotically inconsistent in the sense that the probability of selecting the model with the best predictive ability does not converge to 1 as the total number of observations $n \rightarrow \infty$ ' and are 'inconsistent in the sense that the probability of selecting the optimal subset of variables does not converge to 1 as $n \rightarrow \infty$ '.

⁵ To see this, let $\mathbf{a}$ be a vector in the null space of $\mathbf{B}$, i.e. $\mathbf{a}'\mathbf{B} = 0$. We can always normalize the first element of $\mathbf{a}$ to be 1 and denote $\mathbf{a} = (1, -\tilde{\mathbf{a}}_{-1}')$. From (7) and (8) we then get $g_{1t}^0 = \bar{\alpha} + \tilde{\mathbf{a}}_{-1}' \mathbf{g}_{-1t} + \varepsilon_{1t} - \tilde{\mathbf{a}}_{-1}' \boldsymbol{\varepsilon}_{-1t}$, where $\bar{\alpha} = \mathbf{a}'\boldsymbol{\alpha}$, and $\boldsymbol{\varepsilon}_{-1t} = (\varepsilon_{2t}, \dots, \varepsilon_{Nt})'$. As $\boldsymbol{\varepsilon}_{-1t}$ is correlated with $\mathbf{g}_{-1t}$, the linear projection of $\boldsymbol{\varepsilon}_{-1t}$ on $\mathbf{g}_{-1t}$ gives us Eq. (9), with $\mathbf{a}_{-1} = \tilde{\mathbf{a}}_{-1}' (\mathbf{I}_{N-1} - \text{cov}(\boldsymbol{\varepsilon}_{-1t}, \mathbf{g}_{-1t}) \text{var}(\mathbf{g}_{-1t})^{-1})$, and $\varepsilon_{1t}^* = \mathbf{a}' \boldsymbol{\varepsilon}_t + \tilde{\mathbf{a}}_{-1}' \text{cov}(\boldsymbol{\varepsilon}_{-1t}, \mathbf{g}_{-1t}) \text{var}(\mathbf{g}_{-1t})^{-1} (\mathbf{g}_{-1t} - E(\mathbf{g}_{-1t}))$.

Shao (1993) proposes a consistent leave- n_v -out cross-validation ($CV(n_v)$) method that overcomes the deficiency of AIC. Let γ be a subset of $\{2, \dots, N\}$ with d_γ elements, $1 \leq d_\gamma \leq N - 1$, which corresponds to the choice of the control cities indexed by γ . Let $\mathbf{g}_{\gamma t}$ (and $\mathbf{a}_\gamma$) denote the components of $\mathbf{g}_{-1t}$ (and $\mathbf{a}_{-1}$) that are indexed by γ . Shao's method works as follows: First split the data $\{(g_{1t}, \mathbf{g}'_{\gamma t})\}_{t=1}^{T_1}$ into two parts $\{(g_{1s}, \mathbf{g}'_{\gamma s}), s \in \tau\}$ and $\{(g_{1s}, \mathbf{g}'_{\gamma s}), s \in \tau^c\}$, where τ is a subset of $\{1, 2, \dots, T_1\}$ with n_v elements and τ^c is its complement with $n_c = T_1 - n_v$ elements; Regress g_{1s} on constant and $\mathbf{g}_{\gamma s}$ using the data indexed by τ^c , and get the OLS coefficients $\hat{\alpha}_{\gamma, \tau^c}$ and $\hat{\mathbf{a}}_{\gamma, \tau^c}$; Calculate the average squared prediction error using the data indexed by τ , $n_v^{-1} \sum_{s \in \tau} (g_{1s} - \hat{\alpha}_{\gamma, \tau^c} - \hat{\mathbf{a}}'_{\gamma, \tau^c} \mathbf{g}_{\gamma s})^2$. As there are $\binom{T_1}{n_v}$ ways to divide $\{1, 2, \dots, T_1\}$ into τ and τ^c , Shao randomly draws M of them, $\{\tau_j, \tau_j^c\}_{j=1}^M$, and calculates the average squared prediction error as

$$\hat{\Gamma}_\gamma = \frac{1}{n_v M} \sum_{j=1}^M \sum_{s \in \tau_j} (g_{1s} - \hat{\alpha}_{\gamma, \tau_j^c} - \hat{\mathbf{a}}'_{\gamma, \tau_j^c} \mathbf{g}_{\gamma s})^2. \quad (10)$$

The optimal choice of the control cities is given by the γ that minimizes $\hat{\Gamma}_\gamma$.

We modify the procedure of HCW by using $CV(n_v)$ method for the choice of the control cities. Our procedure for evaluating the effects of the purchase restriction is as follows:

Step 1. From all the possible subsets of $\{2, \dots, N\}$, choose γ^* that minimizes $\hat{\Gamma}_{\gamma^*}$ defined in (10). Following Shao (1993), we choose n_c to be the integer part of $T_1^{3/4}$ and $M = 2T_1$.

Step 2. Regress g_{1t} on constant and $\mathbf{g}_{\gamma^* t}$ using data before the purchase restriction, i.e. $t \leq T_1$, and get the coefficients estimates $\hat{\alpha}_{\gamma^*}$ and $\hat{\mathbf{a}}_{\gamma^*}$.

Step 3. Construct the counterfactuals $\hat{g}_{1t}^0 = \hat{\alpha}_{\gamma^*} + \hat{\mathbf{a}}_{\gamma^*}' \mathbf{g}_{\gamma^* t}$ for $t = T_1 + 1, \dots, T$. A prediction for the effect of the purchase restriction at time t will then be the difference between the actual housing price growth rate and the counterfactual housing price growth rate, i.e. $\Delta_{1t} = g_{1t}^1 - \hat{g}_{1t}^0$, for $t > T_1$.

A model defined by $g_{1t}^0 = \alpha_\gamma + \mathbf{a}_\gamma' \mathbf{g}_{\gamma t} + e_{1t}$ generally falls into one of the following two categories:

Category I. At least one nonzero component of $\mathbf{a}_{-1}$ is not in $\mathbf{a}_\gamma$.

Category II. $\mathbf{a}_\gamma$ contains all nonzero components of $\mathbf{a}_{-1}$.

The models in Category I are incorrect models, and the models in Category II may be inefficient because of their unnecessarily large sizes. The optimal model, denoted by $\mathcal{M}^*$, is the model in Category II with the smallest dimension.

The next theorem shows that our procedure above is consistent in the sense that the probability of selecting the optimal model converges to 1 as $T \rightarrow \infty$. Moreover, the probability of selecting the model with the best predictive ability also converges to 1. We need the following additional assumptions.

Assumption A3. $E[\varepsilon_{jt}|d_{it}] = 0$, for $j \neq i$.

