Chapter 1: Measuring bank competition under binding interest rate regulation: The case of China.

Co-authored with Adrian van Rixtel and Michiel van Leuvensteijn. Shorter version published at *Applied Economics*, vol 48 (49), p.4699-4718, DOI: 10.1080/00036846.2016.1164818 (April, 2016). Earlier versions of this chapter are published as *Bank for International Settlements Working Paper* (No.422/2013) and *Banco de España Working Paper* (No.1404/2014).

Many empirical studies suggest that financial reform promoted bank competition in most mature and emerging economies. However, some earlier studies that adopted conventional approaches to measure competition have concluded that bank competition in China declined during the past decade, despite progressive reforms implemented since the 1980s. In this chapter, we show theoretically and empirically that this apparent contradiction is the result of flawed measurement. Conventional indicators such as the Lerner index and Panzar-Rosse H-statistic fail to measure competition in Chinese loan markets properly due to the system of interest rate regulation. By contrast, the Profit Elasticity (PE) approach that was introduced in Boone (2008) as Relative Profit Differences (RPD) does not suffer from these shortcomings. Using balance sheet information for a large sample of banks operating in China during 1996-2008, we show that competition actually increased in the past decade when the PE indicator is used.

Chapter 2: Do banks extract informational rents through collateral?

Co-authored with Adrian van Rixtel and Honglin Wang. Published as *Bank for International Settlements Working Paper* (No.522/2015), *Hong Kong Institute for Monetary Research Working Paper* (No.01/2016), *Bank of Finland (BOFIT) Discussion Paper* (No.5/2016), and *Banco de España Working Paper* (No.1616/2016). Selected as *Featured Article* by *Banco de España*, 2016.

This chapter investigates if informational monopolies resulting from relationship lending and bank market concentration allow for rent extraction through collateral. Our identification strategy hinges on the notion that informational equalization shocks (such as equity IPOs) erode rent seeking opportunities, while competing theories do not rely on information asymmetries among lenders. Using a unique hand-collected database of 9,288 bank loans obtained by 649 listed Chinese firms, we find that collateral incidence is positively associated with relationship intensity and bank market concentration, while this effect is moderated for post-IPO loans. We also demonstrate important cross-sectional variation among borrowing firms: after IPO, rent extraction through collateral is moderated for less risky firms, but intensified for risky firms. These findings are not driven by alternative theories including: shifts in firm risk around IPO; heterogeneous dynamics of risk shifting around IPO; and concurrent lending and corporate bond underwriting. We further demonstrate that our results are not sensitive to: endogeneity of IPO or relationship lending; unobserved time-invariant firm risks; alternative samples; and the endogeneity of loan contract terms. Our study complements the findings in other studies that banks extract rents by charging higher lending rates from their informational monopolies (Hale and Santos, 2009; Schenone, 2009). Furthermore, we provide the first loan-level evidence on the determinants of collateral in Chinese bank lending markets.

Chapter 3: Collateral and the disruption of firm as non-financial intermediary: Evidence from Chinese Property Law.

Forthcoming as *Bank of Finland (BOFIT) Discussion Paper* and Banco de España Working Paper.

This chapter investigates the effects of collateral law reform on access to external finance and trade credit. By allowing large classes of movable assets to be used as collateral, the Chinese Property Law reform transformed firms' role as non-financial intermediaries in China. We find after the legal reform, firms relied on trade credit financing could substitute to more bank credit. Accordingly, the providers of trade credit reduced significantly their provision of trade credit. In particular, we find the Property Law has disrupted the practice in which firms borrow short-term debts and redistribute them via trade credit. After the reform, instead of providing trade credit, firms started to accumulate more fixed asset investment, which in turn allowed for more long-term borrowing from banks. Our findings are not driven by confounding factors such as liquidity drain due to financial crisis. Our results also cannot be explained by other important reforms which were introduced around the same time. Our paper highlights the importance of looking at other financing channels when investigating the effect of collateral laws.

Measuring bank competition under binding interest rate regulation: the case of China

Bing Xu[a], Adrian van Rixtel[b,] , Michiel van Leuvensteijn[c]

[a] *Universidad Carlos III de Madrid, Calle Madrid 126, 28903 Getafe, Spain*
[b] *Associate Directorate General International Affairs, Bank of Spain, Alcalá 48, 28014 Madrid, Spain.*
[c] *Corporate Strategy&Policy, All Pensions Group (APG), Gustav Mahlerplein 3, 1082 MS Amsterdam, the Netherlands.*

Abstract

Many empirical studies suggest that financial reform promoted bank competition in most mature and emerging economies. However, some earlier studies that adopted conventional approaches to measure competition have concluded that bank competition in China declined during the past decade, despite progressive reforms implemented since the 1980s. In this chapter, we show theoretically and empirically that this apparent contradiction is the result of flawed measurement. Conventional indicators such as the Lerner index and Panzar-Rosse H-statistic fail to measure competition in Chinese loan markets properly due to the system of interest rate regulation. By contrast, the Profit Elasticity (PE) approach that was introduced in Boone (2008) as Relative Profit Differences (RPD) does not suffer from these shortcomings. Using balance sheet information for a large sample of banks operating in China during 1996-2008, we show that competition actually increased in the past decade when the PE indicator is used.

JEL classification: D4; G21; L1

Keywords: Competition; Banking industry; China; Lending markets; Regulation

1. Introduction

There is a continuing debate on which empirical approaches may be the most suitable for measuring competition in specific banking systems. This discussion has assumed growing weight in the bank competition literature, underpinning the rather unsatisfactory observation that the currently available empirical toolkit frequently yields contradictory and inconclusive results for specific countries. For example, one study concludes that "… well-known indicators of bank competition often give conflicting predictions of competitive behaviour across countries, within countries, and over time" and that the "… determination of competition may differ depending on the measure chosen to assess it" (Carbó Valverde *et al.*, 2009, p. 132).

In the light of this discussion, the Chinese banking sector offers an interesting test case. First, Chinese banking has been reshaped profoundly by financial reform. During the past 30 years, the banking landscape in China moved from a government controlled monolithic structure to a pluralistic system comprising various groups of market-oriented banks. The question arises how different measures assess the development of competitive conditions under these fundamental changes. Second, relatively few econometric analyses have concentrated explicitly on bank competition in China. These have applied conventional measures such as the Panzar-Rosse H-statistic (Yuan, 2006; Fu, 2009) or the Lerner index (Fungáčová *et al.*, 2012; Soedarmono *et al.*, 2013; Fu *et al.*, 2014). These studies concluded in general that competition in the Chinese banking sector followed a decreasing trend, despite the comprehensive process of financial reform. This is a rather counterintuitive result and in contrast to the results of a large body of research which indicates that deregulation fostered competitive conditions in many emerging market economies (Claessens, 2009).

This paper contributes to the literature on bank competition by arguing that conventional measures of competition like the Panzar-Rosse H-statistic[1] and the Lerner index may not assess bank competition in China correctly, mainly due to the existence of interest rate regulations. Instead, we argue that the Profit Elasticity (PE) indicator may be better suited to investigate competitive conditions in Chinese loan markets. The PE indicator, whose

[1] See Bikker *et al.* (2012) for further discussion on the shortcomings of Panzar-Rosse H-statistic in the empirical literature.

theoretical foundation is the concept of Relative Profit Differences (RPD), is based on the notion that competition rewards efficiency. In other words, in a more competitive market, firms are punished more harshly (in terms of profits) for being inefficient (Boone *et al*., 2007; Boone, 2008; Van Leuvensteijn *et al*., 2011 and 2013). Boone *et al*. (2007) demonstrates that the PE indicator is more robust from both a theoretical and an empirical point of view when compared with conventional measures. We show that the theoretical foundation of the PE indicator allows for a correct measurement of competition under both binding deposit and lending rate regulation, hence it is particularly suitable to assess competitive conditions in regulated markets such as Chinese loan markets (See Appendix C for theoretical proofs).

As in previous studies, our empirical results for the (elasticity-adjusted) Lerner index show declining competition, and this conclusion holds for alternative specifications. In contrast, the PE indicator shows improving competition in Chinese loan markets over time, especially after 2001, with some retreat in the final years of our sample. Moreover, we are fairly able to explain the specific pattern of the development of competition. Our results for the PE indicator are in line with the development of various indicators of financial reform. Finally, the findings for the PE indicator are robust for several alternative specifications and pass various robustness tests. All in all, we see our a priori theoretical objections to the conventional measures as appropriate gauges to assess Chinese banking competition validated by the empirical results.

The remainder of this paper is organized as follows. Section 2 gives background information on the structure of Chinese banking. Section 3 provides a brief overview of the literature on bank competition in China and sets out our main hypotheses. Section 4 presents the methodological framework of the standard and elasticity-adjusted Lerner indices and the PE indicator. Section 5 shows our data and sample characteristics. Section 6 compares the empirical results for the (elasticity-adjusted) Lerner index and the PE indicator and presents a detailed interpretation of the results from the PE indicator, including their relationship with various financial reform indicators. Section 7 concludes.

2. Chinese banking sector

China's financial system has undergone a comprehensive process of reform during the past 30 years, of which one of the main objectives was to improve competition and efficiency in the banking sector.[2] In this section, we provide insights in the main elements of Chinese banking which are relevant for our analyses. A timeline of selective reforms is summarized in Table 1.

2.1. Market structure

The development of a commercial banking system in China and the entrance of important new players were made possible by the promulgation of the Commercial Bank Law in May 1995 (Fu and Heffernan, 2009). The commercialization of the Chinese banking sector was triggered by mounting problems at the four state-owned banks which experienced a significant deterioration of their asset quality in the early 1990s and were converted into state-owned commercial banks (SOCBs). In addition, 12 so-called joint-stock commercial banks (JSCBs[3]) and more than 100 city commercial banks (CCBs[4]) were established under the Commercial Bank Law.

Hence, since the mid-1990s, three main groups of Chinese commercial banks – SOCBs, JSCBs and CCBs – have become active in Chinese loan markets. Arguably, this expansion of the number of providers of credit may have promoted competitive conditions in these markets. In fact, as is shown in Table 2, the market share of the SOCBs has declined significantly. While their average annual market shares of total assets and loans during 1996-2001 were 86% and 88%, respectively, these shares dropped to 72% and 71%, respectively, during 2002-2008.[5] These declines in market shares have been mirrored in considerable increases in those of especially the JSCBs, and of the CCBs as well.

---

[2] According to the CBRC, financial reform in China can be classified in three major stages (1978-1993, 1994-2002, 2003-present); see Liu (2009). Clear overviews in English are presented in Allen *et al*. (2005) and Matthews and Zhang (2010).

[3] The JSCBs initially offered banking services only regionally, but later they were allowed to operate freely nationwide, competing with the SOCBs for large firms and with the CCBs for small and medium-sized enterprises.

[4] CCBs offer commercial banking services to small and medium-sized enterprises and households in the main cities or in certain provinces, but have been expanding to larger companies that would normally do business with the SOCBs and JSCBs. The requirement for CCBs to operate only within the cities' own administrative districts was lifted from 2007 onwards, allowing them to compete in larger geographical areas.

[5] We present the data for the full sample of 1996-2008 and the two subsamples that we use (1996-2001 and 2002-2008), with 2001 being the year of China's entry into the WTO.

Competition may also have benefited from the growing role of foreign banks. An important catalyst here was China's accession to the WTO in 2001 (see Table 1). Under the conditions of WTO membership, the activities of foreign banks were liberalized profoundly. The foreign liberalization of Chinese banking resulted in a sharp increase in the number of foreign players. Table 2 shows that the market shares of foreign banks in total assets and loans increased during 1996-2008, but remained below 1%.

2.2. Deregulation of credit controls and interest rates

An important reform affecting Chinese loan markets was the replacement of the People's Bank of China's (PBC) binding credit plan system with an indicative non-binding credit target, effective from 1 January 1998, with this target serving only as a reference for commercial banks (see Table 1). Until then, the PBC had controlled the lending of SOCBs through binding credit quotas, which set the lower limit for new loans to be made annually and stipulated their allocation to specific sectors (Wong and Wong, 2001). Hence, since 1998, in principle Chinese banks have become free to lend according to commercial considerations, with the formal abolishment of policy loans that were provided in compliance with state directives or planning targets instead of on the basis of proper credit assessments. This change in policy has been hailed by Chinese monetary authorities as an important step in transforming the credit culture of Chinese banks.

Notwithstanding the significance of the abolishment of the credit plan system, there are clear signs of continuing quantitative controls on bank credit, which potentially may affect the lending policies of banks in China. Various observers emphasize the use by the PBC of quantitative instruments aimed at controlling credit growth, despite the discontinuation of the binding credit plan system (Fukumoto *et al*., 2010; Huang and Wang, 2011; He and Wang, 2012; Ma *et al*., 2011). These include yearly aggregate target levels for new loans and the use of so-called window guidance to influence the development of bank lending. The latter policy can be described as a form of moral suasion aimed at controlling the sectoral direction of lending, although in practice this guidance also may have affected the amount of lending (Okazaki, 2007).[6]

The reform of the credit control system has been followed – in terms of degree of deregulation – by interest rate liberalization, resulting in relatively liberalized bank interest rates in 2004, when the deposit rate floor and the lending rate ceiling were eliminated for the major banks.[7] At the same time, the PBC maintained its control of the deposit rate ceiling and lending rate floor, although in practice, the latter was probably not binding. In fact, during December 2004 and December 2008, only between 19% and 29% of all loans were made at the floor lending rate, suggesting that most loan rates were higher. In contrast, empirical research has suggested that the ceiling on deposit rates has been binding, which put them at levels below equilibrium (Feyzioğlu *et al*., 2009; He and Wang, 2012; Ma *et al*., 2011).

3. Empirical literature on bank competition in China

The theoretical literature on the measurement of competition is generally categorized into two major streams, namely the "Structure Conduct Performance" (SCP) approach and the non-structural approaches promoted within the so-called New Empirical Industrial Organization (NEIO) literature. The former approach includes concentration indicators as proxies for competition, such as the "Hirschman-Herfindahl Index" (HHI) and concentration ratio (CRn) that measures the market shares of the *n* largest banks. The latter approaches include the Panzar-Rosse H-statistic (Rosse and Panzar, 1977; Panzar and Rosse, 1987), the (elasticity-adjusted) Lerner index (Genesove and Mullin, 1998) and a relatively recent method known as the Profit Elasticity (PE) model or the "Boone" indicator (Boone, 2008).

Only a few studies have investigated explicitly and in-depth competition in the Chinese banking sector by using some of the above mentioned indicators, while others discuss it on the sidelines and often adopt more descriptive approaches. One of the first to address this issue was Wong and Wong (2001). It adopts the structural approach and

[6] Moreover, as we shall discuss in more detail in section 3 and especially in section 6, the PBC re-introduced credit quotas for individual banks in 2007 in order to curb lending activities.
[7] The PBC started to widen the floating band on banks' interest rates from 1998 onwards, after it liberalized interbank interest rates. Hence commercial banks got more discretion in setting loan rates (PBC, 2005; Feyzioğlu et al., 2009).

calculates concentration indicators (HHI) which show high degrees of concentration that may inhibit competition during the 1990s, and acknowledges that reforms stipulated under the conditions for China's WTO accession helped to create a more competitive banking system.

Turning to non-structural approaches, Yuan (2006), using the Panzar-Rosse method, investigates bank competition in China during 1996–2000, just before the country joined the WTO in 2001. The paper concludes that the banking system in China was already close to a state of perfect competition before foreign banks began to enter Chinese banking more extensively. Fu (2009) looks at competition in Chinese commercial banking, also by using the Panzar-Rosse method. Based on a sample of 76 banks for the period 1997–2006, it is concluded that China's overall banking market was perfectly competitive in 2001, but featured monopolistic competition thereafter until 2006. Thus, the paper supports the conclusion of Yuan (2006) that the Chinese banking sector was close to a state of perfect competition before China joined the WTO and that WTO membership might not promote overall bank competition further. Bikker and Spierdijk (2008), as part of an investigation of 101 countries with the Panzar-Rosse model, also measure competition in Chinese banking and have results suggesting perfect competition. However, they warn that these results should be interpreted with great caution due to the limitations of the Panzar-Rosse model for China.

A few studies apply the Lerner index to Chinese banking. Fungáčová *et al*. (2012) find that competition in the Chinese banking industry declined, based on a sample of 76 banks during 2002–2011. They also find that competition differs depending on the type of banks, with foreign banks being the most competitive. Also Soedarmono *et al*. (2013) report lower competition in Chinese banking over time, as part of an investigation of 11 Asian banking systems for 1994–2009 (for China covering 103 banks). Fu *et al*. (2014) analyze 14 Asian banking systems for the period 2003-2010 and also come to the conclusion that competition in the Chinese banking sector decreased over time despite financial deregulation. All these studies report Lerner indices for China that lie predominantly between 0.25 and 0.4 for 2002–2008. Ho (2010) uses the Lerner index to analyze deregulation and competition in the banking sector of Hong Kong.

The studies that measure competition using the non-structural approach (Panzar–Ross method and Lerner index) predominantly report that competitive conditions have worsened over time despite continuing financial deregulations in business scope, geographic expansion, foreign entrance and administrative controls in price and quantity. This is in contrast to the considerable empirical evidence that deregulation fostered competitive conditions in banking markets around the globe (Claessens, 2009). We postulate that to some extent prior result of deteriorating competition could be due to a flawed method to measure bank competition in an economy where strict price and quantitative controls are being practiced. In light of the recent theoretical advance by Boone (2008), we revisit the competitive status of Chinese banking markets using the PE indicator, which will be discussed in the next section. Given the discussion of financial reform in Section 2, we suggest that competitive conditions should have improved over the years, particularly so after China's accession to the WTO in 2001. Under the conditions of WTO membership, the activities of foreign banks were liberalized profoundly (for an overview see Table 1).[8] The foreign liberalization of Chinese banking resulted in a sharp increase in the number of foreign players, from 4 banks in 1996 to 26 banks in 2008, although their market share remained low. Despite this low share, the importance of foreign banks in promoting competitive condition should not be underestimated. Various studies suggest that both the threat of foreign entry and its actual realization forced Chinese banks to respond in terms of improving their business models and efficiency and hence the overall degree of market competition (He and Fan, 2004; Leung and Chan, 2006). Some observers claim that foreign banks have snatched significant market shares in key cities from their Chinese competitors (Xu and Lin, 2007), and have penetrated into other parts of China through equity partnerships or less institutionalized forms of cooperation with Chinese banks (He and Fan, 2004; Leung and Chan, 2006). Xu (2011) further provides empirical evidence that foreign bank entry has been supportive of developing a more competitive banking industry.

Taking all these considerations into account, we postulate our first hypothesis as follows:

[8] For example, foreign banks were allowed to provide foreign currency services to Chinese residents and were permitted greater freedom in local currency operations as well. Furthermore, the participation of foreign investors in Chinese banks was promoted, with foreigners being allowed to take equity stakes of up to 25%.

Hypothesis I: Competitive conditions improved significantly after WTO accession in 2001.

As discussed in Section 2, another significant step was taken in October 2004, with the removal of the lending rate ceiling and the deposit rate floor, while the lending rate floor was reduced to a specific range versus the benchmark rates. Although these steps were important milestones in the process of financial reform in China, we believe that a priori their impact on competitive conditions in Chinese loan markets may have been mixed.

First, it may have been the case that the removal of the lending rate ceiling out-weighted the effect of the reduction in the lending rate floor and allowed inefficient banks to increase lending rates above the previously existing binding ceiling. In fact, after the reform, the percentage of loans priced above the reference rate increased sharply, from around 50% in 2004 to 60% in 2005, suggesting that lending rate ceilings were binding before the reform.[9] This increase allowed banks potentially to expand price-cost margins, which would suggest a deterioration in competition, as the relationship between performance and efficiency would have become weaker.

Second, the removal of the deposit rate floor in October 2004 may not have improved competitive conditions significantly as well, as various studies have documented that the deposit rate ceiling was binding, and not the deposit rate floor (see section 2).

Third, we demonstrate theoretically (Appendix C) that removing the lending rate ceiling and the deposit rate floor most probably would not lead to a significant improvement in competitive conditions. In contrast, we have shown that competition most likely will improve by removing the deposit rate ceiling and lending rate floor. However, these two major aspects remained unchanged in the 2004 partial interest rate reform.

Moreover, the Chinese supervisory authorities strengthened various aspects of their regulatory policies in 2004, in addition to the interest rate reforms. An overview of these measures is presented in Table 1 and we shall discuss them in more detail in section 6. These policy actions represented a certain degree of re-regulation, which may have contributed to a decline in competition. They were followed by several other policy steps in subsequent years

[9] The PBC provides only summary data on the percentage of loans issued around the reference rates from December 2004 onwards. Therefore, we lack sufficient data to formally test how binding the lending rate ceiling was prior to that date.

– such as the re-introduction of credit quotas in 2007 – which resulted in an overall increase in financial repression in China, as measured by the financial repression index introduced in Huang and Wang (2011) (see section 6).

Overall, we postulate our second hypothesis as follows:

Hypothesis II: Competitive conditions deteriorated after 2004.

4. Methodology

In this section, we discuss the methodological frameworks of the conventional Lerner and elasticity-adjusted Lerner indices and the Profit Elasticity (PE) indicator.

4.1. Lerner and elasticity-adjusted Lerner indices

4.1.1. Lerner index

The Lerner index reflects firms' ability to set prices over marginal costs. Fierce competition will lower its level, as firms reduce prices towards marginal costs. In the extreme case of perfect competition, the Lerner index will be reduced to zero, while with monopoly it will reach one. The traditional Lerner index has been applied widely in empirical competition literature (Fernández de Guevara *et al.*, 2007; Berger *et al.*, 2009). However, to the best of our knowledge, only Fungáčová *et al.* (2012) conducted an in-depth analysis based on this measure for Chinese banking markets during the post-WTO accession period. Our approach differs in the sense that we do not focus on bank competition in general but instead concentrate on competition in loan markets. Hence, we define the Lerner index as:

$$L_{it} = \frac{p_{it} - mc_{ilt}}{p_{it}} \qquad (1)$$

where p_{it} denotes the price of a loan for bank i at time t, which is defined as total interest income divided by total loans, while mc_{ilt} are marginal costs of loans.

