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**THE HEALTH CONSEQUENCE
OF RISING HOUSING
PRICES IN CHINA**

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The Health Consequence of Rising Housing Prices in China*

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Abstract

China has experienced a rapid boom in real estate prices in the last few decades, leading to a substantial increase in living costs and heavy financial burdens on households. Using an instrumental variable approach, this paper exploits spatial and temporal variation in housing price appreciation linked to individual-level health data in China from 2000 to 2011. We find robust evidence that increases in housing prices significantly raise the probability of residents having chronic diseases. This negative health impact is more pronounced among individuals from low-income families, households that purchased rather than inherited or was allocated the home, and those who migrated from rural to urban areas. We also find evidence that marriage market competition exacerbates these negative health effects, particularly for males and parents with young adult sons. Further empirical results suggest that housing price appreciation induces negative health consequences through increased work intensity, higher mental stress, and reduced sleep time. This paper provides a novel explanation to the increased prevalence of chronic diseases in China.

JEL Codes: G51, I10, I12, R21, R31

Keywords: Housing Prices; Chronic Diseases; Health; Marriage Competition; China

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1 Introduction

The Chinese real estate market has experienced a rapid boom since the early 2000s, with housing prices growing at an annual rate of 8.5 percent,¹ higher rate than both average household disposable income (Chen, Hao and Stephens, 2010; Chen and Wen, 2017) and GDP. China's first-tier cities have experienced an even faster growth in housing prices, with an average annual growth rate of 13.1 percent from 2003 to 2013 (Fang et al., 2016). Recent studies have found that such rapid housing appreciation has affected individuals and households enormously. The rise in housing prices has been found to tighten household consumption (Su, Yin and Tao, 2018; Waxman et al., 2019), reduce female labor force participation (Fu, Liao and Zhang, 2016), increase marriage age (Wrenn, Yi and Zhang, 2019), and raise the probability of inter-generational co-residence (Li and Wu, 2019), among other things. Yet little is known about the health consequences of the rapid growth in housing prices.²

The Chinese housing market was centrally regulated until the late 1990s, when China enacted sweeping reforms privatizing the housing sector. Since then, the real estate market has experienced a remarkable boom. Despite rapid housing appreciation, house ownership rates have been growing steadily, reaching over 90 percent in 2004, as shown in Figure A.1 in Appendix A. Such high home ownership rates originate from Chinese social norms, in which owning a home is preferred to renting. However, this preference for home ownership has imposed heavy financial burdens on households in light of the surging housing prices. Two groups in particular are especially vulnerable to this rapid housing price growth - young unmarried men buying a house to increase their competitiveness in the marriage market (Wei, Zhang and Liu, 2017), and rural migrant workers seeking to settle in urban areas (Hu and Wang, 2019). There is a strongly-rooted marriage culture in China in which men are obliged to obtain a house before marriage, and this makes the strain of needing to buy a home particularly severe for young men on the marriage market (Wrenn, Yi and Zhang, 2019). Also, many cities' housing policies impose additional barriers to home ownership on those who do not have a local household registration, which in many cases affects mainly the migrant population seeking to settle in urban areas; added housing price appreciation makes it even harder to purchase a house in cities.

Increases in housing expenditure and the financial stress induced by it may directly affect individual's mental health. One may also work more intensely to alleviate their financial burdens when housing prices are high, which could lead to both mental and physical health

¹ Author's own calculation based on housing price data (2004-2013) obtained from various issues of China Statistical Yearbook.

² Li and Liu (2018) is an exception, studying the impact of housing prices on mental health of migrants in urban China.

problems. As a response to this increased work intensity, individuals may also get less sleep or reduce their leisure time, resulting in a higher probability of chronic diseases. In this paper, we explore the health consequences of rising housing prices in China, in particular its impact on chronic diseases, by exploiting the temporal and spatial variation in housing market boom since the late 1990s.

We consider the prevalence of chronic diseases to be the main health outcome of interest in this paper. This arises from the concern that individuals' physical health is less likely to react immediately and intensely to changes in housing prices, and more likely to respond gradually and to be reflected by chronic diseases. In the last few decades, China has witnessed a significant increase in the prevalence of chronic diseases. As shown in Figure 1, about 16 percent of individuals had chronic diseases in the early 2000s; by 2011 this number had almost doubled. The share of those with hypertension shows a similar trend. The increased prevalence of chronic diseases could be induced by a wide range of factors, including an unhealthy diet, lack of physical activities, tobacco use, excessive alcohol intake, constant mental stress (Fine et al., 2004), and environmental pollution (Ko and Hui, 2012), among other things. In this paper, we propose that the rapid rise in housing prices could be an additional driving force.

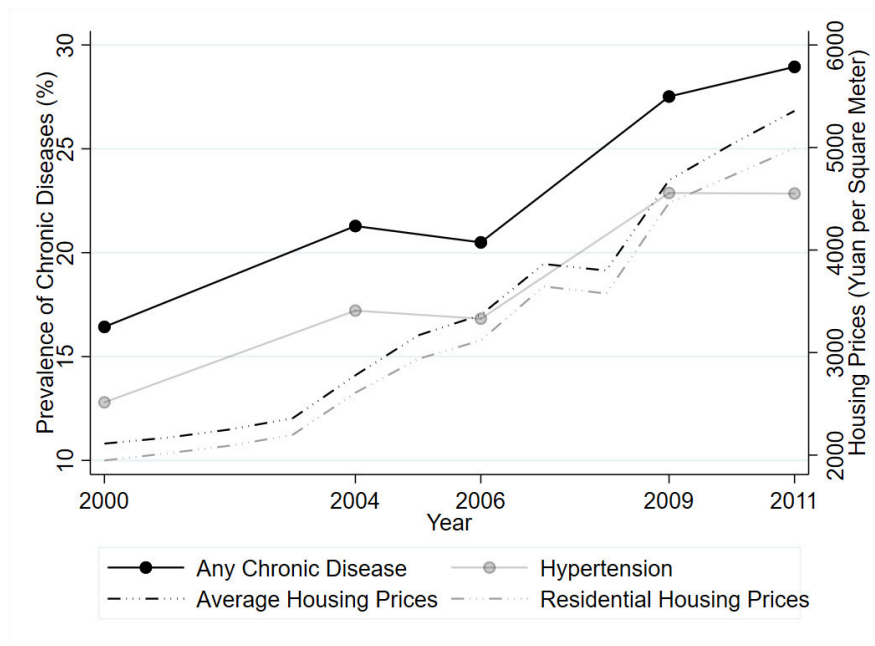


Figure 1: Average Housing Prices and the Prevalence of Chronic Diseases in China, 2000-2011

Data sources: Housing price data are from various issues of *China Real Estate Statistics Yearbook*. The prevalence of chronic diseases and hypertension is calculated based on the CHNS dataset.

We link individual health records with housing prices at the province level to explore whether individuals in areas experiencing larger housing price growth rates are more likely to acquire chronic diseases. Individual-level data is collected from the China Health Nutrition Survey (CHNS) between 2000 and 2011, when China experienced enormous housing price appreciation. The main advantage of this database lies in the fact that it reports rich information on an individual's health conditions and health-related activities, such as diagnoses of over sixteen types of chronic diseases as well as biomarkers of blood pressure and lifestyle and health-related behaviors, including consumption of tobacco and alcohol, sleep time, sports activities, etc. These allow us to precisely and comprehensively measure individuals' health conditions and explore the potential mechanisms behind changes in health outcomes. Additionally, the panel nature of the data allows for further controls for unobserved time-invariant confounders by including individual fixed effects, which could substantially reduce potential omitted variable biases. To identify a causal link between housing prices and health, we employ an instrumental variable approach. We construct the instrumental variable for housing prices by combining exogenous variations in both initial land supply across provinces and national interest rates over the years. The former captures cross-province housing supply capacities and the latter tackles the national trend in housing demand. Our instrumental variable implies that a province initially more constrained by land supply would witness a higher growth in housing prices when confronting a positive housing demand shock nationwide.

Our empirical results show several interesting findings. First, we find robust evidence that individuals from provinces that experienced more rapid growth in housing prices were more likely to have chronic diseases and hypertension relative to others. Such a negative health impact holds when we use alternative measures of housing prices and health outcomes, when we control for additional factors, and when we apply an alternative instrumental variable. We also find that the health consequences of the high housing prices in China are highly heterogeneous among groups. Specifically, the negative health effects were mainly experienced by individuals from relatively low-income households, while those from high-income families were immune to the negative health shock. We also find that only individuals from households that purchased their houses had a higher probability of encountering chronic diseases, while those acquiring houses through other means, such as through inheritance or through house allocation by their workplace until the late 1990s, experienced no such negative effects. Lastly, we find that the health effect was more pronounced among individuals who migrated from rural to urban areas than their counterparts who stayed in urban or rural areas.

The negative health consequences of housing prices are linked to various channels. First, we find evidence suggesting that the marriage culture and the increasing marriage market

competition due to male surplus play a crucial role in strengthening the negative health effects. This is reflected in our findings that the health effect is more pronounced among males than females. In addition, the negative health effects are only significant for parents with at least one adult son during the housing boom, but are not significant for those without sons, those having only old sons who are most likely to be married, or young sons who have not yet entered the marriage market. We also find that increasing working activities among the employed, higher mental stress, and a reduction in sleep time among the young cohort play a role in shaping the health consequences of housing prices.

Our paper contributes to several strands of literature. First, it adds to an emerging set of literature that documents the socio-economic impact of rising housing prices in China. At the individual and household level, rising housing prices in China are shown to, for example, induce higher saving rates (Chamon and Prasad, 2010; Wan, 2015), lead to declines in consumption (Waxman et al., 2019), discourage entrepreneurial activities (Li and Wu, 2014), increase wages (Liang, Lu and Zhang, 2016), delay marriages (Wrenn, Yi and Zhang, 2019), increase intergenerational co-residence (Li and Wu, 2019), and discourage female labor force participation among house owners (Fu, Liao and Zhang, 2016). For firms, the high housing prices have also been found to reduce firms' innovation (Rong, Wang and Gong, 2016) and crowd out fixed capital investment (Han and Lu, 2017). Our paper is among the first to examine the health consequences of rising housing prices in an emerging economy context and contributes to this strand of literature by providing additional evidence on the health impact of Chinese housing market.