Assumption A4. There exists an $\mathbf{a} = (1, -\tilde{\mathbf{a}}_{-1})'$ in the null space of B such that $E(g_{1t}^0 - \bar{\alpha} - \mathbf{a}'_{-1} \mathbf{g}_{-1t})^2$ has a unique minimum in the neighborhood of $\mathbf{a}$.

Assumption A5. When both T_1 and $T_2 = T - T_1$ are large,

$$\begin{aligned} \frac{1}{T_1} \sum_{t=1}^{T_1} \mathbf{g}_{-1t} \mathbf{g}'_{-1t} &= O_p(1), \quad \text{and} \\ \frac{1}{T_1} \sum_{t=1}^{T_1} \mathbf{g}_{-1t} \mathbf{g}'_{-1t} - \frac{1}{T_2} \sum_{t=T_1+1}^T \mathbf{g}_{-1t} \mathbf{g}'_{-1t} &= o_p(1), \end{aligned}$$

as $T_1 \rightarrow \infty$ and $T_2 \rightarrow \infty$.

Assumption A6.

$$\max_{1 \leq j \leq M} \left\| \frac{1}{n_v} \sum_{s \in \tau_j} \mathbf{g}_{-1s} \mathbf{g}'_{-1s} - \frac{1}{n_c} \sum_{s \in \tau_j^c} \mathbf{g}_{-1s} \mathbf{g}'_{-1s} \right\| = o_p(1).$$

Assumption A3 says j th unit's idiosyncratic components are conditionally independent of d_{it} , for $j \neq i$. A4 is an identification assumption. Bai et al. (2014) prove the existence of $\mathbf{a}$ for nonstationary data. A1–A4 are also assumed in HCW; see Li and Bell (2012) for a discussion on these assumptions. Assumptions A5 and A6 are satisfied by some stationary and ergodic processes with finite second moments.

Theorem 1. Under Assumptions A1–A6, we have

- (1) $\lim_{T \rightarrow \infty} P(\text{the selected model is in Category I}) = 0$.
- (2) $\lim_{T \rightarrow \infty} P(\text{the selected model is } \mathcal{M}^*) = 1$.

We then derive the asymptotic properties of the estimated treatment effects $\hat{\Delta}_{1t}$ and $T_2^{-1} \sum_{t=T_1+1}^T \hat{\Delta}_{1t}$ in the next proposition.

Proposition 1. Under Assumptions A1–A6, we have

$$\hat{\Delta}_{1t} = \Delta_{1t} + \varepsilon_{1t}^* + O_p(T_1^{-1/2})$$

for $t = T_1 + 1, \dots, T$, and

$$\frac{1}{T_2} \sum_{t=T_1+1}^T \hat{\Delta}_{1t} - \frac{1}{T_2} \sum_{t=T_1+1}^T \Delta_{1t} = o_p(1),$$

as $T_2 \rightarrow \infty$.

3.2. Monte Carlo simulations

In this subsection we compare the performance of our method with that of HCW through some Monte Carlo studies. We use exactly the same setup as theirs.

We generate model (6) with $N = 21$ cities, $T_1 = 25, 40$ and 60 pre-intervention periods, and $T = T_1 + 10$. We consider the same 1-factor, 2-factor and 3-factor structures as HCW:

1-factor:

$$f_{1t} = 0.95f_{1t-1} + u_{1t},$$

2-factor:

$$f_{1t} = 0.3f_{1t-1} + u_{1t},$$

$$f_{2t} = 0.6f_{2t-1} + u_{2t},$$

3-factor:

$$f_{1t} = 0.8f_{1t-1} + u_{1t},$$

$$f_{2t} = -0.6f_{2t-1} + u_{2t} + 0.8u_{2t-1},$$

$$f_{3t} = u_{3t} + 0.9u_{3t-1} + 0.4u_{3t-2},$$

where u_{it} are generated by $N(0, 1)$, $i = 1, 2, 3$. $\{\varepsilon_{it}\}$ in model (6) are generated by $\sigma^2 N(0, 1)$; b_i is generated by $N(1, 1)$. We try $\sigma^2 = 1, 0.5$ and 0.1 in our simulations. We then use our method described above and HCW's method, respectively, to select the best model to predict g_{it}^0 for the post-intervention period, which has $T_2 = 10$ periods.

We compare the two methods using the post-intervention mean square prediction error

$$PMSE = \frac{1}{T_2} \sum_{t=T_1+1}^T (g_{1t}^0 - \hat{g}_{1t}^0)^2,$$

Table 1

Comparison of different model selection methods.

	$\sigma^2 = 1$			$\sigma^2 = 0.5$			$\sigma^2 = 0.1$		
	CV(n_v)	AICC	AIC	CV(n_v)	AICC	AIC	CV(n_v)	AICC	AIC
<i>1-factor model, $T_1 = 25, T = 35$</i>									
Avg.#	4.83	3.56	10.92	5.60	3.63	11.14	5.31	3.66	10.95
Avg.PMSE	1.8650	2.0967	6.0158	0.9321	1.0540	3.1426	0.2179	0.2101	0.7461
<i>1-factor model, $T_1 = 40, T = 50$</i>									
Avg.#	5.34	3.71	5.91	5.50	3.75	5.94	5.41	3.76	5.97
Avg.PMSE	1.5685	1.7310	1.8935	0.8096	0.8635	0.9543	0.1521	0.1728	0.1915
<i>1-factor model, $T_1 = 60, T = 70$</i>									
Avg.#	5.89	4.14	5.60	5.75	4.20	5.37	6.27	4.19	5.39
Avg.PMSE	1.3537	1.4180	1.4722	0.6890	0.7258	0.7473	0.1455	0.1453	0.1496
<i>2-factor model, $T_1 = 25, T = 35$</i>									
Avg.#	5.05	4.16	11.09	5.12	4.11	10.97	5.43	4.24	11.15
Avg.PMSE	2.1096	2.4358	6.2314	1.0465	1.1658	2.7615	0.1921	0.2372	0.6337
<i>2-factor model, $T_1 = 40, T = 50$</i>									
Avg.#	5.79	3.94	6.25	5.88	4.05	6.28	5.99	3.98	6.35
Avg.PMSE	1.6657	1.8450	1.9634	0.9196	0.9380	1.0153	0.1823	0.1865	0.1970
<i>2-factor model, $T_1 = 60, T = 70$</i>									
Avg.#	6.08	4.35	5.58	6.23	4.50	5.78	6.81	4.54	5.83
Avg.PMSE	1.5310	1.6194	1.6502	0.7809	0.7904	0.7986	0.1483	0.1578	0.1605
<i>3-factor model, $T_1 = 25, T = 35$</i>									
Avg.#	5.10	4.40	11.18	5.36	4.54	11.70	5.48	4.74	11.72
Avg.PMSE	2.5427	2.8493	6.5231	1.3047	1.4640	3.2716	0.2768	0.3051	0.7024
<i>3-factor model, $T_1 = 40, T = 50$</i>									
Avg.#	6.46	4.90	7.10	6.06	5.08	7.22	6.40	5.04	7.24
Avg.PMSE	1.9202	1.9705	2.1136	0.9371	0.9716	1.0804	0.1958	0.1987	0.2095
<i>3-factor model, $T_1 = 60, T = 70$</i>									
Avg.#	6.72	5.12	6.34	6.80	5.18	6.36	7.00	5.24	6.42
Avg.PMSE	1.6658	1.6437	1.6842	0.8331	0.8302	0.8516	0.1613	0.1677	0.1685

where $\hat{g}_{1t}^0$ is generated using HCW's and our method, respectively, for $t = T_1 + 1, \dots, T$. We repeat the experiment 100 times.⁶ The average PMSE is reported in Table 1. We also report the average number of selected control cities.

