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Real estate collateral value and investment: The case of China

Jing Wu [a], Joseph Gyourko [b,c], Yongheng Deng [d,e,*]

[a] Hang Lung Center for Real Estate, Department of Construction Management, Tsinghua University, China
[b] The Wharton School, University of Pennsylvania, United States
[c] NBER, United States
[d] Institute of Real Estate Studies, National University of Singapore, Singapore
[e] Department of Real Estate, and Department of Finance, National University of Singapore, Singapore

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ABSTRACT

Previous research on the United States and Japan finds economically large impacts of changing real estate collateral value on firm investment that amplified the business cycles of those countries. Working with unique data on land values in 35 major Chinese markets and a panel of firms outside the real estate industry, we estimate investment equations that yield no evidence of a collateral channel effect. Further analysis indicates that China's debt is not characterized by the frictions that give rise to collateral channel effects elsewhere. Essentially, financially constrained borrowers appear able credibly to commit to repay debt in China. While there is no impact on investment via the collateral channel, our results should not be interpreted as implying there will be no negative fallout from a potential real estate bust on the Chinese economy. There likely would be, but through different channels.

1. Introduction

In the absence of complete contracting, economists realized that pledging collateral such as owned real estate can allow firms to borrow more, and thus, to invest more (Barro, 1976; Stiglitz and Weiss, 1981; Hart and Moore, 1994). Macroeconomists recognized the implication this had for amplifying the business cycle via a collateral channel effect (Bernanke and Gertler, 1989; Kiyotaki and Moore, 1997). Falling asset values reduce the debt capacity of credit constrained firms, which depresses their investment on the downside of the cycle. An analogous impact occurs on the upside of the cycle when collateral values are increasing for these firms.

Research on the United States and Japan supports this theory and has shown that rises and declines in property values substantially amplify the volatility of investment by non-real estate firms (Chaney et al., 2012; Cvijanovic, 2014; Gan, 2007a, 2007b; Liu et al., 2013). For example, Chaney et al. (2012) report that a one standard deviation increase in underlying real estate collateral value is associated with over one-quarter of a standard deviation higher level of corporate investment. This implies about six cents added investment for every dollar increase in collateral value. Earlier research by Bernanke (1983) concludes that this factor helps account for the extraordinarily large variation in output during America's Great Depression.

The remarkable boom and recent cresting of China's housing and land markets raise the question of whether the amplitude of its economic cycle has been magnified by a collateral channel effect on investment. China is an increasingly important factor in the global economy, so the answer to this question is important. Two new data sources are combined to provide the first estimate of the impact of changing real estate collateral values on the investment behavior of Chinese firms outside the real estate sector. One is a constant quality land price series in 35 major Chinese cities; the other measures real estate collateral value for publicly-traded firms outside the property sector in China.

In stark contrast to the recent findings referenced above for America and Japan, we find no evidence of a collateral channel effect among non-real estate firms' borrowing and investment behavior in China. This conclusion is robust to a wide range of permutations. For example, there is no evidence of asymmetry in the collateral channel effect depending upon whether housing and

* Corresponding author at: Institute of Real Estate Studies, NUS, 21 Heng Mui Keng Terrace, #04-02, Singapore 119613, Singapore.
E-mail address: ydeng@nus.edu.sg (Y. Deng).

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land markets are rising or falling. We also do not see heterogeneity in impact by measures of likely financial constraint. For example, there is no difference in our baseline results depending upon whether the firm is a low versus high dividend payer. Nor is there any meaningful effect (or difference in impact) depending upon whether or not the firm is a state-owned enterprise (SOE). We can also rule out the possibility that our results might be driven by financially constrained firms tending to be located in markets without good investment opportunities (so that they rationally would not want to invest even if collateral value increased substantially to lessen the degree to which they are constrained). Actual growth rates of local GDP were high in absolute terms among the slowest growing of our 35 markets during our sample period, so it seems likely that there are profitable investment opportunities in those places. And, there is no evidence of a positive collateral channel effect among firms headquartered in the markets with the strongest growth trends.

While these are noteworthy finding in their own right, we also show that the analysis provides insight into the nature of China's financial markets more broadly. A well-known theoretical literature tells us that collateral channel effects would not be expected if no firms were credit constrained or if there was 'contract completeness' in the financial markets (Barro, 1976; Stiglitz and Weiss, 1981; Hart and Moore, 1994). There is reason to believe that an important type of company in China, the state-owned enterprise (SOE), is not financially constrained because of its special access to government-controlled bank funding (Allen et al., 2005; Ayyagari et al., 2010). Thus, SOEs have no need to rely on collateral value to fund their investment programs. In contrast, private firms (which we call non-SOEs) are highly likely to be constrained. In an environment with incomplete contracting so that credible commitments to repay debt could not be made, we would expect pledging collateral to ease financial constraints and make investment more plentiful (on the upside of a cycle).

That we find no evidence of a collateral channel effect for either group of firms raises the question of whether Chinese capital markets are fundamentally different in the sense that private firms can credibly commit to repay. Further empirical analysis of variation in collateral channel effects among financially-constrained non-SOEs supports this hypothesis. For example, we look at how estimated impacts differ by whether the local lending market is dominated by the four biggest lenders, each of which is itself a state-owned commercial bank.[1] The underlying hypothesis is that non-SOE firms can credibly commit to repay their lenders because the costs of defaulting on what is effectively an arm of the government in a state dominated by a single party are prohibitively high. Concomitantly, a given borrower is less able to credibly commit to repay if the lender is not one of the dominant SOE banks. We find evidence of collateral channel effects for private firms the lower the share of the 'big four' SOE lenders in the borrower's home market. A similar pattern is found in additional analysis using a variable that measures the transparency of the local market's business law environment. The stronger a city's underlying legal system's protections against unilateral government sanctions again non-party actors, the more we see a collateral channel effect among non-SOE borrowers.

In sum, financially constrained firms do exist in China among the group of non-state-owned enterprises. However, there is no evidence of 'contract incompleteness' in markets dominated by the big four SOE lenders or in markets with weaker legal systems that do not protect entities from government whim. In these cases, the frictions that give rise to collateral channel effects in other countries are absent in China, which is consistent with the claims of Allen et al. (2005). We would not expect meaningful collateral channel effects to occur unless, and until, China develops a more effective and independent legal system that can protect defaulting borrowers from unilateral sanctions by powerful state-supported creditors.

Before getting to that analysis, the next section describes the unique real estate and firm data we bring to bear in our estimation of the collateral channel effect. Section 3 then discusses our estimation strategy and reports initial results. Section 4 delves more deeply into the nature of Chinese financial markets with its analysis of non-SOEs. There is a brief conclusion.

2. Data on land values and listed firms

We bring two new data sources to bear on the question of whether there is a collateral channel effect on Chinese firm investment. Both are unique to the study of the Chinese economy. The first is a panel on land prices across 35 Chinese cities; the second is a panel on firms not directly involved in the real estate industry.

2.1. Land value data

Our land price series is based on sales of raw land by local governments, and is described more fully in Deng et al. (2012). While raw land sales are rarely observed in most countries, this is not the case in China. Local governments own all the urban land in the country and allow private parties to purchase use rights of up to 70 years for residential purposes (i.e., technically, this is a leasehold estate).[2] We treat the upfront lump sum payment as the transactions price for raw land because there are no further rental payments required.