Second, this paper complements the broader literature that documents the health effects of socio-economic changes like globalization and pollution. Fan, Lin and Lin (2020) find negative health effects of input tariff reductions following China's WTO accession, through increased working hours. Export expansion is shown to increase infant mortality due to deterioration of environmental quality (Bombardini and Li, 2020). He, Liu and Zhou (2020) documents that air pollution due to straw burning significantly increases mortality. The findings in this paper underline the role of an important albeit less researched factor - housing prices - in causing health problems.

This paper is also related to the growing literature studying the interplay between housing prices and the marriage market, and particularly marriage competition, in China. Due to a high ratio of males to females in China, the marriage rate has declined in recent years (Huang and Zhou, 2015) and marriage market competition, particularly among unmarried young males, has grown. This has led to increasing housing demand, since owning a spacious house is a way to increase one's competitiveness in the marriage market, a factor which contributed to the appreciation of housing values (Wei, Zhang and Liu, 2017). In this paper, we find that marriage market competition is not only a source of housing value

appreciation, but is also an important channel through which housing prices induce negative health consequences.

The rest of the paper proceeds as follows. Section 2 describes the main data sources used in the paper and discusses measures of housing prices and health outcomes. Section 3 presents our empirical approach and identification strategies. Section 4 reports empirical results, including the main results, a rich set of robustness checks, and heterogeneous effects. Section 5 empirically examines several potential mechanisms, and section 6 concludes.

2 Data and Measures

Our primary dataset is the China Health and Nutrition Survey (CHNS). It is an individual-level longitudinal survey conducted by the Population Center at the University of North Carolina at Chapel Hill since 1989.³ Earlier waves of the survey cover nine provinces, and it was extended to 15 in 2011; the provinces are selected to ensure that the sample is representative for China's eastern, central, and western regions. For consistency concerns, we only consider the nine provinces surveyed in all waves in our analysis. The sample within each province is drawn from a multi-stage, random cluster sampling approach. The raw dataset includes around 3,456 to 7,319 households and 11,860 to 20,914 individuals across years. An important feature of the dataset is that it reports rich information on individuals' health status, sickness history, and health-related behaviours, such as diet, tobacco consumption, alcohol intake and physical activities, etc. This allows for a precise measure of individual's health condition, as well as potential behaviors associated with health status. It also records detailed information on individual's demographic characteristics, labor market participation, and household characteristics, which enables us to control for an extensive number of confounders that may affect individuals' health. Compared to other individual-level databases in China that are often repeated cross-sections, the longitudinal nature of the CHNS allows us to track individuals across time and control for unobserved time-invariant characteristics by including individual fixed effects in the estimation.

The CHNS conducted ten waves until 2015, which well covered the period of China's housing market boom as well as a long period before the boom.⁴ For the purpose of this paper, we mainly use the five waves ranging between 2000 and 2011. The choice of the sample is based on two concerns. First, the Chinese housing market was not liberalized until the end of the 1990s and the data on housing prices is only available since 1998. As we link

³More details about the CHNS could be found from <https://www.cpc.unc.edu/projects/china>.

⁴The ten waves are for the years of 1989, 1991, 1993, 1997, 2000, 2004, 2006, 2009, 2011, and 2015.

the CHNS data with housing prices, our sample could only start from 2000. However, in later discussions, the first four waves prior to 2000 are used for a pre-trend analysis. Second, all waves except 2015 contain detailed information on three common types of chronic diseases - hypertension, diabetes, and cardiovascular diseases (including heart diseases and stroke) - whereas the 2015 wave reports only information on cancer. We therefore exclude 2015 from our main analysis in order to have consistent measures of the outcome variable. The 2015 wave, however, additionally surveyed individuals' mental health status. We use this cross-sectional data to investigate the impact of housing prices on mental health, a potential channel through which housing prices affect chronic diseases.

We link the CHNS data with records of housing prices at the province level using individuals' province location.⁵ Our province-level data on housing prices is from various issues of the *China Real Estate Statistics Yearbook*, and housing prices are measured in Chinese Yuan per square meter. The *China Real Estate Statistics Yearbook* reports the average selling price for each type of commercialized buildings by use, including all residential houses, villas and deluxe apartments, and office and business buildings, as well as the average price of all types of commercial buildings. We use the natural log of average residential housing prices as our main measure of housing prices as they affect the entire population, while other types may only affect certain groups of individuals. However, we consider the price of business and office buildings and the average price level of all types of buildings as alternative measures to check the robustness of our results.

We restrict our sample to working-age population aged 15 to 60. The final sample includes 25,179 observations of 8,472 individuals from nine provinces. Table A.1 in the appendix presents descriptive statistics of our key variables. The average housing price is 2,988.52 Yuan (approximately 426 US dollar) per square meter, and the average residential housing price is slightly lower at 2,787.79 Yuan (approximately 400 US dollar) per square meter. A typical Chinese house is often in the 90 square meters range, which is usually the smallest size that a standard family could dwell in while fulfilling their needs in urban China. The total cost reaches 36,000 US dollars. This is over 18 times the annual income of the average person in our sample. In contrast, the median price of a house in a major US metropolitan area is about 162,000 US dollars, which is less than 3.5 times the average annual income (Dettling and Kearney, 2014). These numbers imply that housing expenditures impose a much higher burden on Chinese than American households.

The main measure of our health outcome is the prevalence of chronic diseases. This arises from the concern that changes in housing prices may not induce health problems immediately, but rather individuals may accumulate health problems gradually, which eventually may develop into chronic diseases. The identification of individuals with chronic

⁵We get access to the province location only for each individual.

diseases is based on either self-reported health information or on their biomarker records. An individual is said to have a chronic diseases if he or she has either hypertension, diabetes, or cardiovascular diseases. We consider the prevalence of hypertension as an alternative, specific measure of the health outcome of interest.⁶ One concern is that diabetes and cancer are relatively new chronic diseases, the prevalence of which may rise with technological changes. However, hypertension is diagnosed at a very low cost, and has been prevalent in China since the late 1980s. Considering hypertension as an alternative measure of adult health could further rule out the increase in chronic diseases due to heterogeneous technology changes in health care across provinces. Table A.1 shows that the overall prevalence of chronic diseases is about 20 percent and the prevalence of hypertension is about 16 percent. Compared to the U.S., where nearly half of the population suffers from at least one type of chronic disease (Raghupathi and Raghupathi, 2018), China has a much lower prevalence.

3 Empirical Methodology

The aim of this paper is to uncover the health consequences of rising housing prices. To this end, we link individuals' health outcomes from the CHNS with province-level housing prices to investigate whether individuals from provinces with a larger growth in housing prices are more likely to have health problems than others. Specifically, we consider the following linear probability model:

$$P(CD_{iht} = 1) = \alpha + \beta HP_{pt} + \lambda \mathbf{X}_{it} + \delta \mathbf{Y}_{ht} + \gamma \mathbf{Z}_{pt} + \mu_i + v_t + \epsilon_{iht} \quad (1)$$

where CD_{iht} is a dummy variable indicating whether an individual i from household h and province p has chronic diseases in year t . HP_{pt} denotes measures of housing prices for province p in year t . The coefficient of interest β measures the average effects of rising housing prices on the probability of having chronic diseases.

\mathbf{X}_{it} is a vector of individual characteristics that may affect one's health conditions, including age, age squared, education level, marital status, etc. \mathbf{Y}_{ht} denotes a set of household characteristics, including household income, household size, and household registration type (rural or urban). We also control for a set of time-variant health-related factors at the province level, denoted by \mathbf{Z}_{pt} . These include the number of medical personnel per 1,000 local residents as a measure of availability of healthcare facilities, the natural log of

⁶ An individual has hypertension if the average systolic blood pressure is over 140 mmHg or the average diastolic blood pressure reading is over 90 mmHg.

waste water emissions (in 10,000 tons) and sulfur dioxide emissions (in tons) to proxy for environmental quality, as well as local GDP per capita to account for the level of economic development. μ_i is individual fixed effects that control for both observed and unobserved time-invariant individual characteristics. As such, the identification of the health effects derives from variations in health conditions across time at the individual level. Notice that if an individual did not change household or province location, the individual fixed effects also capture household-level and province-level time-invariant characteristics that may be related to health. These could capture systematic differences in health conditions, for example, between households with and without genetic diseases, or between individuals from the relatively more developed eastern coastal provinces and those from the less developed inland regions. However, to account for the possibility that some individuals move out of a household (e.g. due to marriage or work changes), we check the robustness of our results by additionally controlling for household fixed effects. In addition, year fixed effects (v_t) are included to account for common macro shocks to all individuals across years. ϵ_{ihpt} is the error term. We cluster standard errors at the community level in our main specification to allow for potential correlations between individuals within communities, but also cluster standard errors at the province level to check the robustness of our results.

The identification of a causal relationship between housing prices and health relies on the assumption that housing prices are uncorrelated with the error term. Such an assumption would be violated if there are unobserved factors that are simultaneously correlated with both housing prices and individuals' health conditions, and the coefficient estimate of housing prices from equation (1) would be biased. One example of such factors is internal migration. Migrant inflows may increase the demand for houses, especially in regions where migrants without a local household registration are allowed to purchase houses, thus inducing higher housing prices.⁷ Meanwhile, migrant workers may have relatively poor health conditions, particularly due to the fact that they often work in occupations with poorer working conditions and earn relatively lower incomes (e.g., Meng, 2012; Han and Li, 2017). As such, rising inflows of migrants would lead to an overestimate of the health effects of housing prices. However, with a lack of accurate records of migrants, and due to the fact that housing policies are often designed at the city level, we are not able to control for the impact of migration. Unobserved factors related to local economic development could also bias our estimates. Local housing prices may be inflated due to regional economic growth. If high prevalence of chronic diseases hinders economic growth, the estimate could be underestimated. However, economic growth could also be associated with higher transportation activities and industrial production, which in turn increases pollutant emission and leads to a higher prevalence of chronic diseases. In this

⁷In many Chinese cities, housing policies prohibit migrants without a local household registration from purchasing houses as a tool to adjust housing prices.

way, the effect of housing prices on health is likely to be overestimated. Although including local GDP per capita in our regressions could capture differential economic growth across provinces, heterogeneity in local infrastructure development and production activities still remains a concern.