Table 1 shows that our method generally has smaller mean square prediction error than HCW except in several occasions, which supports our theoretical findings above.

4. Empirical analysis

In this section, we evaluate respectively the effects of home-purchase restrictions on housing prices in Beijing and the effects of the trial property taxes on housing prices in Shanghai and Chongqing. For each policy, we first find a proper control group and experimental period, and we then select cities to construct the counterfactuals and investigate the policy effect.

4.1. Home-purchase restrictions in Beijing

The purchase restriction policy in Beijing entered into effect on April 30, 2010. Resident households were prohibited from buying more than two units of residential premises and non-resident households can buy at most one unit of residential premises with proof of local tax receipts or social security records of one year.

The policy evaluation period is fixed from May 2010 to November 2011. The timeline of home-purchase restrictions across cities are given in Fig. 2. Starting with Beijing, most major cities have gradually adopted purchase restrictions. We exclude

the cities that adopted purchase restrictions during the policy evaluation period from the control group. On December 10, 2011, the municipality of Beijing relaxed the standards of ordinary housing so that more housing transactions are exempt from business tax, accounting for 2 to 5 percent of the sale prices. This important policy change is not implemented nationwide. Thus, to isolate the effects of purchase restrictions, we choose the period between May 2010 and November 2011 for policy evaluation.

We choose cities without purchase restrictions as the control group. Specifically, we include Tangshan, Qinhuangdao, Baotou, Jinzhou, Jilin, Yangzhou, Bengbu, Anqing, Quanzhou, Jiujiang, Ganzhou, Yantai, Jining, Luoyang, Pingdingshan, Yichang, Xiangfan, Yueyang, Changde, Guilin and Beihai into the control group. We select these cities because they are more likely to share common factors with Beijing. They are either major cities in China or geographically close to Beijing.

We use the monthly growth rates of Newly-Built House Price Indexes (NBHPI) of these 21 cities to construct the counterfactuals of Beijing in the absence of purchase restrictions. All the housing price indexes are from National Bureau of Statistics of China. It is important to note that, contrary to other developed countries, a striking feature of the Chinese urban housing market is that most housing transactions (about 70%) involve newly built homes.

There are three alternative data of housing prices in China. One is Newly-Built Commercial House Price Indexes for 70 cities, reported by NBS. However, NBS stopped reporting this data between 2008 and 2010. The second alternative data is Sale Price Indexes of Residential Property of 70 cities from NBS, which was not reported from 2011. As data during policy experimental periods are not reported, these two price indexes are not appropriate for our studies. The third alternative data is from the National Development and Reform Commission (NDRC) of China that reports monthly average per square meter prices of commercial residential properties across 27 provinces and four municipalities nationwide. However, except for Chongqing, all provincial capital cities have adopted

⁶ Exhausting all the possible subsets of $\{2, \dots, N\}$ is computationally heavy, and the cross validation with $M = 2T_1$ makes it even more burdensome. We expect similar results with more repetitions.



Fig. 2. Home-purchase restriction timeline across cities in China.

home-purchase restrictions at the time of this study. Thus, this data is not appropriate for studying the effects of home-purchase restrictions. Neither is provincial-level data from NDRC appropriate for our approach to evaluate the effects of property taxes because we have to use cities with home-purchase restrictions to estimate the counterfactual prices of Shanghai (adopted home-purchase restrictions in October 2010) in the absence of property tax and use cities without purchase restrictions to estimate the counterfactual prices of Chongqing that has not adopted purchase restrictions at the time of this study. As a result, we choose NBHPI, which have been consistently reported and allow us to study the effects of the two policies.

We calculate the monthly growth rates on a year-on-year basis, i.e. we measure the NBHPI change compared with the corresponding month in the previous year. We do not use the sequential growth rates, i.e. measuring the change since the previous month, because the sequential growth rates of some cities have strong periodicity even after seasonal adjustments.

Using the procedure described in Section 2, we first select cities to construct the counterfactuals. The procedure involves an optimal choice of control cities. Using the $CV(n_v)$ method, we select Jilin, Yangzhou, Bengbu, Anqing, Pingdingshan, Xiangfan and Guilin to construct the hypothetical housing prices growth path of Beijing had there been no implementation of purchase restrictions.

The OLS weights based on 2008:M1–2010:M4 data are reported in Table 2. Fig. 3 plots the actual and constructed growth paths for the period 2008:M1–2010:M4. Fig. 3 shows that the counterfactual path, produced by the control groups, traces closely the actual path of Beijing's housing prices before the implementation of purchase restrictions with an R^2 of 0.99.

Next, we construct the counterfactual housing price growth rates of Beijing without purchase restrictions from 2010:M5 to 2011:M11. The actual housing price growth rates of Beijing and the counterfactuals constructed upon the control groups are reported in Table 3. The estimated purchase restriction effects are simply the difference of the two.

According to Table 3, the estimated average treatment effect from 2010:M5 to 2011:M11 is -7.69 percent, and is significant at 1% level according to the asymptotic distribution derived in Li and Bell (2012). Specifically, the average actual housing price growth rate is 9.28 percent, while the average predicted housing price growth rate without purchase restrictions is 16.97 percent using the group of cities selected. The difference of these two gives that the estimated average treatment effect is -7.69 percent. In other

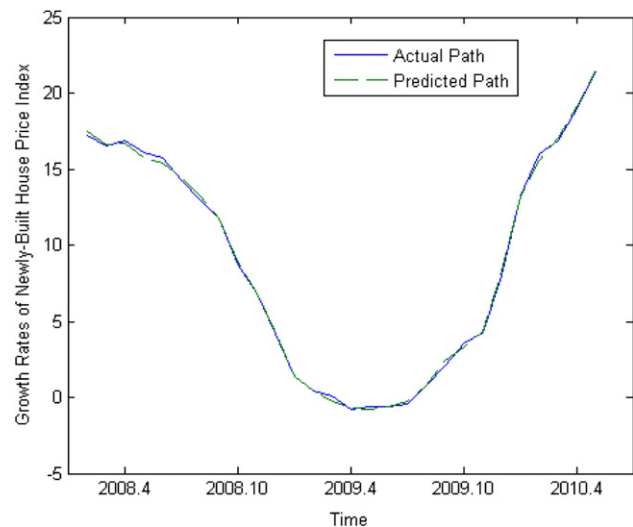


Fig. 3. Actual and predicted path of housing prices in Beijing, 2008:M1–2010:M4.