In order to calculate marginal costs of loans, we first estimate a Translog Cost Function (TCF) using individual bank observations (Van Leuvensteijn *et al*., 2011). This function assumes that the technology of an individual bank can be described by one multiproduct production function. Our TCF has the following form:

$$lnc_{it} = \alpha_0 + \sum_{t=1}^{T-1} \gamma_t d_t + \sum_{h=1}^{H-1} \eta_h d_h + \sum_{j=1}^{K} \delta_j lnx_{ijt} + \sum_{j=1}^{K} \sum_{k=1}^{K} \epsilon_{jk} lnx_{ijt} lnx_{ikt} + v_{it} \quad (2)$$

where the dependent variable c_{it} reflects the production costs of bank i (i = 1, .., N) in year t (t = 1, .., T). d_t are year dummies and d_h are bank type dummies (h = *SOCB, JSCB, CCB*).[10]

The explanatory variables x_{ikt} represent three groups of variables (k = 1, .., K). The first group consists of (K_1) bank output components, such as loans, securities and other services (proxied by other income). The second group consists of (K_2) input prices, such as wage rates, deposit rates (as price of funding) and the price of other expenses (proxied as the ratio of other expenses to fixed assets). The third group consists of (K - K_1 - K_2) control variables, e.g., the equity ratio. In line with Berger and Mester (1997), the equity ratio corrects for differences in loan portfolio risk across banks (Van Leuvensteijn *et al*., 2011). v_{it} is the error term.

The marginal costs of loans are obtained by differentiating the TCF (Equation 2) with respect to loans, namely:

$$mc_{ilt} = \frac{c_{it}}{x_{ilt}} \left(\delta_1 + 2\epsilon_{1l} lnx_{ilt} + \sum_{k=1 \ldots K; k \neq l} \epsilon_{1k} lnx_{ikt} \right) \quad (3)$$

Once marginal costs of loans are obtained, an individual bank's Lerner index is calculated according to Equation (1). The yearly Lerner index L_t is then the average of the

[10] In this section, we assume that cost functions for each bank type are similar, as only the constant term is allowed to vary across bank groups through bank type dummies. The alternative approach is to assume different cost functions for each bank type by allowing bank type dummies to interact with independent variables. We follow this approach in Appendix D as an additional robustness test.

individual L_{it} for each year t, and the subsample Lerner index $L_{subsample}$ is the average of the individual bank's Lerner indices for each subsample. The subsamples are the periods of pre-WTO (1996-2001), post-WTO (2002-2008), and two samples divided by the 2004 interest rate reform, i.e. pre-reform (2002-2004) and post-reform (2005-2008).

4.1.2. Elasticity-adjusted Lerner index

The traditional Lerner index cannot distinguish markets that have high margins due to inelastic demand from markets that have high margins because they are less competitive or perhaps collusive (Corts, 1999, p.31). To overcome this problem, the elasticity-adjusted Lerner index has been developed (Genesove and Mullin, 1998; Corts, 1999; Wolfram, 1999; Van Leuvensteijn, 2008). More precisely, this measure normalizes the Lerner index for the price elasticity of demand.

We estimate the elasticity-adjusted Lerner index following Angelini and Cetorelli (2003). We provide only a very brief introduction of this indicator, since it is rather standard in the literature. Bank i solves the following profit-maximizing problem:

$$\max_{q_i} \Pi = p(Q)q_i - C(q_i, w_i) \tag{4}$$

where $Q = \sum_i q_i$ is the total amount of bank loans in loan markets and q_i is the loan provided by bank i. $C(q_i, w_i)$ is the cost function of bank i, and w_i represents the vector of factor input prices. The corresponding first-order condition is:

$$p_i = C'(q_i, w_i) - \frac{\Theta_i}{\varepsilon} \tag{5}$$

where Θ_i is the conjectural elasticity of total industry output with respect to the output of bank i, and ε is the market semi-price elasticity of demand, namely $\Theta_i = \frac{dQ/dq_i}{Q/q_i}$ and

$\varepsilon = \frac{dQ/dp}{Q}$. In a perfectly competitive market, Θ_i equals zero for all banks, while in a monopoly market Θ_i equals one. The separate identification of these two elasticities requires the simultaneous estimation of a supply and demand equation (Angelini and Cetorelli, 2003).

Appelbaum (1982) suggests that it is sufficient to estimate the ratio $\lambda = \frac{\Theta_i}{\varepsilon}$ if the goal is to evaluate the industry's overall degree of market power.[11] The elasticity-adjusted Lerner index will then be defined as $L = \frac{\lambda}{p}$, where p is the average price of loans. Market power depends on both the elasticity of demand and the degree of competition, measured by conjectural variation.

To identify λ and the elasticity-adjusted Lerner index, we estimate simultaneously the Translog Cost Function (TCF) and the supply equation, imposing cross-equation restrictions. The TCF and marginal costs of loans are defined the same way as Equations (2) and (3), respectively. Substituting the marginal costs Equation (3) into the supply Equation (5), we obtain:

$$p_{it} = \frac{c_{it}}{x_{ilt}}\left(\delta_1 + 2\epsilon_{1l} lnx_{ilt} + \sum_{k=1\ldots K;k\neq l} \epsilon_{1k} lnx_{ikt}\right) + \sum_{t=1\ldots T-1} \lambda_t d_t + \varepsilon_{it} \qquad (6)$$

where d_t is a year dummy and ε_{it} is the error term.

To access the evolution of bank competition, we perform two types of regressions: yearly estimates and subsample estimates. The yearly elasticity-adjusted Lerner index is then derived as $L_t = \frac{\lambda_t}{p_t}$, where p_t is the yearly average loan rate. To obtain subsample estimates of the elasticity-adjusted Lerner index, we regress simultaneously Equations (2) and (6), replacing year dummies d_t with subsample dummies in both equations. The subsample elasticity-adjusted Lerner index is then defined as $L_{subsample} = \frac{\lambda_{subsample}}{p_{subsample}}$, where $p_{subsample}$ is the average loan rate for each subsample. The estimation is carried out with

[11] As a robustness test, we estimated in Appendix D (D.1) explicitly the conjectural variation parameter as a direct measure of competition.

three-stage least squares (3SLS). To control for endogeneity of the cost and quantity variables, we employ one-period lagged variables as instruments; therefore the results are available starting from 1997.

4.2. The Profit Elasticity (PE) model

The PE indicator, developed in a broad set of theoretical models (Boone, 2000, 2001 and 2008; Boone *et al.*, 2007; Boone and Van Leuvensteijn, 2010), is the empirical operationalization of the Relative Profit Differences (RPD) concept proposed by Boone (2008). It is based on the notion, first, that more efficient firms (that is, firms with lower marginal costs) gain higher market shares or profits and, second, that this effect is stronger the higher competition in that market is (Van Leuvensteijn *et al.*, 2011). Boone (2008) shows that there is a continuous and monotonically increasing relationship between RPD and the level of competition. This property is the main advantage of RPD over traditional measures such as the HHI and Lerner index (or price-cost margin (PCM) approaches). Another advantage is that RPD and the PE indicator are not dependent on assumptions about the type of competitive model, such as whether this is Bertrand or Cournot competition. The RPD provides a solid theoretical foundation for the PE indicator.

We first derive the theoretical concept of RPD. Following Boone (2008), and replacing “firms” with “banks”, we consider a banking industry where each bank i produces one product q_i (or portfolio of banking products), which faces a demand curve of the form,

$$p\left(q_i, q_{j\neq i}\right) = a - bq_i - d\sum_{j\neq i} q_j \tag{7}$$

and has constant marginal costs mc_i. This bank maximises profits $\pi_i = (p_i - mc_i)q_i$by choosing the optimal output level q_i. We assume that $a > mc_i$ and $0 < d \leq b$. The first-order condition for a Cournot-Nash equilibrium can then be written as:

$$a - 2bq_i - d\sum_{j\neq i} q_j - mc_i = 0 \tag{8}$$

When N banks produce positive output levels, we can solve the N first-order conditions, yielding:

$$q_i(mc_i) = \frac{(2b/d-1)a-(2b/d+N-1)mc_i+\sum_j mc_j}{\left(2b+d(N-1)\right)(2b/d-1)} \tag{9}$$

We define profits π_i as variable profits excluding entry costs ε. Hence, a bank enters the banking industry if, and only if, $\pi_i \geq \varepsilon$ in equilibrium. Note that Equation (9) provides a relationship between output and marginal costs. It follows from $\pi_i(mc_i) = (p_i - mc_i)q_i$ that profits depend on marginal costs in a quadratic way, i.e.

$$\pi_i(mc_i) = \frac{(2b/d-1)a-(2b/d+N-1)mc_i+\sum_j mc_j}{\left[\left(2b+d(N-1)\right)(2b/d-1)\right]}(p_i - mc_i) \tag{10}$$

The theoretical concept RPD is then defined as $RPD = \frac{\pi(mc^{**})-\pi(mc)}{\pi(mc^{*})-\pi(mc)}$ for any three firms with$mc^{**} < mc^{*} < mc$. In this market, competition can increase in two ways. First, competition increases when the produced services of the various banks become closer substitutes, that is, d increases (keeping d below b). Second, competition increases when entry costs ε decline. Boone (2008) proves that RPD is an increasing function of interaction among existing firms ($\frac{dRPD}{dd} > 0$) and a decreasing function of entry costs ($\frac{dRPD}{d\varepsilon} < 0$). In other words, RPD increases when competition intensifies, i.e. fiercer competition increases (decreases) profits of more efficient firms by larger (smaller) amounts than those of less efficient firms. Hence, competition rewards efficiency, a concept that can be traced back to Demsetz's (1973) efficiency structure hypothesis.

Boone (2008) demonstrates how RPD can measure the level and evolution of competition in practice. Firms are first ranked by their efficiency level. Subsequently, RPD of firm *i* are normalized by calculating its relative profit difference against the profits of the most and the least efficient firms. This procedure yields a normalized RPD curve as a function of normalized relative efficiency differences. The level of competition is then represented by the area under the normalized RPD curve. Since changes in competition move all points on the RPD curve monotonically, shifts in this curve measure the evolution in competition.

Although this procedure is mathematically elegant, it is computationally intensive, as it requires the ranking of firms by efficiency levels (i.e. marginal costs) for each year. Conversely, most empirical studies that adopt Boone's work regress the logarithm of profits on the logarithm of marginal costs to capture the essence of RPD. They refer to the estimated elasticity of profits with regard to marginal costs, i.e. $\frac{dln(\pi)}{dln(mc)}$, as the PE indicator or Boone indicator (Boone *et al*., 2007; Van Leuvensteijn *et al*., 2011 and 2013; Schaeck and Cihák, 2010; Delis, 2012; Tabak *et al*., 2012). This indicator is in theory negative, reflecting the fact that higher marginal costs are associated with lower profits. In addition, its value should be lower the more competitive market conditions are. The PE indicator is based on the same theoretical foundation as RPD, as they both capture the central idea that less efficient firms are punished more in more competitive markets. Boone *et al*. (2007) conduct simulations for the PE indicator and find that changes in competition are correctly identified with this measure. Unlike the computationally intensive RPD, the PE indicator has the advantage that it can be easily estimated in practice and has a rather straightforward interpretation. We therefore employ the PE indicator to measure competition in Section 6.

We note that the PE indicator model, like every other model, is a simplification of reality (Van Leuvensteijn *et al*., 2011). First, efficient banks may choose to translate lower costs either into higher profits or into lower output prices in order to gain market share. Our approach assumes that banks in China compete on efficiency in order to predominantly increase profits and not to expand market share, given quantitative restrictions in the form of explicit lending quotas and informal window guidance (see section 2). Even when some banks would choose to increase profits by lowering their price and increasing their market share, the PE indicator would also measure this effect. Still, we assume that this behavior does not diverge too strongly across banks. Second, the PE indicator model ignores differences in bank product quality and design, as well as the attractiveness of innovations. We assume that banks are forced over time to provide quality levels that are more or less similar. By the same token, we presume that banks have to follow the innovations of their peers. Hence, like many other model-based measures, the PE indicator focuses on one important relationship (that between efficiency and profits), thereby disregarding other aspects. All in all, the PE indicator may be applied in relatively homogeneous product markets where product innovation and differences in quality do not matter too much. Therefore, we focus only on competition in loan markets and not on overall bank competition in China.

Taking into account that we believe that the PE indicator is the most suitable measure to assess competitive conditions in Chinese lending markets, we reformulate our two hypotheses that we presented in section 3 as follows:

HI: Competitive conditions improved significantly after WTO accession in 2001, with:

H_0: $\beta_{pre\text{-}WTO} > \beta_{post\text{-}WTO}$ and H_1: $\beta_{pre\text{-}WTO} <= \beta_{post\text{-}WTO}$ (recall β is negative).

HII: Competitive conditions deteriorated after 2004, with:

H_0: $\beta_{pre\text{-}reform} < \beta_{post\text{-}reform}$ and H_1: $\beta_{pre\text{-}reform} >= \beta_{post\text{-}reform}$.

5. Data

The main data source of our analysis is BankScope. We collect Chinese banks' balance sheet data running from 1996 to 2008. This 13-year period is selected to capture various banking sector reforms, including those related to WTO accession, and to facilitate comparison of our results with those of other papers using the Lerner index. Whenever BankScope does not provide sufficient information, we use various issues of the Almanac of China's Finance and Banking, China Statistical Yearbook and individual banks' annual reports to double-check and fill in missing data.

We focus on state-owned commercial banks (SOCBs), joint-stock commercial banks (JSCBs), city commercial banks (CCBs) and foreign banks (FOREIGN).[12] To exclude irrelevant and unreliable observations, banks are incorporated in our sample only if they fulfill the condition that total assets, loans, deposits, equity and other non-interest income should be positive. We lose 43 observations after applying this criterion, mainly due to negative non-interest income. At the end, we are left with 714 observations covering 1996-2008. Our sample includes extensive information on 127 banks, including all four SOCBs, all 13 JSCBs, 28 foreign banks[13] and 82 CCBs. Table 8 in Appendix A summarizes the distribution of the observations. Table 3 describes the variables used in the translog cost estimations, such as costs, loans, securities and other services, each expressed as a share of total assets, income or funding. Costs are defined as the sum of interest expenses, personnel expenses and other expenses.

6. Results[14]

6.1. Empirical results: Lerner and elasticity-adjusted Lerner indices.

We summarize the results for the traditional Lerner index and elasticity-adjusted Lerner index in Table 4. Panel A reports yearly estimates. They suggest a general increasing level of bank competition up to around 2002 and a decreasing trend afterwards, with the lowest level of competition registered for both indices in 2007. Moreover, the traditional Lerner index indicates a lower level of competition than the elasticity-adjusted Lerner index for most years. Furthermore, the elasticity-adjusted Lerner index is significantly different from zero or one

---

[12] There are other types of financial institutions in China, such as trust and investment corporations, rural commercial banks, savings banks, co-operative banks, investment banks and policy banks. We exclude these institutions from our investigation for several reasons. Firstly, in the 1990s, trust and investment corporations were important financial institutions that operated similarly to commercial banks (Hong and Yan, 1997). However, in the late 1990s, they experienced significant problems and most of them were taken over by commercial banks. Since the primary focus of this paper is to assess bank competition during 1996-2008, we believe it is safe to exclude trust and investment corporations from our analysis. Secondly, most of the other banks that are not included in our investigation capture only very small portions of Chinese lending markets and/or were established with different objectives from commercial banks. Thirdly, there are significant data limitations for especially the large number of small banks that are excluded from the sample.

[13] Banks with more than 50% foreign ownership are classified as foreign banks. We only include foreign banks that provide separate balance sheet data to the PBC. This means that several banks headquartered in Hong Kong SAR and which are classified as foreign banks by the CBRC, but do not provide separate balance sheet data for their operations in mainland China, are excluded from our sample.

[14] The results for the Panzar-Rosse H-statistic are reported in Appendix E.

for all years, rejecting the null hypothesis that Chinese loan markets are in a state of either perfect competition or monopoly.

Panel B and Panel C report subsample results. In Panel B, the sample is divided by the WTO accession in 2001. The results indicate that the Pre-WTO (1996-2001) elasticity-adjusted Lerner index is 0.249 and increases to 0.342 for the Post-WTO period (2002-2008). In Panel C, we further divide the Post-WTO sample into two subsamples, using the year 2004 with the interest rate reforms as the break-year. The Pre-Reform period (2002-2004) has an elasticity-adjusted Lerner index of 0.245; it increases to 0.357 for the Post-Reform period (2005-2008). We find similar results for the traditional Lerner index.

We conduct Chi-squared Wald tests to determine if competitive conditions experienced structural changes among the subsamples. Our results are summarized in the last rows of Table 4. We focus on the elasticity-adjusted Lerner index. The null hypothesis that the Pre-WTO elasticity-adjusted Lerner index is larger or equal than the one for the Post-WTO period is rejected at 1%, indicating a higher level of competition for the former period relative to the latter period. When we divide our estimations into three subsamples, we find that competitive conditions remain unchanged between the Pre-WTO and Pre-Reform periods (H_0: *Elasticity Adj Lerner Pre-WTO>=Elasticity Adj Lerner Pre-Reform* cannot be rejected), while deteriorating significantly for the Post-Reform period (H_0: *Elasticity Adj Lerner Pre-Reform>=Elasticity Adj Lerner Post-Reform* and H_0: *Elasticity Adj Lerner Pre-WTO>=Elasticity Adj Lerner Post-Reform* are rejected), marking the latter period as the least competitive period of all.

Combining these results, we conclude that competition such as measured by the (elasticity-adjusted) Lerner index was the highest for the Pre-WTO period (1996-2001) and deteriorated markedly for the Post-WTO period (2002-2008), most significantly for the Post-Reform period (2005-2008). We conducted various robustness tests including alternative ways to calculate marginal costs and estimating directly the conjectural variation variable (Uchida and Tsutsui, 2005). These tests all yielded similar results.[15]

Our findings are supported by Soedarmono *et al*. (2013) and Fungáčová *et al*. (2012),

[15] These results are reported in Appendix D.

which also document a general decreasing trend of bank competition in China during 2002-2008, based on the elasticity-adjusted Lerner index and traditional Lerner index, respectively. The latter study obtains an average Lerner index of 0.378 for this period, while our elasticity-adjusted Lerner index and traditional Lerner indices for the same period are 0.342 and 0.375, respectively. Comparison with results from studies for other countries show that the values obtained for China are relatively high: Berger *et al*. (2009) find an average Lerner index of 0.22 for 23 industrial countries calculated over the period 1999-2005, while Carbó Valverde *et al*. (2009) obtain a mean of 0.16 for the European Union during 1995-2001. At the same time, our estimates for the pre-WTO period are around 0.25, indicating that competition in Chinese loan markets was not that much lower than that in developed economies during those years. The post-WTO period, however, is significantly less competitive for China than for the other countries.

6.2. Empirical results: Profit Elasticity (PE) indicator.

Similarly to the Lerner index, the empirical estimation of the PE indicator starts with the estimation of marginal costs. In this section, we improve the marginal cost estimation by assuming different Translog Cost Functions (TCF) for each bank type.[16] More specifically, we estimate one separate TCF for the SOCBs, JSCBs, CCBs and the foreign banks (FOREIGN), which should improve the accuracy of the estimation of marginal costs.[17] Given the estimated marginal costs, we are now able to estimate the PE indicator. For China, we use the relationship between the marginal costs of individual banks and their profits:

$$ln\pi_{ilt} = \alpha + \sum_{t=1}^{T-1} \gamma_t d_t + \sum_{t=1}^{T} \beta_t \, d_t lnmc_{ilt} + u_{ilt} \tag{11}$$

where π_{ilt} stands for profits derived from loans, d_t is a time dummy, mc_{ilt} denotes marginal costs, *i* refers to bank *i*, *l* to output type "loans", and *t* to year t; u_{ilt} is the error

[16] Estimating marginal costs from homogenous Translog Cost Function does not change our results. See Appendix D.
[17] The estimation of the TCF is reported in Appendix B (Table 11).

term. This provides us with the coefficient β_t, i.e. the PE indicator (as is given by $PE = \frac{dln(\pi_i)}{dln(mc_i)}$). β_t is negative in theory, reflecting that higher marginal costs reduce profits for all banks.[18] Moreover, the more competitive the market, the lower the value of β_t. In other words, banks are punished more harshly for being inefficient in more competitive markets. Note that the indicator $\beta_t$is time-dependent. Profits derived from loans are defined as:

$$\pi_{ilt} = q_{it}(p_{it} - mc_{ilt}) \tag{12}$$

where q_{it} denotes the total amount of loans and p_{it} is the loan interest rate calculated as interest income over loans.

We expect higher profits to go hand in hand with lower marginal costs, but since our definition of profits is a function of marginal costs, there may be an endogeneity problem. To correct for this, we employ lagged marginal costs as instrument variable and investigate various alternative estimation techniques.

We follow the strategy set out by Angrist and Pischke (2009) and first test whether the instrumental variables are weak. For this purpose, we employ Angrist-Pischke (AP) F-statistics to test for weak identification of individual endogenous regressors. The AP first-stage F-statistics indicate that a particular endogenous regressor is weakly identified if the null hypothesis is rejected.[19] Table 5 reports that nearly all instrumental variables used are strong with F-test values above 16.38, with the exception of the instrumental variables for 1997, 1998, 2002 and 2003, indicating for these years weak instrumental variables.