Our data series begins in 2003 because of an important 2002 ruling by the Ministry of Land and Resources that required local governments to sell land via public auction and to publicly report the winning bidder along with the transactions price. This marks an important break with past practice that has been criticized as open to corruption (Cai et al., 2013), which muddies the interpretation of price data before this change. We also typically observe the land parcel's precise address, designated usage, land conditions upon delivery, and certain planning indicators such as the floor-to-area ratio.

Building upon prior research on the city of Beijing in Wu et al. (2012), we worked with a leading residential real estate data vendor in China (Soufun) to collect information on all residential usage land sales to private parties from 2003 to 2011 in the 35 major markets mapped in Fig. 1. The geographic breadth of our sample is noteworthy. We are not limited to a few coastal-region markets that the media typically report to have the biggest booms. Table 1 reports summary statistics on the sample. We have complete data dated since 2003 for 15 markets, with the rest entering the sample in subsequent years. The number of transactions per market ranges from 25 to 50 depending upon the year.

Land parcels in China are priced in terms of the floor area of housing permitted to be built on the parcel, instead of in terms of the land area. For each parcel, its real price in constant 2009

[1] These firms are Industrial and Commercial Bank of China (ICBC), China Construction Bank (CCB), Agricultural Bank of China (ABC) and Bank of China (BOC). Together, they accounted for just over 42% of the bank loan market in China in 2012. See Deng et al. (in press) for more institutional detail about these four dominant state-owned banks.

[2] Not only does Chinese law facilitate the use of such leasehold estates in urban areas as collateral for borrowing, but the data confirm that they can and will be transferred to the lender if the borrower defaults. For example, 14 of the 16 commercial banks listed on the Shanghai or Shenzhen exchanges regularly report the value and breakdown of repossessed assets seized because of defaulted loans. At the end of 2011, the total book value of their repossessed assets was 10.79 billion yuan RMB, of which the leasehold estates associated with properties accounted for 8.79 billion yuan RMB (or 81.4%). The remainder was comprised of plant and equipment, securities, etc.

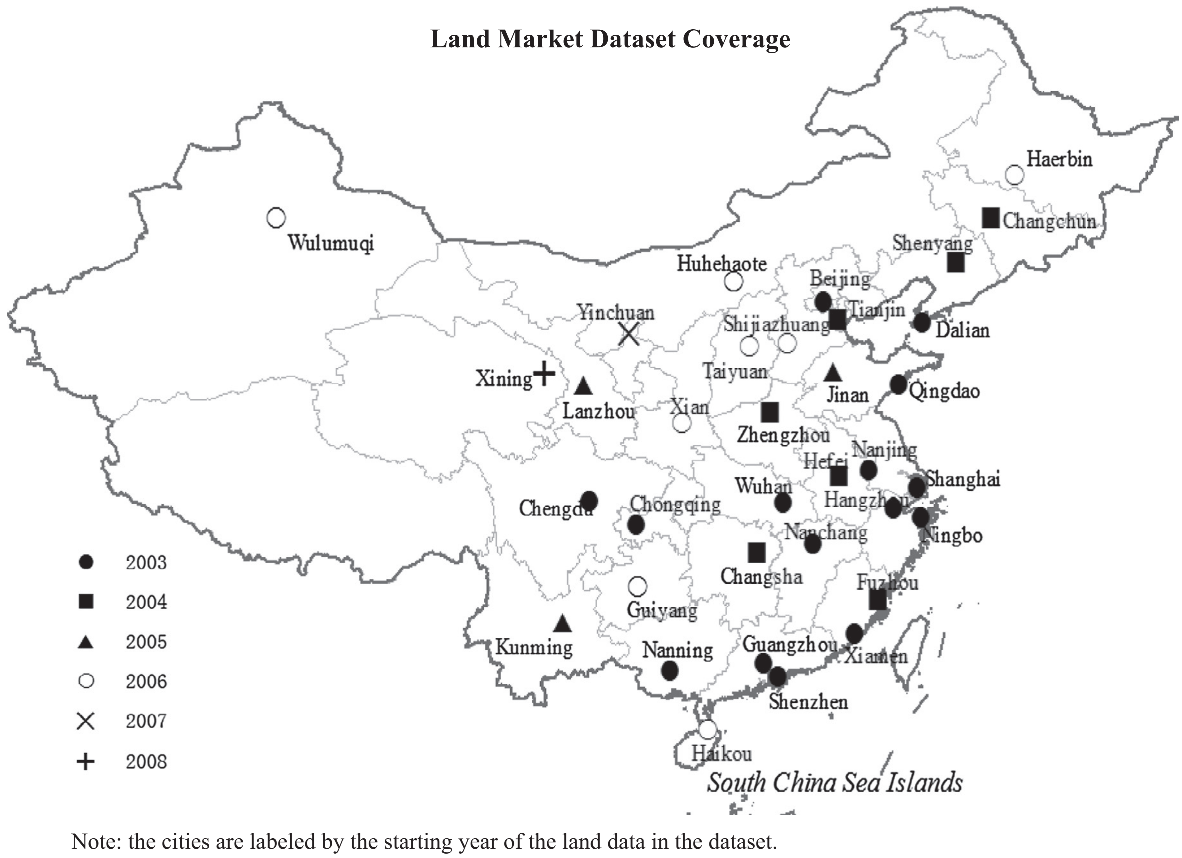

Fig. 1. Land market dataset coverage.

yuan per square meter of permitted space is computed by deflating with the relevant monthly CPI series for each city. We do not work with these unadjusted transactions prices because they may be driven by quality changes over time. Hence, we follow Wu et al. (2012) in creating constant quality land price indexes for each market.[3] Tables 2 and 3 (both from Deng et al. (2012)) report summary statistics on average annual land price growth over time and real compound average constant quality price appreciation rates for each city, respectively. These data show that there clearly was a boom in land prices in most cities in China, although it is incorrect to claim that there is a single national land market, as there is substantial variation in land price appreciation across and within cities over time. Other data from Deng et al. (2012) not reported here for space reasons highlights that land values are much more volatile than house prices and other factors of production in housing. Standard deviations in land prices typically are in the 20–40% range, which is about four times that of house prices, construction sector wages or physical construction costs. This is consistent with standard real estate models, as theory predicts the residual claimant on property value (i.e., land) should be much more volatile.

Table 1
Sample sizes in the land transaction dataset.

