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Spatial Variation in the Determinants of House Prices and Apartment Rents in China

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Abstract This paper provides an examination of China's residential real estate market at the county level using data from that country's 2000 census. The market is a new one, having only been fully established in 1998. The analysis in the paper is in the form of an aggregate (county-level) hedonic model specified in two versions. Global parameters results are estimated using spatial error model specifications while more local effects are estimated by geographically weighted regression. Global results are typical in that structural characteristics such as floor space and contextual characteristics such as level of in-migration are important in residential prices. Local results, however, indicate significant spatial variation in the effect of both structural amenities and locational context on housing prices. In a simpler specification, rents are shown to respond positively to both median house prices levels and the supply of apartments available at market prices, but also with significant spatial variation across China.

Keywords Housing market · China · Hedonic model · Geographically weighted regression · Spatial regression

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Introduction

Most discussions of China's economic reforms focus on production and especially the rapid growth of manufacturing for export that began in the late 1970s. There has been less analysis of the ongoing reforms in housing, which only began in earnest in the 1980s (Chen 1996; Fu et al. 2000). The reforms are largely attributed to dissatisfaction with state provided housing on both the parts of government and residents. Most urban housing in China was nationalized in the 1950s. State provided or subsidized housing, however, was an expensive prospect for the government and its attempts to keep costs low often led to overcrowding. That problem became even more severe when migration from rural to urban areas increased as a result of the growth of China's manufacturing in its coastal provinces. Initial reforms removed the state from urban housing provision directly by transferring much of that responsibility to local enterprises or work units in a system that ironically paralleled the rural communal system. Chen (1996, p. 1080) describes a decrease in central government housing investment from 90% to 16% between 1979 and 1988 while work unit investment rose to 52%, with the rest attributable to private, communal, or local government sources. Private property rights in urban areas were endorsed by the state in 1988 (Fu et al. 2000), and commercial real estate markets were encouraged more strongly in 1993 through a series of appraisal and credit initiatives (Chen 1996). The reforms ultimately led to the effective privatization of the urban residential real estate market in 1998.

China's reforms in manufacturing appear to have led quickly to behavior that would be expected in a market economy. For example, agglomeration economies have quickly emerged as important in the location of production (Fujita and Hu 2001). Real estate reforms also seem to have quickly led to typical market based patterns, at least in the Beijing metropolitan area (Ding 2004; Zheng and Kahn 2008). The purpose of this paper is to more broadly analyze the emerging residential real estate market in China in an effort to identify factors that have emerged as important with respect to housing prices. The analysis utilizes two versions of a type of regional hedonic model that employs county-level housing prices as a function not only of structural characteristics, but also contextual population characteristics. An additional model concerns apartment rents determined largely as a function of house prices. Both the housing and apartment models are specified as spatial error models to estimate global parameters. They are also specified as geographically weighted regressions (GWRs) in order to capture spatial variation in effects across China, a country geographically marked by cultural, economic, and environmental heterogeneity. The next part of the paper provides a brief background to the analysis. The third part contains the development of the GWR models and a description of the data used in the analysis. Part Four contains the results of the analyses, and the paper concludes with a summary of the findings and some suggestions for additional research.

Background

Tong and Hays (1996) described the major issues that led to housing market reform in China. In the pre-reform era, housing construction was funded by the state, but placed in the category of nonproductive investment and given a very low priority. The chronic underfunding that occurred during that period led to significant issues of overcrowding, with average living space dropping from 4.5 square meters in 1949 to just over 3 square meters in 1960. It recovered to the 1949 level by 1983 as reforms began and rose to over 9.1 sq meters per person by 1991. The increase in average living space hid some ongoing problems, however, including high levels of inequality. About one third of China's households were classified as either homeless or severely crowded in 1991. Further, about 14% of existing housing was considered dilapidated. Large proportions of the urban housing stock lacked basic amenities in 1991; only 82% had running water, and less than 43% had indoor toilets. More than 24% of public housing units lacked heat. Despite the shortfall in amenities, living space appears to have been the most important shortcoming in state housing. Zax (1997) reported, for example, that in 1989 the average total living area in China's state-owned housing was 30.2 square meters (s. d.=14.6) and the average number of rooms was 2.44 (s. d.=1.09), while the measures for own-built housing were 52.9 (s. d.=35.2) and 3.6 (s. d.=2.09), respectively. At the same time, state-owned facilities had higher rates of sanitary facilities and heating.

Wang and Murie (1999) described the transition from state control to commercial housing in China as an evolutionary process with an uneven geography. During the 1980s most private construction was sold to enterprises rather than households, with the enterprises in turn renting dwellings to employees. A dual residential real estate market emerged in the 1990s. One was for the developing class of high income individuals including business entrepreneurs and film stars. The other was for ordinary urban households which were generally limited in their opportunities by high prices in the residential market. Even in the 1990s, however, residential purchases by individuals were only about half of residential purchases made by enterprises. There was, however, considerable geographical variation in that proportion. For example, only 20% of residential sales were made to individuals in Beijing but more than 70% were to individuals in Guangdong Province. There was also considerable geographical variation in the total sales, as measured in floor space. Again using Beijing and Guangdong Province as examples, the former had 1,783,000 square meters of residential space sold and the latter nearly 8,499,000 square meters. The source of variation in both instances is likely the level of in-migration in combination with historic supplies of residential space. Beijing is large but old, and with a considerable supply of already existing state and enterprise owned housing, so its residential real estate market was thin in the mid-1990s. Much of Guangdong Province is recently developed because of its role as an export platform, so its residential real estate market was established as a response to large-scale in-migration that occurred after the state had left the housing sector. Across China, more than 50% of all housing was sold to individuals by 1996 (Fu et al. 2000).