Table 2

Weights of control groups for Beijing, 2008:M1–2010:M4.

	Coefficient	St.Error	T-stat
Constant	3.2128	1.4228	2.2581
Jilin	0.2181	0.0729	2.9913
Yangzhou	0.1327	0.3473	0.3820
Bengbu	0.1182	0.1724	0.6859
Anqing	0.7689	0.0797	9.6447
Pingdingshan	-0.7321	0.1592	-4.5978
Xiangfan	0.4968	0.1962	2.5323
Guilin	0.8618	0.1884	4.5746
$R^2 = 0.99$			

words, housing prices growth rate of Beijing is declined by more than 7.69 percent compared with the rate had there been no home-purchase restrictions. Fig. 4 plots the actual and constructed paths for the period 2010:M5–2011:M11. The dotted lines denote the 95% confidence bands of the counterfactuals. From Fig. 4, except the first three months, the purchase restriction effect on housing prices at each month is negative and statistically significant at 5% level.

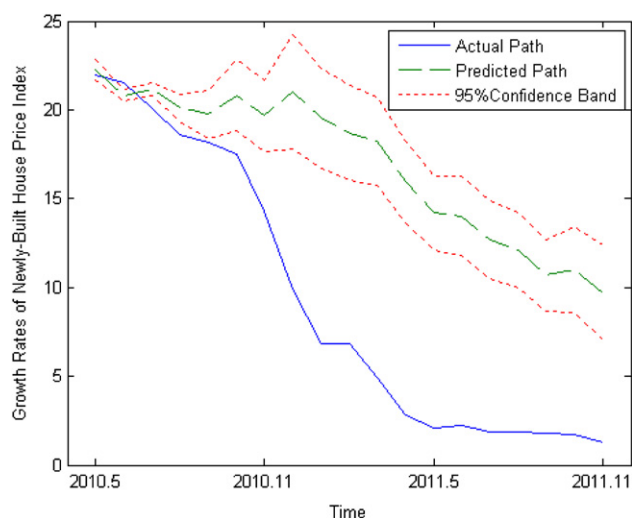
The result is supported by the fact that a relatively large fraction of households is restricted from home purchasing. We compute

Table 3

Treatment effects of Beijing home-purchase restriction, 2010:M5–2011:M11.

	Actual	Control	Treatment
2010.5	22.0	22.2575	−0.2575
2010.6	21.5	20.8147	0.6853
2010.7	20.1	21.1591	−1.0591
2010.8	18.6	20.1271	−1.5271
2010.9	18.2	19.7381	−1.5381
2010.10	17.5	20.8257	−3.3257
2010.11	14.3	19.6680	−5.3680
2010.12	9.9	20.9961	−11.0961
2011.1	6.8	19.5182	−12.7182
2011.2	6.8	18.7061	−11.9061
2011.3	4.9	18.2187	−13.3187
2011.4	2.8	15.9394	−13.1394
2011.5	2.1	14.2006	−12.1006
2011.6	2.2	14.0037	−11.8037
2011.7	1.9	12.7099	−10.8099
2011.8	1.9	12.0981	−10.1981
2011.9	1.8	10.6785	−8.8785
2011.10	1.7	11.0177	−9.3177
2011.11	1.3	9.7134	−8.4134
Average	9.2789	16.9679	−7.6890
T			−3.8437

Note: T-statistic in the last row is calculated using Theorem 3.2 of Li and Bell (2012), and it shows that the average treatment effect here is significantly less than 0 at 1%.

**Fig. 4.** Actual and predicted path of housing prices in Beijing, 2010:M5–2011:M11.

the fraction of restricted households in Beijing using CHFS data. From the sample we extract 303 resident and 107 non-resident households living in Beijing more than 6 months. With information on the number of residential units owned, we find that 31 percent households are restricted in Beijing. In particular, 13 percent of resident households and 78 percent of non-resident households are restricted. As a result, the impact of home-purchase restrictions on the aggregate demand is quite large.

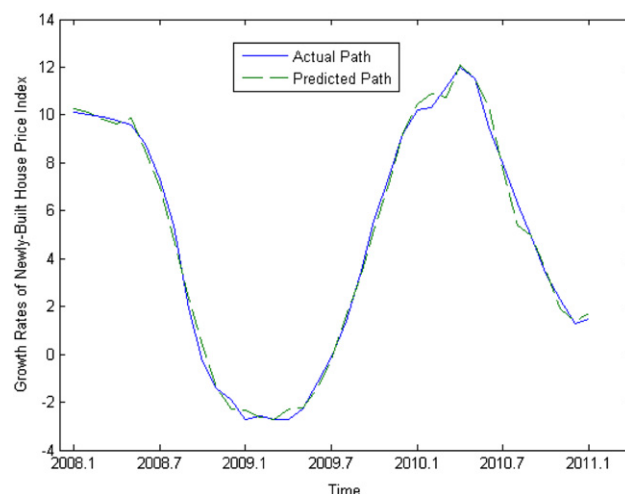
4.2. Trial property tax in shanghai

The trial property tax of Shanghai entered into effect on January 28, 2011. It has no impact on cities without property tax. However, from 2010, the Chinese government imposed home-purchase restrictions in some cities as part of efforts to cool down the overheated real-estate sector and to rein in inflation. To separate the effect of property tax from that of purchase restrictions, we include 23 cities with purchase restrictions implemented about the same time as Shanghai into the control group, namely Shenzhen, Tianjin, Nanjing, Guangzhou, Xiamen, Ningbo, Fuzhou, Haikou, Wenzhou, Hangzhou, Dalian, Taiyuan, Zhengzhou, Wuhan, Hefei,

Table 4

Weights of control groups for Shanghai, 2008:M1–2011:M1.

	Coefficient	St.Error	T-stat
Constant	0.1288	0.1754	0.7344
Hangzhou	0.4492	0.0395	11.3795
Hefei	0.2244	0.0664	3.3808
Fuzhou	0.0583	0.0554	1.0525
Jinan	0.0094	0.0806	0.1165
Guangzhou	0.2080	0.0134	15.5141
Haikou	−0.0225	0.0071	−3.1843
Wenzhou	−0.0679	0.0292	−2.3279
R ² = 0.99			

**Fig. 5.** Actual and predicted path of housing prices in Shanghai, 2008:M1–2011:M1.

Nanchang, Jinan, Qingdao, Kunming, Shijiazhuang, Xi'an, Wuxi and Jinhua.

