Capital Scarcity and Industrial Decline: Evidence from 172 Real Estate Booms in China

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Abstract

In geographically segmented credit markets, local real estate booms can divert capital away from manufacturing firms, create capital scarcity, increase local real interest rates, lower real wages, and cause underinvestment and relative decline in the industrial sector. Using exogenous variation in the administrative land supply across 172 Chinese cities, we show that the predicted variation in real estate prices does indeed cause substantially higher capital costs for manufacturing firms, reduce their bank lending, lower their capital intensity and labor productivity, weaken firms' financial performance, and reduce their TFP growth by economically significant magnitudes. This evidence highlights macroeconomic stability concerns associated with real estate booms.

Key words: Factor price externalities, reverse Balassa-Samuelson-effect, firm growth

JEL codes: D22, D24, R31

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1 Introduction

In geographically segmented credit markets, real estate investments competes with corporate investments for the local household savings. During real estate booms with a strong surge in value of housing investment, the residual capital available for corporate investment can become scarce and expensive—thus undermining the competitiveness and growth potential of the local manufacturing sector, which competes with firms in more capital-abundant locations. Empirically, the negative effect on local corporate growth is difficult to establish because cross-country studies are plagued by many confounding effects and exogenous factors, making real estate investment booms generally hard to identify.

This paper proposes exogenous variations in the land supply policies of Chinese cities to study the role of capital costs on long-run firm growth and competitiveness. The land supply in Chinese cities is shown to be a good predictor of local real estate booms, which channel local household savings into real estate investment rather than corporate investment. We find that the larger the real estate boom in any Chinese city for the period 2002–2007, the more severe the credit and capital scarcity for the local corporate sector became. Such city-level capital scarcity manifests itself in higher interest rates for corporate loans, a lower share of firms with bank credit, significantly lower corporate investment rates, and large negative externalities on the growth and competitiveness of local manufacturing firms.

Real estate booms have many links to local economic conditions. Productivity shocks can boost local output, increase housing demand through migration or income effects, and induce a positive correlation between housing prices and the performance of local manufacturing firms. We therefore need a valid instrument that accounts for exogenous variation in the housing price and the ensuing capital diversion into real estate investment. The institutional features of China's housing market provide such an instrument: Constructible land is supplied monopolistically by the local government, governed by an administrative process, and subject to exogenous constraints on land availability in a city. We define the annual *Adjusted Land Supply* as the surface of new constructible residential land scaled by the size of the existing housing stock and local population density. While this supply measure is a very good predictor for the (log) housing price level, it is itself unrelated to local economic conditions such as local GDP, or local government expenditure or revenue.

We use the *Adjusted Land Supply* as an instrument for explaining the large cross-sectional and intertemporal variation in Chinese real estate prices. Figure 1 sorts 251 prefecture-level cities by their initial real estate price index in 2003 (blue spikes) and shows the large variation of the same price index in 2010 (red spike). We are able to construct panel data on land supply in 172 prefecture-level cities and use firm-level data from these cities for our main analysis. Our first-stage regression can explain a large share of the price variation between 2002 and 2007 by the *Adjusted Land Supply* in each city. The second-stage regression then documents how real estate price variation traced back to land supply variation impacts firm development. We measure corporate capital scarcity directly at the firm level by examining measures of firm bank access and bank credit costs—showing their strong relationship to the instrumented real estate price.

China's highly segmented capital market for small and medium-size firm finance provides an excellent case for studying the equilibrium effects of capital scarcity. Such capital market segmentation at the prefecture-city level has been documented through the crowding out of corporate finance by local government borrowing (Huang, Pagano, and Panizza, 2017, 2018). Yet government investment in infrastructure and public good provision can generate its own growth externalities, which are difficult to separate from the negative effect of higher corporate capital costs. By contrast, real estate booms causing corporate capital scarcity are generally of little benefit to local manufacturing firms and the negative effect of higher capital costs can be identified directly.

Our main finding is the strong economic effects of exogenous variations in real estate prices on corporate capital costs, corporate investment and growth. A 50% relative increase in a city's real estate price due to a shortage in local land supply over the period 2002–2007 increases the borrowing costs of firms by an average 0.9 percentage points annually and reduces the share of firms with bank credit by 3.2 percentage points, which represents a 9% reduction relative to the sample mean. This local credit crunch reduces the average corporate net investment share (net new investment to book capital) by 7.3 percentage points, which represents a 21.4% reduction relative to the sample mean. The long-run relative output decline amounts to a staggering 35.5% of value-added output and total factor productivity features a relative decline of nearly 12% for the average manufacturing firms.

We cross-check these astonishingly large real effects with independent data sourced from the

Chinese customs authorities. Chinese customs data record product-level data on real quantities for each exporting firm. These accounting data are not subject to any price mismeasurement concerns like revenue-based output measures based on industry price deflators. Yet the recorded export quantities (for firms that export more than 75% of their output) also show a staggering 17.3% shortfall in exports for firms in cities with a 50% higher real estate price index. Moreover, export prices show no pass-through effect of real estate prices and so confirm that output deflators are not subject to any systematic measurement biases across cities.

A challange for our instruments is to exclude confounding effects on outcome variables which correlate with housing price variation. We adopt a variety of strategies to convince the sceptical reader: First, we verify that our instrument (i.e., *Adjusted Land Supply*) is unrelated to local economic and fiscal variables except the local real estate price. Second, we highlight that our regression analysis generally includes firm fixed effects so that identification is achieved through the intertemporal variation in local land supply related to random contingencies within the bureaucratic planning process. This should alleviate concerns that unobserved cross-sectional factors drive our results. Third, we develop a structural model of saving diversion into residential housing investment and confirm its predicted effects on local factor prices (i.e., corporate loan costs and firm wages) and many other firm variables (i.e., net investment share, gross investment share, bank loan dummy, log employment, log output, log labor productivity, firm exit dummy, return on assets, log total factor productivity). In particular, the factor price evidence supports the transmission channel through local capital costs and is unlikely to be explained by other channels linking new residential land supply to corporate development. Fourth, we confirm the predicted heterogeneity in firm effects resulting from unequal bank access. Firms with large fixed assets and state-owned enterprises (SOEs) enjoy privileged credit access to the "big five" national banks. We show that this greatly attenuates their exposure to the capital scarcity induced by local real estate booms and reduces their relative competitive decline. For example, the average SOE shows a reduction in the investment share of only 1.9 percentage points for a 50% higher local real estate price relative to an investment shortfall of 8.4 percentage points observed for privately owned firms. Fifth, we show robustness of our results using an alternative instrument for real estate prices—namely the elasticity of new Chinese residential housing construction taken from Wang *et. al.* (2012). This purely cross-sectional approach resembles recent work by Mian and Sufi (2011, 2014), Mian *et al.* (2013), Adelino *et al.* (2015) on the U.S. data.

Finally, we point out that the evidence does not support a "Dutch Disease" effect in which *all* factor prices increase. Instead, we find that real wages fall significantly wherever real estate prices boom and interest rates increase. This reverse factor price dynamics is best captured by a modified Harrod–Balassa–Samuelson (HBS) model we develop in which a construction sector and the industrial sector compete for scarce labor and capital resources. But unlike in the traditional HBS model, the factor price externality of the construction sector operates through inflated capital costs, whereas real wages decrease under competitive pricing both in the model and the data.

The macroeconomic economic literature has recognized that real estate markets and mortgage institutions can have an influence on the savings rate of households (Deaton and Laroque, 2001) and possibly growth. For example, cross-country variations in the loan to value ratios in mortgage markets affect the liquidity constraints of households, influence household saving rates and appear to correlate negatively with corporate investment rates and growth rates (Jappelli and Pagano, 1994). The channel we highlight in this paper focuses not so much on the equilibrium saving rate *per se*, but more directly on savings that are diverted from corporate to housing investments if the latter promise higher returns during real estate booms.

Recent finance research has examined the relationship between real estate booms and corporate investment by U.S. firms. For firms with real estate property, a local property price increase can relax borrowing constraints and increase firm investment (Chaney *et al.*, 2012; Jiminez *et al.*, 2014). For Chinese firms this balance sheet effect may not matter much because of a lack of real estate assets on firms' balance sheets and the state's monopoly of land development. Among Chinese listed firms in 2007, only 35.1% report positive real estate assets and their aggregate value accounts for only 2.56% of aggregate assets. For all firms, including those that do not hold real estate assets, the real estate value share is lower at 1.12% of aggregate assets—suggesting that smaller non-listed manufacturing firms own only negligible amounts of real estate assets. Real estate booms can also increase local consumption through a collateral effect and/or wealth effect (Cloyne *et al.*, 2019). We highlight that any collateral effect related to real estate booms—either for firms or households—should reinforce local firm growth rather than contribute to the industrial decline documented in this paper.

Unlike a collateral channel, the equilibrium effect of corporate underinvestment due to saving diversion concerns *all* firms and has potentially larger economic ramifications. The literature on financial stability has often highlighted real estate booms as a precursor of financial crisis through imprudent bank lending (IMF, 2011). The negative effects of such booms on the real sector through reduced credit and a loss of competitiveness are apparently important features of recent financial crises in southern Europe (Sinn, 2014; Martín *et al.*, 2018) — yet identifying a clear causal link between real estate booms and reduced firm investment has generally been difficult. An exception here is recent evidence by Chakraborty *et al.* (2018) showing that local real estate booms adversely affect the volume and cost of business loans from U.S. banks: A one standard deviation increase in U.S. housing price increases the corporate borrowing costs of financially constrained U.S. firms by 0.53 percentage points and reduces corporate investment rate by an average of 6.2 percentage points. A higher degree of regional credit market fragmentation and large geographic variations in housing booms make China a very good candidate to study sectoral competition for local credit. In China, a one standard deviation increase in the real estate prices implies an average increase of corporate borrowing costs by a much larger 1.1 percentage points and reduces the investment rate by 8.7 percentage points. The influence of real estate booms on China's internal capital allocation has been highlighted by two related working papers. Wu *et al.* (2016) present evidence on a significant and robust negative association between housing prices and corporate investment. Chen *et al.* (2016) emphasize both speculative real estate investment and the crowding out of corporate investment for the same data period.

Our contribution to the literature is threefold: First, we provide a simple neoclassical framework in which a segmented capital market implies different local real effects for local housing supply variation. The simple model generates predictions not only for the local real estate price level and local capital costs, but also for real wages — thus broadening the scope of the analysis. Second, China's state monopoly on land supply provides an ideal empirical setting to test the model predictions for firm output and investment, as well as for additional firm variables like firm profitability, and total factor productivity. We extend the evidence by Chakraborty *et al.* (2018) to the local real effects of real estate booms. Highly segmented capital markets like in China allow better measurement of local real effects, whereas a more integrated capital market like in the U.S. tends to diffuses real effects to the macro level. In other words, real effects are better identified in the Chinese setting, even if the external macroeconomic validity of the channel is similar. Third, we propose different instrumental variable (IV) strategies based on monoplistic land supply or housing supply elasticities. Time variation in the local land supply allows for a new intertemporal identification of real estate price effects where firm fixed effects control for time-invariant firm heterogeneity. Alternatively, we implement a strictly cross-sectional identification strategy based on local housing supply elasticities similar to Mian and Sufi (2011, 2014), and obtain quantitatively similar results.

In Section 2 we develop the two-sector model and contrast the effect of diverging capital costs with the Harrod-Balassa-Samuelson model of diverging labor costs. The testable model implications are spelled out in two propositions in Section 2.2., followed by two additional hypothesis on firm heterogeneity and firm performance in Section 2.3. In Section 3 we explain the data and the estimation strategy based on within-city land supply variation. Our empirical analysis first validates the model implications for factor prices, namely capital costs and wages, in Section 4.1, and then for other firm variables in Section 4.2. The role of firm heterogeneity in capital access is studied in Section 4.3, and we examine additional performance measures in Section 4.4. Robustness is discussed in Section 5 followed by our conclusions in Section 6.

2 Theoretical Framework

One of the best documented stylized facts about relative competitiveness is the Harrod-Balassa-Samuelson effect. Productivity growth in a country drives factor costs and in particular real wage growth. This makes non-tradeable labor-intensive service sectors expensive and non-competitive by international comparison; yet their very non-tradeability implies that high wage costs can be passed on to high prices for non-tradeables. The following section presents a similar two-sector economy in which one booming sector adversely influences the other sector through factor prices. We argue that in a Chinese city with a booming real estate sector and increasing housing prices, local savings are predominantly channeled into real estate investment where rapid price inflation promises a high return. Unlike the Harrod-Balassa-Samuelson world with its perfect capital market, China's corporate credit market is highly segmented so that the large capital demand of the real estate sector dramatically increases the local interest rate. High local capital costs and/or capital scarcity undermine the competitiveness of the local manufacturing sector. Unlike the non-tradeable sector in the Harrod-Balassa-Samuelson world, the manufacturing sector cannot pass on a higher factor cost to a competitive international market price and instead faces industrial decline and demise. We develop this modified or "inverted" Harrod-Balassa-Samuelson world of factor price externalities more formally in the next section before applying it empirically to the Chinese economy.[1]

2.1 A Two-Sector Model

We retain the two-sector structure of the Harrod-Balassa-Samuelson model and replace the non-tradeable sector with a real estate sector.

Assumption 1: Real Estate and Tradeable Sector

Consider a competitive real estate sector (R) producing housing Y_R and a competitive manufacturing sector (T) producing tradeables Y_T. Both sectors compete for capital with inputs K_R and K_T, respectively. The real estate sector requires a governmental land supply S as a complementary factor and a high real estate price P requires proportionately more capital to produce the same amount of housing. The production function for real estate is given by

$$Y_R = A_R \min(S, K_R/P) \tag{1}$$

where land supply S and real capital K_R/P are strictly complementary. The tradeable sector features a Cobb-Douglas production function with labor input L (capital input K_T) and labor (capital) elasticity μ $(1-\mu)$ given by

$$Y_T = A_T L^{\mu} K_T^{1-\mu}. \tag{2}$$

For simplicity, we assume real estate production does not require any labor input. This assumption can be easily relaxed and is not critical for our analysis. More important is the assumption that the capital requirements for real estate production increase linearly in the price of real estate P. This assumption is motivated by the monopolistic land supply S, where local government rations land supply and increases land prices in line with the real estate price. Hence, the same real housing production requires an increasing amount of private capital as

[1] In the context of investment booms triggered by natural resources, negative cross-industry externalities are sometimes referred to as a "Dutch Disease", and consist of rising real wages, that undermine industrial competitiveness. But in the Chinese context the factor price for capital increases, whereas wages decrease in cities with real estate booms. References to a "Dutch Disease" are therefore misleading.

real estate prices increase. This implies that a real estate boom in our model does not require that more real resources are allocated to housing. Yet, inflated costs of new housing reduce the share of private savings available for corporate investment. We assume that the revenue from land sales is consumed by the government (or invested otherwise) and does not relax the limited supply of local (private) capital.[2] In particular, we assume a fixed local factor supply for both labor and capital.