Because the instrumental variables for some years have weak properties, we use only just-identified instruments as they are median-unbiased and not subject to the weak instrumental variable critique. Furthermore, following the suggestion of Angrist and Pischke (2009), we check the two-stage least squares (2SLS) results with estimates from Limited

[18] In practice, a positive β_t is possible (Van Leuvensteijn et al., 2011), which could be the result of extreme collusion, market regulation or banks competing on quality (Tabak et al., 2012).
[19] The Stock-Yogo weak ID test critical value at 10% (maximal LIML size) is 16.38 (Stock et al., 2002).

Information Maximum Likelihood (LIML), as the latter results are less biased. LIML can be seen as a "combinatory estimation" technique where the ordinary least square (OLS) and 2SLS estimations are combined and the weights for the two estimations are determined by the data (see Angrist and Pischke, 2009, for further explanation). We use as instrument variables one-year lagged values of marginal costs and kernel-based heteroskedastic and autocorrelation-consistent (HAC) variance estimations. The bandwidth in the estimation is set at two periods, and the Newey-West kernel is applied. The results of 2SLS and LIML are very similar, in fact almost identical, and therefore we only present the results with LIML.[20]

To assess the evolution of bank competition, we first estimate the yearly PE indicator based on Equation (11). Table 5 reports the results. The estimation results for the years 1996 -1999 are based on a small sample and should therefore be interpreted with caution. The estimations for the subsamples and whole sample estimates are panel estimates and are therefore more reliable. The yearly PE indicators are significantly different from zero for most of the sample years, except for the 1997-2000 period. Competition increased sharply during 2001-2003 and then declined up to 2005. It then intensified again, followed by a slightly decreasing level of competition in 2007 and 2008. In general, the development of the yearly PE indicator suggests that competitive conditions in Chinese loan markets improved, especially after WTO accession in 2001.[21] Admittedly, the insignificant results for the early years in our sample could be caused by the small number of observations for those years, in which case the results could be influenced strongly by outliers. Therefore, we estimate the PE indicator for subsamples to avoid small-sample bias.

We estimate one PE indicator for each subsample and test whether competition changed significantly after WTO accession and after the 2004 interest rate reforms. These point estimates can be interpreted as averages of yearly estimates over their respective sample periods, weighted by the number of observations in each year. Estimations are based on the following equation:

[20] Results with 2SLS are available upon request.

[21] Delis (2012) estimates worldwide bank competition with the PE indicator, including China for a sample covering 1988-2005 and using market shares as performance indicator. Our values are not that different from the values reported for China in that study, especially those reported for 2002-2004.

$$ln\pi_{ilt} = \alpha + \gamma Trend + \beta lnmc_{ilt} + u_{ilt} \quad (13)$$

where *Trend* is a time trend.[22]

Table 6 reports the results. The Kleibergen-Paap rk LM statistics are significant at the 1% level for the whole sample and for each subsample, rejecting the null hypothesis that the model is unidentified. The Kleibergen-Paap Wald rk F-statistic is larger than 16.38 for each sample, suggesting that the estimations do not suffer from weak identification. Both test statistics are robust to heteroskedasticity and autocorrelation. All PE indicators have the correct sign (negative) and are significant at the 1% level, except for the pre-WTO period.

Column (1) of Panel A shows that the PE indicator for the whole sample is -2.388 (significant at 1%). For subsamples, we find that the PE indicator for the pre-WTO period is -1.514 (see Column 2), but is statistically insignificant. Competitive conditions improve significantly for the Post-WTO period (2002-2008), such as indicated by a value of the PE indicator of -3.750*** (Column 3). Next, following our approach for the Lerner index estimations, we divide the Post-WTO period into two subperiods: Pre-Reform (2002-2004) and Post-Reform (2005-2008). Columns (4) and (5) suggest that the Pre-Reform period is the most competitive period (PE indicator of -5.094***), while competitive conditions deteriorated in the Post-Reform period (PE indicator of -3.027***). Our estimates are in general less negative than those of Delis (2012), which could be related to our use of profits as performance indicator instead of market shares. In general, the correlation between marginal costs and profits may be higher than the correlation between marginal costs and market shares.

In Panel B, we conduct formal tests on structural changes. We compare the differences in PE indicators among the various subsamples. Column (1) shows that the difference between the PE indicator for the Pre-WTO period and the one for the Post-WTO period is 2.056, which is significantly larger than zero (chi^2 test statistics is 3.61**). This result confirms Hypothesis I that competitive conditions improved after WTO accession, as the PE

[22] Using year dummies instead of a time trend generates similar results for all estimations reported in this paper. Results are available upon request.

indicator for the Post-WTO period is significantly lower than for the Pre-WTO period. Columns (2) and (4) show that both Pre-Reform and Post-Reform periods are more competitive than the pre-WTO period, with a difference between the PE indicators of 3.58 (chi^2:5.89***) and 1.513 (chi^2: 1.99*), respectively. Column (3) compares the changes in competition before and after the 2004 interest rate reform. In line with Hypothesis II, competitive conditions deteriorated after 2004, as indicated by a positive difference between the PE indicators of 2.067 (chi^2: 3.22**). Some caution with the interpretation of this result is appropriate, because after 2004 the Chinese banking sector was reregulated, as we discuss below using the Financial Repression Index of Huang and Wang (2011).

Finally, our estimates of the yearly PE indicators and of the PE indicators for the whole samples and subsamples (point estimates) are robust to different estimation methods and different specifications of the PE indicator.[23]

6.3. Interpretation results PE indicator

The key in understanding why the results for the PE indicator are so different from the (elasticity-adjusted) Lerner index lies in the system of interest rate regulation in China. We show in Appendix C that the Lerner index yields biased results under binding interest rate regulation, while RPD does not. Whether and to what extent interest rate regulation is binding is an empirical question, which definitely requires more attention in the literature on measuring bank competition. The empirical literature on binding interest rate regulation in the context of China is rather small. However, the general consensus is that: a) the lending rate floor was considered to be non-binding in practice (He and Wang, 2012; see Sections 2 and 3); b) the deposit rate ceiling was binding (Feyzioğlu *et al*., 2009; Ma *et al*., 2011; He and Wang, 2012;[24] PBC, 2009; Yi, 2009); and c) the lending rate ceiling was most likely binding during the pre-WTO period (Yi, 2009).

In this respect, the binding deposit rate ceiling may bias the Lerner index through the

[23] These robustness tests are presented in Appendix D.

[24] He and Wang (2012, p. 34): "… Using the regression results, we can then estimate the equilibrium interest rate by subtracting the effects of financial repression from the observed real interest rate: the equilibrium deposit rate in China was estimated at 4.7% in 2005. This estimated equilibrium deposit rate is significantly higher than the observed real deposit rate of 1.6% in 2005, which means that the deposit-rate ceiling must have been binding in China."

reallocation of output and profits among lenders. This reallocation effect (Boone *et al*., 2007) relates to the fact that more intensive competition due to more aggressive conduct will reallocate output and profits from less efficient banks to more efficient banks. As more efficient banks usually have higher PCMs than less efficient banks, the PCM for the whole market, which is an (output) weighted average of individual banks' PCMs, actually may increase in response to more intense competition. The increase in the market PCM (or aggregate Lerner index) would be interpreted as a decline in competition, while actually it has increased. Boone *et al*. (2007) show that the reallocation effect is particularly strong in concentrated markets. In fact, Chinese loan markets are highly concentrated, with during 2001-2008 the four SOCBs having an average annual market share of around 71%. The binding deposit rate ceiling was particularly important after 2004, when real deposit rates reached negative values, reducing the reliability of measuring competition with the (elasticity-adjusted) Lerner index for this period. Moreover, the binding lending rate ceiling during the pre-WTO period could have resulted in an over-estimation of competition with the Lerner index.[25] In contrast, the PE indicator does not suffer from these shortcomings.

Turning to the specific results we obtained with the PE indicator estimations, we find generally positive and insignificant values for the PE indicator for the early years of our sample, suggesting that during 1997-2000 a negative relationship between efficiency (marginal costs) and profits could not be established (see Table 5). This indicates that during the years when Chinese banking markets were still heavily regulated, there was no reward for being more competitive than one's competitors. Actually, this finding is similar to the results for Japan during the 1990s in Van Leuvensteijn *et al*. (2011), which showed positive (and significant) values for the PE indicator. During those years, market shares in Japanese loan markets were more or less guaranteed under the so-called "convoy system" and competitive forces were largely absent.

Subsequently, we start to find negative and highly significant values for the PE indicator for Chinese loan markets from 2001 onwards, indicating that, as loan markets became more competitive, more efficient banks started to generate more profits than less

[25] See Appendix C and Salvo (2010).

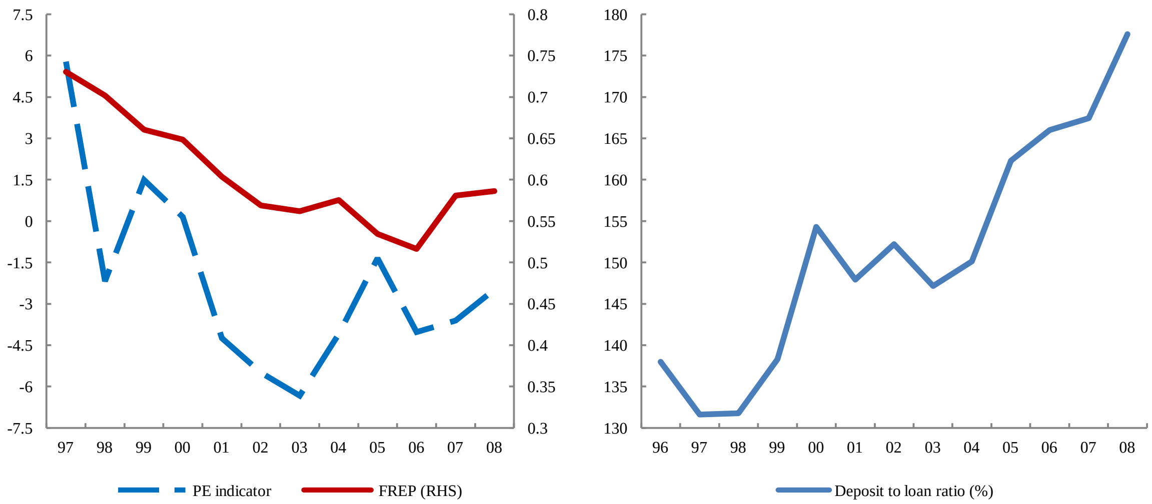
efficient banks. The PE indicator improved further until 2003, when it reached its lowest value of –6.3 (e.g. highest level of competition). From an international perspective, this value is comparable to the most competitive yearly results obtained for several mature economies (Van Leuvensteijn *et al*., 2011).

Then, after 2003, competitive conditions in Chinese loan markets declined (but still with negative and, except for one year, highly significant results), which was the most notable in 2004, 2007 and 2008. We believe that various policy measures and a certain degree of prudential re-regulation may be responsible for this pattern of slightly declining competition (for an overview see Table 1). Perhaps the most important policy change was the interest rate liberalization of 2004, when the PBC removed the lending rate ceiling and deposit rate floor, but maintained the lending rate floor and deposit rate ceiling. The policy move implied that Chinese banks benefited from a more or less guaranteed minimum interest rate spread (due to the remaining floor on the lending rate and ceiling on deposit rate), while they faced no restrictions with respect to its potential maximum width (García-Herrero *et al*., 2005). As a result, the negative correlation between inefficiency (higher marginal costs) and profitability may have been weakened after the reform, because inefficient banks will not be punished as harshly with this guaranteed interest rate spread being in place. The PE indicator is able to pick up this effect by showing a decline in competition after the partial interest rate liberalization in 2004.

Apart from the interest rate reforms, various other policy measures and a certain degree of re-regulation adopted during the 2004-2008 period may also have contributed to the deterioration of competition that we find. In 2004, the China Banking Regulatory Commission (CBRC) adopted new capital adequacy requirements, including the requirement to fully provision their non-performing loans and maintain at least 8% of aggregate capital adequacy, that banks should meet by 2007 (Podpiera, 2006). Further in 2004, the CBRC strengthened other parts of its regulatory policies, including its on-site examinations and monitoring of large exposures, and introduced risk-based supervision for the CCBs. Regulation was tightened regarding non-performing loans (NPLs), with a view to reducing banks' NPL ratios (Liu, 2009). The combined impact of these measures may have affected competitive conditions in Chinese loan markets. In addition, the People’s Bank of China, worried by a possible overheating of the Chinese economy, re-introduced credit quotas in the fall of 2007 that aimed to mitigate bank lending growth (Fukumoto *et al*., 2010). These lending restrictions were kept in place until the fall of 2008 and can be characterized as a major step of re-regulation, as they re-instated elements of the old credit plan system.

The element of re-regulation is picked up nicely by the financial repression index (*FREP*) developed for China in Huang and Wang (2011) (see Figure 1, left-hand panel). It is based on six financial repression variables, including two interest rates, two loan market share variables, reserve requirements and capital account controls. During the years of our sample – 1996-2008 – the index declines, suggesting less financial repression for all years except for 2004, 2007 and 2008, when it increases. After its first rise in 2004, indicating stronger financial repression, it fell to its lowest level ever in 2006, before strongly increasing in 2007, followed by a further pick-up in 2008. The yearly results of the PE indicator, which is depicted for illustrative purposes in Figure 1 (left-hand panel), closely follow the pattern of the financial repression index. The generally increasing re-regulation in the latter part of our sample may be reflected in the rather sharply increasing deposits to loans ratio from 2004 onwards (Figure 2, right-hand panel). Possibly, tightened loan controls and other regulatory steps forced banks to reduce the growth of their loans relative to that of their deposits.

Figure1. Interpretation results PE indicator and financial repression



Given the strong similarity between the pattern of the PE indicator and the financial repression index, we are interested in how this relation looks for other financial liberalization indices. To this end, we employ two additional indicators of financial reform: the overall financial liberalization indices (*Fin_Lib Index*) and interest rate liberalization indices (*Int_Lib Index*) developed by Abiad *et al*. (2010).[26] The former index measures the overall degree of financial liberalization, with values ranging from 0 to 1, with a higher value indicating a more liberalized financial system. The latter index, which takes the values 0, 1, 2 or 3, indicates fully repressed, partially repressed, partially liberalized and fully liberalized interest rates, respectively.

In order to provide a more comprehensive analysis, we calculate the pair-wise correlation coefficients between the three indices of financial reform and the same three measures of competition that we used. The results are reported in Table 7. Should financial reform promote competition, one would expect positive correlations between the financial repression index and the competition measures. In contrast, one would expect negative correlations between the two other financial liberalization indices and the competition measures if financial reform promotes competition. Since financial reform may affect

[26] The values of these indicators are shown in Appendix A, Table 9.

banking behavior with a time lag, we use one-period lagged values of the three indices.[27]

In case a more liberalized financial system is associated with more intense competition, the correlations of the PE indicator show the expected sign with all three indices (positive for the financial repression index and negative for the two others). The correlations are also highly significant at the 1% level. In contrast, the correlations of the other measures that are significant all have the opposite sign, suggesting that increased liberalization is associated with weaker competition.

7. Conclusions

Using balance sheet information for a large sample of banks operating in China during 1996-2008, we show that competition actually increased in the past decade when the Profit Elasticity (PE) measure introduced by Boone *et al*. (2007) and Boone (2008) is used as indicator of competition. We find that the period after China's entry into the WTO in 2001 was characterized by significantly more competitive loan markets than before. This stands in contrast to the results that we obtain by calculating the conventional and elasticity-adjusted Lerner indices.[28] We doubt the latter findings, as they may be distorted by various factors, including restrictions on market shares and interest rates. Our results for both the PE indicator and other measures are robust for a large number of alternative specifications and estimation methods. All in all, our analysis suggests that bank lending markets in China have been more competitive than previously assumed. Another major empirical finding that we report is that significant improvements in competition in Chinese loan markets moved in parallel with progress in financial reform. This result is much in line with those obtained for other emerging economies. Furthermore, our analysis of the interest rate reforms implemented in 2004 shows that removing the lending rate ceilings was especially beneficial for the inefficient banks, as suggested by the reduction in competition that we find using the PE indicator.

---

[27] We also employed the current values of the financial reform indices to account for the possibility that banks may anticipate financial reform measures and adjust their competitive strategies accordingly. The results are similar to the ones we report here.

[28] We find a similar result for the Panzar-Rosse H-statistic. See Appendix E.

Theoretically, we have shown that the theoretical foundation of the PE indicator (or Boone's RPD model) is not biased due to interest rate regulation (see Appendix C). This makes the PE indicator a much better measure to gauge competition in loan markets that are subjected to interest rate regulation than conventional approaches. This is a very general insight that can be useful for investigations of competitive conditions in banking markets in countries where binding regulation of interest rates is a distinctive characteristic. More generally, our findings indicate that the bank competition literature may wish to focus more explicitly on the potential biases in competition measures that result from the existence of binding price regulation. Finally, the policy implication of our analysis is that foreign entry or the threat of foreign entry is an effective way to increase competition in loan markets.

References

Abiad, A., Detragiache, E., Tressel, T., 2010. A new database of financial reforms. *IMF Staff Papers* 57(2), 281-302.

Allen, F., Qian, J., Qian, M., 2005. Law, finance, and economic growth in China. *Journal of Financial Economics* 77(1), 57-116.

Altunbas, Y., Liu, M., Molyneux, P., Seth, R., 2000. Efficiency and risk in Japanese banking. *Journal of Banking and Finance* 24, 1605-1628.

Angelini, P., Cetorelli, N., 2003. The effects of regulatory reform on competition in the banking industry. *Journal of Money, Credit and Banking* 35(5), 663-684.

Angrist, J., Pischke, J.S., 2009. Mostly harmless econometrics. Princeton University Press, Princeton and Oxford.

Appelbaum, E., 1982. The estimation of the degree of oligopoly. *Journal of Econometrics* 19, 287-299.

Berger, A., Klapper, L., Turk-Ariss, R., 2009. Bank competition and financial stability. *Journal of Financial Services Research* 35(2), 99-118.

Berger, A., Mester, L., 1997. Inside the black box: What explains differences in the efficiencies of financial institutions? *Journal of Banking and Finance* 21(7), 895–947.

Bikker, J., Spierdijk, L., 2008. How banking competition changed over time. Working Paper 167, De Nederlandsche Bank.

Bikker, J., Shaffer, S., Spierdijk, L., 2012. Assessing competition with the Panzar-Rosse Model: The role of scales, costs, and equilibrium. *Review of Economics and Statistics* 94 (4) 1025-1044.

Boone, J., 2000. Competition. Discussion Paper Series 2636, CEPR.

Boone, J., 2001. Intensity of competition and the incentive to innovate. *International Journal of Industrial Organization* 19(5), 705–726.

Boone, J., Van Ours, J., Van Der Wiel, H., 2007. How (not) to measure competition. Tilburg Law and Economics Center (TILEC) Discussion Paper No.2007-014, Tilburg University.

Boone, J., 2008. A new way to measure competition. *Economic Journal* 118, 1245 -1261.

Boone, J., Van Leuvensteijn, M., 2010. Measuring competition using the profit elasticity: American sugar industry, 1890-1914. Discussion Paper Series No.8159, CEPR.

Bresnahan, T., 1982. The oligopoly solution concept is identified. *Economics Letters* 10, 87–92.

Carbó Valverde, S., Humphrey, D., Maudos, J., Molyneux, P., 2009. Cross-country comparisons of competition and pricing power in European banking. *Journal of International Money and Finance* 28(1), 115-134.

Claessens, S., 2009. Competition in the financial sector: Overview of competition policies. *The World Bank Research Observer* 24(1), 83-118.

Corts, K., 1999. Conduct parameters and the measurement of market power. *Journal of Econometrics* 88, 227-250.

Delis, M., 2012. Bank competition, financial reform and institutions: The importance of being developed. *Journal of Development Economics* 97, 450-465.

Demsetz, H., 1973. Industry structure, market rivalry, and public policy. *Journal of Law and Economics* 16, 1-9.

Fernández de Guevara, J., Maudos, J., Pérez, F., 2007. Integration and competition in European financial markets. *Journal of International Money and Finance* 26(1), 26–45.

Feyzioğlu, T., Porter, N., Takáts, E., 2009. Interest rate liberalisation in China. Working Paper WP/09/171, International Monetary Fund.

Fu, X., 2009. Competition in Chinese commercial banking. *Banking and Finance Review* 1(1), 1-16.

Fu, X., Heffernan, S., 2007. Cost X-efficiency in China's banking sector. *China Economic Review* 18, 35-53.

Fu, X., Heffernan, S., 2009. The effects of reform on China's bank structure and performance. *Journal of Banking and Finance* 33(1), 39-52.

Fu, X., Lin, Y., Molyneux, P., 2014. Bank competition and financial stability in Asia Pacific. *Journal of Banking and Finance* 38, 64 -77.

Fukumoto, T., Higashi, M., Inamura, Y., Kimura, T., 2010. Effectiveness of window guidance and financial environment: In light of Japan's experience of financial liberalisation and a bubble economy. *Bank of Japan Review*, No.2010-E-4.

Fungáčová, Z., Pessarossi, P., Weill, L., 2012. Is bank competition detrimental to efficiency? Evidence from China. BOFIT Discussion Papers No.31/2012, Bank of Finland.

García-Herrero, A., Gavilá, S., Santabárbara, D., 2005. China's banking reform: An assessment of its evolution and possible impact. Documentos Ocasionales No.0502, Banco de España.

Genesove, D., Mullin, W., 1998. Testing static oligopoly models: Conduct and cost in the sugar industry, 1890-1914. *The RAND Journal of Economics* 29(2), 355-377.

He, D., Wang, H., 2012. Dual-track interest rates and the conduct of monetary policy in China. *China Economic Review* 23(4), 928-947.