	Number of cities covered	Number of land parcels sold
2003	15	378
2004	22	681
2005	24	773
2006	33	1133
2007	34	1413
2008	35	963
2009	35	1564
2010	35	1759
2011	35	1749
Aggregated	–	10,413

Twenty-seven of the 35 markets have experienced real average annual growth rates in constant quality land prices of above 10% for the length of their sample periods. Nine have experienced average compound annual growth rates above 20%. Naturally, this

[3] City-level hedonic models are estimated via ordinary least squares (OLS), with the log of the real transactions price in constant 2009 yuan as the dependent variable. Quality controls on the right-hand side include: (a) the parcel's distance to the center of the corresponding city, which is measured after mapping the precise location of each site with GIS software; (b) the distance to the nearest subway station; this variable is relevant in 10 of the 35 cities with operating subway systems during our sample period; (c) district dummies which control for local/neighborhood-level fixed effects not captured by the two previous location controls; (d) a set of physical attributes including the size of the parcel (in land area), the density permitted on the site when built, and whether the parcel is leveled on delivery; (e) in some cases, a small portion of a residential land parcel is designated for affiliated commercial properties, public establishments, or public housing units; we control for such conditions via a set of dummies; (f) the parcel's transaction form as reflected in whether it was purchased via sealed bidding, regular English auction, or two-stage auction; and (g) year dummies, whose coefficients are used to create the constant quality price index. We also conducted a two-stage Heckman estimation to control for potential bias arising from the fact that there were a total of 614 parcels listed that failed to result in transactions (either because there were no bidders if there was an auction or the bid prices were lower than the local governments' reserve prices, which is relevant for cases involving sealed bids). If these failures were disproportionately concentrated in certain periods such as the financial crisis, selection bias would result in an overestimation of the price index for that period. That said, we could not find any statistically significant impact for the inverse Mills ratio estimated from our first-stage probit model. Finally, correcting for quality changes over time is statistically and economically important. Average annual appreciation in our hedonic price series is about five percentage points higher than in the unadjusted prices series, which indicates that parcel quality has been falling over time on average. The declining quality of location with more sites being in outlying areas as Chinese cities have rapidly urbanized is an important factor, but this varies by time and market. See Deng et al. (2012) for more detail. All underlying results are available upon request.

Table 2
Annual real land price appreciation, summary statistics, 35 major Chinese markets.

	2003–2004	2004–2005	2005–2006	2006–2007	2007–2008	2008–2009	2009–2010	2010–2011
Mean (%)	32.1	12.2	23.5	46.4	−5.3	28.5	31.4	2.6
Standard deviation (%)	21.7	23.1	40.5	42.1	24.0	30.7	29.4	30.2
Max (%)	64.1	47.2	128.8	131.2	38.6	93.1	83.6	108.6
Median (%)	27.8	10.0	20.7	47.7	6.6	29.5	41.5	2.7
Min (%)	4.4	−28.0	−36.1	−29.2	−59.9	−20.2	−31.6	−44.2
Number of cities	15	22	24	33	34	35	35	35
Number with positive appreciation	15	15	17	28	18	27	29	16
Number with negative appreciation	0	7	7	5	16	8	6	19

implies large increases in real estate values, as a 10% compound annual average rate of price appreciation implies more than a doubling in real value between 2003 and 2011 (i.e., $1.1^8 = 2.14$); analogously, a 20% compound annual rate implies that prices were over three times greater at the end of 2011 than at the beginning in 2003 (i.e., $1.2^8 = 4.30$). Thus, there is no doubt that constant quality land prices are higher now in virtually every market than they were in 2003. Hence, if Chinese firms are financially constrained and collateral value is important in obtaining the desired amount of debt, there has been a huge increase in those values over time on average, with economically large variation across markets.[4]

We believe these data are far superior to any alternative, the most prominent of which are two government-provided series on house prices. One is called the Average Selling Price of Newly-Built Residential Units. This reports the simple average of transactions prices as measured by total sales values divided by the total amount of housing square footage in the units. This series makes no attempt to control for quality differences across markets or drift over time. As Wu et al. (2014) show, not effectively controlling for quality leads to substantial biases. The other government-provided house price series, officially termed "Price Indices in 70 Large and Medium-Sized Cities", is a measure of the change in average prices on unit sales within individual housing complexes over time. More specifically, this index is calculated by first computing the average sales price of new units each month in each distinct housing complex. The series reported by the government then is the transactions-volume weighted average of each complex's average price changes over time. As Deng et al. (2012) show, this series has very little variation over time in most markets, including the big coastal region cities that are thought to have had the biggest booms. That makes this data source highly suspect on its face, and Wu et al. (2014) explain why it produces downwardly biased estimates of price growth, with much lower price volatility than exists in reality.

2.2. Firm data

We next collected data on publicly-traded firms listed on the Shanghai or Shenzhen stock exchanges. There are 1291 firms listed during or before 2003 on these two exchanges. We get to our final sample size of 444 firms as follows. First, we exclude firms delisted during or before 2011. We also drop those with measured negative equity and those involved in a major takeover operation during the sample period, as we suspect either data error or some other aspect of firm strategy is likely to complicate any potential relationship between real estate collateral value and investment and other spending behavior. Next, firms operating in the industries of "finance", "real estate", and "construction" are dropped because it seems likely that the relationship between firm investment behavior and real estate price fluctuations may be determined by a mechanism other than the collateral channel effect in these business sectors. Firms in the industries of "agriculture", "mining", "production and supply of electricity, gas and water" and "transportation and storage" also are excluded because they tend to own properties outside of urban areas, and we cannot impute property value price changes outside of the 35 major markets listed above. Thus, our final sample is restricted to firms in the industries of "manufacturing", "information technology", "wholesale and retail", "social service", and "media and culture". According to the official industry codes issued by the China Security Regulatory Commission, there are also a few firms defined as in "multiple industries". These companies are grouped based on their largest sales sector. We also restrict our focus to firms with headquarters in 35 major cities for which we have a reliable land price series that is used to impute the value of real estate collateral over time. This leaves us with a balanced panel of 444 firms with 3996 firm-year observations during 2003–2011.

Determining the market value of these firms' real estate asset holdings obviously is a critical task for our estimation. The nature of Chinese accounting and reporting practices is such that three major categories of assets on the balance sheet are involved in constructing our measure. One is "Buildings" (a sub-entry of "Fixed Assets", the equivalent of "Property, Plant and Equipment" in China's accounting codes); a second is "Land Ownership" (a sub-entry of "Intangible Assets"); and the third is "Investment Properties". Table 4 provides more detail on related accounting codes, including a minor adjustment in 2007. Unfortunately, none of the available listed firm electronic databases in China presently provides full information on all three categories of property assets. Consequently, we manually collected this information from the original version of the firms' annual financial reports.

While we believe this is the first systematic collection of non-real estate firms' property holdings, this is only the starting point for our analysis. As with the Compustat data on U.S. corporations, Chinese company financial reports include values based on historic cost, not current market values. We follow the procedures adopted by Chaney et al. (2012) to translate these book values into market values. From the financial reports, we know both the original book value and the current book value after depreciation. Then, following certain rules on depreciation, the average age of properties can be computed.[5] Finally, the book value is updated to the market value

[4] We use residential land prices because it is not yet feasible to create an analogous index for commercial properties. Theory suggests these two series should be highly correlated, since land is substitutable between these two uses on the margin. As a robustness check, we calculated the correlation coefficient between the average transaction prices of newly-built residential buildings and commercial properties as reported by the National Bureau of Statistics of China in our 35 city sample. It was a strongly positive 0.64. Chaney et al. (2012) report similar findings in their robustness checks using U.S. data.

[5] We use the following strategy to guide us in these calculations. If a firm explicitly describes its depreciation method in the appendix of its financial reports, we adopt that specific rule for that firm. Otherwise, we follow conventional rules on depreciation for China, which reflect an average of the different rules published by the listed firms in our sample: for the items of "Building" and "Investment Properties", it is assumed that the properties are straight-line depreciated with 25 years of depreciable life and 5% remains; for "Land Ownership", the corresponding assumption is straight-line depreciation method, 40 years of depreciable life, and 0 remains. Western readers may be surprised by the depreciation of land, but it does make sense because this technically is a leasehold estate position.