Assumption 2: Factor Supplies

The local capital and labor supply are both price inelastic and fixed; hence

$$K_R + K_T = \overline{K} \tag{3}$$

$$L = \overline{L}. \tag{4}$$

Completely price inelastic local factor supplies in both capital and labor are two simplifying assumptions. However, these are not essential for the qualitative implications of the model. In Appendix A we solve the model for the general case of price elastic factor supplies. We find that all qualitative predictions are robust to this generalization. We also note that housing price inflation can be further accelerated by speculative buying of housing in view of future capital gains; yet we do not explicitly model any additional speculative housing demand here (Chen *et al.*, 2016; Shi Yu, 2017).

The traditional Harrod-Balassa-Samuelson literature generally assumes perfect capital market integration. However, a constrained local capital supply provides a better empirical benchmark for China: its internal capital market appears to be segmented with only limited capital flows compensating for capital demand shocks across cities (Huang, Pagano, and Panizza, 2017, 2018). Many restrictions on banking across various administrative units contribute to the regional segmentation of the corporate credit market. The lack of true capital market integration leaves plenty of scope for geographically diverging real interest rates and capital. In Appendix B, we estimate an error correction model for the median corporate bank loan rate in any city

[2] If local government does not consume (or invest) its gains from land sales, but instead deposits these revenues in local banks, then we do not obtain a local capital scarcity effect under real estate price inflation. A general equilibrium model therefore needs to model government expenditure decisions in addition to private saving decisions.

relative to the median rate of all firms in neigbouring cities. The variation in the median corporate bank loan rate across cities ranges from 3.8% for a city at the 10% quantile to 6.4% for a city at the 90% quantile. The mean reversion of only 13.8% between a city's median loan rate and those of firms in the neighboring cities illustrates the strong geographic segmentation of China's corporate credit markets.

We close the model with a housing demand function of low price elasticity.

Assumption 3: Housing Demand

The (log) housing demand is price elastic and for strictly positive parameters γ_0, γ_1 with $0 \lesssim \gamma_p < 1$, total housing demand follows as

$$\ln Y_R^D(P) = \gamma_0 - \gamma_p \ln P. \tag{5}$$

Under $0 \lesssim \gamma_p < 1$ housing demand features a low price elasticity. As the local housing production is constrained by the land supply S, the equilibrium real estate price follows directly as

$$\ln P = \frac{1}{\gamma_p} \left[\gamma_0 - \ln A_R - \ln S\right], \tag{6}$$

and the capital demand of the real estate sector is given by

$$\begin{aligned} \ln K_R &= \gamma_0 + (1-\gamma_p)\ln P - \ln A_R = \\ &= \frac{1}{\gamma_p}(\gamma_0 - \ln A_R) - \frac{1-\gamma_p}{\gamma_p}\ln S. \end{aligned} \tag{7}$$

An insufficient land supply by local government therefore inflates the real estate price P and at the same time increases the capital demand $\ln K_R$ by the real estate sector.

2.2 Model Implications

To simplify notation, we express all variables in percentage changes relative to steady state log values, that is $\widehat{X} = dX/\overline{X}$. The zero profit condition in the tradeable sector implies the following relationship for changes in the equilibrium wage $\widehat{w}$ and the local interest rate $\widehat{i}$

$$\widehat{A}_T = \mu\widehat{w} + (1-\mu)\widehat{i}, \tag{8}$$

where we abstract from any productivity growth by assuming $\widehat{A}_T = \widehat{A}_R = 0$. Profit maximization in the tradeable sector also implies

$$\widehat{Y}_T = \widehat{w} + \widehat{L} = \widehat{i} + \widehat{K}_T, \tag{9}$$

and the factor supply conditions give $\widehat{L} = 0$ and $\widehat{K}_T \overline{K}_T + \widehat{K}_R \overline{K}_R = 0$. Combining these relationships implies the following proposition.

Proposition 1: Wage and Interest Rate Channel:

Under Assumptions 1–3, and a limited supply of constructible land S, the local interest rate change $\widehat{i}$ (real wage changes $\widehat{w}$) is proportional (is inversely proportional) to real estate prices inflation $\widehat{P}$ with percentage changes characterized as

$$\widehat{i} = \mu \frac{\overline{K}_R}{\overline{K}_T}(1 - \gamma_p)\widehat{P} \tag{10}$$

$$\widehat{w} = -(1 - \mu)\frac{\overline{K}_R}{\overline{K}_T}(1 - \gamma_p)\widehat{P}. \tag{11}$$

Real estate inflation itself is proportional to changes in the local land supply $\widehat{S}$ as

$$\widehat{P} = \widehat{S} \times \eta, \tag{12}$$

with a price elasticity of supply $\eta = -1/\gamma_p$.

The negative effect of the real estate boom on wages distinguishes our model from a so-called "Dutch Disease" scenario, where an investment boom (often in natural resource industries) increases real labor costs and exercises competitive pressures on other firms through a higher local wage level. By contrast, our model predicts a decrease in the real wage level because of corporate underinvestment under high interest rates.

The linear relationship between the real estate price and the land supply in Eq. (12) suggests that land supply should be a good instrument for local real estate inflation. As land supply by local government also conditions housing output, with $\widehat{Y}_R = \widehat{S}$ absent any productivity growth in construction ($\widehat{A}_R = 0$), we can relate changes in the (new) housing supply $\widehat{Y}_R$ and housing sales $\widehat{HS} = \widehat{P} + \widehat{Y}_R$ to change in the local housing price index, hence

$$\widehat{Y}_R = -\gamma_p \widehat{P} \tag{13}$$
$$\widehat{HS} = (1-\gamma_p)\widehat{P}. \tag{14}$$

Figure 2 shows the positive relationship between annual (log) changes in housing sales and the annual (log) changes in the city housing price index. A "near unit" slope $1-\gamma_p \lesssim 1$ shows that the housing demand is very price inelastic so that small changes in the land supply (and consecutively new housing supply) translate into large housing price changes.

The first part of our empirical analysis consists in showing that local factor prices across Chinese cities are indeed related to local real estate inflation $\widehat{P}$ and constructible land supply $\widehat{S}$ as predicted in Proposition 1. The second part of our analysis explores the role of the implied factor price variation for the manufacturing sector summarized in Proposition 2:

Proposition 2: Manufacturing Under a Real Estate Boom

Under Assumptions 1–3 and a limited supply of constructible land S, the local production response in the manufacturing sector to real estate inflation $\widehat{P}$ is characterized by a relative (percentage) adjustment in capital $\widehat{K}_T$, the net investment share $(\widehat{NI/K})_T$, manufacturing output $\widehat{Y}_T$, and labor productivity $(\widehat{Y/L})_T$ given by

$$\widehat{K}_T = -\frac{\overline{K}_R}{\overline{K}_T}(1-\gamma_p)\widehat{P} \tag{15}$$
$$(\widehat{NI/K})_T = \widehat{K}_T = -\frac{\overline{K}_R}{\overline{K}_T}(1-\gamma_p)\widehat{P} \tag{16}$$
$$\widehat{Y}_T = -(1-\mu)\frac{\overline{K}_R}{\overline{K}_T}(1-\gamma_p)\widehat{P} \tag{17}$$
$$(\widehat{Y/L})_T = -(1-\mu)\frac{\overline{K}_R}{\overline{K}_T}(1-\gamma_p)\widehat{P}, \tag{18}$$

where a low price elasticity of housing demand implies $0 \lesssim \gamma_p < 1$.

Our model predicts the direct real effects of real estate booms on firm investment, output, and labor productivity. We do not model financial intermediation and the banks' role in channeling credit into real estate rather than firm investment. For China, we do not dispose any disaggregate data which allows us to document the credit allocation decision at the bank level similar to Chakraborty *et al.* (2018). However, aggregate data suggests that the banking sector allocated an increasing proportion of credit to housing developement: The outstanding individual housing loans increased fivefold from 560 billion Yuan in 2001 to 3 trillion Yuan in 2007. In the last sample year 2007, roughly 13.8% of all new medium and long term bank loans were allocated to the real estate companies compared to only 7.5% for the entire manufacturing (People's Bank of China, 2007).

2.3 Extensions to Firm Heterogeneity

The simple two-sector model presented in Section 2.1 does not allow for firm heterogeneity in capital access. Naturally, some firms are exposed to local capital scarcity more than others. In particular, firms with large fixed assets (available as collateral) and state-owned enterprises with political support should find it much easier to maintain credit access even under local capital scarcity. We therefore add the following testable hypothesis:

> **Hypothesis 1: Heterogeneous Capital Access Within Cities**
>
> Under real estate inflation, firms with large fixed assets or SOEs should find it easier to maintain credit access and *ceteris paribus* experience a relative increase in investment and capital growth, a larger loan growth, and larger growth in output and labor productivity.

Our competitive model also ignores the additional consequences of higher capital costs and underinvestment on (long-term) firm profitability, leverage, and factor productivity. However, firm performance measures are likely to decline if real estate booms increase the capital costs of local manufacturing firms (Dörr *et al.* 2017; Manaresi and Pierri 2018). Lower profitability should predict higher leverage. We summarize these effects in a second testable proposition:

> **Hypothesis 2: Firm Profitability, Leverage, and Factor Productivity**
>
> For tradeable producers, increased local capital costs under real estate inflation imply reduced profitability (lower ROA) and increased leverage. Moreover, credit supply constraint adversely affects total factor productivity (TFP) growth because of underinvestment. Within a city, these effects should be less pronounced for SOEs or firms with large fixed assets enabling easier access to credit.

3 Data Issues

3.1 Data Sources

We use firm data from the annual survey of all industrial firms (ASIF) conducted by China's National Bureau of Statistics over the period of 1998–2007. The ASIF data cover state-owned and private-owned enterprises in the mining, manufacturing, and utility sectors. Private enterprises are covered if their annual operating income exceeds RMB 5 million.[3] The survey consists of a stratified firm sample for 31 provinces, 398 cities, 43 two-digit industries, and 195 three-digit industries. The survey reports detail accounting data, allowing us to construct measures of firm investment, productivity, and financial performance. The location of firm's headquarters is identified so that we can match additional city-level statistics—in particular to the local real estate market.

Three main shortcomings of the data source should be highlighted. First, the firm sample is unbalanced, smaller firms in particular are typically covered only for less than three consecutive years. Second, the survey contains data errors and must be filtered for implausible data points. We provide details of our data cleaning procedure in Appendix C, which produces a final sample of around 900,000 firm-year observations for the period 2002–2007. Third, the survey data do not report any plant-level information. Multi-plant firms can produce in multiple cities with diverging real estate environments. However, the city level represents a relatively large administrative unit with an average population of 3.5 million. Only very large corporations are likely to operate in multiple cities and eliminating large firms from the sample does not appear to influence our main estimation results.

Table 1 gives the statistical description of the firm-level variables. The two important factor prices of a firm are the (log) average annual employee salary $\ln w_{j,t}$ and bank loan rate $i_{j,t}$ measured by the ratio of interest payments to the sum of long-term bank and short-term bank credit, where the latter term is interpolated from the more comprehensive reporting of listed firms.[4] For most manufacturing firms, long-term debt consists almost exclusively of bank credit.

[3] RMB 5 million was equivalent to US$ 603,930 in 1998 and US$ 657,549 in 2007.

[4] For listed manufacturing companies, we calculate the ratio of short-term credit to short term debt annually between 2002 and 2007. For example, in 2002, this ratio is 44.9%. For any sample firm j in the ASIF in 2002, we add short-term credit, $ST\ Credit_{j,t} = 44.9\% \times\ ST\ Debt_{j,t}$, to the reported long-term credit to obtain a firm's total credit.

We denote as $NI/K_{j,t}$ the real net investment rate. The ASIF only reports the book value of fixed assets so that nominal investments are not comparable across firms and reporting years due to inflation. Following Brandt *et al.* (2012), we assume that firms start purchasing fixed assets from the starting periods with a certain pattern so we can deflate the book value to obtain the real terms. As a robustness check, we also calculate the real gross investment rate $I/K_{j,t}$ which does not take into account depreciation. Appendix D reports in detail the procedure we use to calculate the real investment rate (Rudai, 2015). The dummy variable $Loan_{j,t}$ marks as 1 for firms that have long-term debt on their balance sheet. The end of the year (log) employment level is denoted as $\ln L_{j,t}$ and the (log) output $\ln Y_{j,t}$ is measured as value-added output deflated at industry output prices. Labor productivity follows as the log ratio $\ln(Y/L)_{j,t}$ and a firm's return on assets $ROA_{j,t}$ is net profits divided by total firm assets. We define as $Leverage_{j,t}$ the ratio of total liability to total assets. Further, (log) revenue-based total factor productivity $\ln TFP_{j,t}$ is measured based on cost shares. As a robustness check, we also calculated (log) TFP using the Olley and Pakes (1996) method. In addition, we define as ln $Fixed\ Assets_j$ the firm's (log) fixed assets in the year a firm enters the survey and a dummy SOE_j of whether the firm represents a state-owned enterprises.

Productivity research generally infers real quantities by applying industry-specific price deflators to revenue statistics. These deflators are not firm-specific and could potentially introduce a measurement bias if firm-specific output prices and industry-wide averages systematically diverge as a function of local real estate prices. To address this concern, we match the ASIF data with additional Chinese customs data that provide quantity and price information at the firm and product level for the period of 2002–2006. Specifically, we retain all firms that export more than 75% of their output and track their various exported items in time-consistent measurement units, i.e. in number of units, weight, volume, etc. The product-level data (at the six-digit product code) is aggregated for each firm into a maximum of 49 different product categories by quantity and unit price. The aggregate quantity is the sum of items in the same measurement units, and the unit price is the ratio of aggregate value to aggregate quantity. This procedure provides a direct real measure of export quantity that is not subject to any price mismeasurement. For export-oriented firms, such a quantity measure should be a good substitute for real output and informative about firm performance. In conclusion, we define for each firm j one or more product categories i and measure the annual (log) export value ($\ln ExpValue_{i,j,t}$), the (log) export quantity ($\ln ExpQuantity_{i,j,t}$) and the (log) export (unit) price ($\ln ExpPrice_{i,j,t}$). Focusing on export quantity allows a robust analysis without any price distortions.

A supplementary panel of city-level data comes from the China City Statistical Yearbook (CSY) and China's Regional Economic Statistical Yearbook (RESY). The RESY reports the total sales value and surface area of so-called "commercial housing." This term refers to residential housing sold at market prices by a "qualified real estate development company." The latter acquires land usage rights via land leasing, develops the real estate, and then sells it at a profit. The ratio of the sales value of commercial housing to its surface area represents our local (city-level) real estate price index. Table 1, Panel B, reports the (log) price level $\ln P_{c,t}$ and the annul real estate price inflation $\ln P_{c,t}/P_{c,t-1}$. The average annual (log) growth rate of inflation is 9.4% with a large standard deviation of 13%. In the full sample is dominated by boom years: We find annual price declines for only 20.7% of all city-year observation.

We instrument local real estate inflation by the local land supply for residential housing L_c at the city level for the period 2002–2007. Unfortunately, the annual land supply for residential housing is reported only at the province level as $L_{p,t}$. However, we know the city-level supply of non-industrial land, which is composed of residential land and commercial (non-industrial) land supply. To infer the component of the city-level land supply for residential housing, we calculate the ratio $\overline{(L_c^{NI}/L_p^{NI})}$ of non-industrial land supply at the city relative to the province averaged over the period 2003–2007. The city-level land supply for residential housing is then constructed as

$$L_{c,t} = L_{p,t}\overline{(L_c^{NI}/L_p^{NI})}. \tag{19}$$

Underlying this approximation is the assumption that the shares of commercial and residential land supply are constant across cities in the same province. An alternative approach proxies the city-level land supply for residential housing by the city-level non-industrial land supply — thus treating the unobserved variations in the commercial land supply as an error term. This method yields a weaker instrument because it does not use information on residential housing supply at the province level.