A low level of access to publicly subsidized housing was shown by Fu et al. (2000) to be an important factor in the intention to purchase private housing in China. Other important factors included affordability and crowding in current living space, and also personal characteristics such as economic confidence and a high tolerance for risk. Affordability has been a problem for most Chinese, who have relatively low incomes, and a mortgage system was established in 1998 with the intention of expanding the residential real estate market (Wang 2001). In addition to affordability, Huang and Clark (2002) found household characteristics such as size, age, and income to be important factors in the residential purchase decision. They also found that institutional factors such as *hukou* (residential permit) status and job rank were important because they had an impact on access to subsidized housing, but their importance varied across cities. Wu (2002) also found that while there was a strong link between income and housing quality, factors such as occupation and political affiliation were the primary sources of inequalities in housing space.

Li and Yi (2007) examined family cycle and institutional effects in the rent-to-own transition over three periods of housing reform. They found that, with the exception of party membership, institutional factors declined over time and that (p. 364) "… it is likely that the pattern of housing consumption in China will increasingly resemble those in market economies." In analyses of the greater Beijing land market, Ding (2004) found the type of land rent gradients that would be expected in any typical market economy. Zheng and Kahn (2008) also found that the spatial distribution of population and residential use in Beijing conforms to the monocentric model that was developed for market-based economies. Using standard hedonic analysis, they also found that local public goods such as proximity to public transit, good schools, and environmental amenities are capitalized in real estate prices—findings typical, again, of market economies.

Regression Models

As indicated above, several papers have described the quickly emerging market economy characteristics of residential real estate in China. The analysis described in the following has the purpose of further exploring residential real estate patterns in China in order to determine whether such market economy characteristics can be found in broader terms than those already described. Like Zheng and Kahn (2008) we use hedonic analysis, but in a regional rather than in a standard form. Our hedonic analysis also directly accommodates spatial variation in parameter estimates, which is of interest in our application which utilizes contiguous spatial units of observation.

Hedonic real estate models are designed to identify the marginal contribution of a house's or other residence's characteristics to its purchase price (Sirmans et al. 2005). Those characteristics can be divided into two general sets: physical characteristics of the house (and often property) and what are in effect locational characteristics that often include the types of public good access used by Zheng and Kahn (2008). As a regression specification, the typical hedonic model takes the general form:

$$\ln P = \alpha + \beta_1 R + \beta_2 L + \mu \tag{1}$$

where lnP is a vector of the natural logarithm of prices paid for residences, α, the intercept, is the estimated price in the absence of any additional information, R is a matrix of the residences' defining physical characteristics, L is a matrix of the residences' defining locational characteristics, β_1 and β_2 are the vectors of the respective estimated parameters, and μ is a vector of random errors, i.i.d.

Hedonic models are often estimated by ordinary least squares (OLS), but that method is also often inappropriate because of the spatial nature of most real estate data (Pace et al. 1998). Spatially distributed values are likely to violate the basic independence of dependent variable observation assumption of OLS, and its use can lead to incorrect signs on estimated parameters (Pace and Gilley 1997). Further, a non-random distribution of the error terms can result in inefficient parameter estimates and a misreading of their statistical significance. Standard approaches to reducing spatial patterns in errors include the use of geographical dummy variables, the addition of spatially lagged explanatory variables, and direct modeling of their spatial autocorrelation (Anselin 2003; Cohen and Coughlin 2008). For the last approach, a spatial errors model (SEM) takes the basic format:

$$\ln P = \alpha + \beta_1 R + \beta_2 L + \lambda W \mu + \varepsilon \tag{2}$$

where λ is a parameter that accounts for autocorrelation in errors, μ, over spatial structure *W*. Note that SEM models diminish missing variable effects if those effects are spatially variable (Cohen and Coughlin 2008). As in the case of OLS, the SEM is global in scope because the regression parameters are taken as applicable in a constant way across the space of the observations; variation results from the spatial heterogeneity of the explanatory variables alone.

Fotheringham et al. (2002) have developed GWR as an alternative format for spatial analysis that is local rather than global in its analytical design. The GWR version of the generalized hedonic specification in Eq. 1 is:

$$\ln P_i = \alpha(x_i, y_i) + \beta_1(x_i, y_i) R_i + \beta_2(x_i, y_i) L_i + \mu_i \tag{3}$$

where (x_i, y_i) are the spatial coordinates of the *i*th residence and $\beta_1(x_i, y_i)$ and $\beta_2(x_i, y_i)$ are realizations of continuous functions at the place of the *i*th residence. Parameter values are considered to be a continuous surface and the spatial variation in the surfaces are measured at specified points—the residential observations. In GWR there is variability in the parameters as well as in the explanatory variables so each observation may, in fact, be treated individually.