The policy evaluation period is fixed from February 2011 to February 2012. The municipality of Shanghai relaxed the standards of ordinary housing on March 1, 2012 so that more home transactions are exempt from business taxes. Since the exemption can account for 2 to 5 percent of sale prices, we choose to fix the policy evaluation period to February 2012 to isolate the effects of property tax from the effects of such policy. We use the monthly NBHPI of these 23 cities to construct the counterfactuals of Shanghai in the absence of the trial tax. We calculate the year-on-year monthly growth rates.

Using the CV(n_v) method, we select Hangzhou, Hefei, Fuzhou, Jinan, Guangzhou, Haikou and Wenzhou to construct the hypothetical housing prices growth path of Shanghai had there been no the implementation of property tax. The OLS weights based on 2008:M1–2011:M1 data are reported in Table 4. Fig. 5 plots the actual and constructed growth paths for the period 2008:M1–2011:M1. Fig. 5 shows that the control groups produce the counterfactual path tracing closely the actual path of Shanghai before the implementation of property tax with an R^2 of 0.99.

Then, we construct the counterfactual housing price growth rates of Shanghai without the trial property tax from 2011:M2 to 2012:M2. The actual housing prices growth rates of Shanghai and the counterfactuals constructed upon the control groups are reported in Table 5. The estimated property tax effects are simply the difference of the two.

According to Table 5, the estimated average treatment effect from 2011:M2 to 2012:M2 is 0.006 percent and insignificant. In particular, the average actual housing inflation is 1.9 percent, while the average predicted housing prices growth rate without property tax is 1.89 percent using the group of cities selected. The difference

Table 5
Treatment effects of Shanghai property tax, 2011:M2–2012:M2.

	Actual	Control	Treatment
2011.2	2.3	2.3042	−0.0042
2011.3	1.7	2.3255	−0.6255
2011.4	1.3	1.0250	0.2750
2011.5	1.4	1.0395	0.3605
2011.6	2.2	1.2797	0.9203
2011.7	2.5	2.3214	0.1786
2011.8	2.8	3.2416	−0.4416
2011.9	3.1	2.7670	0.3330
2011.10	2.9	3.0404	−0.1404
2011.11	2.4	2.8092	−0.4092
2011.12	1.8	1.9329	−0.1329
2012.1	0.7	0.6746	0.0254
2012.2	−0.4	−0.1411	−0.2589
Average	1.9000	1.8938	0.0062
T			0.0324

Note: *T*-statistic in the last row is calculated using Theorem 3.2 of Li and Bell (2012), and it shows that the average treatment effect here is insignificant even at 10%.

Table 6
Weights of control groups for Chongqing, 2008:M1–2011:M1.

	Coefficient	St.Error	T-stat
Constant	2.9723	1.8832	1.5783
Bengbu	0.6609	0.2058	3.2110
Anqing	−0.8355	0.2677	−3.1215
Quanzhou	−0.0486	0.2789	−0.1742
Ganzhou	1.2605	0.3051	4.1315
Pingdingshan	−0.0579	0.4781	−0.1211
Yueyang	−0.4585	0.2018	−2.2719
Huizhou	0.0610	0.1321	0.4617
Shaoguan	−0.6897	0.3726	−1.8511
Guilin	0.5952	0.2843	2.0941
Nanchong	0.1779	0.3057	0.5819
$R^2 = 0.98$			

of these two gives the estimated average treatment effect of 0.006 percent.

Fig. 6 plots the actual and constructed growth paths for the period 2011:M2–2012:M2. The dotted lines denote the 95% confidence bands of the counterfactuals. From Fig. 6, except three months, the effects of Shanghai property tax on housing prices are not statistically significant at 5% level. The results suggest that the trial property tax of Shanghai has no significant effects on housing prices.

Our result is confirmed by the fact that the fraction of taxable households in Shanghai is very small. We compute the fraction of taxable households in Shanghai using CHFS data. As the survey was conducted in 2011 with questions on when the house/apartment was purchased, we obtain that the fraction of taxable households is only 0.6 percent. We actually extract households living in Shanghai more than 6 months with and without registered permanent residence in Shanghai. With information on the number of residential units owned, the year when the house/apartment was purchased, the household size and the area of homes, we find that only 0.5 percent of resident households and 1 percent of non-resident households are likely to be taxable.⁷

4.3. Trial property tax in Chongqing

The trial property tax of Chongqing entered into effect on January 28, 2011. As Chongqing has not adopted home-purchase restrictions, to separate the effects of property tax from those of home-purchase restrictions, we include 22 cities without

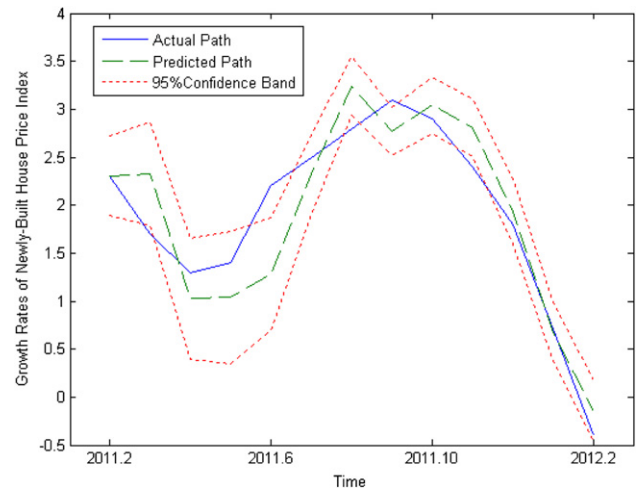


Fig. 6. Actual and predicted path of housing prices in Shanghai, 2011:M2–2012:M2.

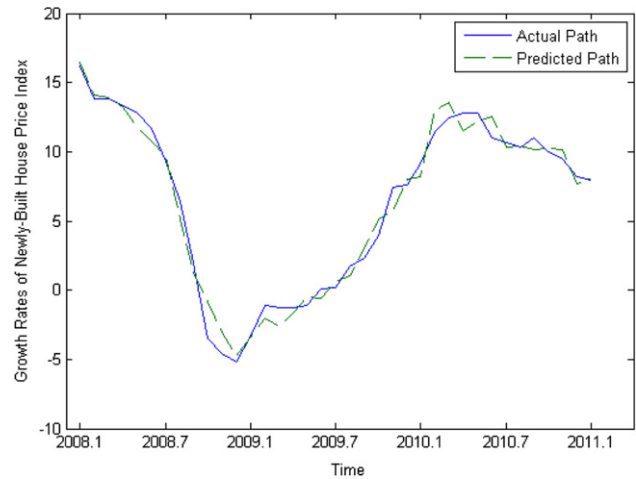


Fig. 7. Actual and predicted path of housing prices in Chongqing, 2008:M1–2011:M1.

purchase restrictions into the control group, namely Yangzhou, Bengbu, Anqing, Quanzhou, Jiujiang, Ganzhou, Jining, Luoyang, Pingdingshan, Yichang, Xiangfan, Yueyang, Changde, Huizhou, Zhanjiang, Shaoguan, Guilin, Beihai, Luzhou, Nanchong, Zunyi and Dali. We use the monthly NBHPI of these 22 cities to construct the counterfactuals of Chongqing in the absence of the trial tax. We calculate the year-on-year monthly growth rates.