To address the potential endogeneity of housing prices, we employ an instrumental variable approach. An ideal instrument should be correlated with housing prices but uncorrelated with the error term. Our identification strategy relies on exogenous shocks that explain variations in housing prices through affecting housing demand and supply. Specifically, our instrumental variable is constructed as the pre-sample land supply at the province level multiplied by national interest rates, with the former measuring housing supply constraints across provinces and the latter capturing national trends in housing demand.

Housing supply elasticity with respect to geography-based measures of land supply is a commonly used instrumental variable in predicting housing prices in the U.S. (e.g., Saiz, 2010; Mian and Sufi, 2011; Chaney, Sraer and Thesmar, 2012; Stroebel and Vavra, 2019). The intuition is that regions with relatively less elastic housing supply would observe a higher growth in housing prices with a positive housing demand shock. In the spirit of this strand of literature, studies on Chinese housing prices often use land supply as the instrument since housing supply elasticity cannot be directly estimated due to data availability constraints.⁸ Land use is strictly regulated in China by the central government and provincial government (Liang, Lu and Zhang, 2016), allowing land supply to be plausibly exogenous for cities and individuals. However, this may not apply in our context, as our housing price measure is at the province level. If provincial government's decision on land supply is correlated with other policies that may affect individual's health, land supply does not meet the exclusion restriction of a valid instrumental variable. There is evidence that income from land sales is an important source of local government revenue (Han and Kung, 2015; Waxman et al., 2019), which in turn determines government's expenditure on public utilities, including health related items like medical and health care, as well as environmental protection. To alleviate such concerns, we use pre-sample land supply in 1997 across provinces, rather than time-variant land supply, to capture the possibility that provinces with higher initial land supply constraints would witness a larger increase in housing prices than others given a housing demand shock. Figure A.2a in Appendix A shows the scatter plots of changes in log housing prices between 2000 and 2011 and the initial land supply for our sample provinces. It shows a clear negative correlation such that provinces with a higher initial land supply, such as Liaoning and Heilongjiang, observed a smaller increase in housing prices.

⁸For example, Waxman et al. (2019) and Li and Wu (2019) use land sales from the local government to real estate companies as a proxy for land supply to predict city-level housing prices and further identify the impacts on consumption and inter-generational co-residence, respectively. Liang, Lu and Zhang (2016) instruments housing prices with land supply quotas to examine the impact on wages.

In order to predict variations in housing prices over time, we interact pre-sample land supply with national long-term interest rates. A similar instrument is adopted by Chaney, Sraer and Thesmar (2012) and Rong, Wang and Gong (2016). Long-term interest rates not only add time variations but also capture a national trend in housing demand, as the mortgage rates are approximately the same as the long-term interest rates in China (Rong, Wang and Gong, 2016). Lower interest rates indicate easier access to mortgage, which stimulates housing demand (Berger et al., 2017). As changes in interest rates are designed by the central government and could hardly be affected by a single province, they are reasonably exogenous to the error term. Specifically, our instrument takes the following form:

$$IV_{pt} = ls_{p,1997} \times i_t \quad (2)$$

where $ls_{p,1997}$ denotes land supply of province p in 1997. We measure land supply using planned urban residential area divided by urban population. Data on planned urban residential area and urban population in 1997 is collected from *China Environment Statistical Yearbook* (1998). i_t is the average level of national long-term interest rate in year t .⁹ Our instrument indicates that regions with limited land supply are disproportionately more affected by a rise in national housing demand due to a lower interest rate, and we therefore expect that our instrumental variable is negatively correlated with housing prices. Figure A.2b in Appendix A displays the scatter plots of changes in log housing prices against changes in our instrumental variable as defined in equation (2) and shows a negative correlation between the two, as expected.

A remaining concern regarding our instrumental variable is that it may violate the exclusion restriction if national interest rates affect individual's health conditions through channels other than housing prices. However, interacting national interest rates with the initial land supply at the province level as in equation (2) ensures that the differential effects of national interest rates across provinces are proportional to initial land supply constraints, thereby ensuring that our instrumental variable could only affect individual's health through housing prices.

In addition, we consider an alternative instrumental variable in line with the strand of literature documenting spatial correlation of housing prices. Spatial diffusion of housing prices is found pervasive in many regions such that housing prices are affected by a price shock emanating from surrounding areas, e.g. in the U.S. (Guerrieri, Hartley and Hurst, 2013; Brady, 2014), the Netherlands (Teye and Ahelegbey, 2017), the UK (Holly, Pesaran and Yamagata, 2011), Taiwan (Chen, Chien and Lee, 2011), and China (Gong, Hu and

⁹In cases where interest rates were adjusted several times within a year, we calculate a weighted average using the share of active days for each rate as weight.

Boelhouwer, 2016). We follow this strand of literature and calculate the average housing prices of all bordering provinces as an alternative instrument for the housing prices of one specific province, and we expect a positive correlation between the two.

4 Empirical Results

4.1 Main Results

Table 1 reports the 2SLS estimation results of equation (1) using initial land supply multiplied by interest rates as the instrumental variable for housing prices and the incidence of any chronic diseases as the dependent variables. We start from a baseline specification controlling for individual characteristics in column (1), and add household and province characteristics in columns (2) and (3), respectively. In all specifications we control for individual fixed effects and year fixed effects, and cluster standard errors at the community level.

Table 1: Average Residential Housing Prices and the Incidence of Chronic Diseases

| | (1) | (2) | (3) |
|---|----------------------|----------------------|----------------------|
| ln Residential Housing Prices | 0.579** (0.240) | 0.573** (0.240) | 0.541** (0.240) |
| <i>First-stage Results</i> | | | |
| Initial Land Supply \times Interest Rates | -0.512*** (0.045) | -0.512*** (0.045) | -0.526*** (0.029) |
| First-stage SW F -Statistics | 128.14 | 127.21 | 334.93 |
| Individual Controls | ✓ | ✓ | ✓ |
| Household Controls | | ✓ | ✓ |
| Province Controls | | | ✓ |
| Individual Fixed Effects | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ |
| Observations | 25,179 | 25,179 | 25,179 |
| R^2 | 0.607 | 0.607 | 0.609 |

Notes: The table reports the 2SLS estimation results of equation (1). Individual controls include age, age squared, marital status, and education level; household controls include ln household gross income, ln household size, and a rural or urban dummy; province controls include healthcare staff per 1,000 local residents, ln waste water emissions in 10,000 tons, ln sulfur dioxide emissions in tons, and GDP per capita. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

The first-stage results show that our instrumental variable is negatively correlated with housing prices and is highly significant in all specifications, indicating that provinces with a tighter initial land supply constraint would observe a larger increase in housing prices when facing a positive demand shock due to lower interest rates. The sufficiently high value of the first-stage F -statistics allows us to reject the null hypothesis of weak instrument.

The second-stage results show that all estimates for the coefficient on average residential housing prices are positive and significant at the 5 percent level, indicating that individuals in provinces with a larger growth rate of housing prices are more likely to have chronic diseases than those in provinces with lower growth rates. The size of the coefficients does not change much across specifications with and without household and province controls. Focusing on the preferred specification in column (3), the estimated coefficient suggests that a 10 percentage rise in housing prices leads to a 5.41 percentage point increase in the prevalence of chronic diseases, holding other variables constant. This impact is economically sizable, as it represents about a quarter of the average prevalence of chronic diseases in China during our sample period (20 percent as in Table A.1 in Appendix A).

To check whether our results in Table 1 are sensitive to model selection, we re-estimate equation (1) using logit regressions where we employ the same instrumental variable as specified in equation (2). Logit regressions with individual fixed effects only take into account individuals that experienced changes in health status during the sample period, thereby resulting in a reduced sample size. The results are reported in Table A.2 in Appendix A. In line with the results in Table 1, the estimated coefficients of residential housing prices are positive and significant in all specifications, indicating a positive effect of the rising housing prices on the incidence of chronic diseases. Our preferred results in column (3) suggest that a 10 percent increase in housing prices is linked to a 0.54 rise in log-odds of chronic diseases holding other variables constant. This translates to a 10 percentage point increase in the prevalence of chronic diseases for an average individual who has a 20 percent probability of having chronic diseases (see Table A.1).¹⁰

4.2 Robustness Checks

In this section, we perform a series of robustness checks on our main results. Specifically, we check whether the negative health effects of housing prices are sensitive to specifications controlling for additional confounders, to alternative measures of housing prices and health outcomes, and to an alternative instrumental variable. We also perform a falsification test to check whether the health consequences of rising housing prices were driven by long-term trends.

¹⁰For an individual with a probability of 10 percent to have chronic diseases, the estimated coefficient in column (3) corresponds to a 6 percentage point rise in the prevalence of chronic diseases with a 10 percent increase in housing prices.

4.2.1 Controlling for Additional Variables

In Table 2 we report the results of a set of robustness checks by controlling for additional confounders. In all specifications, we use the prevalence of chronic diseases as the outcome variable and average residential housing prices (in natural log) as the housing price measure, including a full array of control variables and fixed effects as in column (3) of Table 1. In column (1), we add household fixed effects. This arises from the concern that some individuals might change households over time, for instance, due to marriage or divorce.¹¹ Individuals could face different household environments that could affect their health. Consider the case of an individual moving out of a parents' household due to marriage, creating a new one; he or she will deal with new relationships and play a different role in the new household. As such, unobserved household characteristics related to such a change might result in different health outcomes. As shown in column (1), however, the estimated coefficient on the housing price measure remains robust to the inclusion of household fixed effects.