GWR estimation is conducted by a form of weighted least squares (WLS) as opposed to OLS. GWR estimates are mildly biased because of their assumed response to locational effects. The local parameters are estimated based on values of spatially proximal observations (those located within a defined spatial window, or bandwidth). Nearby observations have more influence than do those farther away from the observation of interest in the WLS procedure used in GWR. Fotheringham et al. (2002) argue, however, that the mild bias is offset by reduced standard errors of the estimates as long as the spatial bandwidth contains a sufficient sample size. Wheeler and Tiefelsdorf (2005) have demonstrated that the WLS solutions can lead to multicollinearity and correlation among the calculated local regression coefficients even when the explanatory variables used in a model are uncorrelated. *If* that is the case, it indicates that while the regression results can be meaningfully relied upon as a suite for inference, individual parameters must be more carefully considered in the interest of hypothesis testing. Further, Griffith (2008) notes that while GWR accounts for much of the spatial pattern between covariates in its spatially varying estimated parameters, spatial autocorrelation of residuals may still be a problem.

Pace and LeSage (2004) indicate that the level of spatial autocorrelation in the errors typically increases with bandwidth size, and is effectively a trade-off with the decreasing parameter variability that occurs with larger sample sizes.

The regression analyses in this paper are conducted at the county level. The source of data analyzed is China's 2000 census as compiled by the All China Marketing Research Co., a licensed affiliate of the State Statistical Bureau of China (Wei et al. 2002). The county-level (counties and equivalent administrative districts) data were made GIS compatible by the University of Michigan's China Data Center in 2005, and provided for the analysis described here by the Harvard Geospatial Library (2008). There are 2,873 observations.

The hedonic GWR results reported below are taken from bandwidths that meet an optimality condition of a minimized corrected (essentially for number of observations) Akaike information criterion statistic (Charlton and Fotheringham 2009). County-unit density is much higher in China's east than in its south, so there is significant variation in the distances encompassing a fixed set of near neighbors that generally follows that spatial trend. The lack of spatial uniformity would also have an impact on a fixed-distance (for example, all neighbors within 150 km) bandwidth sampling strategy, with large variations in the number of neighbors used to estimate local parameters in that case.

Apart from some questions concerning the count of young females, China's 2000 census data are considered generally reliable (Chan 2003.) It is the first of China's censuses to provide significant coverage of housing (Wang 2003). The data used in the analyses below are from a 9.5% sample of the population. That sample has 33,575,685 households, of which 5,269,944 had purchased their residence and 2,904,345 were renters. About 70% of households in cities owned their own home in 2000 (Li and Yi 2007), but the sample used here includes both urban and non-urban observations. The great majority of households in the sample, more than 70%, lived in self-built housing which is not included in the analysis. Nationally, about 37% of China's households lived in self-built housing in 2000.

Purchase price and contract rent data are each given in nine categories in the dataset, so county-level group medians were calculated for use in the hedonic analysis. In both cases, the upper category is unbounded, so the highest price and rent medians were estimated based on the intervals used in the lower price and rent categories. (The price categories in Yuan are under 10,000; 10,000–20,000; 20,000–30,000; 30,000–50,000; 50,000–100,000; 100,000–200,000; 200,000–300,000; 300,000–500,000; 500,000 and over. The rent categories in Yuan are under 20; 20–50; 50–100; 100–200; 200–500; 500–1,000; 1,000–1,500; 1,500–2,000; 2,000 and over.) Minimum county-level purchase prices were about 5,000 Yuan at the median while the maximum median was about 200,000 (Table 1). The mean county-level median was about 13,000 Yuan. Minimum median monthly rents were 10 Yuan with the mean county-level median at 41 Yuan. Moran's I statistics were calculated with 0–1 queen's case weights and indicate fairly strong spatial autocorrelation across both prices and rents that points to the necessity of explicitly incorporating space in their models. There are some exceptions, but the main spatial concentrations of higher prices and rents are around Beijing and then along the coast from around Shanghai south into Guangdong Province. That pattern is not surprising given the population and industrial geography of China.

Table 1 Summary statistics for residential prices and rents at county levels in China in 2000 (in Yuan)

	Minimum	Maximum	Mean	Standard deviation	Moran's I
Median Price	5,138	202,796	13,145	10,725	0.734
Median Rent	10	867	41	36	0.658

The census dataset contains a limited but useful number of residential characteristic and locational context variables applicable to modeling house prices at the county level, including number of rooms per household, floor space per household, number of houses by year of construction before or after 1980, number of houses with kitchen facilities, with tap water, with hot water, with bath facilities, and with lavatory facilities. The first version of the hedonic model used here includes the average floor space per household, the percent of houses built before 1980, the percent of houses with private kitchens, the percent of houses with tap water, and the percent of houses with central hot water for each county level unit. The model also includes the percent of houses in the county-level unit classified as "economic and functional," a designation given to houses intentionally built for low and middle income residents. Expectations are that house prices are positive in floor space, negative in age of the housing stock (houses built before 1980), positive in the kitchen, lavatory, and water amenities, and negative in the stock of lower and middle income construction. Housing stock age, or vintage, is often correlated with structural characteristics in housing markets (Muth 1973; Dubin et al. 1999), but that is not the case at the county scale in China. The highest Pearson correlation with any of the other variables used in the analysis (lavatory facilities) is 0.279. Its correlation with floor space is only 0.072.