A key indentification strategy is that variation in the residentual land supply does not directly influence firm investment and performance through channels other than the residential housing price. In this context we highlight that land supply policies for industrial land do not correlate at economically significant magitudes with residential land supply. The correlation between the (log) non-industrial land supply $\ln L_{c,t}$ and the industrial land supply $\ln L^I_{c,t}$ is very low at 0.03. In addition, industrial land prices feature constantly low prices during our sample period; with industrial land prices being on average only 20% of non-industrial land prices. The correlation between the (log) price of non-industrial land and the (log) price of industrial land is negligible at 0.008. Hence, there is no evidence that industrial land is a scarce production factor in China. Also real estate booms for residential property generally do not spill over into higher rental income for industrial property.

3.2 Land Supply Variations as Instrument

Recent work on the determinants of U.S. growth before and during the Great Recession has used housing supply constraints as instruments for housing price inflation to explore causal effects on household debt and consumption (Mian and Sufi, 2011; Mian *et al.,* 2013). We apply a similar logic to China's housing market: We argue that the local housing price depends on the supply of new constructible land for residential housing in a particular city c. We normalize the new constructible land supply $L_{c,t}$ by the size of the existing housing stock $Stock_{c,t}$. A second scaling variable for the price effect of new land supply is local population density $PD_{c,t}$. We define *Adjusted Land Supply* in city c and year t as

$$Adjusted\ Land\ Supply_{c,t} = \frac{L_{c,t-1}}{Stock_{c,t-1} \times PD_{c,t-3}}, \tag{20}$$

where land supply and the housing stock are lagged by one year and the population density is lagged by three years to reduce endogeneity concerns. Previous work has show that (log) land supply in Chinese cities is linear in population growth or (log) changes in population density (Hsu *et al.*, 2017). The *Adjusted Land Supply* therefore captures variations in the land supply which deviates from the trend growth implied by local population growth[5].

We argue that this adjusted land supply for housing is reasonably exogenous to the economic

[5] We note that a linear regression

$$\ln Adjusted\ Land\ Supply_{c,t} = \alpha_0 + \alpha_1 \ln \frac{L_{c,t-1}}{Stock_{c,t-1}} + \alpha_2 \ln PD_{c,t-3} + \epsilon$$

yields $\alpha_0 \approx 0$ and $\alpha_2 \approx 1$.

fortunes of the local manufacturing sector because of China's unique institutional setting. Land supply is a monopoly of the local government in China. This subjects housing construction to the bottleneck of an bureaucratic planning process characterized by specific political interests. Monopolistic revenue maximization has been alleged to motivate a restrictive land supply by local governments in some locations (Ding, 2003). Other research has linked land sales to the financing of ambitious city development projects motivated by the career and promotion concerns of top city officials (Tian and Ma, 2009; Lichtenberg and Deng, 2009). This implies that idiosyncratic aspects related to local party leaders or city mayors play an important role in land supply. Chen and Kung (2016) argue that ostentatious development projects were occasionally financed by land sales. Hsu *et al.* (2017) find that supply for residential land is associated with corruption and competition-for-promotion motive. Local land supply determinants thus relate to local party politics and therefore largely exogenous to the development of the manufacturing sector.

A second important factor influencing land supply is geography and geographic topology. Liu *et al.* (2005) use satellite images to study the modes of geographic expansion for 13 Chinese mega-cities and find important variation. For example, the urban land of Beijing and Chengdu expanded evenly in all directions in the form of concentric expansion, whereas Guangzhou and Chongqing sprawled along rivers or lakes and their expansion is subject to specific conditions of terrain. Wuhan and Nanjing showed multi-nuclear urban land expansion constrained by their respective terrains and the conditions imposed by city development planning. Generally, the topology of the land surface also matters for urban expansion: a larger share of "flat" land correlates positively with an expansion of the land supply, whereas a higher average slope of the land inhibits the expansion of urban housing.

Our econometric strategy allows for unobservable economic factors influencing the cross-sectional pattern of land supply as we include city fixed effects in all 2SLS regressions. Hence, our identification relies on *intertemporal variation* in the land supply. The intertemporal variation is subject to many exogenous uncertainties of the bureaucratic and administrative approval process. Shenzhen was the first city to adopt land supply plan system in 1988, but most other cities started only after 2000. Typically, planned and implemented land supply show large discrepencies. For example, Beijing delivered only 33% of it planned housing supply in 2005, and 49% of its target in 2006. Such (random) housing supply variation can be traced to a variety of institutional features:

1. **Ineffective intragovernmental coordination:** Implementation of the land supply plan relies on the coordination of various city-level government departments (e.g. Land and Resources, Housing and Urban-Rural Development) and county-level institutions. Implementation of the land supply plans therefore depends on successful intragovernmental bargaining and faces many bureaucratic contingencies that can delay supply (Bo Qu, 2008).

2. **Property right conflicts:** The land supply requires (often conflictious) negotiations over incumbent usage rights and local protest can hold up land deployment. For example, China's Central Television received 15,312 letters on such land conflicts in 2004 (Hui and Bao, 2013). Even if local government can ultimately prevail, legal conflict can inflict considerable delays in implementation.

3. **Policy conflicts:** The central government occasionally interferes with city level development plans by stipulating particular quotas for the types and sizes of housing units that city governments are allowed to approve. Imposed revisions to local land supply policies can also result in supply delay (Bo Qu, 2008).

These three institutional features explain why actual and planned land supply show large discrepancy and justifies why the intertemporal pattern of land supply is a plausible exogenous source of variation.

In Table 2, we further explore this exogenous nature of the variable $\textit{Adjusted Land Supply}_{c,t}$ by correlating it with important economic variables. Column (1) shows that cities with higher GDP or a larger population do not show statistically significant differences in their land supply. Column (2) shows that a larger share of urban area has no effect on the *Adjusted Land Supply*, whereas the share of park areas is significantly negatively correlated with the *Adjusted Land Supply*. This suggests that geographic constraints matter for the local land supply policy more than economic differences across cities. In Column (3), we find that *Adjusted Land Supply* is not significantly correlated with government budget expenditure and revenue, which suggests that governmental financial incentives do not play a role in controlling land supply during our data period. Moreover, Column (4) shows that the *Adjusted Land Supply* does not correlate with the surface area of roads. The latter variable is often regarded as a measure of the development of local public infrastructure. Column (5) includes all variables simultaneouly without consequence for the result. Geographical contraints like the *Park Share* of the city surface are the most important determinants of *Adjusted Land Supply* unlike economic or fiscal city variables. In Column (6) we use one year lagged terms of all independent variables without change to the regression outcome. Only the exogenous variable *Park Share* has explanatory value for the land supply.

A potential concern is that planned land supply anticipates population growth and is thus influenced by local growth expectations. Column (7) includes measures of future population growth in the three consecutive years as a potential determinant of the *Adjusted Land Supply*. We find that current adjusted land supply for residential housing is not predicted by future population growth at any conventional level of statistical significance. Overall, Table 2 supports our instrument choice because the *Adjusted Land Supply* is uncorrelated with various local economic and fiscal variables that could influence simultaneously local factor prices and manufacturing firms' performance.

3.3 Land Supply and Housing Price Inflation

Generally, a more restrictive *Adjusted Land Supply* $S_{c,t}$ stimulates housing price inflation. However, the city-level response is also dependent on local (inverse) land supply elasticity η_c. For example, cities with a larger elasticity experience a greater change in housing prices for the same variation in *Adjusted Land Supply*. Hence, incorporating this elasticity is useful for constructing a stronger instrument for local housing prices. It is straightforward to estimate the local house price elasticities in a panel regression regrouping the $N = 172$ cities in a vector $S_t = (S_{1,t}, S_{2,t}, ..S_{N,t})$ and run the random coefficient regression

$$\ln P_{c,t} = \mu_c + \ln \textit{Adjusted Land Supply}_{c,t} \times \eta_c + \nu_t + \varepsilon_{c,t} \quad , \tag{21}$$

with city fixed effects μ_c, city-specific (inverse) land supply elasticities η_c (the slope parameters), and time fixed effect ν_t. The predicted price then follows as

$$\ln \widehat{P}_{c,t} = \widehat{\mu}_c + \ln \textit{Adjusted Land Supply}_{c,t} \times \widehat{\eta}_c + \widehat{\nu}_t \quad , \tag{22}$$

which we use as our instrumental variable for the observed housing price level in each city. We refer to this as the "city-specific instrument." An alternative specification imposes that the price elasticity of housing is identical across all cities. In this case we can stack the data matrices and estimate a single average (inverse) supply elasticity $\widehat{\eta}_c = \widehat{\eta}$. We refer to this as the "pooled instrument."

Figure 3 compares observed (log) house prices $\ln P_{c,t}$ in each city-year on the y-axis to the (log) *Adjusted Land Supply* on the x-axis. Panel A depicts the pooled elasticity estimation and Panel B uses city-specific elasticities. Both panels show that there is an strong negative relationship between housing prices and adjusted land supply after controlling city and year fixed effects. Yet, the specification with the random effect instruments provides a better fit to the data as shown by the regression statistics reported in Table 3: the F-value increases from around 10 to more than 250. Hence, accounting for city variation in supply elasticity greatly strengthens the quality of the land supply instrument. The housing price elasticity estimate $\widehat{\eta}_c$ flexibly characterizes city-level geographic constraints and might also embody expectations about future residential land supply and housing prices. Such geographic constraints for residential housing development need to be uncorrelated with changes in manufacturing firm performance that do not operate through the real estate price channel for the exclusion restriction to hold.

4 Empirical Analysis

4.1 House Price Inflation and Factor Price Response

The first step in the empirical analysis is to verify the positive effect of the real estate price level $\ln P_{c,t}$ in city c on the capital costs $i_{j,t}$ of firm j in city c and the negative effect on its real wage $\ln w_{j,t}$ as stated in Proposition 1. A linear panel specification consists of the regression

$$i_{j,t} = \alpha_0 + \alpha_p \ln P_{c,t} + \alpha_X X_{c,t} + \lambda_j + \nu_t + \epsilon_{j,t} \tag{23}$$

$$\ln w_{j,t} = \beta_0 + \beta_p \ln P_{c,t} + \beta_X X_{c,t} + \lambda_j + \nu_t + \epsilon_{j,t}, \tag{24}$$

with predicted coefficient $\alpha_p > 0$ and $\beta_p < 0$. The extended 2SLS specification controls for the same macroeconomic variables $X_{c,t}$ at the city level used in Table 3, namely local (log) GDP, (log) population, share of urban area, share of park area, local (log) government expenditure and revenue, and (log) surface road area. We also allow for firm fixed effects λ_j and time fixed effects ν_t. The error term is $\epsilon_{i,t}$ is clustered at city level to address the concern that standard errors among manufacturing firms within the same city are positively correlated. In total the panel includes a cross-section of real estate prices for 172 Chinese cities. Bank loan rates are available for $423,014$ firm-years and the average employee wage is recorded for $916,051$ firm-years.

Table 4, Column (1) shows the OLS estimates for the interest rate on bank loans. The point estimate is positive at 0.009 and marginally significant at the 10% level. Yet, various economic channels may simultaneously influence local interest rates and the real estate price level. For example, local productivity shocks could increase local interest rates and higher interest rates could moderate local housing price inflation. Table 4, Column (2) therefore proceeds to the 2SLS regression that instruments variations in the local real estate price with the land supply interacted with (inverse) land supply elasticity. The first-stage regression corresponds to Table 3, Column (3); the Kleibergen-Paap F-statistics of 97.8 indicate a very strong instrument. Under the 2SLS specification, the point estimate increases to 0.022 and is statistically significant at the conventional 5% level. We also estimate an extended specification that controls for other potential determinants of the the real estate price level. The first-stage regression here follows Table 3, Column (4). The coefficient for the interest rate effect of real estate inflation is similar at 0.022. This coefficient implies that an increase of the local real estate price by 50% increases the capital costs of local firms by approximately 0.9 percentage points [$= 0.022 \times \ln(1.5)$], which is large compared to a mean sample value of 6.1 percentage points ($0.9/0.061 = 14.8\%$ of the sample mean). This represents an economically highly significant factor price effect that deters capital investment.

The factor price effect of real estate prices on wages is documented in Table 4, Columns (4)–(6). The OLS estimate in Column (4) is negative at -0.111 and statistically significant. Various economic channels are likely to push the OLS estimate upward. First, lower local wages reduces household incomes and could have a negative effect on real estate prices. Second, (omitted) economic shock can produce a positive correlation between local wages and local housing prices. To address these issues, we once again use the 2SLS estimator reported in Columns (5)–(6), which features much more negative point estimates at -0.394 and -0.391, respectively. Now, a 50% increase in real estate prices is associated with 15.9% [$= -0.391 \times \ln(1.5)$] decrease in nominal wages. Hence, the wage effects of local capital scarcity induced by high real estate prices is quantitatively large. The lower equilibrium manufacturing wage is a consequence of less capital investment and lower labor productivity as we show in the next section.

4.2 Baseline Results for Firm Response

Having confirmed the predicted factor price response to real estate booms, we now test the additional firm level implications articulated in Proposition 2. Local capital scarcity induced by real estate booms implies lower firm investment, lower levels of bank lending to firms, less output, and lower labor productivity. The corresponding panel regressions are reported in Table 5. Panel A provides the OLS results. Panel B reports the simple 2SLS regressions that instrument the (log) real estate price level $\ln P_{c,t}$ with the *Adjusted Land Supply*. Panel C documents the extended 2SLS regression with macroeconomic control variables. Panel D adds additional industry year fixed effects.

The higher real estate price $\ln P_{c,t}$ has a strong negative effect on net investment rates $(NI/K)_{j,t}$ in all 2SLS regressions. A 50% higher real estate price implies a decrease in the average firm investment rate by 7.3 percentage points [$= -0.180 \times \ln(1.5)$], which is large compared to a mean sample value of 21.5 percentage points. Using firm's gross investment rate $(I/K)_{j,t}$ which does not wipe out depreciation of fixed assets, Column (2) shows that a 50% higher real estate price leads to a decrease in the average firm investment rate of 9.9 percentage points [$= -0.244 \times \ln(1.5)$]—indicating a slightly smaller investment shortfall relative to a sample mean of 33.7 percentage points for $(I/K)_{j,t}$. The magnitude of coefficients in 2SLS regressions double compared with OLS results. This difference is plausibly explained by the following two effects: First, unobserved positive technology and demand shocks can stimulate corporate investment and housing price inflation simultaneously and bias OLS estimates upwardly. Second, better manufacturing firm performance can contribute to a local real estate boom—thus also delivering a higher OLS estimate. Both endogeneity concerns apply equally to the OLS estimates for other firm outcomes.