Ho, C.Y., 2010. Deregulation, competition and consumer welfare in a banking market: evidence from Hong Kong. *Journal of Regulatory Economics* 37, 70–97.

Hong, Z., Yan, Y., 1997. Trust and investment corporations in China. Working Paper No.97/06, Federal Reserve Bank of Cleveland.

Huang, Y., Wang, X., 2011. Does financial repression inhibit or facilitate economic growth? A case study of Chinese reform. *Oxford Bulletin of Economics and Statistics* 73(6), 833-855.

Liu, M. 2009. Thirty years of reform and opening up – China's banking industry. (in Chinese), China Financial Publishing House, Beijing.

Ma, G., Yan, X.D., Xi, L., 2011. China's evolving reserve requirements. Working Paper No.360, Bank for International Settlements.

Matthews, K., Zhang, N., 2010. Bank productivity in China 1997-2007: Measurement and convergence. *China Economic Review* 21(4), 617-628.

Okazaki, K., 2007. Banking system reform in China: The challenges of moving toward a market-oriented economy. Occasional Paper No.194, RAND National Security Research Division.

Panzar, J., Rosse, J., 1987. Testing for 'monopoly' equilibrium. *Journal of Industrial Economics* 35(4), 443–456.

People's Bank of China 2005. Supplement to the China Monetary Policy Report – Report on Steady Progress in Market-Based Interest Rate Reform. People's Bank of China, Monetary Policy Analysis Group, January.

People's Bank of China 2009. Chinese Monetary Policy Report, 2009Q2. (in Chinese), People's Bank of China.

Podpiera, R., 2006. Progress in China's banking sector reform: Has bank behaviour changed? Working Paper No.06/71, International Monetary Fund.

Rosse, J., Panzar, J., 1977. Chamberlin vs Robinson: An empirical study for monopoly rents. Bell Laboratories Economics Discussion Paper No. 90.

Salvo, A., 2010. Inferring market power under the threat of entry: The case of the Brazilian cement industry. *RAND Journal of Economics* 41(2), 326-350.

Schaeck, K., Cihák, M., 2010. Competition, efficiency and soundness in banking: An industrial organisation perspective. Discussion Paper No.2010-20S, Tilburg University European Banking Center.

Schmalensee, R., 1987. Competitive advantage and collusive optima. *International Journal of Industrial Organization* 5, 351-367.

Soedarmono, W., Machrouh, F., Tarazi, A., 2013. Bank competition, crisis and risk taking: Evidence from emerging markets in Asia. *Journal of International Financial Markets, Institutions and Money* 23, 196-221.

Stock, J., Wright, J., Yogo, M., 2002. A survey of weak instruments and weak identification in Generalized Method of Moments. *Journal of Business and Economic Statistics* 20, 518-29.

Tabak, B., Fazio, D., Cajueiro, D., 2012. The relationship between banking market competition and risk-taking: Do size and capitalization matter? *Journal of Banking and Finance* 36(12), 3366-3381.

Uchida, H., Tsutsui, Y., 2005. Has competition in the Japanese banking sector improved? *Journal of Banking and Finance* 29,419-439.

Van Leuvensteijn, M., 2008. The Boone indicator: Identifying different regimes of competition for the American Sugar Refining Company 1890-1914. Tjalling C. Koopmans

Research Institute Discussion Paper No.08-37, Utrecht School of Economics.

Van Leuvensteijn, M., Bikker, J., Van Rixtel, A., Kok-Sørensen, C., 2011. A new approach to measuring competition in the loan markets of the euro area. *Applied Economics* 43(23), 3155-3167.

Van Leuvensteijn, M., Kok-Sørensen, C., Bikker, J., Van Rixtel, A., 2013. Impact of bank competition on the interest rate pass-through in the euro area. *Applied Economics* 7(48), 1359-1380.

Vander Vennet, R., 2002. Cost and profit efficiency of financial conglomerates and universal banks in Europe. *Journal of Money, Credit and Banking* 34 (1), 254-282.

Wolfram, C., 1999. Measuring duopoly power in the British electricity spot market. *American Economic Review* 89(4), 805-826.

Wong, Y., Wong, M., 2001. Competition in China's domestic banking industry. Cato Journal 21(1), 19-41.

Xu, B., Van Rixtel, A., Van Leuvensteijn, M., 2013. Measuring bank competition in China: A comparison of new versus conventional approaches applied to loan markets, Working Paper No.422, Bank for International Settlements.

Yi, G., 2009. Thirty years of reform and opening up, China's interest rate marketization process. *Journal of Financial Research* (Chinese Journal, Jinrong Yanjiu) 1(343).

Yuan, Y., 2006. The state of competition of the Chinese banking industry. *Journal of Asian Economics 17* (3), 519-534.

Table 1: Main banking reforms and policy steps[1]

	Banking reforms and important policies	**Interest rate liberalization**
1995	Commercial Bank Law was promulgated; Central Bank Law was passed.	
1996	Creation of interbank market; start foreign currency business commercial banks.	Liberalization of interest rates in the inter-bank market.
1997	Transformation of state owned banks into "commercial banks without direct administrative controls".	Liberalization of bond repo rates.
1998	Credit allocation based on market principles rather than quotas under credit plan.	Lending rate ceilings increased to 120% of benchmark rate.
1999	Transfer bad debts of SOCBs to four asset management companies.	Lending rate ceiling increased to 130% of benchmark rate.
2000		Liberalization of foreign currency lending rates; liberalization foreign currency deposit rates for accounts over 3 million $.
2001	Accession to World Trade Organization (WTO; foreign banks to be treated equally domestic banks within five years.	
2002	Foreign banks allowed to provide foreign currency services to Chinese residents;	
2003	New bank regulator (China Banking Regulatory Commission or CBRC); CBRC encouraged foreign banks to buy stakes in Chinese banks; Law of Banking Regulation and Supervision adopted.	Liberalization interest rates small-value deposits in Sterling, CHF and Can $; removal lower interest rate limit on small-value foreign currency deposits.
2004	Foreign banks allowed providing local currency services to Chinese enterprises in designated cities; CBRC required banks to fully provision NPLs and maintain min 8% capital ratio, fully binding as of 2007; CBRC strengthened on-site examinations and monitoring of large exposures, and introduced risk-based supervision.	Liberalization foreign currency deposit rates for small accounts with maturity > 1 y; RMB lending rate ceiling increased to 170% of benchmark rate; removal RMB lending rate ceiling, except for urban and rural credit cooperatives; removal of all RMB deposit rates floors; established lending rate adjustment reporting system; lending rate floor was reduced to a specific range versus benchmark rates.
2005	Banks encouraged to list on stock exchanges; five-tier loan classification made fully compulsory for all banks.	
2006	By December, China opened its banking sector fully to foreign banks and eliminated geographic and client restrictions; raised reserve requirements ratio (RRR) to 9% from 6% during 2003-2006.	
2007	PBC re-introduced credit quotas for individual banks to curb lending activities.	
2008	RRR were increased in 16 steps from 9% to 17.5% during 2007-August 2008; RRR were reduced to 15.5% for large banks and to 13.5% for small banks in Q4 2008.	

[1] This overview includes various policy steps after 2004 which correspond to a certain degree of re-regulation and policy tightening after 2004 that we discuss in section 6.

Table 2: Overview of the Chinese banking sector 1996-2008

	Share of total assets (%)				Share of total loans (%)				Share of total deposits (%)			
	SOCB	JSCB	CCB	FOREIGN	SOCB	JSCB	CCB	FOREIGN	SOCB	JSCB	CCB	FOREIGN
Average 1996–2001	86.32	11.83	1.69	0.16	88.03	10.54	1.29	0.14	87.55	11.41	0.90	0.14
Average 2002–2008	72.17	21.25	5.78	0.79	71.30	22.40	5.43	0.87	74.47	21.51	3.43	0.59
Average 1996–2008	78.70	16.91	3.89	0.50	79.02	16.93	3.52	0.53	80.51	16.85	2.26	0.38

Source: BankScope and authors' own calculations.

Table 3: Mean values of key variables by bank group

	SOCB	JSCB	CCB	FOREIGN	All Banks
Total costs/Total assets	4.61	3.71	3.57	3.16	3.6
Loans/Total assets	57.66	55.61	53.38	55.12	54.46
Securities/Total assets	21.21	19.08	22.81	9.81	20.22
Other services/Total income	7.26	6.01	9.98	15.4	9.85
Interest expenses/Total funding	3.4	2.35	2.25	2.78	2.45
Other expenses/Fixed assets	37.16	69.02	56.47	224.22	87.68
Interest income/Total assets	5.19	4.43	4.36	4.14	4.39
Personnel expenses/Total assets [a]	0.54	0.41	0.51	0.59	0.5
Interest income/Total loans	9.07	8.22	8.26	10.36	8.69
Other non-earning assets/Total assets [b]	3.23	2.68	3.18	2.95	3.04
Funding mix [c]	94.3	91.41	89.22	64.96	86.36
Equity/Total assets	3.72	4.74	5.11	29.5	9.49
Other income/Interest income	6.77	6.6	11.8	14.41	10.83

Notes: All data are in percentages.

[a] Personnel expenses to assets ratio serves as a proxy of the wage rate. Ideally, the wage rate is the ratio of personnel expenses to the number of staff. However, many banks do not provide information on the variables. Some researchers replace the missing number of employees by assuming that its growth rate is equal to that of total assets for a given bank (Fu and Heffernan, 2007; Altunbas *et al*., 2000; Vander Vennet, 2002). This approach might not be appropriate for our sample, as very few CCBs report the number of employees, so this growth rate cannot be calculated anyway. We instead follow the approach taken by Van Leuvensteijn *et al*. (2011) and proxy wages by the ratio of personnel expenses to total assets. We adopt the following procedure to approximate missing data on personnel expenses. For banks that provide these data but not for all years, we fill in the missing values of personnel expenses by assuming that they grew at the same rate as non-interest expenses. This is a reasonable assumption, as, according to Chinese accounting standards, non-interest expenses are composed of personnel expenses and non-operating expenses. For banks that do not report personnel expenses at all, we replace missing values by assuming that the ratio of personnel expenses to non-interest expenses equals the average of this ratio for the corresponding bank group, namely $pe_{it} = (pe_{jt}/nie_{jt})nie_{it}$ where pe_{jt}/nie_{jt} is the average personnel expenses to non-interest expenses ratio, by bank type and year; j (j = SOCBs, JSCBs, CCBs, FOREIGN) represents bank groups and i stands for individual banks. Since our sample has almost complete data on non-interest expenses, we can use this approach to back-engineer the missing data on personnel expenses.

[b] Other non-earning assets to total assets ratio is defined as: (total assets minus loans minus other earning assets)/total assets.

[c] The funding mix is defined as: customer deposits/(total funding minus deposits from banks).

Table 4: Lerner Index, elasticity-adjusted Lerner Index and marginal costs

	λ_t	Average loan rate	Average loan deposit spread	Elasticity-adjusted Lerner index	Lerner index	MC_e	MC_l
Panel A: Yearly results							
1997	0.062	0.168	0.085	0.367	0.358	0.095	0.098
1998	0.045	0.118	0.063	0.381	0.359	0.079	0.076
1999	0.029	0.102	0.044	0.288	0.305	0.064	0.061
2000	0.021	0.098	0.049	0.212	0.303	0.056	0.053
2001	0.019	0.086	0.05	0.224	0.303	0.052	0.049
2002	0.015	0.071	0.049	0.214	0.306	0.045	0.041
2003	0.017	0.071	0.049	0.235	0.327	0.047	0.043
2004	0.02	0.068	0.047	0.288	0.34	0.047	0.043
2005	0.023	0.079	0.054	0.287	0.358	0.052	0.048
2006	0.025	0.079	0.052	0.324	0.38	0.052	0.048
2007	0.036	0.079	0.052	0.452	0.434	0.05	0.046
2008	0.041	0.094	0.061	0.439	0.402	0.06	0.055
Panel B: Subsample results-WTO (2001) as breaking point							
Pre-WTO: 1996–2001	0.029	0.115	0.059	0.249	0.32	0.071	0.066
Post-WTO: 2002–2008	0.027	0.079	0.052	0.342	0.375	0.051	0.047
Panel C: Subsample results-WTO (2001) and interest rate reform (2004) as breaking points							
Pre-WTO: 1996-2001	0.029	0.115	0.059	0.249	0.32	0.071	0.066
Pre-Reform: 2002-2004	0.017	0.070	0.049	0.245	0.33	0.047	0.042
Post-Reform: 2005–2008	0.031	0.082	0.054	0.375	0.40	0.054	0.049

H_0: *Elasticity Adj Lerner Pre-WTO>=Elasticity Adj Lerner Post-WTO*: $chi^2(1)=12.13$ p-value = 0.0002

H_0: *Elasticity Adj Lerner Pre-WTO>=Elasticity Adj Lerner Pre-Reform*: $chi^2(1)=0.01$ p-value = 0.54

H_0: *Elasticity Adj Lerner Pre-WTO>=Elasticity Adj Lerner Post-Reform*: $chi^2(1)=23.29$ p-value = 0.0000

H_0: *Elasticity Adj Lerner Pre-Reform>=Elasticity Adj Lerner Post-Reform*: $chi^2(1)=15.14$ p-value = 0.0000

Notes: λ_t are statistically different from zero for all years at 1% significance level; MC_e and MC_l are average marginal costs for elasticity-adjusted Lerner index and traditional Lerner index, respectively.

Table 5: Yearly PE indicator

	PE Indicator	z-value	AP $chi^2(1)$ p-value	AP F (1,433)
1997	5.783	-0.44	0.4866	0.46
1998	–2.177	(–1.23)	0.1021	2.53
1999	1.489	-0.56	0.0000	31.78
2000	0.147	-0.05	0.0000	27.91
2001	–4.250***	(–5.85)	0.0000	31.11
2002	–5.497**	(–2.36)	0.0002	13.1
2003	–6.327***	(–2.64)	0.0147	5.64
2004	–4.092***	(–3.92)	0.0000	58.28
2005	–1.352	(–1.45)	0.0000	67.26
2006	–4.024***	(–4.17)	0.0000	20.73
2007	–3.611***	(–5.36)	0.0000	89.77
2008	–2.482***	(–4.12)	0.0000	28.15
Constant	0.401	-0.23		
Nr obs		457		
F		6.249		
Centered R^2		0.131		

Notes: z-values in parenthesis; AP chi^2 is the Angrist-Pischke (AP) first-stage chi-squared test. AP F is the Angrist-Pischke (AP) F-statistic, which can be compared to Stock *et al*. (2002) critical values for Cragg-Donald F statistic with K1=1. The Stock-Yogo weak ID test critical value at 10% maximal LIML size is 16.38. Year dummies are not reported to save space.
** denotes test statistic significant at the 5% level
*** denotes test statistic significant at the 1% level

Table 6: Point estimates PE indicator: Whole sample and subsamples

Panel A: Estimates of PE indicators

	All: 1996-2008	*Pre-WTO*: 1996-2001	*Post-WTO*: 2002-2008	*Pre-Reform*: 2002-2004	*Post-Reform*: 2005-2008
	(1)	(2)	(3)	(4)	(5)
PE Indicator	–2.388***	–1.514	–3.570***	-5.094***	-3.027***
	(–5.78)	(–1.43)	(–7.74)	(-4.48)	(-6.78)
Time Trend	–0.0332	–0.519**	0.345***	0.3011	0.5813***
	(–0.82)	(–2.37)	(4.71)	(0.98)	(4.71)
Constant	–0.24	4.966**	–8.050***	-12.429***	-9.185***
	(–0.19)	-2.18	(–4.51)	(-2.27)	(-4.71)
Nr. Obs	457	87	370	112	258
F	16.78	2.97	33.67	11.52	34.23
Centered R^2	0.089	0.141	0.18	0.01	0.25
KP-F	211.4	30.98	130.8	25.07	138.50
KP-LM	62.00(0.0000)	13.94(0.0001)	44.22(0.0000)	18.03(0.0000)	31.17(0.0000)

Panel B: Differences in PE indicators

	Pre-WTO – Post-WTO	*Pre-WTO –Post-Reform*	*Post-Reform –Pre-Reform*	*Pre-WTO –Post-Reform*
	(1)	(2)	(3)	(4)
Difference	2.056	3.58	2.067	1.513
Test Difference<=0, chi^2 (p-value)	3.61** (0.029)	5.89***(0.008)	3.22**(0.036)	1.99*(0.079)

Notes: z-values in parenthesis; Since there is only one endogenous variable, we use Kleibergen-Paap rk Wald F (KP-F) and Kleibergen-Paap rk LM (KP-LM) tests to test weak identification and under-identification. The Stock-Yogo weak ID test critical value at 10% maximal LIML size is 16.38.

** denotes test statistic significant at the 5% level

*** denotes test statistic significant at the 1% level

Table 7: Pair-wise correlation coefficients with financial reform indices

	PE	Lerner	Elasticity-adjusted Lerner
FREP	0.6560***	–0.5794***	–0.3104***
Fin_Lib Index	–0.4223***	0.5015***	–0.055
Int_Lib Index	–0.2206***	0.7917***	0.3285***

Notes: FREP is the financial repression index as reported in Huang and Wang (2011). Fin_Lib Index and Int_Lib_Index represent financial liberalization index and interest rate liberalization index, respectively.
*** denotes test statistic significant at the 1% level

Appendix

A. Supplemental tables for the main analysis

Table 8: Distribution of observations

	SOCB	JSCB	CCB	FOREIGN	Observations
1996	4	9	1	4	18
1997	4	10	3	6	23
1998	4	10	5	7	26
1999	4	10	9	7	30
2000	4	10	14	5	33
2001	4	10	17	7	38
2002	4	10	27	8	49
2003	4	10	33	8	55
2004	4	12	40	8	64
2005	4	12	55	10	81
2006	4	13	74	11	102
2007	4	13	73	26	116
2008	4	13	36	26	79
Total observations	52	142	387	133	714
Number of banks	4	13	82	28	127

Table 9: Financial reform indices

	1996	1997	1998	1999	2000	2001	2002	2003	2004	2005
Financial liberalization index	0.179	0.226	0.298	0.345	0.345	0.345	0.393	0.393	0.488	0.488
Interest rate liberalization index	0	0	0	0	0	0	1	1	2	2

Source: Abiad et al. (2010), http://www.imf.org/external/pubs/ft/wp/2008/data/wp08266.zip. A value of 0 indicates a fully repressed financial system, while a value of 1 points at a fully liberalised one. Interest rate liberalization index, which takes values of 0, 1, 2 and 3, indicates respectively a fully repressed, partially repressed, partially liberalised and fully liberalised system.

Table 10: Estimation of elasticity-adjusted Lerner index

	Yearly estimates		Subsample estimates	
Panel A:Cost Equation				
	Coefficient	z-value	Coefficient	z-value
ln(securities)	–0.505***	(–2.76)	–0.285	(–1.51)
(ln(securities)) ^{2}	0.0300***	(3.57)	0.0314***	(3.61)
ln(other services)	0.973***	(5.23)	0.831***	(4.37)
(ln(other services)) ^{2}	0.0426***	(4.05)	0.0288***	(2.74)
ln(wage)–ln(other expenses)	1.270***	(4.51)	1.447***	(5.20)
(ln(wage) –ln(other expenses)) ^{2}	0.151***	(5.36)	0.150***	(5.41)
ln(funding rate)–ln(other expenses)	0.460**	(2.26)	0.285	(1.38)
(ln(funding rate) –ln(other expenses)) ^{2}	0.197***	(4.94)	0.189***	(4.94)
(ln(wage) –ln(other expenses))(ln(funding rate)–ln(other expenses))*	–0.274***	(–4.96)	–0.268***	(–5.05)
*ln(securities) * ln(other services)*	–0.0265	(–1.59)	–0.0220	(–1.32)
ln(securities)(ln(funding rate)–ln(other expenses))*	0.0528**	(2.29)	0.0415*	(1.84)
ln(securities)(ln(wage)–ln(other expenses))*	–0.164***	(–5.31)	–0.133***	(–4.25)
ln(other services)(ln(funding rate)–ln(other expenses))*	–0.00508	(–0.21)	–0.0306	(–1.32)
*ln(other services) *(ln(wage)–ln(other expenses))*	0.147***	(4.66)	0.161***	(5.19)
ln(equity/assets)	–0.0116	(–0.06)	0.0321	(0.17)
(ln(eqauity/asset)) ^{2}	–0.00769	(–0.23)	0.000250	(0.01)
SOCB	0.398***	(3.11)	0.371***	(3.04)
JSCB	0.332***	(4.37)	0.304***	(4.51)
CCB	0.194***	(3.25)	0.189***	(3.44)
constant	4.054***	(4.17)	4.273***	(4.42)
Panel B:Supply Equation				
ln(loans)	0.864***	(6.39)	0.724***	(4.75)
(ln(loans)) ^{2}	0.0263**	(2.52)	0.0298**	(2.54)
*ln(loans) * ln(securities)*	–0.0370**	(–2.35)	–0.0522***	(–3.06)
*ln(loans) * ln(other services)*	–0.0432***	(–3.26)	–0.0226	(–1.55)
ln(loans)(ln(funding rate)–ln(other expenses))*	–0.0366*	(–1.69)	0.00182	(0.08)
ln(loans)(ln(wage)–ln(other expenses))*	0.0374	(1.54)	–0.0135	(–0.52)
1997	0.0616***	(9.17)		
1998	0.0449***	(7.89)		
1999	0.0294***	(5.89)		
2000	0.0208***	(4.84)		
2001	0.0191***	(4.66)		
2002	0.0151***	(4.19)		
2003	0.0167***	(4.84)		
2004	0.0196***	(5.96)		
2005	0.0227***	(7.39)		
2006	0.0255***	(9.18)		
2007	0.0359***	(13.99)		
2008	0.0415***	(13.18)		
1996–2001			0.0287***	(9.38)
2002–2008			0.0269***	(14.12)
Number of observations:	453		453	

Notes: *z-values* in parenthesis; * p<.1, ** p<0.05, *** p<0.01; Time dummies in cost equation not shown to save space.