Three locational context variables are also included in the hedonic model and are the same in both versions: county level percent urban population, percent minority population, and a net migration variable that is measured as a county's total resident population divided by its total registered population. Net in-migration is indicated by values of that ratio greater than one. Urban population is included because China's real estate market reforms were enacted only in urban areas. Housing prices are expected to be positive in the proportion of urban population because they are still strongly controlled in rural areas, if residential markets are in operation at all. Percent minority population is used as an indirect substitute for professional occupational status, which should indicate some measure of ability to pay in the absence of income information. Percent minority population is inversely correlated with professional employment, but does not contribute significantly to multi-collinearity in the estimated model. Housing prices are expected to decrease in percent minority population. Finally, housing prices are expected to increase in the net migration ratio because it is a measure of local demand (Frame 2008). The multicollinearity condition index of the complete hedonic model (version 1) is 28.3, which is just under the level of 30 that raises serious concern.

Although a population variable is not included on the right-hand side of the hedonic model, it could be that the first version is subject to some simultaneity in the determination of housing characteristics (as ratios) and price. The second version of the model avoids that potential issue by using dichotomized measures of the structural characteristics with respect to their means (0 if < mean, 1 if > mean). There is one difference between the two versions of the model on their right-hand sides. Version 2 of the model uses a dichotomized *percent houses with independent lavatory* as a substitute for *percent houses with central hot water* because the mean of the latter is only one percent (Table 2). The use of both variables in the same estimating equation increases model collinearity considerably. Higher levels of independent lavatories are expected to have a positive effect on housing prices. Other expectations with respect to signs of the associated estimates remain the same as in the case of version 1 of the model.

There is considerable spatial variation in the explanatory variables used in the hedonic model (Table 2). Five of the variables are at zero levels at their minimums, while two are at 100% maximums. The average floor space per household was about 78 square meters. At their means a majority of houses in the county-level units of observation were built before 1980 and have private kitchens, but houses with tap water and especially those with central hot water are not typical, nor are those with independent lavatories. The average county-level observation had about 38% of its population in urban areas and 16% of its population classified as minority, while net migration was flat at the mean. With the exception of minority population, the general spatial trend in each of the explanatory variables is an increase from west to east.

Physical characteristics of apartments are not given in the data source so rent is modeled in its response to only two variables: county median housing price and the percentage of county-level rental units that are designated as “commodity” apartments. Linkage between renting and purchasing in China has been documented by Chen (1996). Rents are expected to be positive in housing prices in a competitive residential real estate market. Many apartments are subsidized publicly owned units, but commodity apartments are rented at market rates (Fu et al. 2000). Rent is also expected to be positive in the proportion of commodity apartments in the county market.

Table 2 Summary statistics for county-level residential characteristics and locational context variables used in the hedonic house price model

	Minimum	Maximum	Median	Mean	Standard deviation
Floor space (m^2) per household	17.8	225.4	75.6	78.1	21.4
% houses built before 1980	22.0	99.7	76.5	75.5	11.0
% houses with private kitchens	0.5	99.4	85.3	79.2	19.0
% houses with tap water	0.0	100.0	38.5	46.4	31.4
% houses with central hot water	0.0	43.2	0.3	1.0	2.3
% houses with Independent lavatory	0.0	91.8	7.8	16.6	19.9
% low and mid-income houses	0.0	54.8	1.7	2.9	3.5
% urban population	0.0	100.0	25.2	38.1	31.5
% minority population	0.0	99.8	1.5	16.2	29.0
Net-migration ratio	0.6	10.3	0.9	1.0	0.3

Results

The results of the two global hedonic models are given in Table 3, and the parameters resulting from the SEM regressions are generally as expected with respect to sign. The exceptions are the negative, but not significant, parameter for kitchen facilities in version 2, and positive parameters for age of the housing stock (pre-1980 construction) and proportion of low and middle income houses in both model versions. It could be that the positive parameter for housing stock age is a response to vintage characteristics not otherwise identified in the model. The spatial error coefficient, λ, is positive and significant in both versions of the SEM hedonic model and, as should be expected, the Moran statistics for error dependence is negligible in both as well. The difference in the Akaike information criteria for SEM versions of the hedonic model further indicates that version 2 is the superior specification. As opposed to the version 2 model, the sign of the kitchen facilities parameter is positive in version 1, but its relatively large standard error indicates that it also is not significantly different than zero. All other parameters estimated in the global SEM models have relatively small standard errors and can be considered significant, including the unexpected positive estimated coefficients for the density of low and middle income houses and age of the housing stock. In general the global findings are that housing prices are positive in floor space, pre-1980 construction, private kitchen facilities, tap water availability, and construction for low and middle income purchasers. The global model also indicates that house prices increase in the proportion of urban population in a county and in its net migration, but decrease in its percent minority population. Each of the three parameters for the location characteristics has the expected sign.