Column (3) suggests that booming real estate prices curtail bank lending to manufacturing firms. A point estimate of -0.079 in Panel D implies that a 50% higher real estate price reduces the percentage of firms with bank credit by 3.2 percentage points [$= -0.079 \times \ln(1.5)$] relative to a sample mean of 34.1 percentage point of firms with bank credit. Real estate investment booms therefore dramatically increase the number of credit constrained firms.

Columns (4)–(6) show the effect of real estate prices on (log) labor input $\ln L_{j,t}$, (log) value added output $\ln Y_{j,t}$, and (log) labor productivity $\ln(Y/L)_{j,t}$, respectively. All 2SLS estimations in Panels B, C, and D document a dramatic decrease in both value added output and labor productivity under higher (instrumented) real estate prices $\ln P_{c,t}$. A 50% higher real estate price induces a output decrease of approximately 35.5% [$= -0.876 \times \ln(1.5)$] in Panel D. And labor productivity $\ln(Y/L)_{j,t}$ decreases by a similar magnitude.

Our theoretical framework links (percentage) real wage changes ($\widehat{w}$) to the change in the labor elasticity of tradeable production ($\widehat{\mu}$) and the change in labor productivity of tradeable [$\widehat{(Y/L)}_T$] according to

$$\widehat{w} = \widehat{\mu} + \widehat{(Y/L)}_T.$$

In additional regressions not reported, we find that the labor elasticity of tradeable production has a 2SLS coefficient of approximately 0.6 with respect to housing price changes, thus $\widehat{\mu} = 0.6 \times \widehat{P}$. Table 5, Column (6), Panel C finds for the change in labor productivity $\widehat{(Y/L)}_T = -1.01 \times \widehat{P}$. This implies for the real wage decline a predicted coefficient of $\widehat{w} = -0.4 \times \widehat{P}$, which is close to the point estimate of -0.391 obtained in Table 4, Column (6). The separate 2SLS estimates for the inflation elasticities of all three variables are therefore mutually consistent with each other.

The real effects of capital scarcity induced by local real estate booms for local manufacturing firms are therefore dramatic in their economic magnitude—causing a substantial (relative) industrial decline for firms located in cities with real estate booms. Also, we expect such industrial decline to be reflected in firm exit rates. Firms exit from the ASIF data whenever their sales drops below a threshold of 5 million RMB and we lable them with the dummy *Exit*. These exiting firms tend to have lower productivity and profitability compared with non-exiting firms. The 2SLS estimate in Column (7), Panel B shows statistically significant positive effect of local real estate price on firm exit, but Panels C and D this effect becomes insignficiant once controlling for local economic condtions. A 50% increase in the real estate price increases the probability of firm exit by approximately 2.2 percentage points [$= -0.055 \times \ln(1.5)$], which presents an economically significant effect.[6]

[6]The overall annual firm exit rate in the sample is high at 9%. Firms exit from the ASIF data whenever their sales drops below a threshold of 5 million RMB and this may not always imply firm closure.

As a robustness check, we substitute the level specification with firm fixed effects with a specification in differences. Formally,

$$\Delta y_{j,t} = \beta_0 + \beta_p \Delta \ln P_{j,t} + \nu_t + \epsilon_{j,t},$$

where $\Delta y_{j,t}$ denotes the annual change in the outcome variable and $\Delta \ln P_{j,t}$ the annual change in the local real estate index. The latter is now instrumented by the annual change of (log) the adjusted land supply interacted with (time-invariant) city elasticty, i.e. $\Delta \ln \textit{Adjusted Land Supply}_{j,t} \times \widehat{\eta}_c$. Table A2 in the Appendix reports the results for this alternative 2SLS specification.[7] The results are consistent with the baseline results in Table 5: Larger housing price growth implies lower growth in the investment rate and lower (value-added) output growth.

Both the investment rate measure and the (revenue-based) value-added output rely on industry-level output and intermediate input prices that might be systematically biased downward for cities with higher real estate prices. Any incorrect inflation adjustment could imply that the residuals of the second-stage regression correlate with our instrument. Fortunately, the Chinese customs authorities collect a comprehensive product-level data set on firm exports that accounts separately for product price and quantity of exporting firms. We aggregate similar products (in the same measurement units) into a single product category by value and unit price. For firms that export more than 75% of their output, we consider the export statistics as a good (real) performance measure devoid of any price distortions. One average, these firms export in 3.8 different product categories.

Table 6 analyzes the (real) export performance of Chinese firms as a function of local real estate prices. Columns (1) and (2) replicate the 2SLS regression of Table 5, Panel D, Columns (1) and (5) for the subsample of exporting firms. At a coefficient of -0.359, local real estate inflation shows an even stronger negative effect on the investment rate for exporting firms than in the full sample (-0.180). The negative output effect of real estate inflation is slightly lower at -0.699 compared to the full sample (-0.876). Columns (3)–(5) of Table 6 show the declining export performance of firms in locations with high real estate inflation. The decline in the value of export products ($\ln \textit{ExpPrice}_{i,j,t}$) is economically large and entirely due to a decline in product quantity ($\ln \textit{ExpQuantity}_{i,j,t}$), but not product price ($\ln \textit{ExpPrice}_{i,j,t}$). The

[7] We exclude the city of Sanya as it represents an extreme outlier with respect to annual housing price growth.

relative decline in export quantity amounts to 17.3% [$= -0.426 \times \ln(1.5)$] for a 50% increase in local real estate prices. This finding is entirely based on real quantities and as such robust to any potential price mismeasurement. At the same time, it confirms the large economic effects reported in Table 5.

4.3 Firm Heterogeneity in Credit Access

Credit market frictions in China predict firm heterogeneity in bank credit access. Hypothesis 1 argues that firms with larger fixed assets and SOEs should be less affected by local capital shortages brought about by real estate booms. Previous research has highlighted the privileged capital market access of SOEs in China (Allen *et al.* 2005). Access to credit from the "big five" national banks should greatly reduce the dependence of large (asset rich) firms and SOEs on local credit market conditions.

Table 7 provides evidence to support this conjecture. In Panel A, we interact the real estate price $\ln P_{c,t}$ with a firm's log fixed assets ($\ln \textit{Fixed Assets}_j$) at the beginning of the sample. Panel B interacts the real estate price with a dummy variable marking SOEs (SOE_j). We expect to find higher investment rates for less financially constrained firms as well as lower output and labor productivity decline. Column (3) confirms that firms with more fixed assets (Panel A) and SOEs (Panel B) do indeed face a smaller or no decline in access to bank loans. Accordingly, their investment rates $(NI/K)_{j,t}$ hold up much better under local real estate booms than their more more financially constrained peers in the same industry. For example, the average SOE shows a reduction in investment share of only 1.9 percentage points [$= (-0.206 + 0.158) \times \ln(1.5)$] for a 50% higher local real estate price relative to an investment shortfall of 8.4 percentage points [$= (-0.206) \times \ln(1.5)$] observed for privately-owned firms. We also note that firms with better financial market access feature lower output and labor productivity decline. However, the latter effects are not statistically significant. Finally, we show in Column (7) that market *Exit* for firms located in booming real estate markets is considerably less likely for firms with larger fixed assets.

It is interesting to show the long-run differential performance of privately-owned firms and SOEs as a function of predicted local real estate inflation. Figure 4 shows the average (log) value added output change at the city level for all privately-owned firms (blue crosses) from

SOEs (red squares) from 2001 to 2007, in Panels A and B respectively. The x-axis represents the instrumented log real estate inflation index relative to the initial (log) real estate price in 2002, i.e. $\Delta \ln \widehat{P} = \ln \widehat{P}_{2007} - \ln \widehat{P}_{2002}$, for each of the 172 cities in our sample. The y-axis value shows the average (log) value added output change of all private-owned firms (Panel A) or SOEs (Panel B) in a particular city relative to initial firm performance in 2002. Formally, we define

$$\Delta \ln Y_{c,type} = \frac{1}{N_{c,type}} \sum_{j \in C,\ j \in Type} \ln Y_{j,2007} - \ln Y_{j,2002} \ ,$$

where C represents the set of all firms headquartered in city c, $N_{c,type}$ the number of firms in city c of a particular type, and firm *type* can be a privately-owned firm or a SOE. Subtracted from the (log) firm output change are interacted industry and year fixed effects. The growth experience of privately-owned firms in Panel A in particular shows a strong negative dependence on relative real estate price growth. The growth of SOEs is generally lower, but also negatively affected by higher local real estate inflation.

4.4 Additional Firm Performance Measures

Higher capital costs and underinvestment for firms in locations with real estate booms predict additional negative effects on firm performance measures. Hypothesis 2 conjectures lower firm profitability (ROA), higher leverage (measured by the debt to asset ratio), and lower (log) total factor productivity. Table 8 reports panel regressions for all three firm performance measures. The OLS coefficients are provided in Columns (1), (4), and (7). The 2SLS results for the baseline specification are given in Columns (2), (5) and (8), whereas Columns (3), (6), and (9) add interaction effects less financially constrained (asset rich) firms and SOEs.

Columns (1)–(3) show negative effect of real state investment booms on firm profitability measured by return on assets (ROA). The 2SLS point estimate of -0.145 in Column (2) implies that a 50% higher real estate price reduces ROA by 5.9 percentage points [$= -0.145 \times \ln(1.5)$] relative to the sample mean of only 7.2 percentage points. The negative effect on firm profitability is even stronger for financially constrained firms as shown in Column (3) with a baseline coefficient of -0.215. SOEs are again somewhat less affected as indicated by the positive coefficient of 0.076 for the interaction term $\ln P_{c,t} \times SOE$.

The results of local capital market scarcity for firm leverage (measured by the debt to asset ratio) is provided in Columns (4)–(6). A 2SLS coefficient of 0.094 in Column (5) suggests that for a 50% larger real estate price increase the average firm leverage by 3.8 percentage points $[= -0.094 \times \ln(1.5)]$ compared to a sample mean of 57.7%. Lower profitability therefore translates only into a modest increase in firm leverage under credit constraints. The leverage increase comes mostly from expended trade credit (payables) as access to bank credit becomes less likely [see Table 5, Column (3)].

The effects of high capital costs and relative underinvestment on TFP levels are again economically more significant. The average manufacturing firm features a 2SLS coefficient of -0.298 in Column (8), which implies that a 50% increase in real estate prices translates into a TFP shortfall of 12.1% $[= -0.298 \times \ln(1.5)]$. Hence, firms in locations with real estate booms suffer a considerable decline in industrial competitiveness. The positive interaction coefficients of 0.025 and 0.056 for asset rich firms and SOEs in Column (9) imply that this average effect varies considerably with firm characteristics. But even for a large SOE, the average loss of competitiveness is economically significant: At the 75% quantile of fixed asset size ($\ln \textit{Fixed Assets}_{j,t} = 8.84$), the relative loss in TFP is still 8.5% $[= (-0.487 + 0.025 \times 8.84 + 0.056) \times \ln(1.5)]$ for a 50% higher local real estate price.[8]

Our results on the adverse effect of local credit constrains on relative productivity growth are similar to recent findings by Manaresi and Pierri (2018), who trace the a quarter of the productivity slowdown in Italian firms in 2007–2009 to worsening credit conditions which imply slower IT-adoption, lower export growth, and slower managerial improvements.

5 Robustness

5.1 Housing Supply Elasticity as an Alternative Instrument

Following Mian and Sufi (2011, 2014), Mian *et al.* (2013) and Adelino *et al.* (2015), we also undertake a specification with time-invariant local housing supply elasticity as instrument. The basic idea is that cities with an elastic housing supply experience only modest housing price changes as they can quickly absorb housing demand shocks through new housing construction, while cities with an inelastic housing supply encounter stronger price increases. As a first-stage

[8] Generally, the adverse growth effects of local real estate booms are most pronounced for small firms as shown in Table A6 of the Appendix.

regression, we use

$$\Delta^{02-07} \ln P_c = \mu_0 + \mu_1 \; SupplyElast_c + \varepsilon_c \; . \tag{25}$$

The dependent variable is no longer the yearly log housing price index, but its change over the entire period 2002–2007. As the city-specific elasticity is time-invariant, this specification dispenses with city and year fixed effects. For data on the housing supply elasticity $SupplyElast_c$, we draw on Wang *et al.* (2012), who estimate the response of new housing construction to price shock for 35 major cities in China for the period 1998–2008. Table A3 reports some their elasticity estimates for the five locations with the largest (*Top 5*) and lowest (*Bottom 5*) values. The cities with lower elasticity (inelastic construction supply), such as Shenzhen and Beijing, experience a greater increase in housing prices in 2003–2010 and feature the highest overall price levels in 2010, as shown in Columns (2) and (3), respectively. By contrast, cities with a higher supply elasticity, such as Yinchuan and Changsha, experience a modest increase in housing prices over the same period, and show much lower price levels as of 2010. Figure 5 shows a strong negative relationship between the (log) change of housing prices in the period 2002–2007 and the respective housing supply elasticities. The t-value for the (first-stage) regression line is above 5 and R-squared is above 40%, indicating a reasonably strong instrument.

The second-stage regression is also reduced to a pure cross-sectional specification given by

$$\Delta^{02-07} y_j = \beta_0 + \beta_p \Delta^{02-07} \ln P_c + \beta_X X_c + \xi_{ind} + \epsilon_j \quad , \tag{26}$$

where outcome variables $\Delta^{02-07} y_j$ are factor price changes given by the firm bank loan rate change $[\Delta^{02-07} i_j]$ and the (log) wage change $[\Delta^{02-07} \ln w_j]$; the net investment rate rate change $[\Delta^{02-07} (NI/K)_j]$; the change in firm share with bank loans $[\Delta^{02-07} Loan_j]$, the change in (log) employment $[\Delta^{02-07} \ln L_j]$, in (log) value-added output $[\Delta^{02-07} \ln Y_j]$, and (log) labor productivity $[\Delta^{02-07} \ln (Y/L)_j]$; the change in firm profitability $[\Delta^{02-07} ROA_j]$, the change in leverage $[\Delta^{02-07} Leverage_j]$, and the change in (log) TFP $[\Delta^{02-07} \ln TFP_j]$. The city-level controls X_c include the GDP per capita, population density, employment share of the secondary sector and GDP share of the secondary sector in 2002 to capture differences across cities at the starting date of the sample period. We also control for two-digit industry fixed effects ξ_{ind} to capture heterogeneity by industry. Table A4 reports the results for this alternative specification with different (time invariant) instruments. The number of (cross-sectional) observations decreases considerably because local housing supply elasticities are available for only 32 cities and a smaller number of firms operate in these locations for the full period 2002–2007.

Column (1) confirms that firms in cities with greater housing price increase $[\Delta^{02-07}\ln P_c]$ experience an increase in their bank loan rate $\Delta^{02-07}i_j$ with a similar magnitude as in Table 4, Column (3). For wage growth $[\Delta^{02-07}\ln w_j]$ in Column (2) we confirm the negative coefficient of similar magnitude as in Table 4, Column (6). Column (3) of Table A4 is also consistent with the result in Table 5: housing price inflation lowers firms' net investment shares at high levels of economic and statistical significance. The point estimate of -0.184 in Table A4, Column (3) is close to the comparable coefficient of -0.198 in Table 5, Panel D, Column (1). Column (4) confirms the negative relationship between (instrumented) housing prices and firms' bank loan acquisition even though the coefficient is statistically insignificant. Columns (6) and (7) confirm the negative effect of housing inflation on firm output and labor productivity with similar magnitudes as results in Table 5. Very similar economic effects are obtained for ROA, leverage and TFP, as shown in Columns (8)–(10). Overall, the pure cross-sectional specification confirms the panel estimates in Table 8.