Table 2: Average Residential Housing Prices and the Incidence of Chronic Diseases: Robustness Checks

| | HH FE (1) | Region-year FE (2) | Trade Openness (3) | All Controls (4) | Prov. Clustering (5) |
|-------------------------------------|--------------------|--------------------------|--------------------------|------------------------|----------------------------|
| In Residential Housing Prices | 0.550** (0.241) | 0.380** (0.174) | 0.521** (0.233) | 0.446** (0.191) | 0.446*** (0.077) |
| Control Variables | ✓ | ✓ | ✓ | ✓ | ✓ |
| Individual Fixed Effects | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ | ✓ |
| First-stage SW <i>F</i> -Statistics | 338.36 | 338.76 | 251.65 | 413.77 | 26.60 |
| Observations | 25,139 | 25,179 | 25,179 | 25,139 | 25,139 |
| <i>R</i> ² | 0.610 | 0.612 | 0.609 | 0.614 | 0.614 |

Notes: This table reports the 2SLS estimation results of equation (1). Column (1) adds household fixed effect to the baseline specification; column (2) adds region-year fixed effects; column (3) controls for trade openness; and column (4) include all these additional control variables. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the province level in column (5) and at the community level in columns (1)-(4).

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

Due to the large geographic area of China, there are remarkable regional variations in development. The coastal regions of eastern China are more developed than the central regions, which are in turn better developed than the far western inland regions. To alleviate the imbalanced development across regions, the Chinese government launched various programs, particularly during our sample period, to stimulate and support the development

¹¹ 1.84% of individuals in the sample changed their household between 2000 and 2011. However, this change could only be observed among those who reside within the sample provinces throughout the survey years. If individuals moved outside of the sample provinces, a problem of sample attrition would arise.

of the western and central regions. For example, the Great Western Development Strategy was launched in 2000 to develop the western regions, including infrastructure constructions, improving foreign investment environment, protecting the ecological system, attracting talents, etc. In 2004, the Rise of Central China Plan was implemented to accelerate the development of the central regions. The average individual health condition in different regions may follow heterogeneous trends due to imbalanced development and due to macro policy changes across years. To account for the differential trends across regions and the potential impacts of macro policy changes, in column (2) of Table 2, we include region-year fixed effects. The coefficient of the housing price measure stays positive and statistically significant, whereas the magnitude turns slightly smaller. A 10 percent increase in residential housing prices leads to a 3.80 percentage point higher probability of having chronic diseases.

During the period under investigation, one notable change in the Chinese economy was the rapid pace of openness to the world market following China's accession to the WTO in 2001. It follows that tariff rates dropped substantially and both imports and exports experienced a surge. Trade openness could affect health through various channels, such as changes in income (Dai, Huang and Zhang, 2018), changes in job opportunities due to import competition (Autor and Handel, 2013; Wang, Kis-Katos and Zhou, 2020) or export expansion (Feenstra, Ma and Xu, 2019) and the associated adjustment in work intensity (Fan, Lin and Lin, 2020), and changes in environmental quality (Cherniwchan, 2017; Lin, 2017; Bombardini and Li, 2020). As Chinese provinces are exposed to trade openness differently (Han, Liu and Zhang, 2012), trade openness may have heterogeneous effects on individuals' health across provinces. To control for the impact of trade openness, we include the share of trade over GDP in column (3) as a measure of trade openness. Our results suggest that the impact of housing prices on chronic diseases is not affected much.

In column (4), we add all four additional control variables and the coefficient of housing prices remains positive and significant. Our last robustness check is to examine whether our results are sensitive to the level of clustering of the standard error. Specifically, we cluster standard errors at the province level rather than the community level. This allows individuals to be correlated within provinces. As shown in column (5), the coefficient estimate keeps positive and significant at the 1 percent level. In sum, our results in Table 2 show consistent evidence that rising housing prices in China increases the occurrence of chronic diseases.

4.2.2 Alternative Measures of Housing Prices and Health Outcomes

In our main specification, we use average residential housing prices as the main housing price measure. However, the average residential housing price level is an average over

all types of residences, including villas and deluxe apartments. Rising prices of luxurious houses may not strongly affect health, since only a specific group of high-income individuals, who might not be as sensitive to housing price increases as average-income individuals, are affected. While we do not have price data on residential buildings excluding luxurious houses, the *China Real Estate Statistics Yearbook* separately reports the average price level of villas and deluxe apartments. This allows us to partial out the effects of the more expensive houses by controlling for their average prices in our main regressions. Column (1) of Table 3 reports the 2SLS results. Surprisingly, it shows that higher prices of villas and deluxe apartments are correlated with a lower probability of having chronic diseases. This could be due to the fact that houses in China are often seen as a signal of wealth and social status (Wei, Zhang and Liu, 2017). The rising prices of the high-end apartments may allow some high-income individuals to show wealth and improve social status, which raises the level of happiness. Controlling for the prices of high-grade houses, the estimated coefficient on average residential housing prices remains significantly positive but slightly larger in size compared to the result in column (3) of Table 1. An individual in a province with a 10 percent increase in housing prices has a 5.98 percentage point higher probability of being affected by chronic diseases.

Table 3: Housing Prices and the Incidence of Chronic Diseases: Alternative Measures of Housing Prices and Health Outcomes

| Housing price measures | Panel A: Chronic Diseases | | | Panel B: Hypertension | |
|--------------------------------------|---------------------------|----------------------|----------------------|-----------------------|----------------------|
| | Residential | Business and Office | All Types | Residential | All Types |
| | (1) | (2) | (3) | (4) | (5) |
| ln Housing Prices | 0.598** (0.263) | 0.403** (0.189) | 0.469** (0.209) | 0.485** (0.225) | 0.420** (0.195) |
| ln High-class Residence Prices | -0.091* (0.052) | | | | |
| <i>First-stage Results</i> | | | | | |
| Initial Land Supply × Interest Rates | -0.472*** (0.028) | -0.707*** (0.090) | -0.607*** (0.035) | -0.526*** (0.029) | -0.607*** (0.035) |
| First-stage SW <i>F</i> -Statistics | 280.44 | 61.61 | 300.35 | 335.07 | 300.27 |
| Control Variables | ✓ | ✓ | ✓ | ✓ | ✓ |
| Individual Fixed Effects | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 25,179 | 25,179 | 25,179 | 25,175 | 25,175 |
| <i>R</i> ² | 0.608 | 0.607 | 0.610 | 0.627 | 0.629 |

Notes: The table reports the 2SLS estimation results of equation (1) using alternative measures of housing prices (Panel A) and the incidence of hypertension (Panel B) as dependent variables. Average housing prices in column (3) indicate the average price level of all types of houses, including residential houses, villas and high-grade apartments, office buildings and houses for other business activities. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

In column (2) of Table 3, we consider the price of houses for business use and office buildings. The health effects of increasing prices of business and office buildings are less clear-cut, as some entrepreneurs and firms own buildings while others have to rent, with the former being more likely to have positive health effects while the latter are more likely to have negative effects. The first-stage result shows that our instrumental variable, initial land supply multiplied by national interest rates, could well predict changes in the housing prices of business and office buildings. The second-stage result stays positive and significant, but smaller in size than the effects of residential housing prices, as in column (3) of Table 1. A 10 percent price increase is associated with a 4.03 percentage point rise in the incidence of chronic diseases.

In column (3) of Table 3, we consider the average price of all type of houses, including residential houses, office buildings and houses for business activities, as well as other types. Again, the result reveals a positive association between housing prices and the incidence of chronic diseases. Overall, the results in Table 3 confirm our main finding that rising housing prices lead to negative health consequences.

One additional concern with the health outcome measure is that chronic diseases are self-reported. The rising prevalence of chronic diseases could be due to technological improvements allowing for cheaper, easier detection of chronic diseases, or due to an increasing awareness of health leading individuals to check their health more actively. The inclusion of individual fixed effects and year fixed effects in our main specification could partially account for time-invariant heterogeneity in health values among individuals and technological differences across provinces, as well as common trends in the development of medical equipment across years. The unobserved changes in individual's health values or province-specific technology improvement across time could still bias our estimates. To address these concerns, we consider an alternative, specific type of chronic disease, hypertension, as our measure of health outcomes. The awareness of hypertension in China can be traced back to the 1980s, and detection of hypertension is easily performed at a low cost. The increasing prevalence of hypertension could therefore hardly be driven by technological improvement or rising awareness.

Panel B of Table 3 reports the estimation results using the prevalence of hypertension as the outcome variable, conditional on a full set of control variables. column (4) and (5) use the natural log of average residential housing prices and the average price of all types of houses as the dependent variable, respectively. The estimated coefficients on both housing price measures are positive and significant, indicating that individuals in provinces with a higher level of housing prices are more likely to have hypertension. Compared to the effects on chronic diseases in general (column (3) in Table 1 and Table 3), the effects on hypertension are slightly smaller.

We further investigate the effects on other measures of health, including the incidence of being sick in the past four weeks, as well as three proxies for overweight and obese. The results are set out in Table A.3 in Appendix A. The estimated coefficients on residential housing prices are only marginally significant in column (3) in which we use an indicator of high waist circumference as the dependent variable; an increase in housing prices is associated with a higher likelihood of having high waist circumference. It is worthy mentioning that waist circumference is based on a direct measure, and is not likely to be affected by technological advancement in health care or rising awareness. Considering that high waist circumference is found to be highly correlated with the incidence of chronic diseases (Li et al., 2007), this set of results provide supportive evidence to our main findings.

4.2.3 Alternative Instrumental Variable

Table 4 presents the estimation results using the average residential housing prices of adjacent provinces as an instrument. The identification of a causal relationship hinges on the assumption that housing prices are spatially correlated but the housing prices in other provinces do not directly affect individual's health. In both specifications we cluster standard errors at the province level. The first-stage results show a highly significant correlation between home and neighbouring housing prices. However, the relatively low value of the first-stage F -statistic implies that we cannot reject the null hypothesis that our instrument is weak.