Table 3 Housing price model: parameter estimates and related statistics from SEM regressions using interval (Version 1) and dichotomized (Version 2) housing ratio characteristics

	Version 1	Version 2
Intercept	7.4191 (0.0532)	8.2630 (0.0379)
Floor space (m^2) per household	0.0067 (0.0003)	0.0069 (0.0003)
% houses built before 1980	0.0126 (0.0005)	0.1788 (0.0100)
% houses with private kitchens	0.0001 (0.0003)	−0.0078 (0.0099)
% houses with tap water	0.0025 (0.0002)	0.0756 (0.0109)
% houses with central hot water	0.0099 (0.0017)	ne
% houses with independent lavatory	ne	0.0762 (0.0111)
% low and mid income houses	0.0071 (0.0013)	0.0197 (0.0097)
% urban population	0.0036 (0.0002)	0.0046 (0.0002)
% minority population	−0.0019 (0.0003)	−0.0018 (0.0003)
Net-migration ratio	0.0019 (0.0001)	0.0023 (0.0002)
λ	0.8129 (0.0123)	0.8006 (0.0128)
Akaike information criterion	−898.13	−655.16
Moran's I	−0.070	−0.067

Standard errors are in parentheses; ne indicates not estimated

The optimal bandwidth for the GWR specification of version 1 of the hedonic model is 146 counties. Evaluated at the median, the parameters estimated by GWR are consistent with those estimated in the version 1 global SEM (Table 4), except for the case of kitchen facilities, which is negative at the median. It is positive in the GWR models, however, as in the SEM estimation, in the highest quartile of its range. The more general consistency in sign between the GWR and SEM specifications is true across the interquartile range of most of the parameters, but there are exceptions. The parameter for low and middle income housing is negative—the expected sign—in its lowest quartile and the parameter for minority population is unexpectedly positive at its highest quartile. Even where there is consistency in sign across the interquartile range, the values of the parameters show meaningful variation. For example, the SEM parameter estimate for floor space is 0.0067 with a standard error of 0.0003. Both the first quartile and third quartile estimates for the GWR are much more than two standard errors beyond the SEM global estimate. The optimal bandwidth for the version 2 GWR is 218 counties. The results of the version 2 GWR model are generally consistent with both the version 2 SEM and the version 1 GWR (Table 4). The major exception is the parameter for percent low and middle income

Table 4 Housing price model: parameter estimates from geographically weighted regressions using interval and dichotomous housing characteristics

	1st Quartile interval/ dichotomous	Median interval/ dichotomous	3rd Quartile interval/ dichotomous
Intercept	6.8187	7.2599	7.7309
	7.3325	7.7057	8.0308
Floor space (m^2)	0.0033	0.0063	0.0095
	0.0047	0.0068	0.0095
% houses built before 1980	0.0051	0.0110	0.0170
	0.1084	0.1751	0.2408
% houses with private kitchens	−0.0038	−0.0008	0.0015
	−0.0595	−0.0008	0.0430
% houses with tap water/independent lavatory	0.0007	0.0026	0.0039
	0.0423	0.1135	0.1674
% houses with central hot water	0.0044	0.0160	0.0449
	0.0219	0.0939	0.1612
% low and mid income houses	−0.0020	0.0086	0.0172
	−0.0485	−0.0102	0.0473
% urban population	0.0021	0.0043	0.0062
	0.0043	0.0059	0.0075
% minority population	−0.0040	−0.0010	0.0028
	−0.0031	−0.0009	0.0023
Net-migration ratio	0.0024	0.0046	0.0072
	0.0039	0.0062	0.0092
Moran's I for residuals	0.160		
	0.199		