5.2 Initial Conditions of Capital Supply

The theoretical model assumes that the initial capital supply conditions are similar across cities and that differences in firm development are caused by the diverging evolution in local real estate prices and investment. However, it is plausible that cross-sectional differences in initial financial development (around the year 2000) account for the diverging firm investment and development thereafter, as different savings institutions might support long-run differences in growth (Deaton and Laroque, 2001). In Table A5, we interact the (log) housing prices $\ln P_{c,t}$ with *ex-ante* measures of financial development at the province level. The 2SLS specification in Column (1) repeats the regression in Table 5, Panel B, Column (1). Columns (2)–(5) add different province-level measures of initial financial development as interaction terms with the (log) housing prices $\ln P_{c,t}$. We also include these measures as control variables (without interaction). Column (2) focuses on the ratio of (total) bank loans to GDP as the proxy for capital supply. Its interaction term $[\ln P_{c,t} \times (Loan/GDP)_p^{2000}]$ with housing prices is statistically insignificant, while the main coefficient for housing prices ($\ln P_{c,t}$) remains significantly negative both in the economic and statistical sense. Column (3) adds the ratio of corporate loans to GDP; Column (4) adds the ratio of total deposit to GDP; and Column (5) uses the ratio of household deposit to GDP as alternative proxies for local capital supply. The initial level of financial development does not appear to matter for firm investment dynamics as all interaction terms are statistically insignificant.

6 Conclusion

This paper addresses the important question of whether real estate investment booms can cause local capital scarcity for corporate investment and so impact long-run corporate competitiveness and growth. We argue that China's state monopoly in residential land supply and its geographically segmented market for small and medium-size firm credit represent an ideal empirical setting to address this question: exogenous variations in local land supply policies provide an instrument that can partially account for the large variation of real estate prices across Chinese cities in the period 2002–2007. Real estate inflation traced to exogenous land supply variation can proxy for local capital scarcity as more household savings are channeled into real estate investment rather than corporate investment. The geographic segmentation of China's credit market ensures that the real effects are localized and better identified unlike in countries with more integrated credit markets.

Based on a sample of 172 prefecture-level cities in China, we show that local real estate booms push up the cost of capital for the local manufacturing sector and cause strong underinvestment relative to industry peers located in cities with less real estate inflation. The median corporate bank lending rate for cities at the 90 and 10 percent quantiles differ by an astonishing 2.6 percentage points and is highly persistent with a mean reversion of only 13.8%.[9] For a neo-classcial production function with capital-labor complementarities, higher corporate capital costs imply lower real wages. We confirm this prediction by showing that real manufacturing wages decrease under real estate sector booms — unlike in a Dutch disease scenario.

The ensuing corporate underinvestment virtually implies the demise of the manufacturing

[9] Of cause, the percieved opportunity costs of corporate investment might be much higher if real estate investments promises extremely high returns.

sector in cities/prefectures with the strongest real estate booms. For a 50% higher real estate price the corporate net investment rate drops by 7.3 percentage points (relative to a mean of 21.4 percentage points) and value-added firm output is lower by a staggering 35.5%. These findings highlight that a lack of credit market integration can generate substantial competitive distortions and large real effects.

As a robustness test, we also examine product-level (real) export data from the Chinese customs authority and independently confirm a 17.3% relative decline in export quantities for firms in cities with a 50% higher real estate price index. The large adverse effects of real estate booms are therefore consistent across two independently collected data sets and cannot be attributed to price mismeasurement.

While real estate booms have long been recognized as a major issue for the financial stability of the banking sector, their important adverse effect on real growth and industrial competitiveness is less established. Yet such evidence provides an additional incentive for macroprudential policies to pay close attention to real estate investment booms. The Chinese setting demonstrates that real estate booms can exercise a highly detrimental growth externality on the manufacturing sector in any financially closed economy.

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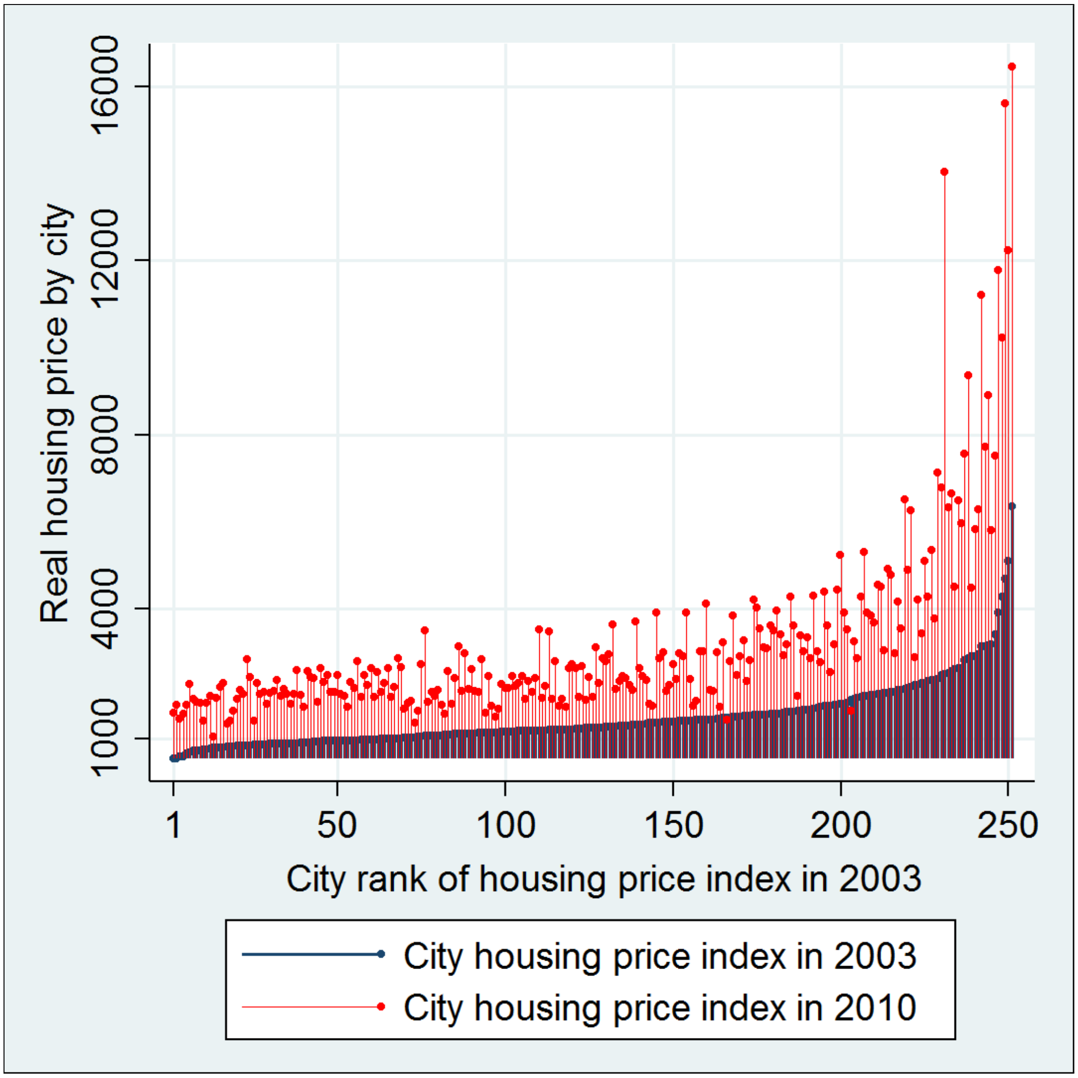

Figure 1: We rank 251 Chinese cities by their local housing price index in 2003 (blue spikes) and compare them the house price index in 2010 (red spikes).

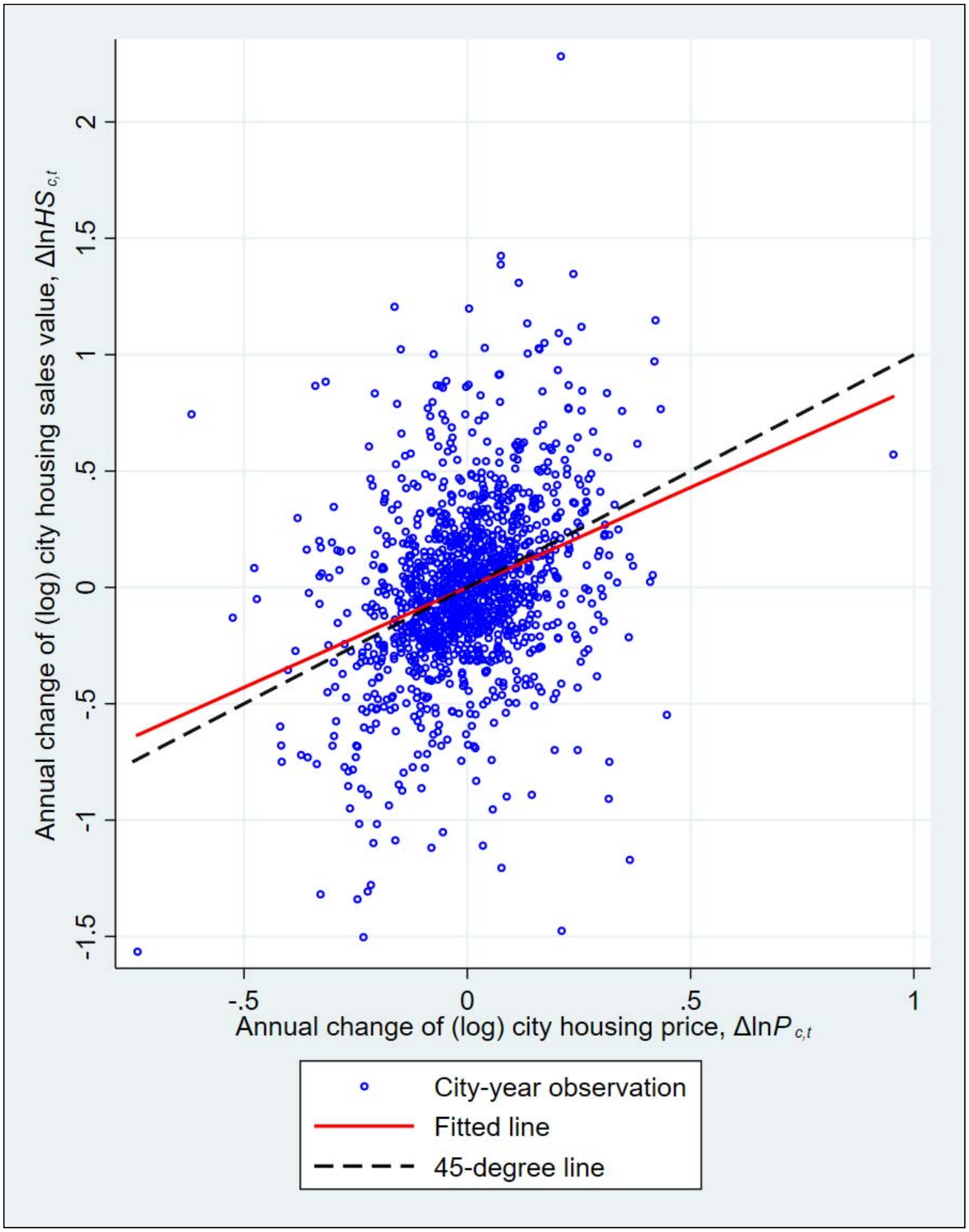

Figure 2: Relationship between the (log) change of housing prices and the (log) change of housing sales over the period 2002–2007. The dashed black line represents the 45-degree line. The red line represents the fitted line.

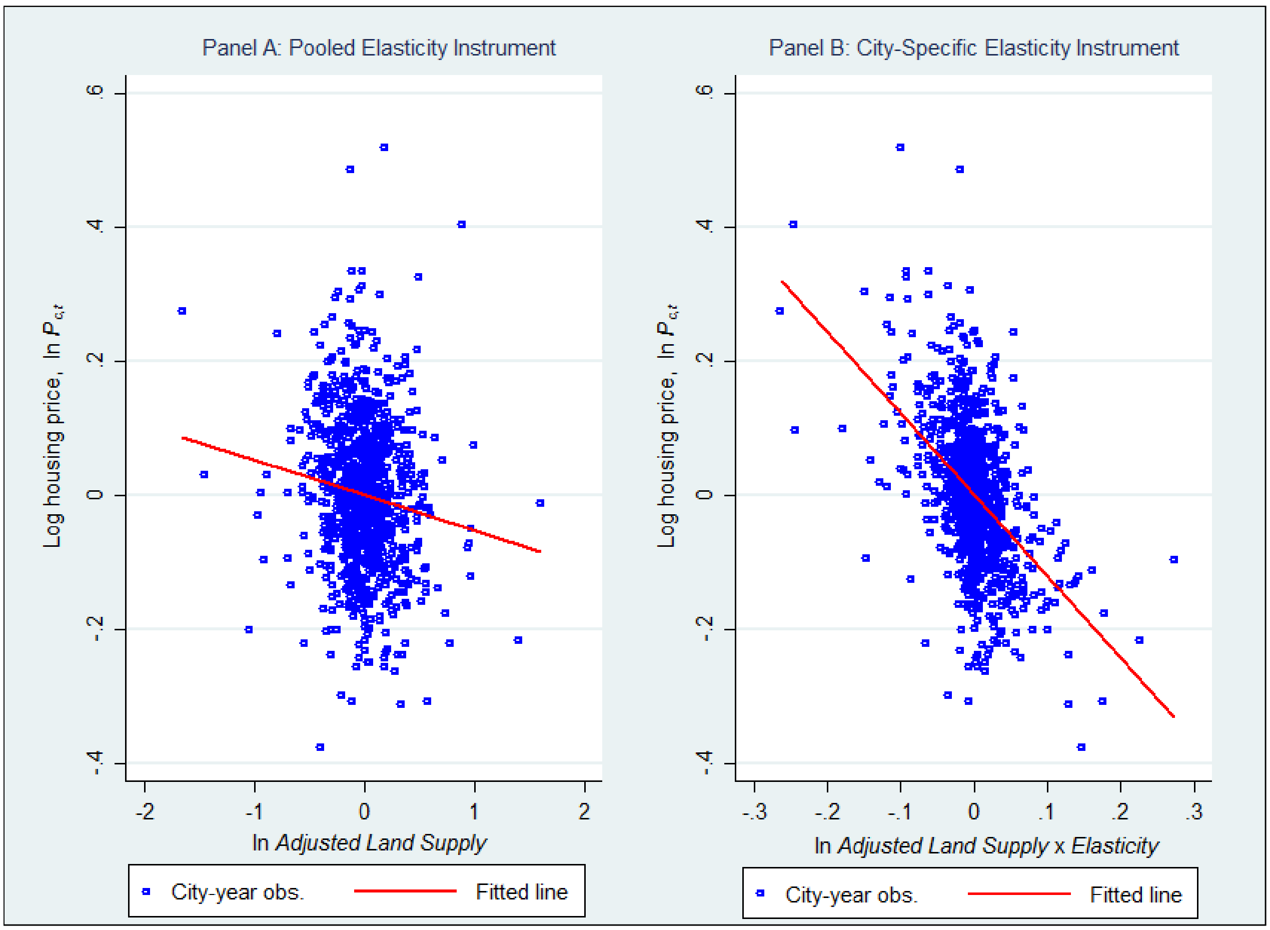

Figure 3: Relationship between (log) housing prices and (log) adjusted land supply. The left-hand panel uses (log) adjusted land supply as the explanatory variable, the right-hand panel uses the interaction of (log) adjusted land supply with elasticity as the explanatory variable. City and year fixed effects are filtered out. One city-year observation (for Sanya in 2002) was excluded from the graph as an outlier.