Table 4: Average Residential Housing Prices and the Incidence of Chronic Diseases: Using Average Housing Prices in Neighboring Provinces as Instrument

| | Chronic Diseases (1) | Hypertension (2) |
|--|-------------------------|---------------------|
| In Residential Housing Prices | 0.422** (0.134) | 0.461** (0.194) |
| <i>First-stage Results</i> | | |
| In Average Prices in Neighboring Provinces | 0.341*** (0.046) | 0.341*** (0.046) |
| First-stage SW F -Statistics | 4.46 | 4.46 |
| AR Wald Test p -Value | 0.0009 | 0.0000 |
| Control Variables | ✓ | ✓ |
| Individual Fixed Effects | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ |
| Observations | 25,179 | 25,175 |
| R^2 | 0.611 | 0.627 |

Notes: This table reports the 2SLS estimation results of equation (1) using ln average residential housing prices in neighboring provinces as an alternative IV. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the province level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

The coefficient estimates on housing prices in the second stage remain positive and significant for both chronic diseases and hypertension as shown in column (1) and (2), respectively. However, due to the presence of weak instruments, we should interpret these results with caution. In Table 4, we also report an Anderson-Rubin Wald test (Anderson and Rubin, 1949) that provides weak instrument robust inference. It tests the null hypothesis that the coefficient of our endogenous variable, average housing prices, equal zero in the structural equation. It is equivalent to estimating equation (1) using the predicted housing prices as the regressor and testing whether the coefficient of this variable equals zero. The p -value shows that the null hypothesis is rejected at the 1 percent level, suggesting that the IV estimates are significantly different from zero. The results in Table 4 therefore suggest a causal relationship between housing prices and the prevalence of chronic diseases and hypertension. Comparing with the results in Table 1 using initial land supply multiplied by national interest rates as an instrument, these results suggest that the negative health effects of rising housing prices are underestimated when using the alternative instrument. In addition, clustering standard errors at the community level as in Table 1 reports insignificant results, implying that the results are sensitive to the level of clustering standard errors. We therefore prefer results using national interest rates interacted with initial land supply as the instrumental variable.

Finally, we conduct a falsification exercise to check whether the negative health consequences of rising housing prices are driven by long-run time trends. This would happen if individuals in provinces that experienced a relatively higher growth in housing prices were more likely to suffer from chronic diseases even before the boom of the real estate market. To this end, we match the health records of individuals from the first four waves of the CHNS collected between 1989 and 1997 with housing prices in 1999-2007 and in 2004-2012, respectively, i.e., 10 and 15 years following the survey. Future housing prices should not be correlated with the prevalence of chronic diseases unless there were some long-run common trends. We repeat the estimation of equation (1) and the results are reported in Table 5. The estimated coefficients of housing price measures are insignificant in both specifications. This rules out the possibility that the effects of housing prices on chronic diseases identified earlier are driven by long-run trends.

4.3 Heterogeneity of the Results

4.3.1 Household Income and Type of House Acquisition

We first examine whether households with different income levels are affected differentially, since low-income households may endure heavier financial burdens due to housing pur-

Table 5: Residential Housing Prices and the Incidence of Chronic Diseases: Falsification Test

| Dependent Variable: Chronic Disease in 1989-1997 | Housing Prices in 1999-2007 (1) | Housing Prices in 2004-2012 (2) |
|---|------------------------------------|------------------------------------|
| ln Residential Housing Prices | -0.222 (0.149) | -0.125 (0.143) |
| Control Variables | ✓ | ✓ |
| Individual Fixed Effects | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ |
| First-stage SW F -Statistics | 27.24 | 65.93 |
| Observations | 24,242 | 24,242 |
| R^2 | 0.547 | 0.550 |

Notes: This table reports the 2SLS estimation results of equation (1) using the prevalence of chronic diseases in 1989-1997 as dependent variable. Column (1) and (2) use residential housing prices in 1999-2007 and those in 2004-2012, respectively. All specifications include individual and household-level control variables. Individual controls include age, age squared, marital status, and education level; household controls include ln household gross income, ln household size, and a rural or urban dummy. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

chases (Fang et al., 2016). In columns (1) to (2) in Table 6, we classify households into low- and high-income based on the average income level across years within provinces. Households with a higher than median income level are classified as high-income, and others are classified as low-income. The empirical results show that housing prices increase the prevalence of chronic diseases for both types of households, with the magnitude of effect being larger for low-income households than for high-income ones. A 10 percent rise in average residential housing prices is associated with a 6.34 percentage point increase in the probability of having chronic diseases for individuals from low-income households. By contrast, those from high-income households experienced an increase of 5.22 percentage points. This suggests that individuals who are relatively financially constrained are more vulnerable to the negative effects of rising living costs induced by housing price growth.

We then check the potential heterogeneity by the type of house acquisition. Prior to the real estate market reform in the late 1990s, houses or living spaces were often provided for free by employers. Some homeowners managed to obtain their property at a low cost during the reform. Also, some households might inherit their houses from older family members. We expect that the negative health effects will be experienced by those who acquired their current house through direct purchase and not through inheritance, since home buyers are directly affected by surging housing prices. Thanks to data on the method of acquisition of individuals' homes provided by the CHNS, we are able to distinguish between households that purchased their current house and those who did not. Such information is only available for 2009 and 2011. We reasonably assume that the type of houses holds for the period of 2000-2006. Indeed, purchasing a home often requires plenty of planning and arrangements before the final acquisition, and the housing price rise between 2000 and 2006 could have already

Table 6: Average Residential Housing Prices and the Incidence of Chronic Diseases: Income and Type of House Acquisition

| | Income Level | | Own Purchase | |
|--------------------------------|--------------------|--------------------|--------------------|------------------|
| | Low (1) | High (2) | Yes (3) | No (4) |
| In Residential Housing Prices | 0.634** (0.266) | 0.522** (0.218) | 0.441** (0.222) | 0.504 (0.330) |
| Control Variables | ✓ | ✓ | ✓ | ✓ |
| Household Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| First-stage SW F -Statistics | 349.29 | 477.03 | 525.82 | 62.64 |
| Observations | 12,522 | 12,506 | 15,529 | 7,241 |
| R^2 | 0.372 | 0.329 | 0.338 | 0.357 |

Notes: The table reports the 2SLS estimation results of equation (1) based on various sub-samples. Households with above-median incomes are classified as high-income, and the others are low-income. Households purchasing own house are identified according to their self-report in the CHNS 2009 and 2011. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

affected potential home buyers. Columns (3) and (4) report the estimation results for both households that purchased their current house and those that did not, the latter including households that inherited their house from senior family members or obtained it from their workplace at a very low cost during the housing market reform in the late 1990s. In line with our expectation, we see that only individuals from households who purchased their current house experienced an increased likelihood of having chronic diseases following the housing price appreciation. By contrast, such a negative health impact is not significant for households that obtained the current house through other ways.

4.3.2 Hukou Registration Status

In China, the household registration (hukou) is classified into urban and rural types. The hukou system is closely linked to local social welfare and restricts the access to social welfare for the migrants without an urban hukou. In order to pursue higher incomes and better quality social welfare, many rural individuals migrate to urban areas, wishing to obtain an urban hukou. However, one condition of applying for a local hukou is having fixed and stable residence in the city and many regions require house ownership as the pre-requisite.¹² Although housing appreciation mainly occurred in urban China (Fang et al., 2016), the health effect could be further extended to rural households that intend to move to urban areas.

¹²The hukou system restricted rural to urban migration before the late 1990s when some regions implemented a hukou reform in small towns and cities that allowed rural migrant workers to obtain a local hukou if certain conditions were satisfied. Since 2001, such reform has been expanded to some medium and large cities, and in 2014, the hukou reform was expanded nationwide.

To investigate potential differential impacts on households with various registration status, we classify households into four types: urban households, rural households, households that changed from a rural to urban type, and those changed from a urban to rural type.¹³ The change from a rural to urban type could happen due to two reasons. First, the whole communities might be reclassified from a rural to urban type administratively during the process of urbanization (Gan et al., 2019). The registration type of households in the affected regions is changed accordingly. The second reason is rural-urban migration due to work or marriage purposes. While the first type does not necessarily induce household changes, the latter is more likely to involve a change of household, e.g. moving out of the parents' household and forming a new one in the city.

Table 7: Average Residential Housing Prices and the Incidence of Chronic Diseases: Household Registration Type

| | Urban (1) | Rural (2) | Rural-Urban (3) | Urban-Rural (4) |
|-------------------------------------|-------------------|-------------------|---------------------|--------------------|
| ln Residential Housing Prices | 0.568* (0.317) | 0.534* (0.272) | 0.877*** (0.315) | 1.039 (0.632) |
| Control Variables | ✓ | ✓ | ✓ | ✓ |
| Household Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| First-stage SW <i>F</i> -Statistics | 182.92 | 305.85 | 143.27 | 31.29 |
| Observations | 7,635 | 12,597 | 3,854 | 1,065 |
| <i>R</i> ² | 0.378 | 0.347 | 0.321 | 0.327 |

Notes: The table reports the 2SLS estimation results of equation (1) based on sub-samples. Rural-urban households are those that changed from a rural to urban hukou registration type during the sample period, and urban-rural households are those that changed from an urban to rural hukou registration type. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

We estimate the health effects for each type of households separately, and the results are set out in Table 7. The estimation results indicate that the negative health effect induced by rising housing prices is prevalent among individuals from urban and rural households as well as those from the rural-urban ones. Individuals changing from an urban to rural hukou type, however, do not observe significant health effects.

The effect on rural households is similar in magnitudes to that on urban households, although the housing price appreciation occurs mainly in urban areas. One potential explanation is that rural-urban migrants who work and live in the urban areas hold a rural hukou. Non-local hukou holders are often constrained to or even prohibited from purchasing houses in many cities (Chen, 2016). Meanwhile, migrant workers earn an average lower income than natives (e.g., Meng, 2012; Han and Li, 2017). With the additional

¹³In our sample, 30.35% of households are always registered as urban, 50.11% always rural, 14.54% changed from a rural to urban hukou type, and only about 5.00% changed from an urban to rural type.

constraints due to their hukou status as well as the relatively lower income level, the rise in housing prices imposes a high pressure on migrants who wish to purchase houses in the urban area (Li and Liu, 2018). Such effects could be further passed through to family members in the rural area, in particular parents, as many Chinese parents consider purchasing a house for their son as an obligation, particularly for marriage purposes. We explore the role of marriage market competition and the intergenerational effects more formally in Section 5.1.