Table 3
Compounded real annual appreciation rates in constant quality land values, 35 major Chinese markets (2009 yuan).

15 Markets, 2003–2011 (8 years)		7 Markets, 2004–2011 (7 years)		2 Markets, 2005–2011 (6 years)		9 Markets, 2006–2011 (5 years)		1 Market, 2007–2011 (4 years)		1 Market, 2008–2011 (3 years)	
Chongqing	27.0%	Hefei	30.1%	Lanzhou	20.7%	Huhehaote	19.7%	Yinchuan	8.9%	Xining	49.9%
Shanghai	23.7%	Changsha	20.3%	Guiyang	12.4%	Haikou	17.8%				
Hangzhou	21.8%	Tianjin	20.2%			Taiyuan	12.2%				
Nanjing	20.5%	Fuzhou	17.7%			Haerbin	10.8%				
Beijing	20.2%	Changchun	13.9%			Jinan	7.2%				
Shenzhen	20.1%	Shenyang	13.7%			Xian	6.9%				
Xiamen	18.7%	Zhengzhou	10.3%			Shijiazhuang	5.1%				
Ningbo	18.5%					Kunming	2.0%				
Chengdu	16.7%					Wulumuqi	−2.9%				
Dalian	15.8%										
Guangzhou	14.7%										
Wuhan	13.3%										
Nanning	12.3%										
Nanchang	7.8%										
Qingdao	5.9%										

Table 4
Real estate asset information in the financial reports.

(A) Before 2006

		Self-occupied and lease-out properties
Self-built properties	Construction in progress	• The lands are listed in the item of "Land Ownerships" as a subentry of "Intangible Assets" • The construction materials, affiliated plants and equipment are listed in the item of "Construction in Progress"
	Completed properties	• The lands are listed in the item of "Land Ownerships" as a subentry of "Intangible Assets" • The structures are listed in the item of "Buildings" as a subentry of "Fixed Assets". The plants and equipment are excluded
Purchased properties		• Both lands and structures are listed in the item of "Buildings" as a subentry of "Fixed Assets". The plants and equipment are excluded

(B) Since 2007

		Self-occupied properties	Lease-out properties
Self-built properties	Construction in progress	• The lands are listed in the item of "Land Ownerships" as a subentry of "Intangible Assets" • The construction materials, affiliated plants and equipment are listed in the item of "Construction in Progress"	
	Completed properties	• The lands are listed in the item of "Land Ownerships" as a subentry of "Intangible Assets" • The structures are listed in the item of "Buildings" as a subentry of "Fixed Assets". The plants and equipment are excluded	• Both lands and buildings are listed in the item of "Investment Properties". The plants and equipment are excluded
Purchased properties		• Both lands and structures are listed in the item of "Buildings" as a subentry of "Fixed Assets". The plants and equipment are excluded	

using the city-level residential land price index described above after 2003, a constant quality newly-built house price index between 2000 and 2002 (Wu et al., 2014), and the city-level CPI index before 2000. Because we do not know the exact address of each property in a firm's portfolio, we follow Chaney et al. (2012) and Cvijanovic (2014) in presuming that a firm's properties are concentrated in the city of its headquarters.[6]

We next develop an estimate of annual change in the value of firms' real estate asset holdings. Our preferred measure is one that reflects changes in the market value over time of real estate assets owned by the firm in the reference year of 2002 at the very beginning of our sample period. Chaney et al. (2012) and Cvijanovic (2014) both do something similar to guard against bias arising from the potentially endogenous decisions of firms to alter real estate holdings in response to (or in conjunction with) market price changes. Thus, our collateral value measure is defined as:

$$RATIO\_REV1_{i,t} = \left[REV_{i,2002} \times \prod_{j=2003}^{t-1} (1 + LPG_{c,j}) \times LPG_{c,t}\right] \Big/ ASSET_{i,t-1}$$

where $REV_{i,2002}$ is the market value of real estate assets owned by firm i at the end of 2002 computed based on the procedures described above, $LPG_{c,j}$ is the annual growth rate in the local land price index for firm i's headquarters city c in year j, and $ASSET_{i,t-1}$ is the total assets of firm i at the beginning of year t (i.e., at the end of the previous year).

As part of our robustness checks described more fully below, we also used a second proxy, which measures the market value change in real estate assets held by the firm at the beginning of each year:

$$RATIO\_REV2_{i,t} = [REV_{i,t-1} \times LPG_{c,t}]/ASSET_{i,t-1}$$

where $REV_{i,t-1}$ is the market value of real estate assets owned by firm i at the beginning of year t (i.e., at the end of the previous year).

[6] Both Chaney et al. (2012) and Cvijanovic (2014) investigate the robustness of this assumption using added information from firm 10-K filings. Unfortunately, similar documents and data are not available in China. We addressed the robustness of this assumption as follows. First, we pared down the sample to firms headquartered in the 32 cities outside of the three national financial centers of Beijing, Shanghai and Shenzhen on the presumption that firms located in the other 32 cities are less likely to be geographically dispersed in their business and, hence, in their real estate asset holdings. All our key results reported below still hold in this "geographically concentrated" group.

Table 5
Definition and summary statistics of variables.

Variable	Definition	Average	Std. dev.
ASSET	Total assets at the beginning of the year; billion yuan RMB	4.882	17.598
RATIO_REV1	Change in the market value of real estate assets held in the reference year 2002, normalized by firm assets (see the text for more details)	0.060	0.151
RATIO_REV2	Change in the market value of real estate assets held at the beginning of each year, normalized by firm assets (see the text for more details)	0.075	0.170
RATIO_INV	Net change in investment on fixed assets, normalized by firm assets (see the text for more details)	0.056	0.056
RATIO_LOAN	Net change in loan balance, normalized by firm assets	0.019	0.074
RATIO_EBITDA	Earnings before interest, taxes, depreciation and amortization, normalized by firm assets	0.089	0.059
MBR	Market-to-book ratio at the beginning of the year	1.627	1.043
LEVERAGE	Leverage level at the beginning of the year	0.501	0.159

We experiment with both measures because it is not obvious *a priori* what the optimal balance is between potential endogeneity bias and measurement error.

In addition to our measures of changes in underlying real estate collateral, we also use a number of variables describing other firm characteristics when estimating collateral channel effects. These are from Wind Info (www.wind.com.cn), which is a supplier of 'Compustat-type' data on Chinese companies. These include the ratio of net investment on fixed assets (property, plant and equipment) to firm asset value (*RATIO_INV*), where the numerator is defined as expenditures on fixed assets less cash inflows from disposing of existing fixed assets over the year and the denominator reflects total assets at the beginning of the relevant year (*ASSET*), the ratio of the net change in firm debt to firm asset value (*RATIO_LOAN*), *RATIO_EBITDA*, which reflects earnings before interest tax, depreciation and amortization (again scaled by firm assets), the market-to-book ratio (*MBR*), and leverage level at the beginning of the year (*LEVERAGE*, defined as total debt on the balance sheet divided by asset value).