Figure 4: We graph the average (log) value added output change of all privately-owned firms (Panel A) and SOEs (Panel B) in each of 172 cities against the instrumented change in the (log) real estate price index from 2002 to 2007. Subtracted from the value-added output growth are interacted industry and year fixed effects.

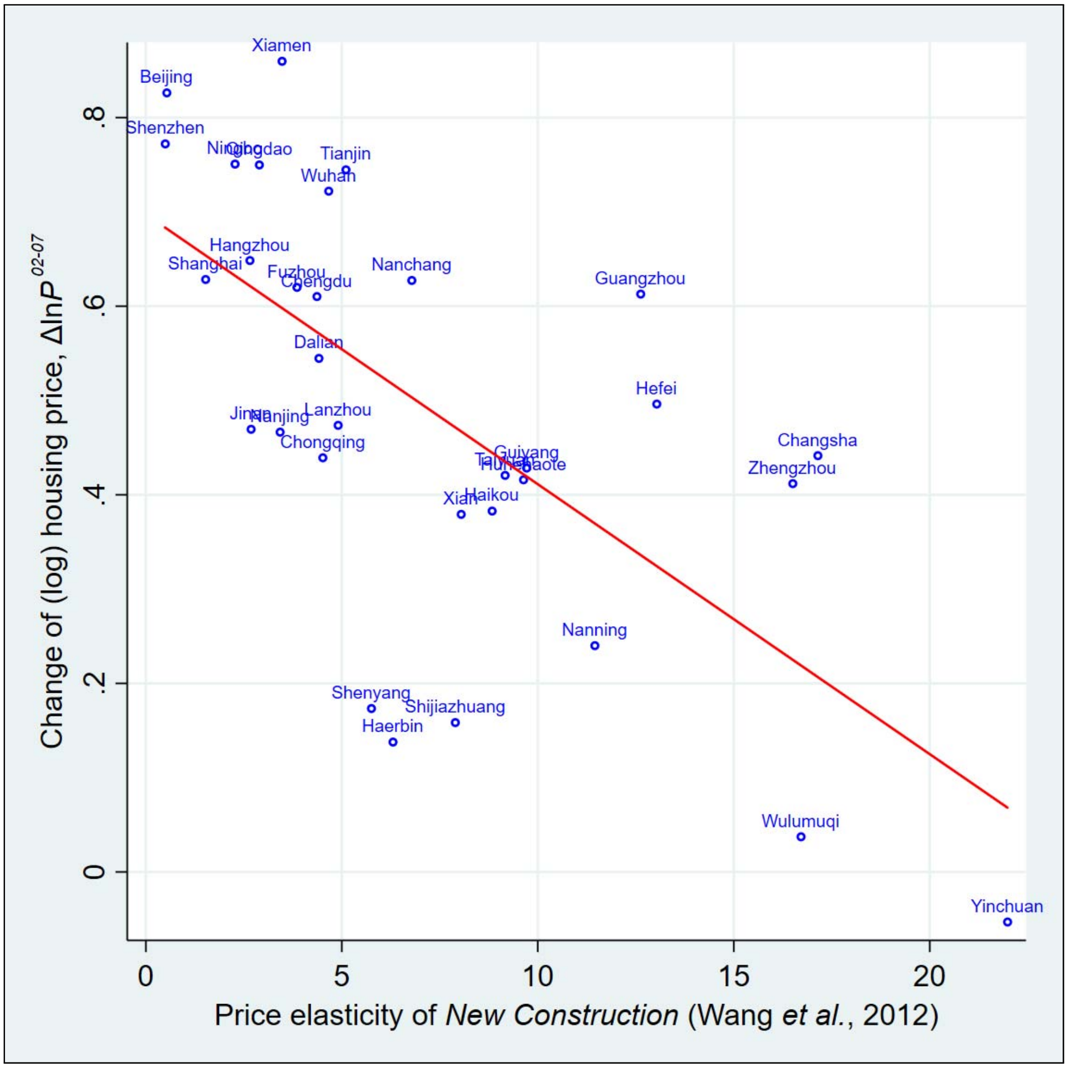

Figure 5: Relationship between the (log) change of housing prices between 2002 and 2007 and housing supply elasticity estimated by Wang *et al.* (2012).

Table 1: Summary Statistics

Summary statistics at the firm level are the (log) average employees' wage ($\ln w_{j,t}$), the firm bank loan rate ($i_{j,t}$), the net investment rate ($NI/K_{j,t}$), the gross investment rate ($I/K_{j,t}$), a dummy for whether a firm has long-run borrowing ($Loan_{j,t}$), the (log) employment size ($\ln L_{j,t}$), the (log) value-added output ($\ln Y_{j,t}$), the (log) labor productivity ($\ln(Y/L)_{j,t}$), a dummy variable for firms reported as "inactive" in period $t+1$ ($Exit_{j,t}$), the return on assets ($ROA_{j,t}$), the firm leverage ($Leverage_{j,t}$), the (log) total factor productivity ($\ln TFP_{j,t}$), a dummy for whether a firm is a state-owned-enterprises (SOE_j) at the beginning of the sample, and the firm's (log) fixed assets ($\ln$ *Fixed Assets*$_{j,t}$). We match additional product-level information from the Chinese customs authorities, which decomposes the annual (log) export value ($\ln ExpValue_{j,t}$) into firm-level export quantity ($\ln ExpQuantity_{j,t}$) and unit price ($\ln ExpPrice_{j,t}$). Summary statistics at the city level are the (log) average real house price ($\ln P_{c,t}$), annual change of (log) average real house price ($\ln P_{c,t}/P_{c,t-1}$) and the log of the *Adjusted Land Supply*$_{c,t}$. We use a generated instrumental variable ln *Adjusted Land Supply*$_{c,t} \times \widehat{\eta}_c$, which is the interaction of the (log) adjusted land supply with the city-level (inverse) land supply elasticity. We also report the median bank loan rate of all firms in a city as $i_{c,t}$, its annual change $i_{c,t} = i_{c,t} - i_{c,t-1}$, and its difference to all firms located in neighboring cities/prefectures $i_{c,t} - i_{c\_nb,t}$.

	Obs. (1)	Mean (2)	SD (3)	Q25 (4)	Q50 (5)	Q75 (6)
Panel A: Firm-level variables						
Firm wage: $\ln w_{j,t}$	916,051	2.588	0.552	2.246	2.561	2.912
Firm bank loan rate: $i_{j,t}$	423,014	0.061	0.042	0.028	0.051	0.084
$NI/K_{j,t}$	690,424	0.214	0.957	−0.094	−0.021	0.164
$I/K_{j,t}$	690,617	0.337	1.031	0	0.054	0.281
$Loan_{j,t}$ (dummy)	935,412	0.337	0.473	0	0	1
$\ln L_{j,t}$	918,040	4.758	1.009	4.043	4.682	5.394
$\ln Y_{j,t}$	914,345	8.788	1.188	7.908	8.644	9.534
$\ln(Y/L)_{j,t}$	915,668	4.023	0.969	3.348	3.959	4.653
$Exit_{j,t}$	935,315	0.095	0.294	0	0	0
$ROA_{j,t}$	916,417	0.072	0.120	0.006	0.036	0.099
$Leverage_{j,t}$	917,449	0.576	0.256	0.390	0.596	0.774
$\ln TFP_{j,t}$	859,742	1.140	0.337	0.981	1.193	1.367
ln *Fixed Assets*$_j$	313,100	7.773	1.804	6.721	7.760	8.840
SOE_j (Dummy)	313,100	0.043	0.203	0	0	0
$\ln ExpValue_{j,t}$	176,883	11.70	3.271	9.393	12.15	14.29
$\ln ExpQuantity_{j,t}$	176,883	10.49	3.527	8.027	10.82	13.19
$\ln ExpPrice_{j,t}$	176,883	1.225	1.736	0.258	1.099	2.050
Panel B: City-level variables						
$\ln P_{c,t}$	1,021	7.429	0.486	7.081	7.340	7.730
$\ln P_{c,t}/P_{c,t-1}$	844	0.096	0.134	0.020	0.093	0.179
ln *Adjusted Land Supply*$_{c,t}$	1,021	−9.617	1.101	−10.31	−9.652	−8.961
ln *Adjusted Land Supply*$_{c,t} \times \widehat{\eta}_c$	1,021	−0.769	0.984	−1.354	−0.773	−0.179
$i_{c,t}$	846	0.050	0.011	0.043	0.049	0.056
$\Delta i_{c,t}$	846	0.001	0.008	−0.002	0.001	0.005
$i_{c,t} - i_{c\_nb,t}$	846	−0.002	0.017	−0.009	−0.001	0.006

Table 2: Determinants of the Adjusted Land Supply

We define ln *Adjusted Land Supply*$_{c,t}$ as the log of land purchases by the real estate sector for residential housing development scaled by the lagged housing stock and predetermined population density in a city c in year t. The explanatory variables include annual city-level statistics for the (log) gross domestic product ($\ln GDP_{c,t}$), (log) population (ln *Population*$_{c,t}$), the percentage of "urban" area within the city territory (*Urban Share*$_{c,t}$), the percentage of park area within the urban area (*Park Share*$_{c,t}$), the annual (log) expenditure (ln *Expenditure*$_{c,t}$) of the city government, its annual (log) revenue (ln *Revenue*$_{c,t}$) and the (log) surface area of the urban road network (ln *Road Area*$_{c,t}$). In Column (6) all independent variables are lagged by one year $t-1$. In Column (7), the explanatory variables are the (log) annual population growth in year $t+1$, $t+2$ and $t+3$. Robust standard errors are in parentheses and are clustered at city level. Robust standard errors are in parentheses and are clustered at city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variable:	ln *Adjusted Land Supply*$_{c.t}$						
	OLS	OLS	OLS	OLS	OLS	OLS	OLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
$\ln GDP_{c,t-k}$	0.045				−0.043	−0.079	
	(0.142)				(0.163)	(0.152)	
ln *Population*$_{c,t-k}$	−0.032				−0.101	0.242	
	(0.233)				(0.202)	(0.307)	
Urban Share$_{c,t-k}$		−0.061			−0.069	−0.237	
		(0.234)			(0.238)	(0.211)	
Park Share$_{c,t-k}$		−0.016***			−0.017***	−0.011**	
		(0.005)			(0.006)	(0.005)	
ln *Expenditure*$_{c,t-k}$			0.154		0.156	0.267	
			(0.244)		(0.251)	(0.163)	
ln *Revenue*$_{c,t-k}$			0.059		0.085	−0.023	
			(0.113)		(0.126)	(0.095)	
ln *Road Area*$_{c,t-k}$				0.037	0.044	−0.073	
				(0.064)	(0.065)	(0.046)	
Pop_growth$_{c,t+1}$							−0.058
							(0.171)
Pop_growth$_{c,t+2}$							0.119
							(0.194)
Pop_growth$_{c,t+3}$							0.140
							(0.154)
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
City fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
$k=$	0	0	0	0	0	1	–
Observations	1,021	1,021	1,021	1,021	1,021	1,006	1,021
R-squared	0.315	0.326	0.318	0.315	0.330	0.347	0.316
Number of cities	172	172	172	172	172	172	172

Table 3: Housing Prices and Adjusted Land Supply

We define ln *Adjusted Land Supply*$_{c,t}$ in city c and year t as the lagged land purchases by the real estate sector for residential housing development scaled by the lagged housing stock and predetermined population density in a city c in year t. The (log) real estate price $\ln P_{c,t}$ is regressed in Columns (1) and (2) on ln *Adjusted Land Supply*$_{c,t}$ assuming an identical real estate price elasticity across cities (pooled elasticity: $\eta_c = \eta$), and in Columns (3) and (4) on the interaction term ln *Adjusted Land Supply*$_{c,t}$ $\times\ \widehat{\eta}_c$ using (estimated) city-level (inverse) land supply elasticity $\widehat{\eta}_c$. The control variables in Columns (2) and (4) include annual city-level statistics for the (log) gross domestic product ($\ln GDP_{c,t}$), (log) population (ln *Population*$_{c,t}$), the percentage of "urban" area within the city territory (*Urban Share*$_{c,t}$), the percentage of park area within the urban area (*Park Share*$_{c,t}$), the annual (log) expenditure (ln *Expenditure*$_{c,t}$) of the city government, its annual (log) revenue (ln *Revenue*$_{c,t}$) and the (log) surface area of the urban road network (ln *Road Area*$_{c,t}$). Robust standard errors are in parentheses and are clustered at city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variable:	$\ln P_{c,t}$			
	Pooled elasticity		City-specific elasticity	
	OLS	OLS	OLS	OLS
	(1)	(2)	(3)	(4)
ln *Adjusted Land Supply*$_{c,t}$	-0.062^{***}	-0.066^{***}		
	(0.018)	(0.018)		
ln *Adjusted Land Supply*$_{c,t}$ $\times\ \widehat{\eta}_c$			-1.215^{***}	-1.214^{***}
			(0.067)	(0.072)
$\ln GDP_{c,t}$		0.022		0.002
		(0.066)		(0.050)
ln *Population*$_{c,t}$		-0.126^{**}		-0.130^{**}
		(0.055)		(0.053)
Urban Share$_{c,t}$		-0.049		-0.036
		(0.065)		(0.068)
Park Share$_{c,t}$		-0.0004		-0.001
		(0.001)		(0.001)
ln *Expenditure*$_{c,t}$		0.069		0.082^{*}
		(0.056)		(0.049)
ln *Revenue*$_{c,t}$		0.077		0.061
		(0.052)		(0.044)
ln *Road Area*$_{c,t}$		-0.020		-0.011
		(0.029)		(0.021)
F-value	11.68	13.06	334.3	270.0
Year fixed effects	Yes	Yes	Yes	Yes
City fixed effects	Yes	Yes	Yes	Yes
Observations	1,021	1,021	1,021	1,021
R-squared	0.709	0.715	0.763	0.768
Number of cities	172	172	172	172