B. Estimation translog cost functions (TCF) for PE indicator

In order to be able to calculate marginal costs, we estimate, for each bank group, a translog cost function (TCF) using individual bank observations. This is done by allowing for bank type dummies d_i^h to interact with the independent variables in the TCF, resulting in the following form:

$$\ln c_{it}^h = \alpha_0 + \Sigma_{t=1,\dots,(T-1)}\gamma_t d_t + \Sigma_{j=1,\dots,K}\delta_{jh} d_i^h \ln x_{ijt} + \Sigma_{j=1,\dots,K}\Sigma_{k=1,\dots,K}\epsilon_{jkh} d_i^h \ln x_{ijt} \ln x_{ikt} + v_{it} \qquad \text{(B.1)}$$

where the dependent variable c^h_{it} reflects the production costs of bank i (i= 1,…, N) in year t (t = 1,…, T). The sub-index h (h = 1,…, H) refers to the type category of the bank (state owned banks, joint-stock banks, city commercial banks, foreign banks). The variable d^h_i is a bank type dummy variable, which is 1 if bank i is of type h and otherwise zero. Another dummy variable is d_t, which is 1 in year t and otherwise zero. The coefficients α_h, δ_{jh} and ϵ_{jkh}, all vary with h, the bank type. The parameters γ_t are the coefficients of the time dummies and v_{it} is the error term. The explanatory variables x_{ikt} follow the same interpretation as in Section 4.1.1. The two standard properties of TCF, linear homogeneity in input prices and cost-exhaustion, hold for each bank type h. Namely, Equation (B.2) holds for each bank type h:

$$\delta_1 + \delta_2 + \delta_3 = 1, \epsilon_{1,j} + \epsilon_{2,j} + \epsilon_{3,j} = 0 \text{ for } j = 1,2,3, \text{ and } \epsilon_{k,1} + \epsilon_{k,2} + \epsilon_{k,3} = 0 \text{ for } k = 4,\dots,K \qquad \text{(B.2)}$$

The marginal costs of output category $j = l$ (of loans) for bank i of category h in year t, mc^h_{ilt} are defined as:

$$mc_{ilt}^h = \partial c_{it}^h / \partial x_{ilt} = \left(c_{it}^h / x_{ilt}\right) \partial \ln c_{it}^h / \partial \ln x_{ilt} \qquad \text{(B.3)}$$

The term $\partial \ln c^h_{it} / \partial \ln x_{ilt}$ is the first derivative of Equation (B.1) of costs to loans. We use the marginal costs of the output component 'loans' only (and not for the other K_1 components) as we investigate the loan markets. We estimate a separate translog cost function for each bank category (SOCB, JSCB, CCB and FOREIGN), allowing for differences in the production structure across bank types. This leads to the following equation of the marginal costs for output category loans (l) for bank i in category h during year t:

$$mc_{ilt}^h = c_{it}^h / x_{ilt} \left(\delta_{1h} + 2\epsilon_{1lh} \ln s_{ilt} + \Sigma_{k=1,\dots,K;k\neq l}\epsilon_{1kh} \ln x_{ikt}\right) d_i^h \qquad \text{(B.4)}$$

Table 11: Estimate translog cost functions by bank type

	SOCB		JSCB		CCB		FOREIGN	
Dependent variable: ln(costs)–ln(other expenses)								
Outputs								
ln(loans)	0.768**	(2.23)	1.332***	(5.15)	1.174***	(8.91)	1.759***	(17.22)
(ln(loans)) ²	–0.0743**	(–2.01)	–0.00285	(–0.07)	0.0595***	(4.11)	–0.0263**	(–2.41)
ln(securities)	0.265	(0.70)	–0.162	(–0.61)	–0.130	(–0.98)	0.0839	(0.96)
(ln(securities)) ²	0.0950***	(4.73)	0.0143	(0.53)	0.0486***	(5.24)	–0.0201***	(–3.81)
ln(other services)	0.945***	(4.76)	–0.411***	(–3.38)	0.142*	(1.82)	–0.0896	(–0.91)
(ln(other services)) ²	0.0144***	(4.21)	–0.00469	(–0.90)	0.00641*	(1.70)	–0.0371***	(–2.90)
Input prices								
ln(wage)–ln(other expenses)	2.907***	(4.78)	–0.698***	(–5.37)	0.352**	(2.04)	1.896***	(13.39)
ln(funding rate)–ln(other expenses)	0.739**	(2.15)	0.966***	(3.76)	–0.0135	(–0.08)	–1.179***	(–9.83)
(ln(wage) –ln(other expenses)) ²	–0.364***	(–8.82)	–0.00712	(–0.60)	0.0872***	(4.08)	0.111***	(5.81)
(ln(funding rate) –ln(other expenses)) ²	–0.0439***	(–3.11)	0.0937***	(3.79)	0.0539***	(2.69)	0.106***	(8.19)
Cross-products between input prices								
(ln(wage) –ln(other expenses))(ln(funding rate)–ln(other expenses))*	0.0831***	(2.82)	–0.0782***	(–3.00)	–0.128***	(–3.50)	–0.225***	(–7.45)
Cross-products between outputs								
*ln(loans) * ln(securities)*	–0.0247	(–0.47)	–0.0163	(–0.25)	–0.0947***	(–4.52)	–0.0467***	(–4.06)
*ln(loans) * ln(other services)*	–0.115***	(–5.40)	0.0454*	(1.91)	–0.0269**	(–2.17)	–0.00174	(–0.12)
*ln(securities) * ln(other services)*	–0.00459	(–0.31)	–0.0176	(–0.97)	0.0122	(0.96)	0.0810***	(5.53)
Cross-products between outputs and input prices								
ln(loans)(ln(funding rate)–ln(other expenses))*	–0.0784**	(–2.30)	–0.00700	(–0.15)	0.0481*	(1.93)	0.216***	(9.49)
ln(loans)(ln(wage)–ln(other expenses))*	–0.745***	(–10.57)	0.123***	(5.19)	0.0975***	(3.80)	–0.130***	(–4.88)
ln(securities)(ln(funding rate)–ln(other expenses))*	0.111***	(4.26)	0.0174	(0.46)	–0.0177	(–0.99)	0.0360**	(2.18)
ln(securities)(ln(wage)–ln(other expenses))*	0.472***	(12.95)	–0.0769***	(–3.06)	–0.0632***	(–3.21)	–0.0811***	(–3.98)
ln(other services)(ln(funding rate)–ln(other expenses))*	–0.0328**	(–2.34)	–0.0119	(–0.60)	0.0222**	(2.17)	–0.198***	(–9.34)
*ln(other services) *(ln(wage)–ln(other expenses))*	–0.126***	(–8.14)	–0.0134	(–0.86)	–0.00528	(–0.43)	0.144***	(5.56)
Control variables								
ln(equity/assets)	–2.490***	(–22.49)	0.105	(0.90)	–0.0254	(–0.13)	0.795***	(5.19)
(ln(eqauity/asset)) ²	–0.371***	(–22.37)	0.0256	(1.45)	0.00136	(0.04)	0.163***	(4.96)
Constant	–0.00271	(–0.86)	–0.0657***	(–3.16)	0.000664	(0.02)	1.03e–13	(0.00)
F	1760657.7		86663.1		18374.9		13849.3	
Adj-R^2	0.9997		0.9998		0.9990		0.9987	

C. Competition measures under interest rate regulations: Theoretical proofs

To understand the direct effect of binding deposit rate regulation on the Lerner index and RPD, we consider the simple model described in Section 4.2. Binding deposit rate regulation in China affects the level of marginal costs of all banks and redistributes market share between efficient and inefficient banks. We show below that this redistribution of output can result in both increasing and decreasing competition as indicated by the Lerner index, which makes it an inconsistent measure of competition under binding deposit rate regulation. On the other hand, RPD is continuous and monotone in competition in a market with binding deposit rate regulation. In the following exercise, we assume that the slope of the loan demand function does not change after exogenous movements in input prices. To keep it simple, we also assume that deposit rate regulation does not affect the number of banks operating in the market, e.g. we do not allow market exit and entry due to changes in deposit rate regulation.

Imposing a deposit rate ceiling should reduce the level of competition because more efficient banks cannot undercut less efficient rivals by setting deposit rates above the ceiling. Less efficient banks then are protected by the ceiling and are less likely to be forced out of the market. Abolishing or raising deposit ceilings should increase competition because more efficient banks can expand market share at the expense of their less efficient rivals.

We assume that deposit rate regulation has a homogeneous impact on each bank's marginal costs. Then, under regulation, a bank's marginal cost of loans becomes $mc_i(\varepsilon) = mc_i - \varepsilon$ ($i = 1,\dots, N$). ε is a regulation parameter, which measures the extent to which deposit rate regulation is binding. $\varepsilon \in (-\hat{\varepsilon}, mc)$, where mc is the marginal cost of the most efficient bank and $\hat{\varepsilon}$ is some positive number that allows the least efficient bank to remain profitable and stay in the market. A positive ε reflects a binding deposit rate ceiling, while a negative ε corresponds to a binding deposit rate floor. Higher values of ε lead to less competition. This parameter can be time-variant to reflect changes in regulation across time. We focus here on deposit rate regulation. Nevertheless, this general setup can also be applied to other regulations (or technology shocks) that impact homogenously upon the cost side of banks. From equations (7), (8) and (9), and imposing ε, we derive the effect of binding deposit rate regulation on optimal output:

$$\begin{aligned} f(\varepsilon) &= q_i(\varepsilon) - q_i \\ &= \frac{\left(\frac{2b}{d} - 1\right)\varepsilon}{\left(2b + d(N-1)\right)\left(\frac{2b}{d} - 1\right)} \end{aligned} \tag{C.1}$$

where q_i is the optimal output without deposit rate regulation. Given $0<d \leq b$, $f(\varepsilon)$ is increasing in ε and takes the same sign as ε. Hence, under a deposit rate ceiling (floor), each bank's optimal output increases (decreases) by the same amount. We write the Lerner index for bank i as a function of regulation-free optimal output, marginal costs and the regulation parameter ε:

$$\begin{aligned} L_i(\varepsilon) &= \frac{bq_i(\varepsilon)}{\left(bq_i(\varepsilon) + mc_i(\varepsilon)\right)} \\ &= \frac{b\left(q_i + f(\varepsilon)\right)}{\left(b\left(q_i + f(\varepsilon)\right) + mc_i - \varepsilon\right)} \end{aligned} \tag{C.2}$$

Taking the derivative with respect to ε and using $f'(\varepsilon) = f(\varepsilon)/\varepsilon$, we obtain:

$$\begin{aligned} sign\left(\frac{dL_i(\varepsilon)}{d\varepsilon}\right) &= sign\left(b\left(q_i + \frac{f(\varepsilon)}{\varepsilon} mc_i\right)\right) \\ &> 0 \end{aligned} \tag{C.3}$$

Hence, a higher value of ε increases an individual bank's Lerner index, indicating less competition, as theory would suggest. However, the aggregate Lerner index – i.e. for the whole market – might not give a consistent value because the market shares of efficient banks decrease due to deposit rate regulation. To see this, define the market share of bank i as $s_i(\varepsilon) = q_i(\varepsilon)/\Sigma_j q_j(\varepsilon)$, and define bank k as the bank that produces at market average marginal costs, namely $mc_k = \Sigma_j mc_j/N$. Market share under deposit rate regulation can then be written as:

$$s_i(\varepsilon) = \frac{1}{N}\frac{\left(\frac{2b}{d} - 1\right)a - \left(\frac{2b}{d} + N - 1\right)mc_i + \Sigma_j mc_j + \left(\frac{2b}{d} - 1\right)\varepsilon}{\left(\frac{2b}{d} - 1\right)a - \left(\frac{2b}{d} + N - 1\right)mc_k + \Sigma_j mc_j + \left(\frac{2b}{d} - 1\right)\varepsilon} \tag{C.4}$$

Taking the derivative with respect to ε yields:

$$sign\left(\frac{ds_i}{d\varepsilon}\right) = sign\left(\left(\frac{2b}{d} - 1\right)\left(\frac{2b}{d} + N - 1\right)\left(mc_i - mc_k\right)\right) \tag{C.5}$$

It is immediately clear that the market share of bank i increases with a higher ε if, and only if, $mc_i > mc_k$. Therefore, regulation reallocates market share from efficient banks to less efficient banks (eg banks with marginal costs above the market average). The effect of binding deposit rate regulation on the aggregate Lerner index is then:

$$\frac{dL}{d\varepsilon} = \sum_{i=1}^{k} \frac{ds_i}{d\varepsilon} L_i + \sum_{i=k+1}^{N} \frac{ds_i}{d\varepsilon} L_i + \sum_{i=1}^{N} s_i \frac{dL_i}{d\varepsilon} \tag{C.6}$$

Denote banks $i = 1,\dots, k$ as low-efficiency banks, which will see their market shares increase. In contrast, the market share of high-efficiency banks $i = k+1,\dots, N$ will decrease. All in all, this leaves the sign of $dL/d\varepsilon$ undetermined. Specifically, if deposit rate regulation reallocates sufficient market share from efficient to less efficient banks (resulting in $dL/d\varepsilon<0$), then competition such as measured by the Lerner index can increase instead of decrease. This simple example shows that the aggregate Lerner index cannot consistently measure competition under deposit rate regulation.[29]

In contrast, RPD is not biased due to interest rate regulation. As described in Section 4.2, RPD is defined as the ratio of the profit differences between any three banks in the market. Banks can be ordered by their efficiency level (marginal costs), with more efficient banks providing more loans. Suppose we take three banks – A, B, C – with $mc_A<mc_B<mc_C$, then RPD is defined as $RPD = (\pi_A - \pi_C)/(\pi_B - \pi_C)$. Using the model presented in Section 4.2, profits can be written as a quadratic function of outputs. Then, after imposing deposit rate regulation, $RPD(\varepsilon)$ can be rewritten as:

$$RPD(\varepsilon) = \frac{\left(q_A + f(\varepsilon)\right)^2 - \left(q_C + f(\varepsilon)\right)^2}{\left(q_B + f(\varepsilon)\right)^2 - \left(q_C + f(\varepsilon)\right)^2} \tag{C.7}$$

We show that $RPD\,(\varepsilon)$ is decreasing in ε by taking the first-order derivative:

$$sign\left(\frac{dRPD}{d\varepsilon}\right) = sign\left(\frac{2f'(\varepsilon)(q_B - q_A)}{\left(q_B + q_c + 2f(\varepsilon)\right)^2}\right) \tag{C.8}$$
$$< 0$$

Given that $q_B - q_A<0$ and $f'(\varepsilon)>0$, the above equation has a negative sign, suggesting that higher binding regulation (ie a higher ε) will lower competition, consistent with theory. Hence, RPD is a consistent measure of competition in case of binding deposit rate regulation.

We show below the two main problems with the (elasticity-adjusted) Lerner index when lending rate regulation is binding. First, this index mainly measures variation in

[29] Boone (2000) provides another example where an individual firm's Lerner index increases after competition intensified. Applying that model with a slight modification, it can be shown that the necessary condition for an individual bank's Lerner index to be increasing in ε is that the marginal cost of this bank is lower than the market average. Proof is available upon request.

competition resulting from changing regulation; it cannot detect competition resulting from shifts in demand. Second, ignoring binding price regulation leads to inconsistent estimates of the elasticity-adjusted Lerner index (see also Salvo, 2010).

Consider the simple case of a monopoly bank serving the entire market under a lending rate ceiling.[30] If this ceiling is not binding (see Panel A of Figure 2), the bank will choose the optimal price and quantity combination by equating marginal cost (*MC*) to marginal revenue (*MR*). When the demand curve shifts from *D1* to *D2* ($da > 0$),[31] the equilibrium combination of prices and output moves from point *E1* to *E2*, resulting in a higher Lerner index, or lower competition. Hence, changes in competition resulting from exogenous shifts in the demand curve can be correctly picked up by the Lerner index. However, this is not the case if the lending rate ceiling is binding, as demonstrated in Panel B. This ceiling (P_c) prevents the bank from choosing the optimal price-output combination according to $MR = MC$. In contrast, following profit-maximising behaviour, it will choose the quantity at the kink of the demand curve (points *E1* and *E2* of Panel B), leaving the price unchanged at the ceiling. Therefore, changes in competition due to exogenous shifts in demand cannot be indentified by the Lerner index, because both prices and costs do not change in relation to the change in demand.

In case both the demand curve and binding lending rate ceiling change, the Lerner index can pick up only variations in competition due to changes in the latter, but not those due to shifts in the former. In Panel B, suppose that the demand curve shifts to *D2* and the lending rate ceiling increases to P_c', both of which will decrease competition. The optimal combination of prices and output moves from *E1* to *E2'* and hence the Lerner index increases. It is immediately clear that changes in this index reflect only changes in the lending rate ceiling but not in the demand curve, because the new Lerner indices are the same with or without demand curve shifts (comparing *E2'* and *E1'*). All in all, in the case of a binding lending rate ceiling, the Lerner index provides only an incomplete assessment of changes in competition.

The above analysis also applies to the elasticity-adjusted Lerner index because it estimates the price-cost margin of an average bank. This conclusion is closely related to the analysis in Salvo (2010), which proves theoretically and empirically that ignoring price

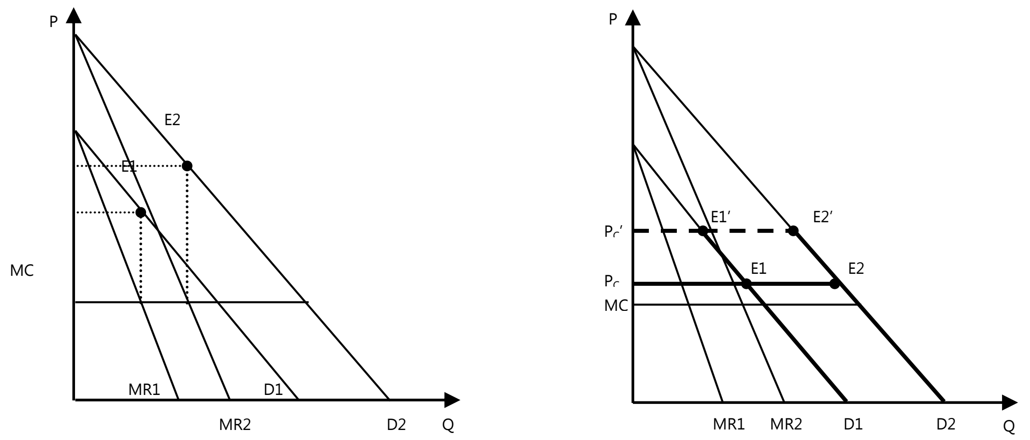
[30] Competition is a concept closely related to market power, and in most of the literature they are considered in a similar fashion. Even for a monopoly, the issue of market power is relevant. We use a monopoly here for reasons of simplicity. The example is also valid for a market with multiple firms. See Koetter et al. (2008) for more details.

[31] For a full proof that $da > 0$ leads to lower competition, please refer to Boone (2000).

ceilings result in an over-estimation of competition by the elasticity-adjusted Lerner index in the context of the Brazilian cement industry. When prices are unconstrained, the traditional joint estimation approach (eg Bresnahan, 1982) can effectively distinguish between monopoly and perfect competition, as demand shifts will lead to price changes in the case of a monopoly but not under perfect competition. In contrast, when prices are regulated (for example, a price ceiling is put in place), demand shifts do not affect prices in the case of both a monopoly and perfect competition.

Figure 2: Lerner index and price ceilings

Panel A Panel B



Thus, unless marginal costs are observed, one cannot tell whether the observed price-quantity combination is established under a monopoly or perfect competition. If one were to ignore the existence of a price ceiling and hence conclude that prices remain stable after a shift in demand, one would falsely reject collusion and argue in favor of competition. In general, if binding price ceilings are not properly accounted for, the underlying structural model will be misspecified. Hence, the orthogonality condition that is required for a consistent estimation of the related parameter will not be met. Salvo (2010) further shows that ignoring price ceilings may lead to an over-estimation of competition, in line with our argumentation.

Overall, we conclude that the (elasticity-adjusted) Lerner index is a biased measure of competition when price ceilings are binding. We suspect that this may account for the very high level of competition that it obtains for the pre-WTO period in China. It is generally acknowledged that the lending rate ceiling was most likely binding during this period.