Table 5 reports the summary statistics on the variables, with each having been winsorized at the 5th percentile to eliminate extreme outliers in the data series. Winsorizing at different cutoff points (including not dropping outliers) does not materially change the results. One noteworthy feature is the large magnitude of the annual market value change of the listed firms' real estate assets. On average, it is equivalent to about 6% of a firm's total assets if we only take properties owned in the reference year into account, and is about 7.5% if all real estate assets are included. The fact that the average value of *RATIO_REV2* is larger than *RATIO_REV1* implies that the listed firms generally are expanding their real estate holdings over our sample period.

It also is the case that these firms have ample amounts of secured and unsecured debts, with the share of secured loans being higher. For example, from 2007 to 2011 the average annual share of their long-term debt that is secured is about 78%. The analogous figure for short-term debt (<1 year) is about 65%.

Table 6 then reports the number of firms in our sample broken down by whether or not they are state-owned enterprises (SOEs). This firm characteristic also comes from the Wind Info data source.[7] SOEs account for about three quarters of these 444 firms, although that proportion declines over time due.

Table 7 compares the values of these variables across the two types of firms. SOEs and non-SOEs differ in several aspects. SOEs tend to be much larger than non-SOEs. But they are less profitable, and have lower market-to-book ratios. However, both these groups experience almost the same degree of change in the market values of their real estate assets during the sample period. And, the differences in their fixed assets investment and net loan change are both statistically insignificant.

We also collected data on a number of other firm financial traits. These include *RATIO_CASH* which reflects the net change in cash holdings divided by assets, *RATIO_SALARY* which is defined as total salary payments divided by firm assets, *RATIO_DIVIDEND* which is total dividend payments scaled by firm assets, *RATIO_FAINV* which is the ratio of net investment on financial assets such as stocks and bonds to firm asset value, and *EMP*, which is the number of employees per million yuan of firm assets.

Information also was collected on the total amount of government subsidies received by a firm during each year. This also comes from Wind Info which includes this as a sub-entry of "non-operating income" in each firm's income statement. This includes both explicit monetary subsidies and implicit subsidies from discounted tax rates. As with the other variables, this also is normalized by the total assets of the firm (*RATIO_SUBSIDY*). It

Table 6
Distribution of sample by ownership structure.

Year	Number of SOEs	Number of non-SOEs
2003	353	91
2004	347	97
2005	343	101
2006	329	115
2007	323	121
2008	325	119
2009	320	114
2010	318	126
2011	318	126

Table 7
Summary statistics of variables by ownership structure groups.

	SOEs		Non-SOEs		*t* Stat.
	Average	Std. dev.	Average	Std. dev.	
ASSET	5.477	20.089	3.147	5.650	3.655***
RATIO_REV1	0.060	0.153	0.060	0.144	0.018
RATIO_REV2	0.076	0.173	0.075	0.163	0.062
RATIO_INV	0.057	0.055	0.055	0.057	1.052
RATIO_LOAN	0.019	0.073	0.021	0.079	1.071
RATIO_EBITDA	0.086	0.057	0.096	0.065	4.442***
MBR	1.552	0.972	1.847	1.201	7.817***
LEVERAGE	0.502	0.158	0.499	0.163	0.455

* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

[7] The ownership structure is defined according to the dominant controller of the firm (*shi ji kong zhi ren*) based on the official standard used by the China Security Regulatory Commission. More detailed information is available, as we can tell whether a SOE is directly controlled by the central government or by some type of sub-national government. See Deng et al. (in press) for more details on the distinction between these two groups. Because all our key conclusions are robust to this breakdown, we only report results for all SOEs versus non-SOEs.

serves as a proxy of government support in some of the analysis discussed below.

Finally, we follow Fan et al. (2007) in collecting information from the resumes of each firm's board members and CEO to proxy for the firm's political connections with government. The dichotomous dummy variable *PC* equals 1 if the CEO or any board member in position at the beginning of the year meets either of the following conditions: he/she is or was a government bureaucrat on or above the county level; he/she is or was a parliament member (*ren da dai biao* or *zheng xie wei yuan*) on or above the prefectural level.

3. Empirical implementation and results on collateral channel effects

Collateral channel effects typically are estimated with a reduced form investment specification as given by Eq. (1), with the underlying model and assumptions from which it is derived well described in the on-line theory appendix to Chaney et al. (2012).[8]

$$RATIO\_INV_{i,c,t} = \alpha + \beta * RATIO\_REV_{i,t} + \gamma * LP_{c,t} + \delta_t + \eta_i + \varphi * OtherControls_{i,t} + \epsilon_{i,t} \quad (1)$$

In this equation, i indexes the firms, c denotes the city of their headquarters, and t reflects the year of the observation. The variables *RATIO_INV* and *RATIO_REV* are as defined above, *LP* is the land price index in the city where the firm is located, δ and η capture year and firm fixed effects, respectively, and *OtherControls* includes standard firm measures of leverage (total debt/asset value), firm value (market-to-book value) and cash flow (EBITDA more specifically) used in these types of regressions. The investment and collateral value measures are scaled to control for firm size differences.[9] Time and firm fixed effects are included so that identification effectively comes from variation over time within firm. One of the two real estate variables is the log of the land price index (*LP*) in the city where the firm is headquartered. This variable is intended to control for broader real estate market changes that could influence investment behavior independent of the value of an individual firm's collateral.

The coefficient of interest is β which captures how changes in the value of a firm's real estate collateral (*RATIO_REV*) affect investment (*RATIO_INV*). Theory does not allow us to sign it. Controlling for firm value (which is done via the market-to-book variable discussed above), investment and collateral value are negatively correlated for unconstrained firms and positively correlated for constrained firms.[10] Overall, the estimate of β using a panel of firms reflects the combination of how many of them are financially constrained, how binding those constraints are, and how easy (or necessary) it is to pledge the underlying collateral to increase debt capacity.

While it is relatively straightforward to generate a specification such as Eq. (1) from a model of investment with financial constraints, it is more challenging to obtain unbiased estimates of β. The typical worry is that OLS yields upwardly biased results on the collateral channel effect. The productivity of a firm is inherently unobservable, and if it is correlated with initial collateral value, the estimate of β will be biased. Reverse causality is perhaps the most obvious problem if property prices and productivity are correlated. Consider the case of a firm that is large enough in its market to affect factor prices, including local land values. Similar effects could occur in markets in which multiple firms from the same industry co-locate. In that case, common shifts in investment patterns not actually driven by collateral value could be captured in the estimate of β from Eq. (1) if the firms' investment behavior bids up local land values. Upwardly biased estimates of β also would result if large land-holding firms are especially sensitive to local demand shocks (for whatever reason) and our real estate variables proxy for local demand to any significant extent (which seems likely).

The recent literature on collateral channel effects on U.S. corporate investment exploits differences in local market supply elasticity to deal with this bias (Chaney et al., 2012; Cvijanovic, 2014). The basic strategy is to instrument for real estate values using the interaction of a demand shifter (e.g., mortgage rates) with the local supply elasticity, along with city and time fixed effects. The underlying logic is as follows. Demand shifters should show up in higher prices the more inelastic is local supply. If supply were perfectly elastic, prices should not change at all. The measure of supply elasticity used (typically from Saiz, 2010) is based on fixed geographic factors such as the amount of water and the slopes of land plots in the market area, so it provides plausibly exogenous variation in real estate values due to changes in demand. Using this type of instrumental variables estimator, Chaney et al. (2012) and Cvijanovic (2014) recently report economically large collateral channel effects on investment among U.S. corporations.