Table 4: Factor Price Response to Housing Price Inflation

We regress local factor prices on the log of the local real housing price ($\ln P_{c,t}$) in 172 Chinese cities for the period 2002–7. The dependent variables are the firm bank loan rate ($i_{j,t}$) in Columns (1)–(3), and the log average firm wage ($\ln w_{j,t}$) in Columns (4)–(6). Columns (1) and (3) report OLS results, Columns (2) and (4) the corresponding 2SLS results. Our instrument is the interaction term ln $\mathit{Adjusted\ Land\ Supply}_{c,t} \times \widehat{\eta}_c$, where $\widehat{\eta}_c$ denotes the city-leval (inverse) land supply elasticity. The 2SLS regressions in Columns (5) and (6) add in the second stage additional city-level controls, namely city (log) GDP ($\ln GDP_{c,t}$), (log) population (ln $\mathit{Population}_{c,t}$), the percentage of "urban" area within the city territory ($\mathit{Urban\ Share}_{c,t}$), the percentage of park area within the urban area ($\mathit{Park\ Share}_{c,t}$), the annual (log) expenditure (ln $\mathit{Expenditure}_{c,t}$) of the city government, its annual (log) revenue (ln $\mathit{Revenue}_{c,t}$) and the (log) surface area of the urban road network (ln $\mathit{Road\ Area}_{c,t}$). All regressions control year and firm fixed effects. Robust standard errors are provided in parentheses and are clustered at city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variables:	*Firm bank loan rate:* $i_{j.t}$			*Firm wage:* $\ln w_{j.t}$		
	OLS	2SLS	2SLS	OLS	2SLS	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)
$\ln P_{c,t}$	0.009^{***}	0.022^{**}	0.022^{***}	-0.111^{**}	-0.394^{***}	-0.391^{***}
	(0.002)	(0.009)	(0.008)	(0.053)	(0.120)	(0.117)
$\ln GDP_{c,t}$			−0.003			0.080
			(0.005)			(0.099)
ln $\mathit{Population}_{c,t}$			−0.003			-0.102^{***}
			(0.004)			(0.025)
$\mathit{Urban\ Share}_{c,t}$			0.002			−0.003
			(0.003)			(0.040)
$\mathit{Park\ Share}_{c,t}$			−0.00005			−0.002
			(0.0001)			(0.002)
ln $\mathit{Expenditure}_{c,t}$			-0.006^{*}			−0.029
			(0.003)			(0.091)
ln $\mathit{Revenue}_{c,t}$			0.003			0.076
			(0.003)			(0.049)
ln $\mathit{Road\ Area}_{c,t}$			0.001			0.021
			(0.001)			(0.034)
Kleibergen-Paap *F-stat*		97.8	98.1		106.3	100.2
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	423,014	423,014	423,014	916,051	916,051	916,051
Number of cities	172	172	172	172	172	172

Table 5: House Prices and Firm Production Choices

Different measures of firm production are regressed on the local housing price level $\ln P_{c,t}$. The 2SLS regressions use the interaction term ln *Adjusted Land Supply*$_{c,t} \times \widehat{\eta}_c$ as instruments, where $\widehat{\eta}_c$ denotes the city-level (inverse) land supply elasticity. Panel A reports the OLS regression, Panel B the baseline 2SLS regression, Panel C augments the 2SLS regression by macroeconomic control variables, and Panel D further controls for interacted industry and year FEs. The dependent variables are the net investment to capital share ($NI/K_{j,t}$) in Column (1), the gross investment rate ($I/K_{j,t}$) in Column (2), a dummy variable of whether firm $_j$ has long-run bank lending ($Loan_{j,t}$) in Column (3), the (log) firm employment ($\ln L_{j,t}$) in Column (4), the (log) value-added firm output ($\ln Y_{j,t}$) in Column (5), the (log) labor productivity ($\ln(Y/L)_{j,t}$) in Column (6), and a dummy variable for firms exit from the sample in period $t+1$ ($Exit_{j,t}$) in Column (7). The macroeconomic controls in Panel C including the percentage of park area within the urban area (*Park Share*$_{c,t}$), the city's (log) GDP ($\ln GDP_{c,t}$), the (log) population size (ln *Population*$_{c,t}$), the percentage of "urban" area within the city territory (*Urban Share*$_{c,t}$), the percentage of park area within the urban area (*Park Share*$_{c,t}$), the annual (log) expenditure (ln *Expenditure*$_{c,t}$) of the city government, its annual (log) revenue (ln *Revenue*$_{c,t}$) and the (log) surface area of the road network (ln *Road Area*$_{c,t}$). All regressions control for year and firm fixed effects. Standard errors are in parentheses and are clustered at the city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variables:	$NI/K_{j,t}$	$I/K_{j,t}$	$Loan_{j,t}$	$\ln L_{j,t}$	$\ln Y_{j,t}$	$\ln(Y/L)_{j,t}$	$Exit_{j,t}$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: OLS regressions							
$\ln P_{c,t}$	-0.099^{**}	-0.120^{**}	-0.017	0.045	-0.231^{**}	-0.267^{***}	-0.021
	(0.040)	(0.046)	(0.024)	(0.037)	(0.088)	(0.081)	(0.032)
Panel B: 2SLS regressions							
$\ln P_{c,t}$	-0.204^{***}	-0.265^{***}	-0.107^{**}	0.166	-0.968^{***}	-1.108^{***}	0.076^{**}
	(0.045)	(0.057)	(0.045)	(0.124)	(0.181)	(0.114)	(0.036)
Kleibergen-Paap *F-stat*	97.2	97.0	106.5	105.9	106.0	106.0	107.6
Panel C: 2SLS regressions with macroeconomic controls							
$\ln P_{c,t}$	-0.190^{***}	-0.254^{***}	-0.096^{*}	0.150	-0.889^{***}	-1.010^{***}	0.052
	(0.040)	(0.055)	(0.049)	(0.114)	(0.172)	(0.122)	(0.037)
Kleibergen-Paap *F-stat*	87.9	87.8	100.4	99.8	99.8	99.7	99.9
Panel D: 2SLS regressions with macroeconomic controls and industry × year FEs							
$\ln P_{c,t}$	-0.180^{***}	-0.244^{***}	-0.079^{**}	0.132	-0.876^{***}	-0.978^{***}	0.055
	(0.045)	(0.057)	(0.036)	(0.127)	(0.184)	(0.115)	(0.036)
Kleibergen-Paap *F-stat*	91.5	91.4	103.5	102.3	102.6	102.6	102.7
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	690,424	690,617	935,412	918,040	914,345	915,668	935,315

Table 6: Real Firm Performance Measures at Product Level

For a subsample of exporting firms with an export share larger than 75% of output, we use product-level export statistics from the Chinese customs authorities to decompose the yearly (log) export value of a firm's exported products ($\ln ExpValue_{j,t}$) into (directly reported) export quantity ($\ln ExpQuantity_{j,t}$) and export (unit) price ($\ln ExpPrice_{j,t}$). We repeat the 2SLS regression in Table 5, Panel B, for this subsample in Columns (1)–(2), and the new export performance measure in Columns (3)–(5). All regressions control for city level macroeconomic variables, interacted industry and year fixed effects and firm fixed effects. Standard errors are in parenthesis and are clustered at the city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

	Export firms		Product level firm performance		
Dependent variables:	$NI/K_{j,t}$	$\ln Y_{j,t}$	$\ln ExpValue_{j,t}$	$\ln ExpQuantity_{j,t}$	$\ln ExpPrice_{j,t}$
	2SLS	2SLS	2SLS	2SLS	2SLS
	(1)	(2)	(3)	(4)	(5)
$\ln P_{c,t}$	-0.359^{**}	-0.699^{***}	-0.421^{**}	-0.426^{**}	0.005
	(0.157)	(0.184)	(0.204)	(0.209)	(0.089)
Kleibergen-Paap *F-stat*	31.2	34.3	37.2	37.2	37.2
Macroeconomic controls	Yes	Yes	Yes	Yes	Yes
Industry $\times$ year fixed effects	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	53,558	65,399	175,001	175,001	175,001

Table 7: House Prices and Firm Heterogeneity in Bank Access

Different measures of firm production are regressed on the (log) local housing price level $\ln P_{c,t}$ and the interaction terms of the housing price level with a proxy for bank access. Panel A uses the ex-ante (log) fixed assets ($\ln$ *Fixed Assets*$_j$) as a measure of collateral availability. Panel B creates interaction terms with the state-ownership dummy (SOE_j), because Chinese SOEs enjoy privileged bank access. All regressions control year and firm fixed effects. Standard errors are in parentheses and are clustered at the city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variables:	$NI/K_{j,t}$	$I/K_{j,t}$	$Loan_{j,t}$	$\ln L_{j,t}$	$\ln Y_{j,t}$	$\ln(Y/L)_{j,t}$	$Exit_{j,t}$
	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A: Interaction with firm's fixed assets							
$\ln P_{c,t}$	-1.088^{***}	-1.238^{***}	-0.272^{***}	0.208^{*}	-1.084^{***}	-1.227^{***}	0.319^{***}
	(0.174)	(0.224)	(0.075)	(0.120)	(0.170)	(0.136)	(0.051)
$\ln P_{c,t} \times \ln$ *Fixed Assets*$_j$	0.114^{***}	0.125^{***}	0.022^{***}	-0.006	0.015	0.016	-0.032^{***}
	(0.020)	(0.025)	(0.006)	(0.009)	(0.020)	(0.014)	(0.004)
Kleibergen-Paap *F-stat*	85.8	85.5	76.6	80.9	81.5	77.9	84.4
Panel B: Interaction with *SOE* dummy							
$\ln P_{c,t}$	-0.206^{***}	-0.269^{***}	-0.109^{**}	0.167	-0.974^{***}	-1.113^{***}	0.078^{**}
	(0.043)	(0.054)	(0.043)	(0.122)	(0.187)	(0.117)	(0.037)
$\ln P_{c,t} \times SOE_j$	0.158^{***}	0.200^{***}	0.112^{*}	-0.006	0.324	0.356	-0.126
	(0.055)	(0.062)	(0.064)	(0.189)	(0.420)	(0.229)	(0.090)
Kleibergen-Paap *F-stat*	7.23	7.30	5.17	5.83	5.25	5.36	6.12
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	690,424	690,617	935,412	918,040	914,345	915,668	935,315

Table 8: House Prices and Firm Performance

Three different measures of firm performance, namely return on assets ($ROA_{j,t}$) in Columns (1)-(3), the debt to asset ratio ($Leverage_{j,t}$) in Columns (4)-(6), and (log) total factor productivity $\ln TFP_{j,t}$ in Columns (7)-(9) are regressed on the local (log) housing price level $\ln P_{c,t}$ and interaction terms of $\ln P_{c,t}$ with two different proxies for firm bank access, namely (log) fixed assets ($\ln$ *Fixed Assets*$_j$) as a measure of collateral availability and a dummy for state-owned enterprises (SOE_j). The house price and interaction terms are instrumented as in Table 7 by ln *Adjusted Land Supply*$_{c,t} \times \widehat{\eta}_c$. All regressions control year and firm fixed effects. Standard errors are in parentheses and are clustered at the city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dep. variables:	$ROA_{j,t}$			$Leverage_{j,t}$			$\ln TFP_{j,t}$		
	OLS	2SLS	2SLS	OLS	2SLS	2SLS	OLS	2SLS	2SLS
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$\ln P_{c,t}$	-0.039^{***}	-0.145^{***}	-0.215^{***}	0.021^{**}	0.094^{***}	0.134^{***}	-0.093^{***}	-0.298^{***}	-0.487^{***}
	(0.013)	(0.023)	(0.034)	(0.010)	(0.014)	(0.031)	(0.021)	(0.035)	(0.054)
$\ln P_{c,t}$ ×			0.009^{***}			-0.005			0.025^{***}
ln *Fixed Assets*$_j$			(0.003)			(0.004)			(0.004)
$\ln P_{c,t}$ ×			0.076			-0.048			0.056^{*}
SOE_j			(0.052)			(0.045)			(0.030)
Kl.-Paap *F-stat*		107.0	3.51		105.6	3.76		104.2	3.36
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	916,417	916,417	916,417	917,449	917,449	917,449	859,742	859,742	859,742

Internet Appendix

A. Model Generalization to Price Elastic Factor Supplies

The benchmark model presented in Section 2.1 assumes a fully price inelastic capital and labor supply. Here we relax this assumption and allow for a price elastic supply in both factors with positive elasticity parameters λ_i and λ_w, respectively. The factor supply constraints in Eqs. (3) and (4) generalize to

$$K_R + K_T = \overline{K}(1 + \lambda_i i) \tag{A1}$$
$$L = \overline{L}(1 + \lambda_w w), \tag{A2}$$

where $\lambda_i = \lambda_w = 0$ reprsents the benchmark case of fully price inelastic factor supplies. Linearizing eqs. (A1) and (A2) implies

$$\overline{K}_T \widehat{K}_T + \overline{K}_R \widehat{K}_R = \left(\frac{\lambda_i \bar{i}}{1 + \lambda_i \bar{i}}\right)\widehat{i} \tag{A3}$$
$$\widehat{L} = \left(\frac{\lambda_w \overline{w}}{1 + \lambda_w \overline{w}}\right)\widehat{w}, \tag{A4}$$

where $\overline{X}$ respesents the steady state value and $\widehat{X} = dX/\overline{X}$ the percentage change of any variable.

The zero-profit condition for tradeable sector implies

$$A_T L^{\mu} K^{1-\mu} - wL - iK = 0.$$

and the Talyor expansion gives

$$\widehat{A}_T = 0 = \mu\widehat{w} + (1 - \mu)\widehat{i}. \tag{A5}$$

Profit maximization in the tradable sector and constant factor shares further implies

$$\widehat{w} + \widehat{L} = \widehat{i} + \widehat{K}_T = \widehat{Y}_T \tag{A6}$$

For the generalized supply structure and $\widehat{K}_R = \widehat{P} + \widehat{S} = (1 - \gamma_p)\widehat{P}$, we directly obtain Proposition 1:

$$\widehat{i} = \mu \frac{\overline{K}_R}{(1 + B_0)\overline{K}_T}(1 - \gamma_p)\widehat{P} \tag{A7}$$
$$\widehat{w} = -(1 - \mu)\frac{\overline{K}_R}{(1 + B_0)\overline{K}_T}(1 - \gamma_p)\widehat{P}, \tag{A8}$$

where we define

$$B_0 = (1 - \mu)\frac{\lambda_w \overline{w}}{1 + \lambda_w \overline{w}} + \mu(1 + \lambda_i \bar{i})\frac{\lambda_i}{1 + \lambda_i}\frac{\overline{K}}{\overline{K}_T} \geq 0.$$

The variables $\overline{K}_R$, $\overline{K}_T$, $\bar{i}$, and $\overline{w}$ represent the steady state values for capital in the two sectors and for the factor prices. For $\lambda_i = \lambda_w = 0$, we obtain $B_0 = 0$. Because $B_0 \geq 0$, local interest rate changes

$\widehat{i}$ (real wage changes $\widehat{w}$) are again proportional (inversely proportional) to real estate prices inflation $\widehat{P}$.