Using the $CV(n_v)$ method, we select Bengbu, Anqing, Quanzhou, Ganzhou, Pingdingshan, Yueyang, Huizhou, Shaoguan, Guilin and Nanchong to construct the hypothetical housing prices growth path of Chongqing had there been no the implementation of property tax. The OLS weights based on 2008:M1–2011:M1 data are reported in Table 6. Fig. 7 plots the actual and constructed growth paths for the period 2008:M1–2011:M1. Fig. 7 shows that the control groups produce counterfactual path tracing closely the actual path of Chongqing before the implementation of property tax with an R^2 of 0.98.

Then, we construct the counterfactual housing price growth rates of Chongqing without the trial tax from 2011:M2 to 2012:M1.⁸ Except the political scandal of Bo Xilai, there is no major housing policy changes during this period. The actual

⁷ There are 399 resident households and 99 non-resident households in the sample.

⁸ At the time of this study, data are available until November 2012.

Table 7

Treatment effects of Chongqing property tax, 2011:M2–2012:M11.

	Actual	Control	Treatment
2011.2	6.2	5.2866	0.9134
2011.3	5.6	7.4988	−1.8988
2011.4	5.3	9.0804	−3.7804
2011.5	5.3	5.7708	−0.4708
2011.6	5.8	5.0611	0.7389
2011.7	5.6	5.6167	−0.0167
2011.8	4.2	5.8100	−1.6100
2011.9	2.2	5.9048	−3.7048
2011.10	1.2	4.2539	−3.0539
2011.11	0.1	3.8068	−3.7068
2011.12	−0.6	3.0193	−3.6193
2012.1	−0.7	3.2306	−3.9306
2012.2	−1.1	2.1803	−3.2803
2012.3	−1.5	1.4333	−2.9333
2012.4	−1.8	1.9552	−3.7552
2012.5	−1.9	2.2377	−4.1377
2012.6	−1.9	2.4340	−4.3340
2012.7	−1.5	2.5378	−4.0378
2012.8	−0.9	3.0689	−3.9689
2012.9	−0.5	1.9466	−2.4466
2012.10	−0.1	1.6278	−1.7278
2012.11	0.6	1.2181	−0.6181
Average	1.3455	3.8627	−2.5172
T			−1.5955

Note: *T*-statistic in the last row is calculated using Theorem 3.2 of Li and Bell (2012), and it shows that the average treatment effect here is significantly less than 0 at 10%.

housing prices growth rate of Chongqing and the counterfactuals constructed upon the control groups are reported in Table 7.

According to Table 7, the estimated average treatment effect from 2011:M2 to 2012:M2 is −2.52 percent. In particular, the average actual housing inflation is 1.35 percent, while the average predicted housing price growth rate without property tax is 3.86 percent. The difference of these two gives the estimated average treatment effect of −2.52 percent. Fig. 8 plots the actual and constructed growth paths for the period 2011:M2 to 2012:M11. The dotted lines denote the 95% confidence bands of the counterfactuals. It is notable that except several months, the effects of Chongqing property tax on housing prices are not statistically significant at 5 percent level, but significant at 10 percent level.

Based on data released by Chongqing Municipal Finance Bureau, the fraction of taxable households in 2011 is around 10 percent, which is much larger than the estimated 0.6 percent in Shanghai. Unlike the trial property tax of Shanghai, the trial property tax of Chongqing is levied on both old and new independent houses. Moreover, Chongqing's property tax rates, varying from 0.5 to 1.2 percent, are more aggressive than Shanghai where tax rates vary from 0.4 percent to 0.6 percent.

5. Conclusions

In this paper, we respectively evaluate the effects of the home-purchase restrictions and the trial property taxes on housing prices using a counterfactual analysis. We exploit the dependence of housing prices among different cities and construct the counterfactuals using data from cities without home purchase restrictions and trial property tax. We propose to use the leave- n_v -out cross-validation for the optimal choice of the control cities. We find that home-purchase restrictions reduced the growth rate of housing prices in Beijing by 7.69 percent and the trial property tax of Chongqing reduced the growth rate housing prices by 2.52 percent, while the trial property tax of Shanghai had no significant effect on housing prices. The results suggest that in order to curb

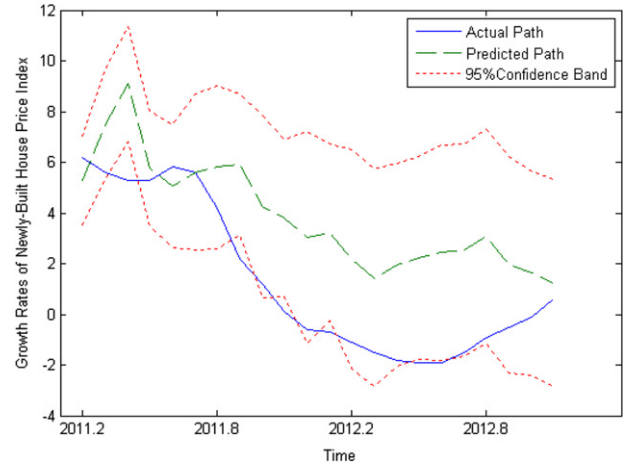


Fig. 8. Actual and predicted path of housing prices in Chongqing, 2011:M2–2012:M11.

the soaring housing prices home-purchase restrictions should not be, at least in the short run, replaced by property taxes. The trial property tax of Chongqing with higher proportion of taxable homes is more effective in curbing housing prices than the trial property tax of Shanghai.