A strong instrument does not exist in the Chinese data, so we report OLS estimates of Eq. (1). Given that the most likely case is for β to be biased upward, finding an insignificantly small or negative coefficient (without too large a standard error) strongly suggests that there is no economically meaningful collateral channel effect in operation. The first three columns of Table 8 report our baseline findings. The precise specification estimated is very similar to Eq. (1), except that it also includes interaction terms of the initial firm controls with local land prices.[11]

The first column reports estimates using the full sample of firms, regardless of type. The estimate of β on our measure of the real estate collateral variable, *RATIO_REV*1, is a very small and statistically insignificant 0.0045.[12] This average could be masking important heterogeneity across types of firms, as state-owned enterprises, which constitute the bulk of our firm sample well could be unconstrained. If so, they would not be expected to exhibit any

[8] It is entitled "A Simple Model of Real Estate Prices and Investment" and is available at www.princeton.edu/dsraer/theoryRE.pdf.

[9] Note that we use asset value in the denominator rather than the more typical measure in the literature of overall property, plant and equipment (which are called "fixed assets" in China). This is due to the nature of the available Chinese balance sheet data. As depicted earlier in Table 4, part of a firm's real estate holdings are not included in the item "Fixed Assets" on its balance sheet. In particular, the 2007 adjustment of accounting codes separated the leased-out properties from "Fixed Assets", and put them as part of a new, independent item called "Investment Properties" on the balance sheet. This makes the fixed assets series inconsistent over our sample period. Hence, we scale by total assets. The 2007 adjustment did not apply to the cash flows, so it does not affect our measure of *RATIO_INV*.

[10] The reasons, which are discussed more fully in the proof of Proposition 1.2 in the on-line theory appendix to Chaney et al. (2012) referenced above, are as follows. If two unconstrained firms have identical market values, but the first has higher collateral value, then it must also have lower productivity and investment than the second firm because the greater collateral value raises liquidation value. Thus, productivity and investment are lower in the first firm to compensate. Next, consider a completely constrained firm. Its investment is independent of its productivity because it is determined by a binding budget constraint set by collateral value (by assumption). However, this constrained firm's productivity must be lower to hold firm value constant, even though this does not affect its investment program. Hence, there still is a positive correlation between such value and investment for this type of firm, even when firm value is controlled for in the regression.

[11] This helps control for another source of potential upward bias. As discussed in Chaney et al. (2012), upward bias in β might also result from potential endogeneity arising from the decision to own real estate in the first place. If firms that are more likely to own real estate also are especially sensitive to local demand shocks, Eq. (1) will overestimate the collateral channel effect. Our inclusion of the firm traits and their interaction with local land prices helps control for any fixed firm-level correlation between investment and real estate values. We have no good instrument to deal with variation that may not be fixed, but this is not costly for us, as we do not find a meaningful collateral channel effect in any event.

[12] The standard error about this estimate is small enough that the upper bound impact presuming a standard deviation higher estimate remains economically small.

Table 8
Do Chinese firms invest and borrow more when collateral value increases?

	Dependent variable: *RATIO_INV*			Dependent variable: *RATIO_LOAN*		
	Full sample	SOE's	Non-SOE's	Full sample	SOE's	Non-SOE's
$RATIO\_REV1_{i,t}$	0.0045	0.0062	−0.0003	−0.0000	−0.0052	0.0190
	(0.0061)	(0.0072)	(0.0123)	(0.0104)	(0.0116)	(0.0203)
$\text{Log}(LP_{it})$	−0.0148	−0.0154	0.0217	0.0110	0.0141	0.0451
	(0.0110)	(0.0122)	(0.0245)	(0.0162)	(0.0172)	(0.0460)
$MBR_{i,t}$	0.0032^{***}	0.0030^{**}	0.0050^{**}	0.0002	−0.0011	0.0069
	(0.0011)	(0.0013)	(0.0023)	(0.0020)	(0.0025)	$(0.0037)^{*}$
$RATIO\_EBITDA_{i,t}$	0.1743^{***}	0.1761^{***}	0.1776^{***}	0.0329	0.0306	0.0185
	(0.0186)	(0.0237)	(0.0375)	(0.0299)	(0.0376)	(0.0571)
$LEVERAGE_{i,t}$	-0.0624^{***}	-0.0662^{***}	-0.0629^{***}	−0.1467	−0.1497	−0.1733
	(0.0108)	(0.0123)	(0.0192)	$(0.0175)^{***}$	$(0.0196)^{***}$	$(0.0356)^{***}$
Initial controls $*$ $\text{Log}(LP_{i,t})$	Yes	Yes	Yes	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Number of observations	3974	2962	1012	3974	2963	1011
R2	0.473	0.488	0.499	0.232	0.248	0.284

Note: (1) the observations are clustered by city-year.
(2) Standard errors in parentheses.
* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

collateral channel effect on their investment behavior. Hence, columns 2 and 3 of Table 8 report estimates of the same specification on subsamples of SOEs and non-SOEs. The results are not meaningfully different, and we cannot reliably discriminate between the coefficients across the two types of firms.[13]

This conclusion about the absence of a collateral channel effect among listed firms in China is robust to many alternative specifications not reported here for space reasons, but available upon request. For example, it could be that our desire to minimize upward bias due to endogeneity by measuring real estate exposure with the quantity of firm holdings at the end of 2002 ends up generating attenuation bias in our estimate of β because of measurement error. However, we obtain very similar results if we use *RATIO_REV2* instead of *RATIO_REV1* as our measure of real estate collateral.

We also investigated whether there were differences in the relationship between firm investment behavior and real estate collateral value depending upon whether the property market was improving or declining. Results from a specification adding an interaction term of *RATIO_REV1* with a dummy variable that takes a value of one if the relevant underlying land market was declining in value found no evidence of any important asymmetry in impacts on investment behavior. Nor did including this added term affect the coefficients on the other right-hand side regressors in any material way.

Another robustness check performed arose out of a concern that unobserved firm-level default risk could be biasing down our estimate of β. This could occur if the firms with the largest real estate holdings were also perceived by lenders as being very risky so that they could not borrow to finance additional investment even in the face of rising property values. To investigate this, we began by estimating a corporate default risk instrument at the company level based on a proportional hazard modeling framework (Deng, 1997; Deng et al., 2000). We then included the inverse Mills ratio, or the ratio of the probability density function to the cumulative distribution function of the corporate default distribution, from that corporate default model specification as a proxy to control for unobserved heterogeneity in default risk in our investment equation. Adding this corporate default risk factor to our baseline model yielded virtually no change in the estimated impact for non-SOEs. That for SOEs did increase a bit, but it still remains immaterial in economic and statistical terms. Thus, the absence of a collateral channel effect in China does not appear to be due to some type of specification bias arising from an omitted firm trait such as default risk.