Compared to the households which always hold an urban or rural hukou, a highly significant increase in the prevalence of chronic diseases with the housing price growth is observed among those changing from a rural to urban hukou, as shown in column (3). An individual would experience a 8.77 percentage point increase in the probability of having chronic diseases with a 10 percent rise in residential housing prices. The sizable effects could be due to the fact that those individuals are relatively young without much saving but often having to purchase a residence.¹⁴ They are thereby more vulnerable. An additional explanation is related to the way of house acquisition. It is likely that some urban households obtained their houses through housing allocation or through inheritance whereas it is hardly the case for the rural-urban migrants.

5 Examining Potential Mechanisms

Rising housing prices may cause an increase in the prevalence of chronic diseases through various mechanisms. We show in this section that the potential mechanisms for such a negative health effect include the following: (i) marriage market competition and intergenerational passing-through, which is particularly relevant for households with unmarried sons that wish to purchase houses; (ii) increased work intensity as a way to improve incomes; (iii) rising mental stress due to higher financial burdens; and (iv) changing lifestyles, such as less sleeping and leisure time due to higher work intensity. We empirically examine these potential mechanisms in the following discussions.

5.1 Marriage Culture and Marriage Market Competition

Historically, the marriage culture in China has led to many males feeling obligated to provide a living space for newly formed households, and owning a house is still a pre-requisite to marrying a potential wife in many regions (Hu and Wang, 2019). Additionally, houses in

¹⁴Those who moved from rural to urban areas are on average two years younger than those who always stayed in urban areas.

China are not only assets or consumption goods, but are also status goods that signal the wealth of a family (Wei, Zhang and Liu, 2017). Owning a spacious house could increase one's competitiveness in the marriage market among unmarried young males (Wei, Zhang and Liu, 2017; Hu and Wang, 2019) and also raise the probability of marrying a wife who is from a household with a similar wealth level (Sun and Zhang, 2020). High housing prices, along with increased competition in the marriage market, however, have placed much pressure on unmarried young males as a result. We therefore expect that males are more likely to suffer from health problems compared to their female counterparts.

The potential negative health effects of the increasing housing prices on young males could also be passed down to parents, as Chinese parents are often obliged to provide financial support for house purchasing to improve their son's success in the marriage market. This originates from a strongly-rooted culture of ensuring the continuity of the family lineage and clan. An old Chinese saying even acknowledges that "there are three ways to be unfilial; having no sons is the worst". The rising housing prices therefore could particularly affect families with unmarried sons.

To investigate the role of marriage culture and marriage market competition, we first examine whether housing prices affect males and females differently. Columns (1) and (2) of Table 8 report the results by sex. It is evident that housing prices significantly raise the likelihood of getting chronic disease for males; a 10 percent increase in residential housing prices is associated with a 6.78 percentage point increase in the incidence of chronic diseases. On the contrary, the effects on females are smaller and only marginally significant. The stronger effects on males could be related to the high societal pressure put on males to purchase a home for marriage purposes.

We further examine the potential passing-through effects on parents. To this end, we classify parents into those who had at least one adult son aged 15 to 40 when interviewed and those who did not, and compare whether housing prices affect them differently. The results for the two types of parents are reported in column (3) and (4), respectively. It shows that only parents with at least one adult son who was about to enter or who had already entered the marriage market experienced the negative housing price induced health effects, whereas households with only daughters, or with older sons who have exited the marriage market, or with young sons who have not yet entered marriage market, are not affected significantly. The estimated coefficient in column (3) suggests that a 10 percent increase in residential housing prices leads to a 6.76 percentage point increase in the incidence of chronic diseases for parents having at least one adult son. In sum, the results in Table 8 suggest that the marriage culture and marriage market competition among males are an important channel through which housing prices cause negative health consequences in China.

Table 8: Average Residential Housing Prices and the Incidence of Chronic Diseases: The Role of Marriage Market Competition

| | Sex | | With Adult Son | |
|--------------------------------|--------------------|-------------------|---------------------|------------------|
| | Male (1) | Female (2) | Yes (3) | No (4) |
| In Residential Housing Prices | 0.678** (0.335) | 0.418* (0.214) | 0.676*** (0.203) | 0.214 (0.402) |
| Control Variables | ✓ | ✓ | ✓ | ✓ |
| Individual Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| First-stage SW F -Statistics | 340.46 | 310.68 | 433.69 | 288.49 |
| Observations | 12,064 | 13,115 | 14,292 | 6290 |
| R^2 | 0.584 | 0.633 | 0.601 | 0.615 |

Notes: The table reports the 2SLS estimation results of equation (1) based on sub-samples. Parents with young adult son is defined as those having at least one male child aged 15 to 40 when interviewed. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

5.2 Work Intensity

As a response to the rising living costs with higher housing prices, one might increase work intensity to improve income and alleviate financial stress. This could be reflected by longer hours with an expectation of higher payment, or by finding a second job to broaden income sources. It is documented in the literature that long working time may cause both mental and physical health problems through reducing leisure time, raising inactivity, affecting sleeping quality (e.g., Sato et al., 2012; Afonso, Fonseca and Pires, 2017; Cygan-Rehm and Wunder, 2018; Wong, Dawson and Van Dongen, 2019). It is therefore likely that the high occurrence of chronic diseases following housing price growth is due to increased work intensity.

We examine the role of work intensity by estimating equation (1), replacing the outcome variable with various measures of work intensity. We specifically consider three measures of work intensity: weekly working days, daily working hours, and total weekly working hours (all in natural logs) and constrain our sample to those who reported to be working in the last week. The 2SLS regression results are reported in Table 9.

The results reveal a positive association between housing prices and work intensity in all specifications, suggesting that individuals from provinces that experienced a higher housing price growth worked more than those from other provinces. A 10 percent rise in housing prices increases weekly working days by 7.47 percent, daily working hours by 11.30 percent, and total weekly working hours by 23.26 percent. For an individual working an average of 5.46 days a week, 7.18 hours a day, and 39.49 hours a week, the estimated effects translate to working 0.41 more days a week, 0.81 more hours a day, and 9.19 more hours a week.

Table 9: Average Residential Housing Prices and Work Intensity

| | In Weekly Working Days (1) | In Daily Working Hours (2) | In Weekly Working Hours (3) |
|--------------------------------|----------------------------------|----------------------------------|-----------------------------------|
| In Residential Housing Prices | 0.747*** (0.277) | 1.130*** (0.373) | 2.326** (0.945) |
| Control Variables | ✓ | ✓ | ✓ |
| Household Fixed Effects | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ |
| First-stage SW F -Statistics | 432.72 | 433.22 | 378.30 |
| Observations | 16,935 | 16,776 | 15,570 |
| R^2 | 0.381 | 0.437 | 0.421 |

Notes: The table reports the 2SLS estimation results of equation (1) using measures of work intensity as the dependent variable. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

In Appendix B, we use the China Family Panel Survey or CFPS database (Institute of Social Science Survey, 2015) to cross-check this channel. Compared to the CHNS, the CFPS database reports information on whether an individual has a second job, supporting an investigation of changes in work intensity along the extensive margin. It also allows for a distinction between those with and without mortgage. The results in Table B.1 (columns (1)-(3)) show a positive impact of residential housing prices on the probability of having a second job, but the coefficients are insignificant. This suggests that adjustments of working intensity following housing price appreciation were mainly along the intensive margin. Overall, the above results imply that the negative health consequences of high housing prices in China could be due to rising working intensity.

5.3 Mental Stress

Mental health and physical health are closely related (Vaillant, 1979; Frerichs et al., 1982). Our results in Table 9 show strong evidence that individuals work more hours when housing prices are higher. The increased work intensity in turn may aggravate mental stress (Wong, Dawson and Van Dongen, 2019). Indeed, housing prices could affect mental health directly, even among those who are not working. This is particularly the case for the disadvantaged groups that have relatively poor financial conditions but a high demand for housing, like migrants, for instance (Li and Liu, 2018). To explore the role of mental stress in shaping the effects of housing prices on physical health, we regress housing prices on measures of mental health conditions.

Unfortunately, the CHNS did not collect data on individual mental health conditions until 2015, when a new module on mental stress was added to the survey. We therefore use self-reported memory conditions as an indirect measure of mental health for the sample of year 2000-2006, and a direct measure of mental health for the 2015 cross section.¹⁵ Specifically, the 2015 survey asked questions on the frequency of feeling stressful, frustrated, and anxious in the past month, with answers ranging from one to five, with one indicating “never” and five indicating “very often”. We define individuals who often or very often experienced any of these distressing feelings as mentally stressed. The regression results are reported in Table 10.

Column (1) reports the 2SLS results based on the 2000-2006 sample. The negative coefficient indicates that a rise in housing prices is associated with memory lapses. This provides suggestive evidence that housing prices might worsen mental health. Column (2) directly checks this relationship based on the 2015 survey. The dependent variable indicates whether an individual experienced stress “often” or “very often”, as opposed to “never”, “rarely”, or “only sometimes”. The results shows that individuals from provinces with high housing prices were more likely to feel stressful. It is worth noting that the result in column (2) is estimated from an OLS regression based on the 2015 cross section, and may be subject to omitted variable biases. Thus the coefficients on housing prices should be interpreted as a simple correlation.

Table 10: Residential Housing Prices, Memory and Mental Stress

| | Panel A: 2000-2006 (2SLS) | Panel B: 2015 (OLS) |
|---|-----------------------------------|------------------------|
| | Better Memory than Last Year (=1) | Any Mental Stress (=1) |
| | (1) | (2) |
| In Residential Housing Prices | -0.940* (0.503) | 0.065* (0.037) |
| <i>First-stage Results</i> | | |
| Initial Land Supply \times Interest Rates | -1.394*** (0.097) | |
| First-stage SW <i>F</i> -Statistics | 206.29 | |
| Control Variables | ✓ | ✓ |
| Household Fixed Effects | ✓ | |
| Year Fixed Effects | ✓ | |
| Observations | 1,499 | 7,515 |
| R^2 | 0.523 | 0.011 |

Notes: This table reports the 2SLS and OLS estimation results of equation (1) using measures of memory as the dependent variable for the sample from 2000 to 2006 (Panel A), and using measures of mental stress as the dependent variable for the sample of the 2015 cross section (Panel B), respectively. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

¹⁵Neither the data on memory nor on mental health is available for 2009 and 2011.