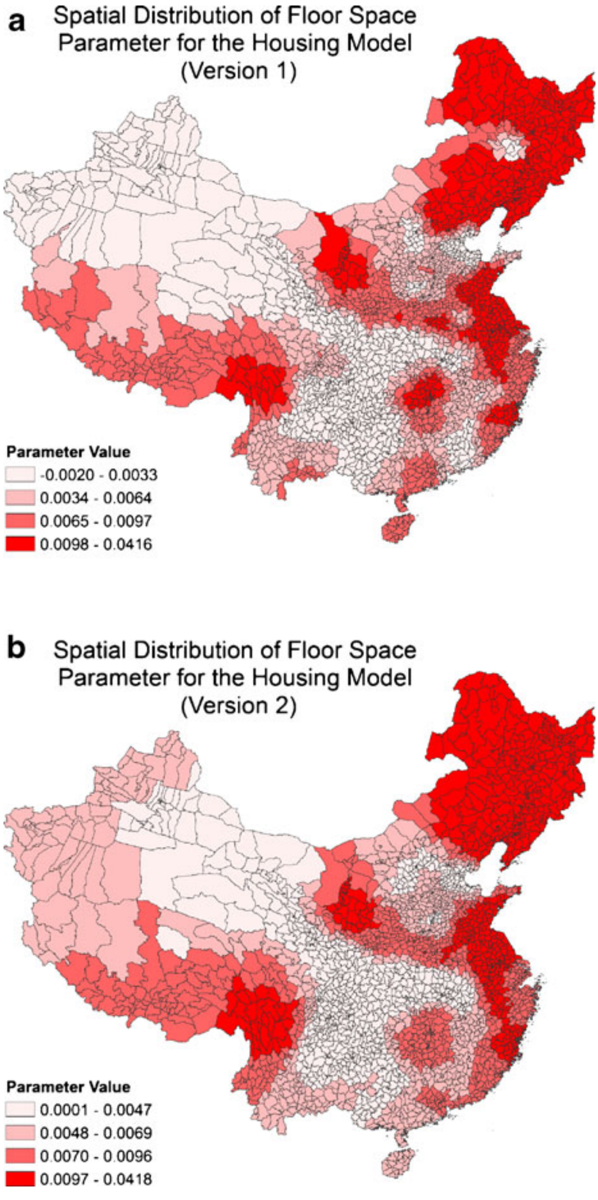
houses, which is positive in the global model but is negative at the median in GWR version 2 (but positive in version 1).

Both version 1 and version 2 have relatively small levels of spatial autocorrelation in their errors, Moran's I values are 0.160 and 0.199 respectively. Individual parameter estimates have generally weak correlations. In version 1 of the model, only the floor space parameter and the intercept have a Pearson correlation with absolute value greater than 0.5 (−0.509). Those two parameters have a much stronger correlation in version 2 of the model (−0.822), and the intercept is also correlated with the net migration parameter in that version of the model (−0.664). Monte Carlo tests indicate there is significant geographical variability in each of the estimated parameters with at least 98% confidence across both versions of the model, and broad spatial patterns are identifiable. While dominant spatial patterns in the parameters seem explicable in terms of China's regional growth, smaller regions of inconsistency and spatial outliers require additional investigation as to their causes.

The floor space and net migration parameters provide examples. A significantly positive relationship between house price and floor space is predominant in most of China (Fig. 1), as indicated by local values of the regression parameter from the GWR version 1 model. The major exceptions are in Xinjiang in the northwest and in south central China from around southern Shaanxi into Guangxi. Note, however, that Guangdong Province, which is the eastern neighbor of Guangxi, is marked by county-level units with relatively high parameter values. That is true of most of the coastal region of China, with some exception in the south in Fujian Province and nearer Shanghai in Shandong Province. Map patterns of the floor space parameter are largely consistent between versions 1 (Fig. 1a) and 2 (Fig. 1b) of the model. There is however a distinction in Northeastern China, where there is a small region of negative parameter values in version 1 that does not appear in version 2. The distribution of net migration parameter values also indicates a coastal bias (Fig. 2), but in this case larger parameters tend to be located farther north including Shandong Province, with weaker parameters down the southern coast. Other regions with higher parameters are in the west of China including Xinjiang and Tibet, indicating some of the stronger effects on housing prices in their response to in-migration are actually where the real estate market is quite thin. There are areas of China where the link between house prices and net migration is unexpectedly negative or very weak, including counties in Guangdong Province—a center of in-migration. As in the case of the floor space parameter there is general agreement between the map patterns of version 1 (Fig. 2a) and version 2 (Fig. 2b) results for the migration parameter.

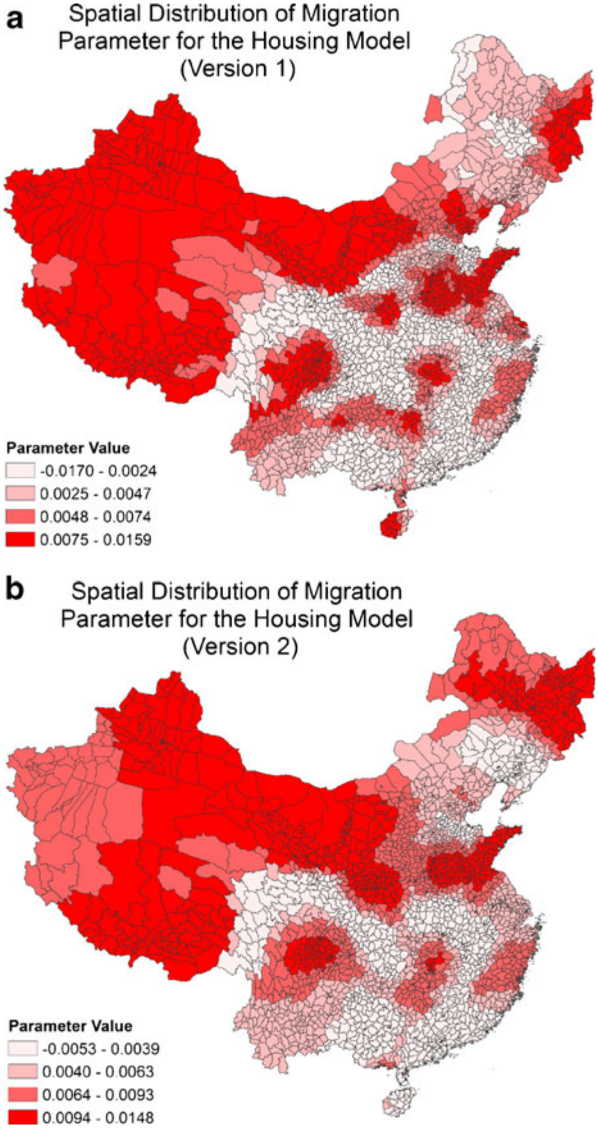
County-level goodness-of-fit (local R^2) for both versions of the model has a marked regional pattern, with a major cluster of high local R^2 values in southern coastal China anchored by Guangdong Province (Fig. 3). As in the cases of the individual parameter maps, the map results for version 1 (Fig. 3a) and version 2 (Fig. 3b) are quite similar. There is a smaller cluster of high local R^2 values ringing Shanghai and another, larger, one that appears centered on greater Beijing. In general it appears that the model's goodness-of-fit is generally lower away from the coast in eastern China, and in the country's northeast and west where populations are lower and market reforms in the housing sector were not well established in 2000.