We also investigated whether the small and insignificant estimates of β might be due to a lack of investment opportunities for some firms. It is not. First, there is no evidence that amassing more real estate at the beginning of the sample period is negatively correlated with growth in EBITDA or asset accumulation in general during the following years. We also reestimated our baseline model on subsets of firms broken down by whether they were headquartered in high, average or low growth cities. These classifications were based on local market GDP growth rates computed by the Chinese government. The results for each group were statistically and economically indistinguishable from those for the combined sample reported in Table 8. It turns out that the average local GDP growth rate in the group of lowest growth markets is quite high (at about 11.7% annually), so there are plenty of investment opportunities in those places, too. We also estimated the baseline model on subsets of firms based on their own asset growth rates. The point estimates for β were larger for both SOEs and non-SOEs among the highest third in terms of asset growth, but they were never statistically significantly different from zero themselves or from the point estimates for the lowest third of firms in terms of annual asset growth.

Another possible explanation for the absence of any estimated collateral channel effect could be that lenders recognize the strong mean reversion in Chinese land price growth suggested above in Table 2 and only respond to longer-run, not annual, changes. However, our estimates using 2- and 3-year moving averages for both the land price and investment variables never yield a statistically significant positive relationship either. In addition, in cities with sufficient land sales each year, we experimented with ARMA models based on their land price series and used them to predict the land price change the following year. The results were not

[13] We also experimented with two instrumental variables. One used the housing supply elasticity estimates from Wang et al. (2012); the other used a coastal dummy and/or region dummies to instrument for local land prices. Both yielded slightly smaller (including barely negative) estimates of β, which is consistent with the discussion above the OLS likely yields an upwardly-biased result. However, these results are not statistically different from those reported in the first three columns of Table 8. Moreover, the first-stage showed neither to be a strong instrument according to standard metrics. That, plus the fact that the one factor we were concerned might bias down our estimates proved not to be a problem (see just below in the text for more on that), reinforced our preference to report OLS estimates.

meaningfully changed by using that imputed land price change to calculate the market value of the change in a local firm's property portfolio. We also experimented using the official housing price indicator to calculate the collateral value change. Again, no statistically or economically significant collateral channel effect is detected, so this null collateral channel effect is not due to the use of our new land price data in lieu of the government series.

The final three columns of Table 8 provide additional evidence consistent with there being no collateral channel effect in China. Those regression results, which substitute the net change in firm debt scaled by firm asset value at the beginning of each year as the dependent variable, document that there also is no meaningful empirical correlation between changes in firm debt and changes in real estate collateral value. The collateral channel works through borrowing, so if we saw firm borrowing responding to collateral value even if investment did not, the case for no collateral channel effect would not be as strong. These results show no correlation with firm debt, not just firm investment.

4. The nature of Chinese financial markets

That there is no collateral channel effect operating for SOEs is readily explainable in terms of their not being financially constrained. Indeed, our findings support the claims by many that SOEs are specially favored within the Chinese economy (e.g., see Lin and Tan, 1999; Allen et al., 2005; Poncet et al., 2010; Deng et al., in press), and have no need to rely on increasing collateral value to secure financing. However, that is not credible for non-SOEs which appear to be financially constrained by any reasonable metric as suggested by Allen et al. (2005) and Ayyagari et al. (2010).

This raises the question of whether there is something special about the nature of the Chinese financial system that can explain the absence of a collateral channel effect even among credit constrained firms. The economic theory referenced above tells us that if complete contracting is possible, then none of the frictions that lead to a collateral channel effect exist. This could result if default were prohibitively expensive. In that case, a borrower could credibly commit to repay debt. The question is whether such a situation seems remotely possible in China, and then whether one could test for it.

China is characterized by a single party government which dominates the financial system, a judiciary that is not completely independent, and a legal system generally not well developed enough to be able to protect well-prescribed borrower rights in the event of default. In that situation, a major SOE lender has the potential to impose large costs on defaulting borrowers outside of any pledged property collateral, possibly by 'blackballing' the borrower with other important government-connected lenders or by utilizing other government linkages to have sanctions imposed outside of the specific debt contract.

We do not observe the individual lenders on a given borrower's projects. However, we have collected information on the market shares of the four largest SOE lenders in each province through 2009.[14] As noted above, those firms are the Industrial and Commercial Bank of China (ICBC), China Construction Bank (CCB), Agricultural Bank of China (ABC), and Bank of China (BOC). We create a measure of the degree of concentration of these top four SOE lenders.

Table 9 reports results including an interaction term of the share of non-top 4 SOE lenders (which equals one minus the share of big 4 SOE lenders) in each market with our standard collateral value measure ($RATIO\_REV1_{i,t} * FMC_{i,t}$ in the second row of Table 9).

Table 9
Do Chinese firms invest more when collateral value increases in markets less dominated by the four largest SOE lenders?

	Dependent variable: *RATIO_INV*		
	Full sample	SOE's	Non-SOE's
$RATIO\_REV1_{i,t}$	−0.0018	0.0068	−0.0245
	(0.0121)	(0.0167)	(0.0216)
$RATIO\_REV1_{i,t} * FMC_{i,t}$	0.0045	−0.0043	0.0377
	(0.0078)	(0.0115)	(0.0169)**
$\text{Log}(LP_{it})$	−0.0070	−0.0110	0.0725
	(0.0144)	(0.0156)	(0.0372)*
$MBR_{i,t}$	0.0049	0.0040	0.0072
	(0.0016)***	(0.0018)**	(0.0030)**
$RATIO\_EBITDA_{i,t}$	0.1764	0.1809	0.1961
	(0.0225)***	(0.0278)***	(0.0495)***
$LEVERAGE_{i,t}$	−0.0884	−0.0885	−0.0908
	(0.0136)***	(0.0161)***	(0.0262)***
Initial controls $* \text{Log}(LP_{i,t})$	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes
Number of observations	3086	2327	759
R2	0.511	0.528	0.538

Note: (1) the observations are clustered by city-year.
(2) Standard errors in parentheses.
(3) Data are for 2003–2009 only.
* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

Not surprisingly, this has no impact on the investment behavior of SOE borrowers (column 2). However, this is not the case for non-SOE borrowers, as reported in the third column of Table 9. The coefficient from the first row confirms the finding from Table 8 that, on average, there is no statistically significant collateral channel effect for this group of constrained firms. However, the interaction term reported in the second row is statistically significant and indicates the presence of a collateral channel effect in markets where the 'big 4' share is smaller. We do not believe much meaning can be attached to the magnitude of this coefficient, as it almost certainly is biased down because we are using an imperfect (noisy) proxy of the degree to which 'contract completeness' is relaxed in each market. That we are able to find any evidence of a collateral channel effect in these data suggests that non-SOEs are indeed financially constrained, but able to credibly commit to repay loans in places where the lender is more likely to be a major state-owned bank.

The robustness of that conclusion is confirmed by Table 10's findings. Using data from an annual national survey of firm managers conducted by the China Society of Economic Reform, a think tank associated with the central government, those specifications use a proxy for the ability of the underlying market's legal system to protect borrower's rights (Fan et al., 2011). The higher the value of the variable, the greater the degree of legal protection according to the managers surveyed, and thus, the less able are borrowers to credibly commit to repay lenders. If that hypothesis is correct, this interaction term ($RATIO\_REV1_{i,t} * LAW_{i,t}$) also should be significantly positively correlated with investment, indicating the presence of a collateral channel effect among constrained non-SOEs in those markets. That is precisely what the results in column 3 show, providing further evidence of a collateral channel effect among constrained non-SOEs in conditions where the costs of default are not likely to be prohibitively high.[15]

[14] These data, which were collected from "Yearbook of Finance, China", are not reported after 2009.