Proposition 2 generalizes to the following expressions:

$$\widehat{K}_T = -\frac{(1+B_1)\overline{K}_R}{(1+B_0)\overline{K}_T}(1-\gamma_p)\widehat{P} \tag{A9}$$

$$(\widehat{NI/K})_T = -\frac{(1+B_1)\overline{K}_R}{(1+B_0)\overline{K}_T}(1-\gamma_p)\widehat{P} = \widehat{K}_T \tag{A10}$$

$$\widehat{Y}_T = -(1-\mu)\frac{(1+B_2)\overline{K}_R}{(1+B_0)\overline{K}_T}(1-\gamma_p)\widehat{P} \tag{A11}$$

$$(\widehat{Y/L})_T = -(1-\mu)\frac{\overline{K}_R}{(1+B_0)\overline{K}_T}(1-\gamma_p)\widehat{P}, \tag{A12}$$

where we define

$$\begin{aligned} B_1 &= (1-\mu)\frac{\lambda_w \overline{w}}{1+\lambda_w \overline{w}} \geq 0 \\ B_2 &= \frac{\lambda_w \overline{w}}{1+\lambda_w \overline{w}} \geq 0. \end{aligned}$$

For $\lambda_i = \lambda_w = 0$, we obtain $B_1 = 0$ and $B_2 = 0$. The capital stock change $\widehat{K}_T$, the net investment change $(\widehat{NI/K})_T$, output change $\widehat{Y}_T$, and labor productivity change $(\widehat{Y/L})_T$ of the manufacturing sector are still negative for a positive local housing price inflation $\widehat{P} > 0$.

B. Persistence of Corporate Loans Rate Differences Across Cities

We identify the loan rate of corproate bank debt from balance sheet data: the annual interest expenses are divided by the outstanding (bank) debt at the beginning of the year. For all firms with bank debt in a given city we calculate the median city-level bank loan rate $i_{c,t}$ and compare it to the median bank loan rate of all firms in the neigboring cities/prefectures denoted by $i_{c\_nb,t}$. Summary statistics are reported in Table 1 of the paper. To evaluate persistence of loan rate differences across cities we estimate an error correction model

$$\Delta i_{c,t+1} = i_{c,t+1} - i_{c,t} = \alpha + \beta \left(i_{c,t} - i_{c\_nb,t}\right) + \epsilon_{c,t},$$

by pooling the data across 172 cities. We report the results in Table A1.

C. Sample Construction

Our data source is the Annual Survey of Industrial Firms (ASIF) over the period 1998–2008. This survey covers industrial firms in the three sectors of mining, manufacturing, and utilities (water, electricity, gas production and supply). This section describes the data-cleaning procedures:

1. We drop firm-year observations before 2001 and after 2007 to focus on the period 2002–2007. Only in the calculation of net investment rates in 2002 do we use a firm's capital stock in 2001.

2. We drop all firms in the mining and utility sector from the sample and focus on manufacturing firms. Tthe corresponding two-digit industry codes are 13–43 according to GB/T 4574–2002.

3. We discard firms in Tibet because of data quality concerns.

4. We drop firm-year observations reported not to have an "operation status."

5. We drop firm-year observations whenever the gross operating income is below RMB 5 million.

6. We discard all firm-year observations reporting fewer than eight employees.

7. We drop firm-year observations reporting negative value-added.

8. We drop firm-year observations reporting a firm age of more than 100 years.

9. We do not use firm-year observations where the corresponding city-year information on housing prices and other macroeconomic control variables is not available.

10. We drop firms in the (i) tobacco and (ii) waste resources and waste materials recycling and processing industry. These two industries have too few observations to allow us to control for industry-year fixed effect in the panel regression.

The raw data comprise 2,635,787 firm-year observations, corresponding to 689,010 distinct firms. After these cleaning procedures, the gross sample has 991,487 firm-year observations belonging to 313,100 distinct firms. To mitigate the role of reporting errors, we also discard firm-year observations in the top 1% and bottom 1% percentiles for each variable. For the real interest rate i we only use firm-year observations in the 10% to 90% percentiles because this variable is estimated as interest cost over outstanding debt. Since a firm-year observation represents an outlier in one regression specification, but does not feature in another, the sample size can vary depending on regression specification.

D. Real Investment and Capital Stock Calculation

Under price inflation, the purchasing year of new capital matters for the real book value of a firm's asset. To obtain the real value of capital, net new investment should be deflated with the deflator corresponding to the investment year. Formally, the real book value follows as

$$K_T = \sum_{s=t}^{T} Deflator_s \times \widetilde{NI}_s + K_t$$

where K_t is the (begining-of-year) initial real book value at the year of entering survey, $\widetilde{NI}_s$ the nominal net new investment in period s, and $Deflator_s$ the price deflator for investment goods in period s. The intial real book value is unobservable and can at best be proxied by an adjustment to the nominal book value $\widetilde{K}_t$. We assume that a firm's (unobservable) nominal book value before entering the survey (prior to t) follows the growth rate of nominal fixed assets in the firm's the two-digit industry, hence

$$\widetilde{K}_t = \widetilde{K}_0 \prod_{\tau=0}^{t} (1 + r_\tau),$$

where $\widetilde{K}_0$ is the nominal book value when the firm starts operation in period 0, and r_τ the industry-level growth rate of nominal fixed asset in period τ. For the evolution of the nominal book value $\widetilde{K}_0, \widetilde{K}_1, ..., \widetilde{K}_{t-1}, \widetilde{K}_t$ in the industry, the initial real book value K_t can be approximated as

$$K_t = Deflator_0 \times \widetilde{K}_0 + \sum_{\tau=0}^{t} Deflator_\tau \times [\widetilde{K}_{\tau+1} - (1 - \delta)\widetilde{K}_\tau],$$

where δ is the depreciation rate of fixed assets.

The real net investment rate NI_s/K_s in period s then follows as

$$NI_s/K_s = \frac{Deflator_s \times \widetilde{NI}_s}{K_s}$$

where $\widetilde{NI}_s = \widetilde{K}_{s+1} - (1 - \delta)\widetilde{K}_s$. A simpler way to calculate the real net investment rate is to divide the real net investment by the total nominal assets so that we do not need to estimate the firm's real initial capital stock. This simplification gives similar estimation results.

Table A1: Persistence of the Corporate Loan Rate

We estimate an error correction model for the average corporate loan rate in each city relative to firms in neighboring cities. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variable:	$\Delta i_{c,t+1}$		
	(1)	(2)	(3)
$i_{c,t} - i_{c\_nb,t}$	-0.105^{***}	-0.138^{***}	-0.138^{***}
	(0.022)	(0.032)	(0.032)
City fixed effects	No	Yes	Yes
Year fixed effects	No	No	Yes
R-squared	0.047	0.133	0.203
Observations	846	846	846

Table A2: Using Time Difference as an Alternative Specification

As a robustness check, we use annual change of firm outcome variables $\Delta y_{j,t}$ as the dependent variables and annual change of (log) housing prices $\Delta \ln P_{c,t}$ as the sole explanatory variable. Our instrument is the interaction term $\Delta \ln \textit{Adjusted Land Supply}_{c,t} \times \widehat{\eta}_c$, where $\Delta \ln \textit{Adjusted Land Supply}_{c,t}$ denotes the annual change of (log) adjusted land supply and $\widehat{\eta}_c$ denotes the city-level (inverse) land supply elasticity. All regressions control for year fixed effects. Robust standard errors are in parentheses and robust standard errors clustered at city level are in brackets. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variables:	$\Delta(NI/K)_{j,t}$	$\Delta(I/K)_{j,t}$	$\Delta \ln L_{j,t}$	$\Delta \ln Y_{j,t}$	$\Delta \ln(Y/L)_{j,t}$
	2SLS	2SLS	2SLS	2SLS	2SLS
	(1)	(2)	(3)	(4)	(5)
$\Delta \ln P_{c,t}$	-0.083^{***}	-0.111^{***}	0.096	-0.688^{***}	-0.777^{***}
	(0.030)	(0.040)	(0.093)	(0.144)	(0.157)
Kleibergen-Paap *F-stat*	43.4	43.6	39.2	38.8	38.8
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	402,105	402,144	671,718	671,836	671,848

Table A3: Housing Supply Elasticity of Major Cities

This table reports the local housing supply elasticity of major cities in China provided by Wang *et al.* (2012) in the top 5 and bottom 5 (Column 1), their corresponding increase of house prices over the period 2003–2010 (Column 2), and the level of house prices in 2010 (Column 3).

City	Housing supply elasticity (1)	Housing prices growth (2003–2010) (2)	Housing price level (in 2010) (3)
Top 5			
Shenzhen	0.49	201.78%	19,169.89
Beijing	0.53	280.83%	17,782.01
Shanghai	1.52	183.01%	14,400.10
Ningbo	2.27	294.74%	11,223.62
Hangzhou	2.65	267.61%	14,133.06
Bottom 5			
Hefei	13.3	185.06%	5,904.47
Zhengzhou	16.5	143.71%	4,957.34
Wulumuqi	16.71	95.40%	4,443.26
Changsha	17.14	118.71%	4,418.11
Yinchuan	21.98	88.79%	3,928.93

Table A4: Using Housing Supply Elastcities as an Alternative Instrument

As a robustness check, we use the housing supply elasticities reported by Wang *et al.* (2012) as an alternative instrument in a cross-sectional specification at the firm level with two-digit industry fixed effects. All regressions control city-level (log) GDP per capita, (log) population density, employment share of the secondary sector and GDP share of the secondary sector in 2002. Robust standard errors are in parentheses and robust standard errors clustered at city level are in brackets. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variables:	$\Delta^{02-07} i_j$	$\Delta^{02-07} \ln w_j$	$\Delta^{02-07} (NI/K)_j$	$\Delta^{02-07} Loan_j$	$\Delta^{02-07} \ln L_j$
	2SLS	2SLS	2SLS	2SLS	2SLS
	(2)	(1)	(3)	(4)	(5)
$\Delta^{02-07} \ln P_c$	0.022	-0.217	-0.184	-0.013	-0.174
	(0.009)**	(0.061)***	(0.087)**	(0.050)	(0.072)**
	[0.015]***	[0.253]	[0.064]***	[0.117]	[0.253]
Kleibergen-Paap *F-stat*	7.1	5.4	3.6	5.4	5.3
Controls	Yes	Yes	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	9,717	27,339	20,307	28,359	27,448

Dep. variables:	$\Delta^{02-07} \ln Y_j$	$\Delta^{02-07} \ln(Y/L)_j$	$\Delta^{02-07} ROA_j$	$\Delta^{02-07} Leverage_j$	$\Delta^{02-07} \ln TFP_j$
	2SLS	2SLS	2SLS	2SLS	2SLS
	(6)	(7)	(8)	(9)	(10)
$\Delta^{02-07} \ln P_c$	-1.289	-0.968	-0.344	0.183	-0.333
	(0.114)***	(0.102)***	(0.020)***	(0.028)***	(0.042)***
	[0.596]**	[0.587]*	[0.228]	[0.113]	[0.145]**
Kleibergen-Paap *F-stat*	5.3	5.3	5.3	5.4	5.3
Controls	Yes	Yes	Yes	Yes	Yes
Industry fixed effects	Yes	Yes	Yes	Yes	Yes
Observations	26,908	25,482	27,471	27,661	25,066

Table A5: Do Initial Conditions Matter for the Investment Effect of Real Estate Booms?

As a robustness check, we verify whether the initial conditions at province level matter for investment regressions. All regressions control year and firm fixed effects. Robust standard errors are in parentheses and are clustered at city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dependent variables:	$NI/K_{j,t}$				
	2SLS	2SLS	2SLS	2SLS	2SLS
	(1)	(2)	(3)	(4)	(5)
$\ln P_{c,t}$	-0.204^{***}	-0.244^{***}	-0.248^{***}	-0.222^{***}	-0.313^{***}
	(0.045)	(0.077)	(0.064)	(0.071)	(0.114)
$\ln P_{c,t} \times (Loan/GDP)_p^{2000}$		0.026			
		(0.029)			
$\ln P_{c,t} \times (Corporate\ Loan/GDP)_p^{2000}$			0.206		
			(0.134)		
$\ln P_{c,t} \times (Deposit/GDP)_p^{2000}$				0.007	
				(0.016)	
$\ln P_{c,t} \times (Household\ Deposit/GDP)_p^{2000}$					0.139
					(0.111)
Kleibergen-Paap *F-stat*	97.2	34.3	36.9	39.0	36.4
Year fixed effects	Yes	Yes	Yes	Yes	Yes
Firm fixed effect	Yes	Yes	Yes	Yes	Yes
Observations	690,424	690,424	690,424	690,424	690,424

Table A6: Factor Price and Firm Performance Response by Firm Size

As a robustness check, we verify the factor price and firm performance responses by firm size. Specifically, small firms are firms with less than 200 employees, medium-size firms are have between 201 and 1000 employees, and large firms have more than 1000 employees. All regressions control for year and firm fixed effects. Robust standard errors are in parentheses and are clustered at the city level. We use ***, **, and * to denote statistical significance at the 1%, 5%, and 10% level, respectively.

Dep. variables:	$i_{j.t}$	$\ln w_{j.t}$	$NI/K_{j,t}$	$Loan_{j,t}$	$\ln Y_{j,t}$	$\ln(Y/L)_{j,t}$	$Exit_{j,t}$	$ROA_{j,t}$	$Leverage_{j,t}$	$\ln TFP_{j,t}$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Panel A: Full sample										
$\ln P_{c,t}$	0.022**	−0.394***	−0.204***	−0.107**	−0.968***	−1.108***	0.076**	−0.145***	0.094***	−0.298***
	(0.086)	(0.120)	(0.045)	(0.045)	(0.181)	(0.114)	(0.036)	(0.023)	(0.014)	(0.035)
Kleib.-Paap *F-stat*	97.8	106.3	97.2	106.9	106.0	106.0	107.6	107.0	105.6	104.2
Observations	423,014	916,051	690,424	935,412	914,345	915,668	935,315	916,417	917,449	859,742
Panel B: Small firms										
$\ln P_{c,t}$	0.022***	−0.465***	−0.222***	−0.086**	−1.150***	−1.161***	0.110**	−0.172***	0.112***	−0.311***
	(0.008)	(0.132)	(0.058)	(0.035)	(0.181)	(0.124)	(0.043)	(0.028)	(0.017)	(0.037)
Kleib.-Paap *F-stat*	88.7	107.4	101.6	108.6	106.8	107.2	108.0	108.3	107.0	104.5
Observations	297,700	684,531	493,600	698,702	688,212	685,222	698,631	685,059	686,358	640,334
Panel C: Medium-size firms										
$\ln P$	0.022**	−0.239**	−0.150***	−0.076	−0.712***	−0.995***	−0.019	−0.094***	0.064***	−0.260***
	(0.010)	(0.120)	(0.046)	(0.059)	(0.173)	(0.131)	(0.029)	(0.016)	(0.015)	(0.046)
Kleib.-Paap *F-stat*	103.1	94.5	83.3	94.3	95.1	93.6	96.5	94.1	93.1	92.9
Observations	107,165	202,489	170,559	206,870	202,448	201,703	206,850	202,083	201,968	192,164
Panel D: Large firms										
$\ln P$	0.010	−0.047	0.108	−0.062	−0.492*	−0.767***	−0.107*	−0.039***	−0.030	−0.195***
	(0.010)	(0.138)	(0.089)	(0.059)	(0.268)	(0.186)	(0.060)	(0.010)	(0.025)	(0.055)
Kleib.-Paap *F-stat*	96.7	89.9	82.2	88.0	92.4	87.2	91.4	88.0	87.1	92.5
Observations	18,149	29,031	26,265	29,840	23,685	28,743	29,834	29,275	29,123	27,244