In contrast, RPD uses relative profits and therefore they can pick up changes in competition due to demand shifts under price ceilings. For illustrative purposes, we use a simplifying assumption for the additional demand that may result from a binding price ceiling. Specifically, we assume that the extra output will be shared among banks according to their market share without the price ceiling. This so-called repartition rule relates to Schmalensee (1987). It should be noted that our proof does not depend on any specific repartition rule, as long as it allows more efficient banks to take on relatively more additional output after the price ceiling is imposed. For simplicity, we assume that $b=d$, meaning that the products provided by different banks are perfect substitutes. Denote aggregate loans that are provided under the price ceiling as $Q^* = \frac{a-\bar{P}}{b}$; without the ceiling, it is Q. If the ceiling is binding, $Q^* \geq Q$. Moreover, banks share the additional output Q^* - Q according to their original market share s_i when there is no price ceiling. Then the optimal output for bank i is:

$$q_i^* = q_i + s_i\left(Q^* - Q\right) = s_i Q^*, \text{ where } s_i = \frac{1}{N}\frac{a-\left(N+1\right)mc_i + \Sigma_j mc_j}{a-\left(N+1\right)mc_k + \Sigma_j mc_j} \tag{C.9}$$

Again mc_k are the marginal costs of producing loans for an average bank. Then profits of bank i are $\pi_i^* = \left(\bar{P} - mc_i\right)s_i Q^*$. We focus on the demand shift parameter a and prove that an increasing a leads to lower competition under the price ceiling when measured by RPD. We reiterate that in this case the Lerner index would not detect any changes in competition. The RPD of any three banks under price ceiling is:

$$RPD\left(a\right) = \frac{\left(\bar{P} - mc_A\right)s_A - \left(\bar{P} - mc_C\right)S_c}{\left(\bar{P} - mc_B\right)s_B - \left(\bar{P} - mc_C\right)S_c} \tag{C.10}$$

Taking the derivative with respect to a, and using $mc_A<mc_B<mc_C$, it can be shown that:

$$sign\left(\frac{dRPD}{da}\right) = \underset{<0}{sign\left(\left(N+1\right)\left(mc_C - mc_B\right)\left(mc_A - mc_C\right)\left(mc_B - mc_A\right)\right)} \tag{C.11}$$

Hence, RPD correctly picks up changes in competition due to demand shifts when a price ceiling is put in place. This is its main advantage (and of the PE indicator as well) when compared with the Lerner index. Both RPD and the PE indicator can measure competition correctly under price ceilings, while the Lerner index can only measure changes in competition resulting from changed ceilings, but not those resulting from shifts in demand.

Finally, the existence of binding interest rate regulations can exarcerbate other shortcomings of conventional competition measures such as the Lerner index. A case in point is the reallocation effect identified in Boone et al. (2007). This relates to the fact that more intensive competition due to more aggressive conduct will reallocate output and profits from less efficient banks to more efficient banks. As more efficient banks usually have higher PCMs than less efficient banks, the PCM for the whole market, which is an (output) weighted average of individual banks' PCMs, actually may increase in response to more intense competition. The increase in the market PCM (or aggregate Lerner index) would be interpreted as a decline in competition, while actually it has increased. Boone et al. (2007) show that the reallocation effect is particularly strong in concentrated markets. As a matter of fact, Chinese loan markets are highly concentrated markets, where during 2001-2008 the four SOCBs had an average annual market share of around 71%. It can be demonstrated that the reallocation effect is more profound when the regulation of interest rates is binding. Hence, this should make the Lerner index even less appropriate as an indicator to measure competition in Chinese loan markets.

D. Additional robustness tests

In this section, we present a number of tests to check the robustness of our results for alternative specifications and estimation methods. The robustness checks show that alternative definitions of competition indicators do not change our results significantly. Specifically, we test whether the main results are sensitive to: 1) alternative Lerner index (conjectural variation); 2) alternative definition of PE indicator; 3) alternative calculation of marginal costs.

D.1. Alternative Lerner index (conjectural variation)

In Section 6.1 we calculated the elasticity-adjusted Lerner index L by first estimating λ, i.e. the ratio of conjectural variation Θ to the elasticity of demand ε. Subsequently we could estimate L as λ/p, with p the average price of loans (average lending rate). An alternative approach is to estimate explicitly the conjectural variation Θ by simultaneously estimating the TCF (Equation 2), the supply equation (Equation 5) and an inverse loan demand function. Then the conjectural variation parameter Θ can serve as a direct measure of competition. In a perfectly competitive market, Θ_i equals to zero for all i, while for a monopoly it equals to one. This approach is adopted in Uchida and Tsutsui (2005) for Japanese banking market. Following this approach, we find that the estimated inverse demand elasticity is very stable across all years, which implies that conjectural variation follows a similar pattern to the evolution of the elasticity-adjusted Lerner index. Subsample estimations show that the conjectural variation is 0.068 and 0.087 for the pre-WTO respectively the post-WTO period, with the former being more competitive than the latter at a 1% significance level. We conclude that our main results obtained with the elasticity-adjusted Lerner index hold if conjectural variation is employed as a direct measure of competition. The full estimation process and results are not reported here to save space, but are available from the authors upon request.

D.2. Alternative definition of PE Indicator

We calculated the PE indicator by using the logarithm of π_{ilt} or profits obtained from loans as the dependent variable (see Section 6.2). This is a more accurate measure of profits generated by loan business. Alternatively, as a robustness check, we follow Boone et al. (2004) and use the logarithm of variable profits as the dependent variable. This approach has the advantage that it avoids potential estimation errors, as variable profits can be obtained directly from accounting data. At the same time, it has the disadvantage that variable profits capture not only profits from loans but also those from other activities. Variable profits are defined here as the difference between total income and the sum of interest expenses and other non-interest expenses.[32] We find that they are highly correlated with the definition of profits that we used in Section 6.2, with a Pearson correlation coefficient of 0.9607.

Similar to the other estimations, we estimate yearly and subsample PE indicators, which are reported in Panels B of Table 12 and Table 13. Again, competition follows the same pattern that we reported for the initial results. The structural break test for the point estimates for the two subsamples again supports our finding that the pre-WTO period is less competitive than the post-WTO period.

D.3. Calculation of marginal costs

For the (elasticity-adjusted) Lerner index, we assumed that the Translog Cost Function (TCF) for each bank group (SOCB, JSCB, CCB, FOREIGN) is the same, as only the constant term is allowed to vary across bank groups through bank type dummies (Equation 2). For the PE indicator, we improved the estimation by imposing different cost functions on different bank groups and estimated a separate TCF for each bank group. Both ways of treating cost functions for specific bank groups are generally accepted in the literature. Nevertheless, this difference could potentially generate different marginal costs. As for both the Lerner index and the PE indicator marginal cost estimations directly affect their values, it is important to test whether the contradictory results that we find could be driven by differences in the estimated marginal costs.

To this end, we conduct the following two robustness tests. First, we re-estimate the (elasticity-adjusted) Lerner indices assuming different cost functions for each bank group. Second, we re-estimate the PE indicator using the marginal costs that we estimated for the elasticity-adjusted Lerner index (MC_e), i.e. assuming similar translog cost functions for bank groups.

When re-estimating the (elasticity-adjusted) Lerner indices, we use different TCFs for each bank group by allowing for bank type dummies to interact with the independent variables in the

[32] An alternative definition of variable profits is *interest income - (interest expenses + other non-interest expenses)*. Our main conclusions are not sensitive to this alternative definition. Results are available upon request.

TCF. We calculate again yearly and subsample values, which are shown in Table 14. The modification in the TCF turns out to change the elasticity-adjusted Lerner index only very marginally for both the yearly and subsample estimations[33]. Moreover, the traditional Lerner index also resembles closely our previous results. This confirms that our previous findings are robust to different calculations of marginal costs.

The results for the re-estimation of the PE indicator using the marginal costs that we estimated in order to obtain the elasticity-adjusted Lerner index (MC_e) are shown in Panel A of Table 12 for the yearly results and of Table 13 for the subsample results. The former follows a very similar pattern to our previous results. Moreover, also our conclusion that the pre-WTO period had much lower competition than the post-WTO era remains intact.

[33] Underlying estimations of elasticity-adjusted Lerner index not shown to save space. Results are available upon request.

Table 12: Yearly estimates of alternative PE indicator

	Panel A: Independent variable ln(MCe)				Panel B: Dependent variable ln(variable profits)			
	PE Indicator	z-value	AP $chi^2(1)$ p-value	AP F(1,440)	PE Indicator	z-value	AP $chi^2(1)$ p-value	AP F(1,442)
1997	–2.314	(–1.53)	0.0000	18.08	6.656	(0.49)	0.4866	0.46
1998	–1.769	(–1.00)	0.0101	6.27	–2.183	(–1.26)	0.1021	2.53
1999	3.609	(1.3)	0.0000	33.08	–0.627	(–0.25)	0.0000	26.78
2000	–1.379	(–0.44)	0.0127	5.88	–0.667	(–0.37)	0.0000	17.13
2001	–5.748***	(–4.26)	0.0000	29.48	–3.086***	(–4.11)	0.0000	31.54
2002	–6.826**	(–2.20)	0.0009	10.46	–3.594***	(–2.64)	0.0000	20.2
2003	–3.754**	(–2.49)	0.0000	65.05	–4.391**	(–2.57)	0.0027	8.57
2004	–3.810**	(–2.25)	0.0000	72.28	–2.937***	(–3.13)	0.0000	58.35
2005	–1.605	(–1.41)	0.0000	95.18	–1.1	(–1.58)	0.0000	67.33
2006	–4.633***	(–2.87)	0.0001	14.46	–3.090***	(–3.28)	0.0000	20.59
2007	–3.669***	(–4.27)	0.0000	74.47	–3.264***	(–5.47)	0.0000	89.25
2008	–3.584***	(–3.93)	0.001	10.27	–1.959***	(–3.26)	0.0000	28.18
Constant	–2.511	(–1.00)			1.983	(1.13)		
Nr. Obs	464				466			
F	4.649				4.685			
Centered R^2	0.132				0.0961			

z-values in parenthesis; ** represent significance level of 5%, *** represent significance level of 1%; AP chi2 is the Angrist-Pischke (AP) first-stage chi-squared test: AP F is the Angrist-Pischke (AP) F-statistics. Test statistic can be compared to Stock-Yogo (2002, 2005) critical values for Cragg-Donald F statistic with K1=1. The Stock-Yogo weak ID test critical values at 10% maximal LIML size are 16.38. Year dummies are not reported here to save space.

Table 13: Subsample estimates of alternative PE indicator

	Panel A Independent variable ln(MC_e)			Panel B Dependent variable: ln(variable profits)		
	1996–2008	1996–2001	2002–2008	1996–2008	1996–2001	2002–2008
PE Indicator	–1.928***	–1.522	–3.717***	–2.023***	–1.487	–2.870***
	(–3.81)	(–1.01)	(–5.65)	(–5.66)	(–1.64)	(–7.07)
Time Trend	–0.0142	–0.508*	0.367***	0.0087	–0.461**	0.296***
	(–0.34)	(–1.67)	(4.9)	(0.24)	(–2.20)	(4.45)
Constant	1.069	4.889	–8.492***	0.516	4.540**	–5.236***
	(0.73)	(1.63)	(–3.67)	(0.46)	(2.34)	(–3.37)
H_0:prewto –postwto<=0 (p-value)		2.14* (0.071)			2.34* (0.063)	
Nr. Obs	464	91	373	466	91	375
F	7.226	1.815	21.24	16.01	2.349	29.25
Centered R^2	0.0247	0.0495	0.104	0.0691	0.101	0.141
K-P rk Wald F	336.7	77.97	163.2	227.9	34.97	142.4
K-P rk LM (p-value)	73.24 (0.000)	15 (0.000)	50.52 (0.000)	64.78 (0.000)	13.59 (0.000)	45.18 (0.000)

z-values in parenthesis; * represents significance level of 10%, ** represent significance level of 5%, *** represent significance level of 1%. K-P rk Wald F is the Kleibergen-Paap rk Wald F statistic. K-P rk LM is Kleibergen-Paap rk LM statistic. The Stock-Yogo weak ID test critical values at 10% maximal LIML size are 16.38

Table 14: Lerner indices assuming different TCFs for each bank type

	λ_t	Elasticity adjusted Lerner index	Lerner index	MC_e	MC_t
1997	0.077	0.458	0.330	0.080	0.104
1998	0.048	0.410	0.317	0.080	0.079
1999	0.030	0.294	0.244	0.062	0.066
2000	0.022	0.221	0.284	0.055	0.055
2001	0.020	0.237	0.228	0.051	0.055
2002	0.016	0.223	0.292	0.044	0.042
2003	0.017	0.240	0.298	0.046	0.045
2004	0.020	0.287	0.311	0.047	0.045
2005	0.023	0.288	0.330	0.051	0.050
2006	0.026	0.332	0.349	0.052	0.051
2007	0.036	0.457	0.417	0.047	0.047
2008	0.040	0.426	0.410	0.059	0.056
1996–2001	0.027	0.235	0.284	0.071	0.069
2002–2008	0.026	0.335	0.355	0.051	0.049

H_0: Elasticity Adj Lerner prewto>=Elasticity Adj Lerner postwto : $chi^2(1)=7.93$ p-value = 0.0024

λ_t are statistically different from zero for all year at a 1% significance level; MC_e and MC_t are average marginal cost estimated from elasticity-adjusted Lerner index and traditional Lerner index, respectively.

E. Panzar-Rosse H-statstic model

The so-called H-statistic developed by Panzar and Rosse has been employed in a small number of empirical studies on bank competition in China (Yuan, 2006; Fu, 2009).[34] The H-statistic is defined as the sum of the elasticities of a bank's total revenue with respect to that bank's input prices (Rosse and Panzar, 1977; Panzar and Rosse, 1987). Under monopoly, the revenues of the banks in question are independent of the decisions made by their actual or potential rivals. Panzar and Rosse proved that in this situation an increase in input prices will increase marginal costs, reduce equilibrium output and subsequently reduce revenues. Therefore, in this situation the H-statistic should be smaller than or equal to zero. In contrast, in the models of monopolistic competition and perfect competition, the revenue function of individual banks depends upon the decisions made by its actual or potential rivals (Bikker and Haaf, 2002). Under monopolistic competition, the change in input price is greater than the change in revenue and the H-statistic should lie between 0 and 1. Finally, under perfect competition, the H-statistic is equal to one because increases in input prices are passed on to output prices (in our case the lending rate). Higher input prices raise both marginal and average costs without, under certain assumptions, changing the optimal output of any individual bank. As some banks exit the market, the demand facing the remaining banks will increase, resulting in higher output prices and revenues equivalent to the rise in costs. Overall, a larger H-statistic indicates a higher degree of competition.

E.1. Recursive least squares

Following Bikker and Haaf (2002), we estimate the H-statistic based on the following revenue equation:

[34] Bikker et al. (2007) and Bikker and Spierdijk (2008) include China in Panzar-Rosse based investigations of bank competition in large country samples as well.

$$\ln\left(II_{it} / TA_{it}\right) = \alpha + \beta \ln\left(AFR_{it}\right) + \gamma \ln\left(PPE_{it}\right) + \delta \ln\left(PCE_{it}\right) + \eta_1 \ln\left(LNS\_TA_{it}\right) \\ + \eta_2 \ln\left(ONEA\_TA_{it}\right) + \eta_3 \ln\left(DPS\_F_{it}\right) + \eta_4 \ln\left(EQ\_TA_{it}\right) + \eta_5 OI\_II_{it} + \Sigma_{h=1 \ldots H-1} \varsigma_h d_i^h + error_{it} \tag{E.1}$$

The dependent variable $\ln(ll_{it} / TA_{it})$ is the logarithm of the ratio of interest income to total assets.[35] Hence, we employ the so-called *scaled* version of the Panzar-Rosse model, in order to be able to compare our results with those of Yuan (2006) and Fu (2009). We use the ratio of interest expenses to total funding as a proxy for the average funding rate (*AFR*). The ratio of personnel expenses to total assets is adopted as a proxy for the wage rate or price of personnel expenditure (*PPE*). Furthermore, the ratio of non-interest expenses to fixed assets is used as a proxy for the price of capital expenditure (*PCE*). The H-statistic, or the sum of the elasticities of a bank's total revenue with respect to that bank's input prices, is then defined as $H=\beta+\gamma+\delta$.

We follow the standard approach to include several bank specific variables as control variables to capture bank differences in risk, size and business structure. As the H-statistic assesses market structure by evaluating the relationship between costs and revenues, bank-specific characteristics need to be controlled for. We take the following variables into account: The ratio of loans to total assets (*LNS_TA*); the ratio of other non-earning assets to total assets (*ONEA_TA*) reflects the composition of assets; the ratio of customer deposits to the sum of customer deposits and short-term funding (*DPS_F*) captures the features of the funding mix; the ratio of equity to total assets (*EQ_TA*) is employed to reflect risk; the ratio of other income to interest income (*OI_II*) proxies the specific business structure. The variable d_i^h is the bank type dummy. As we have four types of banks in our sample (SOCB, JSCB, CCB and FOREIGN), we drop the *CCB* dummy to avoid over identification. The respective data are summarised in Table 3.

[35] Bikker et al. (2007) and Bikker et al. (2012) demonstrate that taking interest income as share of total assets, or the inclusion of scale variables as explanatory variables, may lead to overestimate competition and distorted tests results. Instead, they suggest using unscaled variables, ie using interest income, as the dependent variable. However, we use the scaled version of the H-statistic in order to be able to compare our results with those of Yuan (2006) and Fu (2009). As a robustness check, we also have estimated unscaled H-statistic. For more details see Appendix E (E.2).

The coefficient for *LNS_TA* is expected to be positive, as more lending potentially generates more interest income. The coefficient for *ONEA_TA* may be negative, as a higher ratio may be associated with lower interest income. *OI_II* is likely to have a negative coefficient, because generating other income might come at the expenses of interest income. For the signs of the coefficients for the other control variables, no prior expectations are offered by theory.

An important limitation of the H-statistic is that the market must be in long-run equilibrium, ie the return on total assets (*ROA*) should not be significantly correlated with input prices. The underlying motivation is that competitive markets will equalise the risk-adjusted rates of return across firms to such an extent that, in equilibrium, their correlation with input prices will be zero (Gutiérrez de Rozas, 2007). As is standard in the literature, we test the long-run equilibrium condition based on a regression in which the dependent variable is *ln(ROA)*, while the independent variables are the same as in Equation (E.1):

$$\begin{aligned}&\ln\left(ROA_{it}\right)=\alpha+\beta\ln\left(AFR_{it}\right)+\gamma\ln\left(PPE_{it}\right)+\delta\ln\left(PCE_{it}\right)+\eta_1\ln\left(LNS\_TA_{it}\right)\\&+\eta_2\ln\left(ONEA\_TA_{it}\right)+\eta_3\ln\left(DPS\_F_{it}\right)+\eta_4\ln\left(EQ\_TA_{it}\right)+\eta_5 OI\_II_{it}+\Sigma_{h=1\ldots H-1}\varsigma_h d_i^h+error_{it}\end{aligned}\tag{E.2}$$

where *ROA* is defined as net income over total assets. With this specification, $E=\beta+\gamma+\delta=0$ indicates long-run equilibrium, while $E<0$ represents disequilibrium.

Estimations are carried out with recursive least squares.[36] This approach does not impose any parametric structure on the evolution of the H-statistic and has the advantage of allowing for the assessment of bank competition for various time windows in our sample. We do not employ the commonly used yearly estimation of the H-statistic, as in Fu (2009) and Yuan (2006), because the test statistics based on a small number of banks in the early years of our sample might not be reliable. Another advantage of recursive least squares is that this approach can avoid the erratic pattern of the H-statistic which is often obtained with yearly

[36] Bikker and Spierdijk (2008) employ a parametric approach by incorporating time variant coefficients in the revenue equation. We use this approach as one of the robustness tests in Appendix E (E.3). We also tested 3-year rolling-window regressions and found similar results to recursive least squares. Results are available upon request.

estimations (Bikker and Spierdijk, 2008). We estimate Equation (E.1) recursively, starting with a window of two years and expanding the sample by one year at a time. In total we obtain 12 estimation windows. The results are summarised in Table 15, Panel A. To ensure standard errors and statistics that are robust to arbitrary heteroskedasticity and autocorrelation, kernel-based heteroskedastic and autocorrelation consistent (HAC) variance estimations are employed. The long-run equilibrium condition tests are provided for each time window, which are summarised in Panel B of Table 15. To save space, the coefficients of the control variables are not reported. Nevertheless, the signs of the coefficients of the control variables confirm our prior expectations.

The estimated H-statistic show a slightly increasing level of bank competition for the early time windows, but with an increasing time span, bank competition generally follows a declining pattern. This result is rather similar to those obtained by Yuan (2006) and Fu (2009). However, it should be noted that the differences between the H-statistic across all time windows are not statistically different. Wald F-tests on the sum of the input price elasticities reject both H=1 (perfect competition) and H=0 (monopoly), indicating that all time windows can be characterised by monopolistic competition. Long-run equilibrium condition tests are rejected for all time windows except for one.

To assess whether bank competition experienced structural changes, we estimate H-statistic for the whole sample and two subsamples. The break year for the subsamples is 2001, the year of WTO accession, resulting in the pre-WTO period 1996–2001 and post-WTO period 2002–2008.[37] The results for the H-statistic are reported in Table 16, while the long-run market equilibrium condition tests for the whole sample and sub-samples are reported in Table 17. The H-statistics for each sub-sample and for the whole sample again suggest that Chinese banking markets were in a state of monopolistic competition. When comparing the H-statistic for each subsample, we cannot reject the null hypothesis that they are equal across the

[37] The selection of 2001 as break year in the dataset is supported by formal structural break tests.

subsamples for any conventional significance level, suggesting no significant structural change. Table 17 shows that the long-run equilibrium condition (E=0) is rejected for the whole sample period and both subsample periods. This is likely to be related to the ongoing process of financial reform in China, which makes it unlikely that the banks have fully adjusted to market conditions. Hence, inferring competitive conditions from these results for China are likely to be biased.[38]

E.2. Unscaled Panzar-Rosse H-statistic

In our estimation of the Panzar-Rosse H-statistic (Appendix E.1), we used the scaled approach, i.e. the logarithm of interest income to total assets as the dependent variable, in order to be able to compare our results with those of Yuan (2006) and Fu (2009). However, we know from the literature that this approach is biased. Bikker et al. (2007) and Bikker et al. (2012) demonstrate that taking interest income as a share of total assets as the dependent variable, instead of the absolute amount of interest income (unscaled version), overestimates the degree of competition. In addition, when using this specification, results indicating both a monopoly and a situation of perfect competition will be distorted. The inclusion of scale variables as explanatory variables in the revenue function has a similar distorting effect.