[15] The variation in real estate collateral values imputed from changes in our land price index is critical to finding these two statistically significant effects. If we use the variation in the two government-provided housing series to create alternative versions of *RATIO_REV1*, we never find any evidence of heterogeneity in collateral channel effects by the degree of 'big four' SOE lender concentration or perceived independence of the local legal system.

Table 10
Do Chinese firms invest more when collateral value increases in markets with more transparent local legal systems.

	Dependent variable: *RATIO_INV*		
	Full sample	SOE's	Non-SOE's
$RATIO\_REV1_{i,t}$	0.0007	0.0076	−0.0202
	(0.0116)	(0.0157)	(0.0207)
$RATIO\_REV1_{i,t} * LAW_{i,t}$	0.0003	−0.0087	0.0392
	(0.0075)	(0.0107)	(0.0157)**
$\text{Log}(LP_{it})$	−0.0078	−0.0120	0.0736
	(0.0144)	(0.0152)	(0.0371)**
$MBR_{i,t}$	0.0049	0.0040	0.0073
	(0.0016)***	(0.0018)**	(0.0029)**
$RATIO\_EBITDA_{i,t}$	0.1764	0.1810	0.1987
	(0.0226)***	(0.0278)***	(0.0496)***
$LEVERAGE_{i,t}$	−0.0884	−0.0883	−0.0933
	(0.0136)***	(0.0161)***	(0.0262)***
Initial controls $* \text{Log}(LP_{i,t})$	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes
Firm fixed effects	Yes	Yes	Yes
Number of observations	3086	2327	759
R2	0.511	0.528	0.539

Note: (1) the observations are clustered by city-year.
(2) Standard errors in parentheses.
(3) Data are for 2003–2009 only.
* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

Table 11
Does the collateral channel effect vary by degree of management's political connections & government favor.

	Dependent variable: *RATIO_INV*		
	Full sample	SOE's	Non-SOE's
A. Heterogeneity by degree of political connection			
$RATIO\_REV1_{i,t}$	0.0091	0.0094	0.0075
	(0.0064)	(0.0073)	(0.0143)
$RATIO\_REV1_{i,t} * PC_{i,t}$	−0.0118	−0.0089	−0.0165
	(0.0099)	(0.0115)	(0.0171)
B. Heterogeneity by amount of government subsidy			
$RATIO\_REV1_{i,t}$	−0.0077	−0.0025	−0.0216
	(0.0080)	(0.0097)	(0.0188)
$RATIO\_REV1_{i,t} * RATIO\_SUBSIDY_{i,t}$	0.8347	−0.2140	2.1283
	(0.9947)	(0.9812)	(2.2173)

Note: (1) all models are estimated with additional control variables (see Table 8 for the full specification).
(2) The observations are clustered by city-year.
(3) Standard errors in parentheses.
* $p < 0.1$.
** $p < 0.05$.
*** $p < 0.01$.

Tables 9 and 10 provide important new evidence consistent with the absence of frictions in credit markets being able to account for why there are no signs of collateral channel effects on average in China, but cannot rule out all other potential explanations. Just as our very small average collateral channel effect estimate for non-SOEs (column 3 of Table 8) masked important variation across markets in degree to which a borrower could credibly commit to repay debt, perhaps something similar is occurring with respect to discrimination against certain types of non-SOEs based on their own political connections, whether they are in an industry favored by the central government or operating in a city with a binding loan quota set by the central government.

To further investigate these other potential mechanisms, we turn first to the information on firm political connections as reflected in whether any board member or CEO of the 444 firms in position at the beginning of each year were former government officials or were presently serving as a member of parliament. In this case, we interact our measure of political connections with our collateral channel variable ($RATIO\_REV1_{i,t} * PC_{i,t}$) and add that term to the baseline specification reported in Table 8. The top panel of Table 11 reports the results for the two collateral channel effect terms from this specification. Note that neither the average effect nor the interaction term is large or statistically significant for non-SOEs (column 3), so we can rule out the possibility that our zero collateral channel effect for this group of firms from Table 8 is masking important heterogeneity associated with the firm management's own political connectedness.

The next panel in Table 11 shows that the same conclusion holds with respect to whether the firm operates in an industry favored by the government. For this estimation, we use the data described above that measures the total amount of subsidies received by the firm (scaled by firm asset) and interacted that with our collateral channel measure ($RATIO\_REV1_{i,t} * RATIO\_SUBSIDY_{i,t}$). Once again, we do not find a statistically significant relationship for any group of firms.

We also experimented with specifications that included the interaction of the increase in the total loan balance (as the proxy of loan quota) for the city in which the firm is headquartered with *RATIO_REV1*. As before, we find no evidence of a collateral channel effect for non-SOEs (or for other firms).

In sum, the only evidence consistent with the existence of any type of collateral channel effect is when firms have borrowed in markets not dominated by the four largest state-owned lenders or in markets perceived to have the strongest legal protections in China. We do not see any such heterogeneity in collateral channel effects associated with variation in the degree of firm management's political connections, in the degree to which the firm is favored by government as reflected in its subsidy receipts, or by whether loan quotas have been changed. This pattern of results is consistent with the nature of Chinese credit markets being such that the typical frictions associated with an inability to credibly commit to repay debt in developed markets especially are absent.[16]

5. Conclusions

The dramatic growth of Chinese property markets has been critical component of that country's extraordinary economic rise. Because housing markets go down, not just up, it is important to ask whether we should expect to see an economically important collateral channel effect akin to what other research has found for the United States and Japan. Bringing new data to bear on this question allows us to provide a first answer. That is 'no'. An important reason is that important Chinese firms such as state-owned enterprises are not financially constrained and thus have no need to pledge collateral to fund their desired investment programs. However, we do not find meaningful collateral channel effects for constrained private firms. The nature of the Chinese financial markets appears to account for this.

We caution that this does not mean a housing bust would have no seriously deleterious consequence for the Chinese economy. There are strong reasons to believe it would (e.g., through an employment channel as construction falls and via spillovers

[16] We also investigated differences between SOEs and non-SOEs by changes in wage expenditures (*RATIO_SALARY*), change in holding of cash (*RATIO_CASH*), investment on financial assets such as stocks and bonds (*RATIO_FAINV*), dividend payment policy (*RATIO_DIVIDEND*), and employment policy (*EMP*). There were some modest differences, but they are best suited to report in future research on differences between SOEs and purely private firms in China.

to many raw and processed materials industries because housing is a large demander of their products), just not from a standard collateral channel effect that amplified the investment cycle.

Finally, our research uses data on listed firms only. Future work should investigate whether the findings generalize. At present, it is not possible to replicate our analysis on non-listed firms. Information on them is completely unaudited and what is available does not include data on their real estate holdings. A more likely exception would be among local government-sponsored enterprises (local SOEs) charged with developing urban infrastructure. These entities typically are capitalized with land grants from a local government. That land, which essentially serves as the entity's equity capital, can be used to help raise debt from banks to complete the financing of infrastructure. Unfortunately, these entities are not publicly traded, so there is no comparable firm-level information available with which to replicate the type of empirical work reported above. It may be possible to aggregate data to the city level, but we leave that potentially interesting exercise to future work.

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