In Appendix B, we use the CFPS to cross-check the causal effects of rising housing prices on mental stress in which we use a depression score to measure mental health, and the results in Table B.1 show consistent findings. In all, these results support the hypothesis that mental health is a channel through which rising housing prices raise the incidence of having chronic diseases.

5.4 Changes in Lifestyle

As a response to increased work intensity and mental stress, individuals might also change their lifestyle. For example, one may increase the consumption of tobacco and alcohol to relieve stress (Ayyagari and Sindelar, 2010). Longer working hours may reduce time for relaxing activities, such as sleeping and exercise; such changes in lifestyle could increase the probability of chronic diseases. In Table 11, we examine whether housing prices change individuals' consumption of tobacco and alcohol, as well as their average time devoted to sports activities and sleep.

Table 11: Housing Prices and Changes in Lifestyle

| | Tobacco (=1) (1) | Alcohol (=1) (2) | Sport (=1) (3) | Sleep Hours (4) |
|---------------------------------------|------------------------|------------------------|----------------------|-----------------------|
| Panel A: Pre-1965 Cohort | | | | |
| In Average Residential Housing Prices | 0.198 (0.184) | -0.068 (0.183) | 0.013 (0.146) | 0.999 (0.861) |
| First-stage SW F -Statistics | 310.66 | 382.80 | 385.12 | 393.35 |
| Observations | 4,580 | 14,242 | 14,141 | 14,017 |
| R^2 | 0.580 | 0.602 | 0.426 | 0.353 |
| Panel B: Post-1965 Cohort | | | | |
| In Average Residential Housing Prices | 0.275 (0.193) | -0.218 (0.210) | 0.072 (0.162) | -1.996** (0.993) |
| First-stage SW F -Statistics | 355.85 | 552.24 | 552.23 | 564.48 |
| Observations | 2,835 | 10,868 | 10,809 | 10,710 |
| R^2 | 0.527 | 0.566 | 0.432 | 0.361 |
| Control Variables | ✓ | ✓ | ✓ | ✓ |
| Household Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ |

Notes: This table reports the 2SLS estimation results of equation (1) using an indicator variable indicating whether an individual consumed tobacco, alcohol, and did sports as dependent variable in columns (1)-(3), respectively. The dependent variable in column (4) is daily hours of sleep. All regressions include a full set of control variables as in column (3) of Table 1. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

Panels A and B display results for individuals born before 1965 and those born after 1965. This is motivated by the fact that the younger cohort may adjust their lifestyle and habits relatively easily (Tseng and Lin, 2008). We do not find evidence that individuals respond to the housing price appreciation by changing consumption of alcohol or tobacco or altering sports activities for either cohort. However, we do observe a cohort difference in the effect of housing prices on sleeping time. While the older cohort did not adjust their sleeping hours with housing price growth as shown in Panel A, the younger cohort significantly reduced their sleeping time (Panel B); a 10 percent increase in residential housing prices corresponds to a 0.20 hour (approximately 12 minutes) reduction in sleeping time. Changes in sleeping time are consistent with the results in Table 9 where housing prices increased work intensity. Results based on the CFPS as shown in Table B.1 in Appendix B confirm the negative effects of housing price appreciation on sleeping hours.

6 Conclusions

Housing prices have been increasing enormously in China since the late 1990s, when the market-oriented reforms to the real estate market were implemented. Recent studies find that the rapid housing price appreciation has profound effects on socio-economic outcomes, such as household consumption, labor market participation, fertility, marriage, etc. This paper widens the spectrum of this strand of literature by examining the health consequences of rising house prices in China, and investigates various potential mechanisms.

Using individual-level data from the CHNS between 2000 and 2011, which is further linked with province-level data on housing prices, we employ an instrumental variable approach to identify a causal relationship between housing prices and the prevalence of chronic diseases. We find robust evidence that the rise in housing prices in China significantly increased the prevalence of chronic diseases. Our baseline results imply that a 10 percent increase in residential housing prices induced a 5.41 percentage point rise in the probability of having chronic diseases. Such negative health effects are stronger for individuals from relatively low-income families, those who changed household types from a rural to an urban one, and those who purchased their current home. We carefully examine several potential channels through which rising housing prices induce negative health consequences. Our results suggest that the negative effects on health are exacerbated by the marriage related cultural norms, which lead to males feeling pressured to purchase a household before marriage, as well as the strong competition in the marriage market among young males. Such a negative health effect due to the stress of purchasing houses before marriage could be passed through to parents who have an adult son. Lastly, we find evidence that individuals respond to such an increase in housing prices by increased work intensity, higher mental stress, and

a reduction in sleep time. These findings not only comprehensively illustrate the health-related changes at the individual level, but also highlight the role of culture and social norms in shaping the health effects of rising housing prices.

The rapid real estate appreciation is not unique to China, and is pervasive in many other countries across the globe. Our findings based on the Chinese context shed light on the potential downside of the real estate market growth in other countries. While the prosperity of the real estate industry could contribute to economic growth, the unintended, negative health consequences should be considered in the design of housing policies.

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Appendices

A Additional Figures and Tables

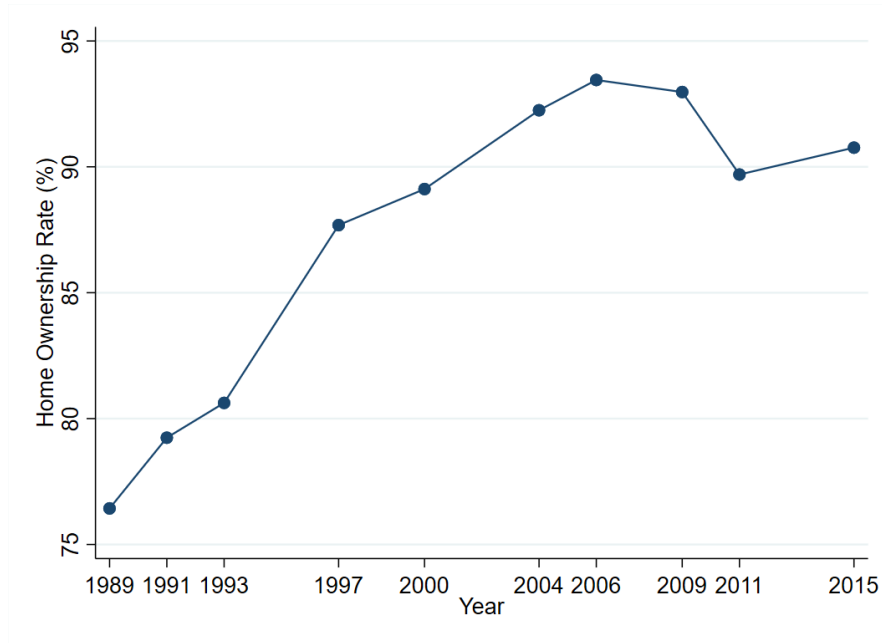


Figure A.1: Home Ownership Rate in China, 1989-2015

Data source: Own calculation based on the CHNS dataset. Home ownership rate is calculated as the share of households owning as opposed to renting the current home in total number of households.

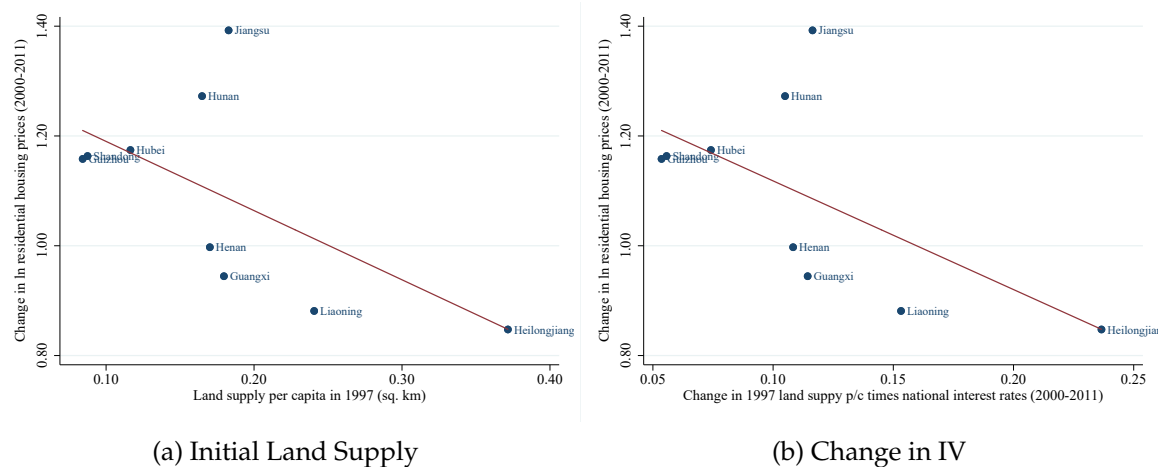


Figure A.2: Changes in Housing Prices, Initial Land Supply, and the Instrumental Variable

Notes: Figure A.2a shows the scatter plots of changes in ln housing prices between 2000 and 2011 and initial land supply by province. Initial land supply is measured as planned urban residential area in 1997 divided by urban population. Figure A.2b shows the scatter plots of changes in ln housing prices and the instrumental variable by province. The instrumental variable is defined as equation (2).