Fig. 1 The spatial distribution of the floor space parameter for the housing model. **a** Model version 1. **b** Model version 2



As expected, global SEM results indicate that rent is positive in median house price and in the density of commodity apartments at the county level (Table 5). Both versions of the model have negligible spatial autocorrelation in their errors. Both versions have optimal bandwidths of 146 county-level units in their GWR models. In addition parameter estimates in both version 1 and version 2 (Table 6) bear strong similarity to the results in their respective global specifications. There is no strong correlation (absolute value of $r \geq 0.5$) between any parameters in either specification. The strongest correlation coefficient among the parameters is

Fig. 2 The spatial distribution of the migration parameter for the housing model. **a** Model version 1. **b** Model version 2



−0.328, between the intercept and the commodity apartment coefficient in version 2. Based on Monte Carlo testing, the intercept and both regression coefficients appear to have significant spatial variation in version 1 and in version 2. The local R^2 values indicate that the goodness-of-fit of the GWR rent models are highest in coastal areas, northwest of Beijing, and in Northeast China and part of Sichuan Province (Fig. 4). Goodness-of-fit is fairly poor in the western part of the country. As in the case of house prices, the general applicability of the rent model is strongest in those parts of China that have experienced the most growth and where real estate reforms were targeted.

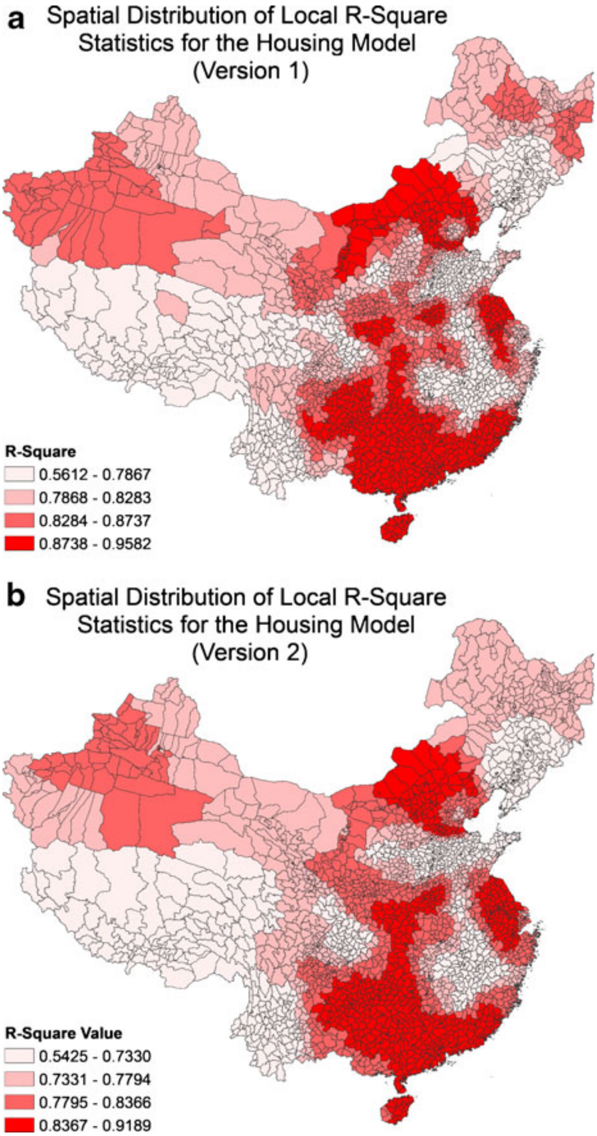


Fig. 3 The spatial distribution of local R-square statistics for the housing model. **a** Model version 1. **b** Model version 2