As a sensitivity test, we estimate an unscaled version of the H-statistic using ln(interest income) as dependent variable. The results show even a more pronounced different pattern before and after China joined the WTO: The H-statistic indicate that Chinese loan markets were characterised by perfect competition before WTO accession and moved to monopolistic competition afterwards. Yuan (2006) and Fu (2009) reached similar conclusions, although with the scaled approach. Hence, the results of the theoretically better founded unscaled version of the Panzar-Rosse model show that Chinese loan markets were already in

[38] To test for monopolistic or perfect competition, it is necessary for the observations to be generated in long-run equilibrium (Panzar-Rosse, 1987). This equilibrium may not have been achieved yet in transitional economies, doubting its usefulness to assess competition in these markets (Shaffer, 1994; Northcott, 2004).

a state of perfect competition before further important financial reforms were implemented in the context of WTO accession in 2001 and that since then competition only declined. We hold the view that applying the more preferable unscaled version actually reinforces the shortcomings of the H-statistic for China (results are avaible upon request).

E.3. Parametric approach

Bikker and Spierdijk (2008) employed a parametric approach by incorporating time variant coefficients in the revenue equation, which allows for formally testing the evolution of bank competition over time. As a robustness test, we also estimated the H-statistic assuming a parametric structure of the evolution of competition, with the following specification:

$$\begin{aligned}\ln\left(II_{it}/TA_{it}\right) &= \alpha + \left(\beta\ln\left(AFR_{it}\right) + \gamma\ln\left(PPE_{it}\right) + \delta\ln\left(PCE_{it}\right)\right)^{*}\exp\left(\zeta^{*}t\right) + \eta_1\ln\left(LNS\_TA_{it}\right) \\ &+\eta_2\ln\left(ONEA\_TA_{it}\right) + \eta_3\ln\left(DPS\_F_{it}\right) + \eta_4\ln\left(EQ\_TA_{it}\right) + \eta_5 OI\_II_{it} + \Sigma_{h=1\ldots H-1}\varsigma_h d_i^h + error_{it}\end{aligned} \qquad (E.3)$$

where t is time, and the H-statistic is defined as $H_t=(\beta+\gamma+\delta)\exp(\zeta^{*}t)$. With this specification, if $\zeta=0$, the competitive structure is constant over time, while $\zeta>0$ ($\zeta<0$) indicates an increasing (decreasing) level of competitiveness over time. Estimation is carried out with nonlinear least square. Our results show a significantly negative time coefficient ζ of –0.0041 (p-value 0.0000), suggesting an annual decrease in the level of competition for the whole sample period. Wald F-tests on the sum of the input price elasticities reject the H-statistic being 1 (perfect competition) and 0 (monopoly) at a 1% significance level, indicating that all years could be characterised by monopolistic competition. Furthermore, a Wald F-test on the long-run equilibrium condition rejects E=0 at a 1% significance level for each year which suggests that Chinese loan markets were in disequilibrium. These results confirm that our results for the H-statistic are not sensitive to specific estimation methods. Results are available upon request.

Feyzioğlu et al. (2009) and Bikker et al. (2007) indicate that the H-statistic probably picks up the co-movement of regulated deposit and lending rates in China. So, instead of measuring the degree of pass-through of input prices to output prices that would measure the degree of competition in a liberalised market, it measures the degree in which the regulator sets deposit and lending rates jointly. The H-statistic may be biased upwards due to the high correlation between the ceilings on deposit and loan rates, which may have been especially relevant for the earlier sample years when interest rate deregulation had hardly started. The high values of the H-statistic for the pre-WTO period reported in previous studies (Yuan, 2006; Fu, 2009) and in our own estimates in the previous section likely are driven by this effect. The ceiling on the lending rate was abolished in 2004, which may have reduced the impact of this bias in subsequent years. This conclusion is supported by the findings reported in Table 16, where the coefficient of the average funding rate (*AFR*) is much higher in the pre-WTO period, while dropping considerably later on when the lending rate ceiling was abolished.

To conclude, using similar specifications as Yuan (2006) and Fu (2009), we find that the market structure indicated by our results is that of monopolistic competition. Moreover, the level of competition does not change significantly across time. Finally, it should be noted that the long-run equilibrium condition underlying the Panzar-Rosse model generally is not satisfied.

Table 15: H-statistic and long run equilibrium condition: Recursive least squares

Panel A: H-Statistic

	ln(AFR)	ln(PPE)	ln(PCE)	H	H_0: H=1 chi^2(1)	H_0: H=0 chi^2(1)	Nr. obs	F	Adj R^2
1996–1997	0.717***	0.0736**	–0.0647*	0.7254	7.77***	54.24***	25	113.09	0.836
1996–1998	0.778***	0.0652**	–0.0588*	0.7840	7.48***	98.52***	39	66.13	0.864
1996–1999	0.715***	0.0771**	–0.0493	0.7424	20.64***	171.51***	60	64.37	0.821
1996–2000	0.689***	0.0828***	–0.026	0.7461	26.48***	228.61***	84	92.34	0.852
1996–2001	0.650***	0.0986***	–0.024	0.7246	38.79***	268.57***	112	74.17	0.852
1996–2002	0.550***	0.113***	0.00124	0.6642	57.99***	226.82***	144	44.51	0.858
1996–2003	0.535***	0.136***	0.0113	0.6818	53.43***	245.21***	184	51.39	0.837
1996–2004	0.517***	0.129***	0.0303	0.6757	52.01***	225.9***	223	51.58	0.826
1996–2005	0.512***	0.134***	0.0164	0.6627	60.64***	234.14***	277	62.03	0.823
1996–2006	0.507***	0.120***	0.0097	0.6364	81.79***	250.59***	350	74.11	0.799
1996–2007	0.522***	0.131***	0.0121	0.6643	74.54***	291.9***	432	86.4	0.795
1996–2008	0.532***	0.126***	0.0183	0.6765	82.5***	360.86***	493	96.5	0.777

Panel B: Long-run equilibrium condition test

	ln(AFR)	ln(PPE)	ln(PCE)	H	H_0: E=1 chi^2(1)	Equilibrium	Nr. obs	F	Adj R^2
1996–1997	–0.0189	–0.0589	–0.0948	–0.1726	0.31	A	24	13.17	0.528
1996–1998	1.186***	–0.164	–0.163	0.8590	6.14**	R	38	16.72	0.585
1996–1999	0.852***	–0.121	–0.128	0.6026	9.71***	R	59	9.904	0.364
1996–2000	0.795***	–0.0983	–0.0735	0.6232	15.83***	R	83	11.82	0.389
1996–2001	0.566***	–0.0499	0.0414	0.5573	15.83***	R	111	8.406	0.345
1996–2002	0.341***	0.00301	0.112	0.4556	10.92***	R	141	7.702	0.307
1996–2003	0.362***	–0.0416	0.0621	0.3825	7.25***	R	181	6.391	0.263
1996–2004	0.311***	–0.0413	0.0174	0.2871	4.00**	R	219	4.969	0.203
1996–2005	0.283***	–0.0695	0.0494	0.2625	3.9**	R	273	5.17	0.167
1996–2006	0.235***	–0.0917	0.0616	0.2049	2.74*	R	345	5.86	0.145
1996–2007	0.267***	–0.0789	0.131**	0.3193	7.34***	R	427	7.847	0.167
1996–2008	0.286***	–0.05	0.155**	0.3907	14.9***	R	486	9.661	0.182

* represents significance level of 10%, ** represent significance level of 5%, *** represent significance level of 1%. A and R represent "Accepting" and "Rejecting" the null hypothesis that E=0 (equilibrium) at a 10% significance level.

Table 16: H-statistic point estimates: Whole sample and subsamples

	1996–2008		1996–2001		2002–2008	
ln(AFR)	0.532***	(24.63)	0.650***	(20.09)	0.537***	(21.81)
ln(PPE)	0.126***	(4.89)	0.0986***	(3.85)	0.145***	(4.03)
ln(PCE)	0.0183	(1.27)	–0.024	(–0.96)	0.0149	(0.92)
lnLNS_TA	0.0920*	(1.68)	0.0293	(0.60)	0.0905	(1.41)
LnONEA_TA	–0.0191***	(–3.75)	–0.0545***	(–4.31)	–0.0140***	(–2.62)
lnDPS_F	0.117***	(2.61)	–0.0378	(–1.03)	0.179***	(3.88)
lnEQ_TA	0.0846***	(3.62)	0.120***	(3.55)	0.0841***	(3.08)
lnOI_II	–0.0737***	(–9.56)	–0.0736***	(–5.56)	–0.0760***	(–8.48)
SOCB	–0.0779***	(–2.83)	–0.100**	(–1.98)	–0.0485	(–1.59)
JSCB	–0.0137	(–0.58)	0.0870**	(2.10)	–0.0595*	(–1.94)
FOREIGN	–0.204***	(–4.29)	–0.402***	(–3.59)	–0.163***	(–3.34)
Constant	–0.361**	(–2.23)	–0.252	(–1.30)	–0.207	(–0.99)
H-statistic	0.6765		0.7246		0.6974	
H_0: $H=0$ $chi^2(1)$	360.86***		268.57***		226.37***	
H_0: $H=1$ $chi^2(1)$	82.50***		38.79***		42.63***	
$H_{prewto}=H_{postwto}$	$chi^2(1)=0.22$ p–value=0.6357					
Nr. Obs	493		112		381	
F	96.50***		74.17***		83.00***	
Adj R^2	0.777		0.852		0.768	

z-values in parenthesis; * represents significance level of 10%, ** represents significance level of 5%,

*** represents significance level of 1%

Table 17: Long-run equilibrium condition: Whole sample and subsamples

	1996–2008		1996–2001		2002–2008	
ln(AFR)	0.286***	(3.42)	0.566***	(3.25)	0.341***	(3.70)
ln(PPE)	–0.05	(–0.61)	–0.0499	(–0.36)	–0.0621	(–0.62)
ln(PCE)	0.155**	(2.57)	0.0414	(0.38)	0.119**	(1.98)
lnLNS_TA	–0.137	(–0.76)	–0.875*	(–1.67)	–0.00647	(–0.03)
LnONEA_TA	–0.111***	(–5.35)	–0.165***	(–2.83)	–0.107***	(–4.98)
lnDPS_F	0.142	(1.01)	–0.113	(–0.77)	0.403***	(3.07)
lnEQ_TA	0.355***	(4.25)	0.473***	(2.65)	0.355***	(3.77)
lnOI_II	–0.00345	(–0.13)	–0.0633	(–1.47)	–0.00194	(–0.07)
SOCB	–0.211*	(–1.76)	–0.614**	(–2.41)	0.0405	(0.36)
JSCB	–0.0959	(–1.08)	0.0739	(0.43)	–0.234**	(–1.98)
FOREIGN	–0.486***	(–2.64)	–1.676***	(–2.82)	–0.21	(–1.41)
Constant	–3.155***	(–5.62)	–2.802***	(–2.94)	–2.880***	(–4.24)
H_0: $E=0$ $chi^2(1)$	14.90***		15.83***		10.22***	
Nr obs	486		111		375	
F	9.661***		8.406***		8.950***	
Adj R^2	0.182		0.345		0.208	

z-values in parenthesis; * represents significance level of 10%, ** represents significance level of 5%, *** represents significance level of 1%.

Do banks extract informational rents through collateral?[1]

Bing Xu
University Carlos III of Madrid

Adrian van Rixtel[2]
Bank of Spain

Honglin Wang
Hong Kong Monetary Authority

March 2017

Abstract

The use of collateral is one of the defining characteristics of loan contracts. This paper investigates if relationship lending and market concentration permit the extraction of informational rents through collateral. We apply equity IPOs as informational shocks that erode rent seeking opportunities. Using unique loan-level data for China, we find that collateral incidence increases with relationship intensity and bank market concentration for loans obtained before the IPO, while this effect is moderated after the IPO. We also demonstrate that the degree of rent extraction declines for less risky firms after the IPO; at the same time, it increases for more risky firms. These results are not driven by differences or changes in firm-specific financial risks. To our knowledge, our paper is the first to investigate the determinants of collateral for China using loan-level data.

JEL Classification: G21; L11.

Key Words: Informational rents; collateral; relationship lending; market structure; IPOs; China**.**

[1] We thank Hongyi Chen, Michael Chui, Ben Cohen, Dong He, Anil Jain, Chen Lin, Yue Ma, Jun Qian, Hao Zhou, Hong Zhang and participants at the 2015 China conference of the Hong Kong Institute for Monetary Research (HKIMR), the 7th IFABS international conference, the XX LACEA Annual Meeting and seminars at the Bank of Finland, Bank for International Settlements (BIS), Fudan University, HKIMR, People's Bank of China (PBC) (Shanghai), Bank of Spain, and the PBC School of Finance at Tsinghua University for very constructive comments and helpful suggestions. All remaining errors are ours. The views expressed in this paper are the authors' and do not necessarily reflect those of the BIS, Bank of Spain and Hong Kong Monetary Authority.

[2] Adrian van Rixtel was at the time of writing in the Financial Markets Group, Monetary and Economic Department, BIS.

1. Introduction

Information asymmetries among lenders may be used to "hold up" borrowers (Sharpe, 1990; Rajan, 1992). Banks obtain private inside information about their customers through lending, giving inside banks an informational advantage relative to outside banks (Santos and Winton, 2008). Adverse selection problems facing outside banks make it difficult for borrowers to switch lenders. Hence, inside banks are in a position to request harsher loan conditions than would prevail were all banks symmetrically informed, allowing them to extract informational rents. Empirical validations of informational rent extraction mainly have focused on lending rates (see e.g. Hale and Santos, 2009; Schenone, 2010), while rent extraction operating through non-price terms, such as collateral requirements, has been left largely unexplored.

In this paper, we intend to fill this gap by examining if inside information, obtained through both relationship lending and concentrated market structures, allows banks to extract informational rents through collateral. In doing so, we employ the equity IPOs of borrowers as information releasing shocks that erode information based rent-seeking opportunities. Using a unique hand-collected loan level dataset from China, our evidence suggests proprietary information does allow rent extraction through collateral.

The crucial precondition for information rent extraction is the existence of information asymmetries among lenders. We focus on two sources that create such information asymmetries: relationship lending and market structure. Banks accumulate proprietary information about borrowers through lending relationships (e.g. Sharpe, 1990; Rajan, 1992). In addition, concentrated bank market structures may facilitate information asymmetries among lenders as well (e.g. Dell'Ariccia et al., 1999; Dell'Ariccia, 2001).

Our identification strategy is very intuitive: informational rent extraction through collateral should be moderated after some exogenous shock that reduces information asymmetries existing between inside and outside banks. In other words, the degree of collateralization should decrease for the informed banks after the information releasing shock. Equity IPOs of borrowing firms present an ideal case of an information-releasing shock (see e.g. Schenone, 2010).[3] In the course of the public offering and after being listed, previously privately-held information about the firm will be released through compulsory listing requirements and subsequent regular financial reporting, public auditing, financial analysts' research and movements in its stock price. As this new information is made public to all banks, the informational monopoly position of inside banks is eroded and the adverse selection problem facing outside banks is alleviated, making rent extraction less likely for loans granted after the IPO than for loans granted before the IPO. Furthermore, we postulate that once the IPO has reduced information asymmetries among lenders, rent extraction will decline for safer firms, but not for risky ones. This because outside banks will be less inclined to lend once the borrower is revealed as risky, leaving inside banks in a better position to charge rents (see e.g. Rajan, 1992).

One crucial part of the methodology is to control for shifts in firm risk around the IPO and for differences in risks between listed and unlisted firms, so that changes in the degree of loan collateralization can be attributed to changes in information asymmetries instead of differences in credit risk. We control for this by introducing a wide range of firm risk characteristics both before and after the IPO, and later perform additional robustness tests.

We test our hypotheses on a unique hand-collected data set with information on individual loans from China. The unique settings of Chinese banking markets and the existing public listing procedure make loan markets in China an ideal case for our purposes (section 2.2 presents more detail). Our sample is composed of loans borrowed by firms listed at the Shenzhen Stock Exchange, both before and after their IPO. Focusing on this sample will bias *against* finding informational rent extraction, as one would expect that the "hold-up" problem is particularly pronounced for smaller firms, while our sample exists of large and relatively transparent firms. Our loan level dataset contains data on around 9,000 loans granted by a differentiated group of Chinese banks to 649 listed non-

[3] A similar approach has been followed by Santos and Winton (2008) and Hale and Santos (2009) using corporate bond IPOs as informational equalization shocks. These papers together with Schenone (2010) investigate informational rent extraction through lending rates.

financial firms.

Our main results can be summarized as follows. First, all else equal, both high relationship intensity and concentrated banking market structures are associated with a higher incidence of collateral, and these effects are less pronounced for transparent firms. We further find that there exists a boundary transparency level beyond which informational rent extraction becomes less feasible. Second, when applying equity IPOs as an informational shock, we find for pre-IPO originated loans that the likelihood of collateralization is increasing with relationship intensity, while this effect is greatly moderated for post-IPO loans. In some specifications, relationship intensity is no longer significant in predicting collateral incidence for loans originated after the IPO. Third, the likelihood of collateral incidence increases with the degree of banking market concentration both before and after the equity IPO, but the effect is moderated for post-IPO loans. This finding supports the hypothesis that concentrated banking markets facilitate the existence of information asymmetries among lenders and hence are associated with a higher likelihood of rent extraction through collateral. Unlike relationship intensity, the impact of market structure on collateral remains significantly positive and economically large for post-IPO loans. This lends some support to the idea that pure market power stemming from concentrated market structures may allow banks to charge rents, regardless of the level of information asymmetries existing among banks (Hainz, 2003; Berlin and Butler, 2002). Fourth, using a novel measure of firm risk – whether a firm's first IPO application was rejected by the Chinese market regulator (China Securities Regulatory Commission or CSRC) or not – we find that once information about firm risk is made public after the IPO, rent extraction through collateral is moderated for safe firms, but intensified for risky firms. This result is in line with the theoretical prediction of Rajan (1992) that informed banks are more able to extract rents from risky firms than from safer ones. Our finding further complements Hale and Santos (2009) who report similar results using lending rates as the rent extraction mechanism.

Overall, our findings are largely consistent with the informational rent extraction hypothesis, but with two important caveats. First, our results may be explained by alternative theories. Second, both listing status and relationship lending could be endogenous, which could bias our results.

Regarding the first caveat, we contrast the informational rent hypothesis with three alternative explanations. Firstly, various theories suggest that relationship lenders require less collateral for financially healthier firms (Dewatripont and Maskin, 1995; Longhofer and Santos, 2000). If listed firms are financially sounder than non-listed firms and our analysis has not fully controlled for this difference, the moderated effect of relationship lending on collateral incidence for post-IPO loans could also be explained by these theories. We show that our findings are not driven by the risk differential between listed and unlisted firms by comparing observed risk proxies and by employing propensity score matching to re-estimate our baseline model with a sample that is matched by the listing status of borrowers.

The second alternative explanation is related to heterogeneous risk dynamics around the IPO for relationship dependent and non-dependent firms. If “relationship dependent” firms improve their credit qualities more (or deteriorate less) than “relationship non-dependent” firms after the listing, then this pattern could explain the moderated effect of relationship lending on collateral incidence for post-IPO loans. To address this concern, we perform difference-in-difference tests for observed risk proxies. We do not find that the dynamics of risk proxies around IPOs differ according to relationship dependence.

The third alternative explanation that we explore is that banks exchange better loan conditions (i.e. a lower likelihood of required collateral) for corporate bond underwriting business. To isolate this alternative explanation, we re-estimate the baseline model on samples of loans that were originated before the bond IPO. Also here, our results hold.

Regarding the second caveat, we employ recursive bivariate Probit models to test if the listing status and relationship dependency are endogenous and if our results change after controlling for the endogeneity of the respective variables. In both cases, we find appropriate instrumental variables, so that the identification does not rely solely on the non-linearity of the functional form. Again, our main results remain valid.

Finally, we investigate if our results are robust to the inclusion of firm fixed effects, the endogeneity of other loan contract terms and to alternative samples.[4] These tests do not change our results.

Our study complements previous studies such as Schenone (2010) and Hale and Santos (2009) that banks price their information monopolies. Our findings suggest that private inside information allows informed banks to charge rents through collateral, a channel that has not been explored before. Understanding this channel is of particular relevance for countries with less developed financial markets, where the pricing of credit risk is generally more difficult and hence charging collateral is a particularly important mechanism to reduce the risk of debt. Furthermore, we provide additional insights to Rajan (1992) and Hale and Santos (2009), supporting their finding that rent extraction is more severe for risky firms. From a methodological perspective, our paper is, to the best of our knowledge, the first to apply equity IPOs as the identification strategy to test if banks charge informational rents through collateral.

Our findings contribute to the discussion on the role of collateral in bank loan contracts.[5] We also contribute to the literature on the structure of banking markets and information asymmetries among lenders (e.g. Dell'Ariccia, 2001; Marquez, 2002; Dell'Ariccia and Marquez, 2006; Hauswald and Marquez, 2006). Our empirical evidence shows that market structure is an important source of information asymmetries, extending previous understandings on how banking market structure affects loan conditions. Our findings may also provide insights about the functioning of other credit markets characterized by asymmetrically informed lenders. For example, lenders providing trade credit are generally more informed about buyers than other lenders and hence may be able to exploit

[4] In a set of unreported robustness tests, we further investigate if our results hold when using alternative relationship lending measures and controlling for regional legal and institutional differences that potentially may determine the likelihood of collateral incidence. These tests do not change our results.

[5] Literature identified three main important roles played by collateral in a loan contract: it mitigates ex-post borrower moral hazard problems (e.g. Boot et al., 1991; Brick and Palia, 2007; Berger et al., 2011; Cerqueiro et al., 2016); signals credit quality of the borrower and mitigates adverse selection problems (e.g. Bester, 1985; Besanko and Thakor, 1987; Berger et al., 2011; Jiménez et al., 2006); and minimizes expected loan losses given a borrower's default (Berger and Udell, 1990).

informational advantages (Smith, 1987). Our findings also contribute to the literature on bank lending markets in emerging markets in general and in China in particular. To our knowledge, the paper is among the first to investigate collateral incidence in Chinese bank lending markets using loan-level data.[6]

The remainder of this paper is organized as follows. Section 2 details our methodology and data. Section 3 presents the main empirical results. Section 4 compares our conclusions with alternative theories. Section 5 controls for endogeneity problems related to IPOs and relationship lending. Section 6 reports the results of further robustness tests. Finally, Section 7 concludes. Additional results are reported in an Internet Appendix to this paper.