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Appendix

Proof of Theorem 1. We first introduce some notations. For the choice of the control cities indexed by γ , we consider fitting the model $g_{1t}^0 = \bar{\alpha}_\gamma + \mathbf{a}'_\gamma \mathbf{g}_{\gamma t} + \varepsilon_{1t}$. Let 1_{T_1} and 1_{T_2} denote a $T_1 \times 1$ and a $T_2 \times 1$ vector of 1's, respectively. Define

$$\mathbf{Y}_{1:T_1} = (g_{11}^0, g_{12}^0, \dots, g_{1T_1}^0)',$$

$$\boldsymbol{\varepsilon}_{1:T_1}^* = (\varepsilon_{11}^*, \varepsilon_{12}^*, \dots, \varepsilon_{1T_1}^*)',$$

$$\mathbf{X}_{1:T_1} = (1_{T_1}, (\mathbf{g}_{-1,1}, \mathbf{g}_{-1,2}, \dots, \mathbf{g}_{-1,T_1})'),$$

$$\mathbf{X}_{T_1+1:T} = (1_{T_2}, (\mathbf{g}_{-1,T_1+1}, \mathbf{g}_{-1,T_1+2}, \dots, \mathbf{g}_{-1,T})'),$$

$$\mathbf{X}_{1:T} = (\mathbf{X}_{1:T_1}', \mathbf{X}_{T_1+1:T}')',$$

$$\mathbf{X}_{1:T_1,\gamma} = (1_{T_1}, (\mathbf{g}_{\gamma,1}, \mathbf{g}_{\gamma,2}, \dots, \mathbf{g}_{\gamma,T_1})'),$$

$$\mathbf{X}_{T_1+1:T,\gamma} = (1_{T_2}, (\mathbf{g}_{\gamma,T_1+1}, \mathbf{g}_{\gamma,T_1+2}, \dots, \mathbf{g}_{\gamma,T})'),$$

$$\boldsymbol{\beta} = (\bar{\alpha}, \mathbf{a}'_{-1})'.$$

With these notations, we have $\mathbf{Y}_{1:T_1} = \mathbf{X}_{1:T_1} \boldsymbol{\beta} + \boldsymbol{\varepsilon}_{1:T_1}^*$, and the OLS estimates

$$\begin{aligned} \hat{\boldsymbol{\beta}}_\gamma &= (\hat{\bar{\alpha}}_\gamma, \hat{\mathbf{a}}'_\gamma)' = (\mathbf{X}_{1:T_1,\gamma}' \mathbf{X}_{1:T_1,\gamma})^{-1} \mathbf{X}_{1:T_1,\gamma}' \mathbf{Y}_{1:T_1} \\ &= (\mathbf{X}_{1:T_1,\gamma}' \mathbf{X}_{1:T_1,\gamma})^{-1} \mathbf{X}_{1:T_1,\gamma}' (\mathbf{X}_{1:T_1} \boldsymbol{\beta} + \boldsymbol{\varepsilon}_{1:T_1}^*) \\ &:= A_\gamma (\mathbf{X}_{1:T_1} \boldsymbol{\beta} + \boldsymbol{\varepsilon}_{1:T_1}^*), \end{aligned}$$

with A_γ implicitly defined.

Next we prove the following equality for the mean squared prediction error using the control cities indexed by γ ,

$$\begin{aligned} MSE_{\gamma} &= E \left[\frac{1}{T_2} \sum_{t=T_1+1}^T (g_{1t}^0 - \widehat{g}_{1t}^0)^2 \middle| \mathbf{X}_{1:T} \right] \\ &= \sigma^2 + \sigma^2 \frac{1}{T_1} d_{\gamma} + \frac{1}{T_1} \boldsymbol{\beta}' \mathbf{X}_{1:T_1}' \\ &\quad \times (I_{T_1} - \mathbf{X}_{1:T_1, \gamma} A_{\gamma}) \mathbf{X}_{1:T_1} \boldsymbol{\beta} + o_p(1), \end{aligned} \quad (11)$$

where $\widehat{g}_{1t}^0 = \widehat{\alpha}_{\gamma} + \widehat{\mathbf{a}}_{\gamma}' \mathbf{g}_{\gamma t}$.

Notice that

$$\begin{aligned} MSE_{\gamma} &= E \left[\frac{1}{T_2} \sum_{t=T_1+1}^T (\bar{\alpha} + \mathbf{a}_{-1}' \mathbf{g}_{-1t} + \varepsilon_{1t}^* - \widehat{g}_{1t}^0)^2 \middle| \mathbf{X}_{1:T} \right] \\ &= \sigma^2 + E \left[\frac{1}{T_2} (\boldsymbol{\beta}' \mathbf{X}_{T_1+1:T}' - \widehat{\boldsymbol{\beta}}_{\gamma}' \mathbf{X}_{T_1+1:T, \gamma}') \right. \\ &\quad \times (\mathbf{X}_{T_1+1:T} \boldsymbol{\beta} - \mathbf{X}_{T_1+1:T, \gamma} \widehat{\boldsymbol{\beta}}_{\gamma}) \middle| \mathbf{X}_{1:T} \left. \right] \\ &= \sigma^2 + E \left[\frac{1}{T_2} (\boldsymbol{\beta}' \mathbf{X}_{T_1+1:T}' - (\boldsymbol{\beta}' \mathbf{X}_1 : \mathbf{T}_1' + \boldsymbol{\varepsilon}_{1:T_1}^*) A_{\gamma}' \mathbf{X}_{T_1+1:T, \gamma}') \right. \\ &\quad \times (\mathbf{X}_{T_1+1:T} \boldsymbol{\beta} - \mathbf{X}_{T_1+1:T, \gamma} A_{\gamma} (\mathbf{X}_{1:T_1} \boldsymbol{\beta} + \boldsymbol{\varepsilon}_{1:T_1}^*)) \middle| \mathbf{X}_{1:T} \left. \right] \\ &= \sigma^2 + E \left[\frac{1}{T_2} \boldsymbol{\varepsilon}_{1:T_1}^* A_{\gamma}' \mathbf{X}_{T_1+1:T, \gamma}' \mathbf{X}_{T_1+1:T, \gamma} A_{\gamma} \boldsymbol{\varepsilon}_{1:T_1}^* \middle| \mathbf{X}_{1:T} \right] \\ &\quad + \frac{1}{T_2} (\boldsymbol{\beta}' \mathbf{X}_{T_1+1:T}' - \boldsymbol{\beta}' \mathbf{X}_{1:T_1}' A_{\gamma}' \mathbf{X}_{T_1+1:T, \gamma}') \\ &\quad \times (\mathbf{X}_{T_1+1:T} \boldsymbol{\beta} - \mathbf{X}_{T_1+1:T, \gamma} A_{\gamma} \mathbf{X}_{1:T_1} \boldsymbol{\beta}), \end{aligned}$$

where $\sigma^2 = E(\varepsilon_{1t}^{*2})$. Under [Assumption A5](#), we have

$$\begin{aligned} MSE_{\gamma} &= \sigma^2 + \sigma^2 \frac{1}{T_1} \text{trace} (A_{\gamma}' \mathbf{X}_{1:T_1, \gamma}' \mathbf{X}_{1:T_1, \gamma} A_{\gamma}) \\ &\quad + \frac{1}{T_1} (\boldsymbol{\beta}' \mathbf{X}_{1:T_1}' \mathbf{X}_{1:T_1} \boldsymbol{\beta}) - \frac{1}{T_1} \boldsymbol{\beta}' \mathbf{X}_{1:T_1}' \mathbf{X}_{1:T_1, \gamma} A_{\gamma} \mathbf{X}_{1:T_1} \boldsymbol{\beta} + o_p(1) \\ &= \sigma^2 + \sigma^2 \frac{1}{T_1} d_{\gamma} + \frac{1}{T_1} \boldsymbol{\beta}' \mathbf{X}_{1:T_1}' \\ &\quad \times (I_{T_1} - \mathbf{X}_{1:T_1, \gamma} A_{\gamma}) \mathbf{X}_{1:T_1} \boldsymbol{\beta} + o_p(1). \end{aligned}$$

Therefore we have proved (11).