Table A.1: Summary Statistics of Key Variables

| | Mean | S.D. | Min. | Max. |
|--|------------|------------|-----------|------------|
| Housing Price Measures | | | | |
| Average residential housing prices | 2,787.790 | 1,079.143 | 1,181.430 | 6,145.200 |
| ln average residential housing prices | 7.859 | 0.389 | 7.074 | 8.723 |
| Prices of high-class residences | 5,403.959 | 2,341.454 | 1,923.895 | 11,643.360 |
| ln prices of high-class residences | 8.495 | 0.457 | 7.562 | 9.362 |
| Prices of business and office buildings | 5,219.048 | 1,735.704 | 2,619.710 | 9,504.770 |
| ln prices of business and office buildings | 8.505 | 0.335 | 7.871 | 9.160 |
| Average housing prices of all types | 2,988.518 | 1,105.142 | 1,385.020 | 6,554.410 |
| ln average housing prices of all types | 7.936 | 0.365 | 7.233 | 8.788 |
| Health Outcomes | | | | |
| Any chronic disease (=1) | 0.200 | 0.400 | 0.000 | 1.000 |
| Hypertension (=1) | 0.159 | 0.366 | 0.000 | 1.000 |
| Individual-level Characteristics | | | | |
| Female (=1) | 0.521 | 0.500 | 0.000 | 1.000 |
| Age | 43.684 | 10.342 | 17.000 | 60.000 |
| Married (=1) | 0.893 | 0.309 | 0.000 | 1.000 |
| Education level | | | | |
| Below primary | 0.144 | 0.351 | 0.000 | 1.000 |
| Primary school | 0.196 | 0.397 | 0.000 | 1.000 |
| Lower middle school | 0.378 | 0.485 | 0.000 | 1.000 |
| Upper middle school | 0.151 | 0.358 | 0.000 | 1.000 |
| Technical or vocational degree | 0.072 | 0.259 | 0.000 | 1.000 |
| University of college | 0.058 | 0.233 | 0.000 | 1.000 |
| Master or above | 0.001 | 0.027 | 0.000 | 1.000 |
| Household-level Characteristics | | | | |
| ln household income | 10.028 | 1.039 | 1.609 | 14.058 |
| ln household size | 1.268 | 0.385 | 0.000 | 2.565 |
| Urban (=1) | 0.387 | 0.487 | 0.000 | 1.000 |
| Household registration type | | | | |
| Urban (=1) | 0.304 | 0.460 | 0.000 | 1.000 |
| Rural (=1) | 0.501 | 0.500 | 0.000 | 1.000 |
| Rural-Urban (=1) | 0.145 | 0.353 | 0.000 | 1.000 |
| Urban-Rural (=1) | 0.050 | 0.218 | 0.000 | 1.000 |
| Own purchase of current house (=1) | 0.682 | 0.466 | 0.000 | 1.000 |
| With at least one adult son (=1) | 0.694 | 0.461 | 0.000 | 1.000 |
| Province-level Characteristics | | | | |
| Health care staff per capita | 51.166 | 13.185 | 23.098 | 72.808 |
| ln waste water emission (10,000 tons) | 12.383 | 0.600 | 10.917 | 13.302 |
| ln sulfur dioxide emission (tons) | 13.800 | 0.397 | 13.102 | 14.510 |
| GDP per capita | 21,977.893 | 12,957.851 | 4,317.000 | 62,290.000 |

Notes: Data on housing price measures is from various issues of the *China Real Estate Statistics Yearbook*. Health outcomes, individual- and household-level characteristics are abstracted from the CHNS database. Province-level characteristics are from various issues of *China Statistical Yearbook*. The number of observations is 25,179.

Table A.2: Average Residential Housing Prices and the Incidence of Chronic Diseases: Logit Bootstrapping 2SLS Regressions

| | (1) | (2) | (3) |
|-------------------------------|---------------------|---------------------|--------------------|
| ln Residential Housing Prices | 5.847*** (1.876) | 5.826*** (1.962) | 5.380** (2.097) |
| Individual Controls | ✓ | ✓ | ✓ |
| Household Controls | | ✓ | ✓ |
| Province Controls | | | ✓ |
| Individual Fixed Effects | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ |
| Observations | 7,451 | 7,451 | 7,451 |
| Pseudo R^2 | 0.096 | 0.098 | 0.100 |

Notes: This table reports the 2SLS logit model estimation results of equation (1) using national interest rates interacted with provincial land supply in 1997 as the instrumental variable (equation (2)). Individual controls include age, age squared, marital status, and education level; household controls include ln household gross income, ln household size, and a rural or urban dummy; province controls include healthcare staff per 1,000 local residents, ln waste water emissions in 10,000 tons, ln sulfur dioxide emissions in tons, and GDP per capita. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parenthesis.

Table A.3: Housing Prices and Additional Health Outcomes

| | Sick in the Past 4 Weeks (=1) (1) | Overweight (=1) (2) | High Waist Circumference (=1) (3) | High Waist-hip Ratio (=1) (4) |
|--------------------------------|--|---------------------------|--|--|
| ln Residential Housing Prices | 0.181 (0.174) | 0.137 (0.152) | 0.229* (0.124) | 0.002 (0.238) |
| Control Variables | ✓ | ✓ | ✓ | ✓ |
| Household Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| Year Fixed Effects | ✓ | ✓ | ✓ | ✓ |
| First-stage SW F -Statistics | 385.47 | 401.15 | 404.75 | 405.02 |
| Observations | 25,134 | 23,601 | 23,534 | 23,274 |
| R^2 | 0.284 | 0.441 | 0.407 | 0.349 |

Notes: This table reports the 2SLS estimation results of equation (1). Sick in the past four weeks is measured using self reports. One is overweight if the body mass index exceeds 25. Waist circumference is identified as being high if it is over 102cm for males and over 88cm for females. Waist-hip ratio is identified as being high if waist circumference exceeds 0.90 of hip circumference for males, and 0.85 for females. All regressions include a full set of control variables as in column (3) of Table 1, household fixed effects and year fixed effects. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

B The Role of Work Intensity, Mental Stress, and Sleep Time: Evidence from the CFPS

In this section, we cross-check the role of mental stress, sleep time and work intensity using an alternative dataset, the China Family Panel Survey (CFPS) (Institute of Social Science Survey, 2015) . It is a longitudinal survey of Chinese families and individuals starting from 2010, and four waves until 2016 are available. In each wave, more than 10,000 households were interviewed. It reports rich information on individual and household characteristics, with the focus on the economic and subjective well-being of the Chinese population. Compared to the CHNS, the CFPS covers a larger number of provinces, which allows for larger variations in housing appreciation across regions.¹⁶ The CFPS includes information on individual's mental health, time use, and whether one has a second job throughout the four waves. Additionally, it includes information on whether one has a mortgage or not. This allows us to examine whether housing prices affect individuals with and without mortgage differently.

Compared to the CHNS, it does not report much information on health, but it covers a wider range of provinces. More importantly, CFPS measures the adult psychological well-being using scales of the Center for the Epidemiological Studies of Depression (CES-D) (Radloff, 1977) . The CES-D is a validated screening tool for detecting depressive symptoms, and is suitable to be used across different racial, gender and age groups (Weissman et al., 1977) . The standard CES-D covers a total of eight to 20 items on mental health, and individuals are asked about the frequency of experiencing each item during the past week. The scale of each item ranges from one to five, indicating frequency from never, rarely, sometimes, often, to very often. The CFPS covers different numbers of CES-D items across years, and we calculate the average score to harmonize the variation in item numbers.¹⁷ Moreover, a few items are positively stated while most of the items are negatively stated. We reverse the scales of positive items to ensure that a higher scale indicates a worse mental well-being throughout the items. In addition, CFPS supports a distinction between those with and without mortgage.

We match the CFPS dataset with provincial housing prices and estimate equation (1) using measures of mental health, sleep time, and whether one has a second job as the dependent variable, allowing the distinction between mortgage payers and non-mortgage payers. Similar to the analysis using CHNS, we restrict the sample to individuals aged between 15 and 60 years old. As the CFPS covers a shorter panel than the CHNS, we control for household instead of individual fixed effects to allow for more variations. To address

¹⁶CFPS covers 25 provinces in 2010, 28 in 2012, 29 in 2014, and 30 in 2016.

¹⁷CFPS has six items on CES-D in 2010, 20 in 2012, 13 in 2014, and nine in 2016.

the endogeneity of housing prices, we use the same instrumental variable as specified in equation (2), i.e. initial land supply interacted with national interest rates. The 2SLS estimation results are set out in Table B.1.

Table B.1: Housing Prices, Mental Health, Sleep Time and the Probability of Having a Second Job: Evidence from the CFPS

| | Second Job (=1) | | | Depression | Sleep |
|---|----------------------|----------------------|----------------------|----------------------|---------------------|
| | All (1) | No Mortgage (2) | With Mortgage (3) | Score (4) | Hour (5) |
| In Residential Housing Prices | 0.054 (0.081) | 0.019 (0.099) | 0.251 (0.240) | 0.362** (0.170) | -2.859* (1.623) |
| <i>First-stage Results</i> | | | | | |
| Initial Land Supply × Interest Rates | -0.147*** (0.042) | -0.124*** (0.042) | -0.178* (0.093) | -0.518*** (0.091) | -0.083** (0.033) |
| First-stage SW <i>F</i> -Statistics | 12.48 | 8.48 | 3.67 | 32.10 | 6.45 |
| AR Wald Test <i>p</i> -Value | 0.438 | 0.824 | 0.183 | 0.028 | 0.003 |
| Individual Controls | ✓ | ✓ | ✓ | ✓ | ✓ |
| Household Controls | ✓ | ✓ | ✓ | ✓ | ✓ |
| Household FE | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year FE | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 54,375 | 42,871 | 10,223 | 50,198 | 66,033 |
| <i>R</i> ² | 0.254 | 0.275 | 0.333 | 0.529 | 0.290 |

Notes: This table reports the 2SLS estimation results of equation (1) using depression scores, sleep hours and a second job dummy as dependent variables. Depression is an average score of different measures of mental depression, each measured with 1-5 scale. The higher the score the worse the mental health. Individual controls include sex, age, age squared, marital status, and level of education; household controls are household income categories, ln household size, and an urban or rural dummy. Robust standard errors are clustered at the community level in all specifications.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Standard errors in parentheses.

The first-stage results in Table B.1 indicate that our instrument is negatively correlated with housing prices, with the *F*-statistics of weak instrument test ranging from 6.45 to 33.60. The *p*-values of the Anderson-Rubin tests confirm causal effects of residential housing prices. Housing prices have a positive but insignificant impact on the probability of having a second job as shown in column (1). In columns (2) and (3) we examine the impact for mortgage payers and non-mortgage payers separately. Similar to the result in column (1), neither mortgage payers nor non-mortgage payers experienced a higher probability of having a second job with housing price growth. These results suggest that adjustments of work intensity along the extensive margin were not a main channel. The coefficient estimates in column (4) and (5) indicate that housing prices led to higher depression scores as measured by CES-D and less sleeping time, which is consistent with our earlier findings based on the CHNS in Tables 10 and 11.

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