Summary and Conclusion

China's market reforms were extended to residential real estate in urban areas beginning in 1979 but the process was only completed in 1998. The analysis in this paper relies on data from 2000, so it concerns a very young housing market that is still undergoing development. In fact most of China's population has yet to enter the residential market and still lives in self-built or other non-market housing. Despite its youth, the analysis in this paper indicates that China's residential market has many characteristics that can be considered typical of long established ones. For example,

Table 5 Rent model: parameter estimates and related statistics from SEM regressions using interval (Version 1) and dichotomized (Version 2) apartment ratio characteristic

	Interval	Dichotomized
Intercept	3.1964 (0.0240)	3.0670 (0.0236)
Median house price	0.000015 (0.000001)	0.000021 (0.000001)
% commodity apartments	0.0480 (0.0026)	0.3462 (0.0176)
λ	0.6288 (0.0185)	0.6060 (0.0192)
Akaike information criterion	2742.39	2707.98
Moran's I	−0.048	−0.042

Standard errors are in parentheses

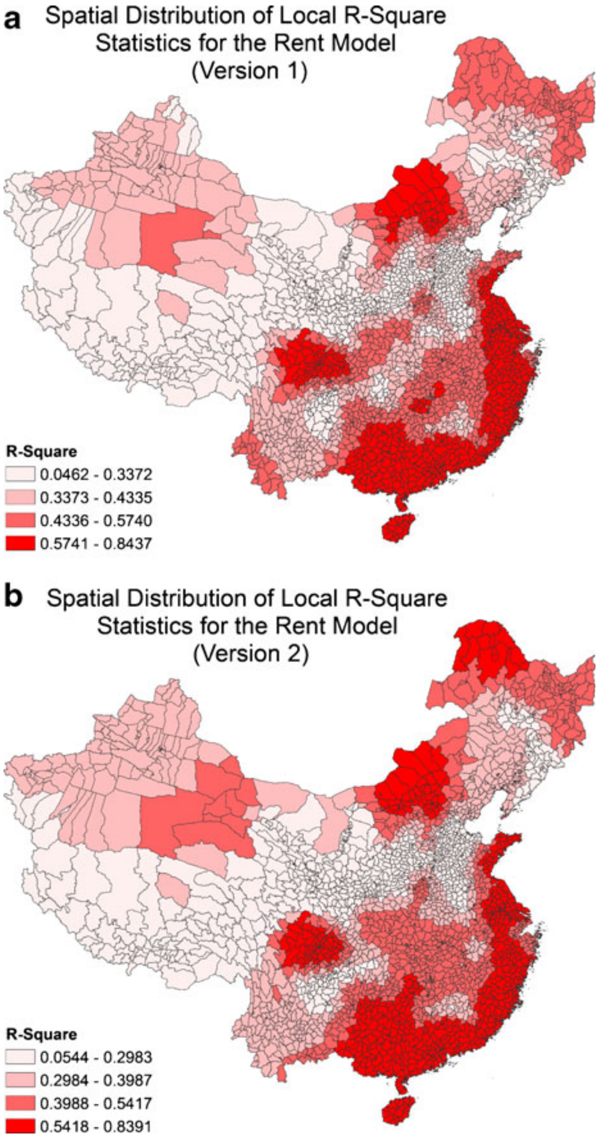
prices were found to be positive in floor space and several residential amenities. Prices were further found to react as expected under market conditions to several locational characteristics that are effective indicators of local demand, including in-migration. Using an abbreviated model, rents were also found to behave as expected in a typical residential market. The price and rent models, however, also indicated that there is significant variation in their responses across county-level units in China. Each parameter, including intercepts, in each GWR specification of both price and rent models had significant spatial variation, with that variation extending to sign in some cases. Such variation is indicative of the lack of maturity in China's residential real estate market that is likely a result of what amounts to a spatially targeted system of reforms that largely have their incidence in the eastern—particularly coastal—part of the country.

While most of the results in this paper do point toward typical market behavior, they are certainly far from conclusive. Parameters estimates on density of affordable housing and kitchen facilities had unexpected signs, for example. Hedonic models

Table 6 Rent model: parameter estimates from geographically weighted regressions using interval and dichotomous housing characteristics

	1st Quartile interval/ dichotomous	Median interval/ dichotomous	3rd Quartile interval/ dichotomous
Intercept	2.9325	3.1062	3.3677
	2.8190	2.9971	3.2318
Median house price	0.000003	0.000017	0.000029
	0.000016	0.000028	0.000039
% commodity apartments	0.0449	0.0692	0.0954
	0.2110	0.3163	0.4434
Moran's I	0.118		
for residuals	0.116		

Fig. 4 The spatial distribution of local R-square statistics for the rent model. **a** Model version 1. **b** Model version 2



are often affected by missing variable bias and the one used in the analysis reported here is likely not an exception. Further research should include environmental characteristics such as air and water quality which are often found to be important in residential price models (Zheng and Kahn 2008). More importantly, the analysis of housing prices and rents in this paper is conducted at the fairly gross, county-level, scale. That scale is useful for an initial evaluation of the national spatial characteristics of the residential market, but future research using individual residences would obviously add considerable insight into residential market behavior in China as it continues its maturation.

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