[Theorem 1](#) (1) then follows from similar arguments as [Theorem 2\(a\)](#) of [Shao \(1993\)](#), and (2) follows from [Theorem 2\(b\)](#) and (c) of [Shao \(1993\)](#). ■

Proof of Proposition 1. For $t > T_1$, we have

$$\begin{aligned} \widehat{\Delta}_{1t} &= g_{1t}^1 - \widehat{g}_{1t}^0 \\ &= g_{1t}^1 - g_{1t}^0 + g_{1t}^0 - \widehat{g}_{1t}^0 \\ &= \Delta_{1t} + \bar{\alpha} + \mathbf{a}_{-1}' \mathbf{g}_{-1t} + \varepsilon_{1t}^* - \widehat{\alpha}_{\gamma^*} - \widehat{\mathbf{a}}_{\gamma^*}' \mathbf{g}_{\gamma^* t} \\ &= \Delta_{1t} + \varepsilon_{1t}^* + \bar{\alpha}_{\gamma^*} - \mathbf{a}_{\gamma^*}' \mathbf{g}_{\gamma^* t} - \widehat{\alpha}_{\gamma^*} - \widehat{\mathbf{a}}_{\gamma^*}' \mathbf{g}_{\gamma^* t} + O_p(T_1^{-1/2}) \\ &= \Delta_{1t} + \varepsilon_{1t}^* + O_p(T_1^{-1/2}), \end{aligned}$$

where the second to last equality follows from (11). Similarly, we get

$$\begin{aligned} \frac{1}{T_2} \sum_{t=T_1+1}^T \widehat{\Delta}_{1t} - \frac{1}{T_2} \sum_{t=T_1+1}^T \Delta_{1t} &= \frac{1}{T_2} \sum_{t=T_1+1}^T \varepsilon_{1t}^* + \bar{\alpha}_{\gamma^*} - \mathbf{a}_{\gamma^*}' \mathbf{g}_{\gamma^* t} \\ &\quad - \widehat{\alpha}_{\gamma^*} - \widehat{\mathbf{a}}_{\gamma^*}' \mathbf{g}_{\gamma^* t} + O_p(T_1^{-1/2}) \\ &= o_p(1), \end{aligned}$$

where the last equality follows from [Assumption A1](#) and the properties of OLS estimator. ■

References

- Akaike, H., 1973. Information theory and an extension of the maximum likelihood principle. In: Petrov, B.N., Csaki, F. (Eds.), *Proceedings of the second International Symposium on Information Theory*. Akadémiai Kiadó, Budapest, pp. 267–281.
- Akaike, H., 1974. A new look at the statistical model identification. *IEEE Trans. Automat. Control* 19, 716–723.
- Bai, C.E., Li, Q., Ouyang, M., 2014. Property taxes and home prices: A tale of two cities. *J. Econometrics* 180 (1), 1–15.
- Bai, J., Ng, S., 2002. Determining the number of factors in approximate factor models. *Econometrica* 70 (1), 191–221.
- Berger, A.N., Udell, G.F., 1992. Some evidence on the empirical significance of credit rationing. *J. Polit. Econ.* 100 (5), 1047–1077.
- Cebula, R.J., 2009. Are property taxes capitalized into housing prices in savannah, georgia? An investigation of market mechanism. *J. Hous. Res.* 18 (1), 63–75.
- Deng, Y.H., Morck, R., Wu, J., Yeung, B., 2011. Monetary and Fiscal Stimuli, Ownership Structure, and China's Housing Market, NBER Working Paper, 16871.
- Fischel, W.A., 1992. Property taxation and the tiebout model: Evidence for the benefit view from zoning and voting. *J. Econ. Lit.* 30, 163–169.
- Hamilton, B.W., 1976. Capitalization of intrajurisdictional differences in local tax prices. *Am. Econ. Rev.* 66, 743–753.
- Howard, D.H., 1977. Rationing, quantity constraints, and consumption theory. *Econometrica* 45 (2), 399–412.
- Hsiao, C., Steve Ching, H., Ki Wan, S., 2012. A panel data approach for program evaluation: Measuring the benefits of political and economic integration of hong kong with mainland china. *J. Appl. Econometrics* 27 (5), 705–740.
- Hurvich, C.M., Tsai, C.L., 1989. Regression and time series model selection in small samples. *Biometrika* 76 (2), 297–307.
- Lang, K., Jian, T.L., 2004. Property taxes and property values: Evidence from proposition 2½. *J. Urban Econ.* 55 (3), 439–457.
- Li, K., Bell, D., 2012. Estimation of Average Treatment Effects with Panel Data: Asymptotic Theory and Implementation, Working Paper.
- Mackay, R.J., Whitney, G.A., 1980. The comparative statics of quantity constraints and conditional demands: Theory and applications. *Econometrica* 48 (7), 1727–1744.
- Netzer, D., 1966. *Economics of the Property Tax*. Brookings Institute, Washington.
- Oates, W.E., 1969. The effects of property taxes and local public spending on property values: an empirical study of tax capitalization and the tiebout hypothesis. *J. Polit. Econ.* 81 (4), 1004–1008.
- Rosen, H.S., Fullerton, D.J., 1977. A note on local tax rates, public benefit levels, and property values. *J. Polit. Econ.* 85 (2), 433–440.
- Shao, J., 1993. Linear model selection by cross-validation. *J. Amer. Statist. Assoc.* 88 (422), 486–495.
- Shao, J., 1996. Bootstrap model selection. *J. Amer. Statist. Assoc.* 91 (434), 655–665.
- Simon, H.A., 1943. The incidence of a tax on urban real property. *Quart. J. Econ.* 57 (3), 398–420.
- Steijvers, T., Voordeckers, W., 2009. Collateral and credit rationing: A review of recent empirical studies as a guide for future research. *J. Econom. Surv.* 23 (5), 924–946.
- van den Noord, P., 2005. Tax Incentives and House Prices Volatility in the Euro Area: Theory and Evidence. In: *Economie Internationale*, vol. 101. CEPII research center, pp. 29–45.
- Wei, S.J., Allen, R., 2010. Beyond the crisis: Growth prospects and critical challenges for major economies over the next ten years. In: *The Chazen Institute Report on World Business & the Economy*. Columbia University.
- Wu, J., Gyourko, J., Deng, Y.H., 2012. Evaluating conditions in major chinese housing markets. *Reg. Sci. Urban Econ.* 42 (3), 531–543.