2. Methodology and data

2.1. Methodology

The methodology of the main analysis contains four parts. First, we investigate if the likelihood of collateral increases with relationship lending and market concentration, after controlling for a broad range of other determinants. The second part attempts to find evidence that the increasing likelihood of collateral is at least partially due to information asymmetries between inside and outside banks. To this end, we test if the effects of relationship lending and market concentration on collateral are less pronounced for transparent firms, using various information transparency proxies. The third part investigates if informational rent extraction is moderated for post-IPOs loans relative to pre-IPOs loans. Finally, we investigate if this moderated effect for post-IPOs loans varies with firm risk. We discuss the methodologies related to alternative explanations, the possible endogeneity of key variables and further robustness tests in Sections 4, 5 and 6, respectively.

[6] Very few studies have investigated the determinants of collateral in China. Notable exceptions include Firth et al. (2012) and Chen et al. (2013). However, none of these studies investigates the determinants of collateral at the loan-level and pays attention to the importance of relationship lending and market structure for the incidence of collateral, as well as how changes in information asymmetries among lenders may affect these linkages.

2.1.1. Relationship lending and market structure as determinants of collateral incidence

We start by testing whether relationship lending and banking market structure are important determinants of collateral incidence. One strand of literature suggests that as relationships between borrower and lender intensify, relationship lenders accumulate inside information, which develops mutual trust and reduces the risk of moral hazard, allowing the inside bank to reduce collateral requirements. In essence, this "information accumulation view" considers relationship lending and collateral as substitutes, and therefore predicts a negative correlation between relationship intensity and collateral (e.g. Petersen and Rajan, 1995; Berger and Udell, 1995; Bharath et al., 2011). On the other hand, the proprietary information obtained through lending relationships can create adverse selection problems for outside banks, allowing inside banks to hold-up borrowers and charge harsher loan conditions (e.g. Sharpe, 1990; Rajan, 1992; Degryse and Van Cayseele, 2000). In other words, proprietary information obtained through relationship lending allows for informational rent extraction. This argument implies a positive correlation between relationship lending and collateral.

Besides information asymmetries, relationship lending affects collateral incidence also through other channels. For instance, Longhofer and Santos (2000) suggest that pledging collateral improves the seniority of a bank's debt claims, which incentivizes the bank to engage in ongoing, long-term, valuable lending relationships. Borrowers benefit from this, because bank seniority induces relationship lenders to provide support to distressed borrowers, as the senior debtors benefit the most from a turn-around of the firm.[7] Dewatripont and Maskin (1995) highlight another potential cost of relationship lending, which hinges on the observation that relationship lenders have an incentive to extend further credit in the hope of recovering loans granted previously when a borrower is in financial stress. Anticipating the ex-post realization of this "soft budget constraint", the borrower is not sufficiently incentivized to make an effort ex-ante to prevent such an adverse outcome. Collateral is therefore more likely to be requested when a bank-firm relationship intensifies to solve this soft budget constraint problem (Boot, 2000). Both theories suggest that, as borrower risk increases,

[7] See Elsas and Krahen (2000) for further discussion and empirical testing of this argument. Their results indicate that house banks require more collateral as compensation for their active involvement in the restructuring of distressed borrowers.

relationship lenders are more likely to request collateral, because the likelihood of engaging in a future rescue increases, in other words the soft budget constraint problem intensifies. Lastly, Menkhoff et al. (2006) suggest that banks may extend relationship length (intensity) to minimize the per unit fixed costs associated with evaluating and monitoring collateral ("cost minimization incentive"), which *de facto* produces a positive correlation between collateral and relationship duration (intensity). These theories imply that finding a positive correlation between relationship lending and collateral does not automatically confirm the validity of informational rent extraction.

In light of these discussions, we postulate the following hypothesis:

H.1: If relationship lending is negatively related to collateral incidence, the information accumulation view holds. In contrast, a positive correlation would reject this.

Banking market structure affects collateral incidence through at least two channels: the information channel and the market power channel. The first relates banking market structure to the information distribution among lenders, which in turn interact with banks' strategic behavior in determining lending policies and standards (e.g. Dell'Ariccia, 2001; Marquez, 2002; Dell'Ariccia and Marquez, 2006; Hauswald and Marquez, 2006).[8] We review the related literatures briefly in order to develop our hypothesis. First, information extraction is likely to be less effective in markets composed of many small banks instead of a few large banks (Marquez, 2002). Concentrated markets also allow for better protection of proprietary information, preventing spillovers to competitors, as banks with larger market shares have higher incentives and capacity to maintain this informational advantage. Therefore, concentrated lending markets not only consolidate market shares, but also protect proprietary information about borrowers. Second, different market structures associated with different implied levels of competition may also affect the incentive of banks to accumulate information.

[8] We restrict ourselves to theories that relate bank market structure to information asymmetry among lenders. Other theories (not crucially related to information asymmetry among lenders) also provide predictions. For instance, Manove et al. (2001) propose a "lazy bank" model in which banks choose between screening the borrower or ask for collateral. They argue that intensified competition would favor bank laziness by reducing screening and requesting more collateral. Hainz et al. (2013) propose that bank competition makes screening more effective. Hence, collateral – an alternative to screening – is less common in competitive markets. Inderst and Muller (2007) develop an inside lenders'–based model of collateral which does not assume the existence of information asymmetries on the borrower's side. These authors predict that the incidence of collateral is higher in more competitive markets.

Increased competition reduces the rent that banks can extract, reducing the incentive to generate information through credit evaluation (Hauswald and Marquez, 2006). More outside borrowing options for firms in less concentrated markets also inhibit the (re)usability of information and diminishes its value, as firms can switch banks easily, therefore banks are incentivized to invest less in information production (Boot and Thakor, 2010; Chan et al., 1986; Berlin and Mester, 1999).[9] Third, because of limited outside options, firms are likely to borrow more often from the same lenders in concentrated markets, which allow these banks to accumulate more private information. Lastly, the consolidation of proprietary information in concentrated markets deters the entry of new banks, as new entrant banks face larger adverse selection problems. Thus, information consolidation further increases the degree of market concentration and reinforces the information monopoly of incumbent banks (Dell'Ariccia et al., 1999; Dell'Ariccia, 2001). To sum up, these arguments suggest that concentrated markets allow for a more efficient extraction of private information and provide stronger incentives to obtain it; offer better protection from this information spilling over to competitors (outside banks); and deters competitors from entering the market which reinforces information monopolies. A straightforward implication is that concentrated markets may also facilitate informational rent extraction.

However, a positive association with collateral would not unequivocally suggest informational rent extraction. Sheer market power in concentrated markets could allow banks to request more collateral, independently from the imbedded information structure. This is the market power channel (e.g. Hainz, 2003; Berlin and Butler, 2002).

Following the arguments above, we postulate our second hypothesis:

H.2: Concentrated markets allow for a higher probability of collateral incidence, either because of the existence of informational monopolies, more market power or both.

To test these hypotheses, we estimate the following Probit model:

[9] If increased competition makes differentiation from outside banks more important, inside banks should acquire information more intensely (Boot and Thakor, 2000 and 2010).

$$P(Collateral_{il}) = F\left(\beta_0 + \beta_1 Sizeconcen_{il} + \beta_2 ACR4_{il} + \sum_{j=1} \sigma_j Relcontrols_{il} + \rho IPO_{il} + \sum_{j=1} \varphi_j FC_{il} + \sum_{j=1} \theta_j LC_{il} + \sum_{j=1} \gamma_j MC_{il} + \sum_{j=1} \delta_j RC_{il} + \sum_{j=1} \alpha_j FE_{il}\right) \quad (1)$$

where *i* indexes for firm, *l* for loan number, and *F(.)* is the cumulative distribution function of the standard normal distribution. The dependent variable $Collateral_{il}$ is a binary variable that equals one if loan *l* extended to firm *i* is collateralized and zero otherwise. IPO_{il} is a dummy equals 1 if a loan is issued after the borrower's IPO.

Following Schenone (2010), we measure bank-firm relationships by the intensity with which the borrower turns to the same lender.[10] This measure, which we call $Sizeconcen_{il}$, is defined as the amount of loans that firm *i* has borrowed from its current lender as a proportion of the total amount of loans which the firm has obtained prior to the current loan.[11] By definition, $Sizeconcen_{il}$ takes values of between zero and one. Borrower *i* is more dependent on the lender if $Sizeconcen_{il}$ is closer to one. This measure of relationship lending essentially takes into account the relative importance of a lender to the borrower, compared to other lenders. The next set of controls $Relcontrols_{il}$ accounts for additional features of relationship lending that can affect collateral incidence, including: the number of different lenders that firm *i* has borrowed from prior to the current loan, $Numlender_{il}$; whether the current loan is the first loan borrowed from the lender, $First_{il}$; and whether the current lender is

[10] The strength of bank-firm relationships is traditionally measured by relationship duration, defined as the time difference between the first loan obtained and the current one (see e.g. Petersen and Rajan, 1995; Berger and Udell, 1995). As suggested in Schenone (2010), duration may not fully capture how dependent a firm is on its current lender or how "locked in" the firm is in the lending relationship.

[11] We employ another relationship measure, $Numconcen_{il}$, defined as the *number* of loans that firm *i* borrowed from its current lender as a proportion of the total *number* of loans which the firm obtained prior to the current loan, as a further robustness check. Our main results are not sensitive to this alternative measure (results are available on request). The implicit assumption of $Numconcen_{il}$ is that the inside lender is more informed than outside lenders if the firm borrows more times from its current lender, while the amounts borrowed are irrelevant for the accumulation of information. As it is expected that banks devote more efforts in assessing firms that borrow larger amounts and subsequently accumulate more firm-specific information if the loan is relatively large, $Sizeconcen_{il}$ is probably a more precise measure of firm-bank relationships.

different from the previous lender, $Switch_{il}$. $Numlender_{il}$ controls for the fact that the same value of $Sizeconcen_{il}$ does not preclude that a firm borrows from different number of banks. For instance, a loan associated with a value for $Sizeconcen_{il}$ of 0.5 can be the result of borrowing from two banks, with each accounting for half of the total loans, or borrowing from five banks, with the largest loan accounting for half of the total loans. The first loan from lender ($First_{il}$) might be subject to different collateral requirement. Finally, we include $Switch_{il}$ to control for the possibility that banks may condition their collateral requirements depending on whether they can provide subsequent loans, for instance to minimize the costs of collateral evaluation. For all these variables, loans originated by either the parent bank or a subsidiary are treated as loans from the same lender, since it is likely that the information available about the borrowing firm is shared within all subsidiaries.

Banking market structure is measured by the concentration ratio $ACR4_{il}$, which is defined as the share of total assets of the four largest banks as a percentage of the total assets of all banks in each province at the time of one semi-accounting year prior to the current loan.[12] We treat each province as a separate banking market.

The set of variables FC_{il} accounts for firm characteristics that are likely to affect collateral. These include the age of the firm in (log) months, Age_{il}; (log) total assets, $Size_{il}$; current assets over total assets, $Liquidity_{il}$; return on total assets, ROA_{il}; tangible assets over total assets, $Tangibility_{il}$; and firms ownership dummy FT_{il} (equals 1 if the Chinese State is the majority owner and 0 if majority ownership lies in the private sector). Following Berger and Udell (1990), we also control for the ratio of loan size relative to total outstanding debt ($Loanconcen_{il}$), as a higher ratio suggests more important loans, which are more likely to be collateralized. These variables are obtained from the semi-annual financial reports that are published the closest to the moment before the loan was originated. This procedure ensures that in our estimations, banks use the most recent publicly available accounting information at the time that the loan is issued. All variables in monetary term are deflated to 2006

[12] For our purposes, market structure should be measured at the regional level. The concentration ratio is the only measure available of regional market structures. Market structure is closely related to competition. For a discussion of bank competition in China and the results for various competition measures see Xu et al. (2013).

RMB.

The set of controls LC_{il} covers loan characteristics, such as the maturity of loan l in (log) months, $Maturity_{il}$; its (log) size in real terms (deflated to 2006 RMB), $Loansize_{il}$; and the difference between its lending rate and the benchmark deposit rate of a corresponding maturity, $Spread_{il}$. We also control for monetary policy and regional macro-economic factors (MC_{il} and RC_{il}, respectively) that potentially can influence the pledging of collateral (e.g. Boot et al., 1991; Kiyotaki and Moore, 1997; Jiménez et al., 2006). Monetary policy controls include the reserve requirements ratio, RRR_{il} and the 7-day repo rate, $Repo_{il}$. These variables are matched to the month when the loan was originated. Regional macro-economic controls are the provincial real GDP growth rate (deflated with national CPI), $Realgdpindex_{il}$; provincial non-performing loan ratio, $NPLratio_{il}$; and the provincial consumer price index, CPI_{il}. These variables are matched to one semi-accounting year before the loan was originated. All these data come from the CEIC database.

The last set of controls are fixed effects (FE_{il}) for time (*Time*), bank-type (*Banktype*), province (*Prov*) and industry-type (*Indu*). These fixed effects capture systematic differences related to: business or credit cycles at the national level; bank type specific propensities in requiring collateral; provincial collateral policies; and differences in technology, production, market conditions, and government industry policies across different industries. In total 7 time dummies, 31 provincial dummies, 7 bank type dummies, and 51 industries dummies are introduced.

2.1.2. Informational rent and borrower transparency

This subsection attempts to find evidence that the increasing likelihood of collateral related to relationship lending and market concentration is at least partially due to informational hold-up. To this end, we test if the effects of relationship lending and market concentration on collateral are less pronounced for transparent firms, because information about these firms is more widely distributed among all lenders. Specifically, we test the following specification:

$$P(Collateral_{il}) = F\left(\beta_0 + \beta_1 Sizeconcen_{il} + \beta_2 ACR4_{il} + \sum_{j=1} \sigma_j Relcontrols_{il} + \rho IPO_{il} + \sum_{j=1} \varphi_j FC_{il} + \sum_{j=1} \theta_j LC_{il} + \sum_{j=1} \gamma_j MC_{il} + \sum_{j=1} \delta_j RC_{il} + \sum_{j=1} \alpha_j\, FE_{il}\right) \quad (2)$$

where an informational transparency measure $Infor_{il}$ (higher value representing more transparent) is interacted with the relationship lending and market structure variables ($Sizeconcen_{il}$ and $ACR4_{il}$, respectively). If $\beta_1 > 0$ and $\beta_3 < 0$, or respectively $\beta_2 > 0$ and $\beta_4 < 0$, it would lend some support to the idea that relationship lending respectively concentrated markets facilitate informational rent extraction, and that rent extraction is relatively more difficult if borrowers are transparent.

We apply two sets of transparency measures ($Infor_{il}$): transparency based on firm characteristics, and transparency resulting from stock market information production. The first set of transparency measures includes: listing board ($Listmain_{il}$); firm ownership (FT_{il}); and firm size ($Medianta_{il}$). $Listmain_{il}$ is a dummy variable that equals one if the firm is listed at the main board of the Shenzhen Stock Exchange, and zero if the firm is listed either at the small and medium-sized firms' board (SME board) or the China Next board (ChiNext board)[13]. Firms listed at the latter two boards are typically smaller or high-tech firms, which should be more informational opaque. Since nearly all banks in China are fully or partly state-owned, it is expected that banks are better informed about state-owned firms than about private firms. Finally, firm size is a standard measure of informational transparency, with smaller firms considered to be more informational opaque. We define a dummy $Medianta_{il}$ that equals one if the firm's total assets are above the provincial median,

[13] The listing boards are unknown for loans obtained before the listing. However, both firms and banks should have some idea about which listing board will be the most likely outcome when the firm applies for an IPO, given the characteristics of the firm. The lengthy approval process of the CSRC also suggests that firms need to decide at which board they will list long before the actual listing. As a robustness check, we reproduce the *Listmain* regression using loans issued only after listing. Our results hold for this alternative sample as well. Results are available upon request.

and zero otherwise.

The second set of transparency measures is related to stock market information production. Specifically, we postulate that firm transparency increases with the number of financial analysts (*Numalst*$_{il}$) following the firm, and the percentage of shares held by non-bank institutional investors (*Instishare*$_{il}$). We further investigate if informational spillovers from the stock market generate a boundary transparency level beyond which inside and outside banks are equally informed, and inside banks can no longer extract informational rents. As these information production variables are available only after being listed, we restrict the sample exclusively to post-IPO loans.

However, since these informational transparency proxies are also correlated with the probability of firms' financial distress or bargaining power, this identification strategy cannot fully differentiate the "hold-up" problem from competing theories (see section 2.1.1). For instance, under the assumption that larger firms are less likely to face financial stress than smaller firms, these firms have less incentive to pledge collateral to relationship lenders in exchange for a possible future rescue, leading to a smaller impact of relationship intensity on collateral incidence on larger firms. The implicit guarantee enjoyed by state owned firms may render collateral irrelevant in exchange for a future rescue from a relationship lender, which can lead to a lower impact of relationship intensity on collateral incidence for these firms. Similarly, as larger firms or state owned firms may have greater bargaining power, market structure could affect their collateral pledging less than that of smaller or private firms. The stock market information production measures could also be positively related to firm size or the financial health of firms. Namely, more analysts are required for larger firms, or non-bank institutional investors target financially healthy firms. These arguments suggest that the coefficients of the interaction terms should be negative, which can be a result independent of the informational rent extraction hypothesis. To better test this hypothesis, in the next sections we use equity IPOs as an informational shock that reveals informational to all banks, and therefore reduces the capacity of inside banks to extract informational rents.

2.1.3. Equity IPOs as strategy to identify informational rent extraction

This subsection formulates the methodology applying equity IPOs to identify informational rent extraction. This strategy hinges on the following observations. Before an IPO, inside banks enjoy superior information obtained from lending relationships, which allows for rent extraction through collateral. After an IPO, the constant release of information and market monitoring prevents any inside bank from obtaining or maintaining an informational monopoly position, therefore alleviating the adverse selection problems facing outside banks. Furthermore, a secondary effect might be at work which reinforces the direct effect of an IPO in reducing information asymmetries among inside and outside banks. Because an IPO will reveal information to all banks, inside banks are less incentivized to acquire additional but costly information to maintain their informational monopoly. This may be caused by a decreasing return on investment in information or an increasing cost of accumulating additional information in markets where all banks are well informed. Banks may also free-ride when costly information production can be conducted and disseminated by the stock market. With less investment in information after an IPO, information asymmetries among banks are reduced further. These arguments suggest that the informational monopolies of inside banks are greatly reduced after IPOs, making rent extraction through collateral less likely.

Similar arguments apply to banking market structure. As discussed in section 2.1.1, when borrowers lack a credible channel for disseminating information, such as before an IPO, concentrated markets permit: more efficient information extraction (Marquez, 2002); better reusability of information (Boot and Thakor, 2010; Chan et al., 1986; Berlin and Mester, 1999) and protection of information from spilling over to outside banks; and deters entry of competitors which self-reinforces the information monopolies (Dell'Ariccia et al., 1999; Dell'Ariccia, 2001). After an IPO, information is made public to outside banks through regularly published financial statements, public auditing, financial analysts' research and movements in stock prices. Hence, the role of market concentration in facilitating information asymmetry among lenders becomes less important, which erodes the possibility of informational rent extraction.

We formulate the following hypotheses:

H.3: If relationship lenders extract informational rents through collateral, this will be more likely for loans originated before the IPO and less likely for those originated after the IPO. If this moderated effect for post-IPO loans is not supported by the empirical results, alternative theories should explain the positive correlation between relationship lending and collateral incidence.

H.4: The positive correlation of market concentration with collateral should be mitigated by the informational shock of an IPO. If this result is not established, the positive impact of market concentration on collateral incidence is attributed to market power.

To test these hypotheses, we introduce interaction terms between the relationship intensity and market structure variables respectively, with IPOs in *Equation (1)*, which yields *Equation (3)*:

$$\begin{aligned} P(Collateral_{il}) &= F\Bigg(\beta_0 + \beta_1 Sizeconcen_{il} + \beta_2 ACR4_{il} + \beta_3 Sizeconcen_{il} * IPO_{il} + \beta_4 ACR4_{il} * IPO_{il} \\ &+ \sum_{j=1} \sigma_j Relcontrols_{il} + \sum_{j=1} \mu_j Relcontrols_{il} * IPO_{il} + \rho IPO_{il} + \sum_{j=1} \varphi_j FC_{il} + \sum_{j=1} \theta_j LC_{il} \\ &+ \sum_{j=1} \gamma_j MC_{il} + \sum_{j=1} \delta_j RC_{il} + \sum_{j=1} \alpha_j FE_{il}\Bigg) \end{aligned} \tag{3}$$

Informational rent extraction by relationship lenders is identified if $\beta_1 > 0$ and $\beta_3 < 0$. Similarly, market concentration facilitates informational rent extraction if $\beta_2 > 0$ and $\beta_4 < 0$. If $\beta_3 < 0$ or $\beta_4 < 0$ is rejected, the positive coefficients of β_1 and β_2 should be explained by other theories as discussed in section 2.1.1. We include the interaction term $Relcontrols_{il} * IPO_{il}$ to control for the possible heterogeneous impact of other relationship characteristics on collateral incidence before and after an IPO.

Two important caveats must be kept in mind. First, the moderated effect of relationship lending on collateral for the post-IPO loans could be explained by theories other than informational rent extraction. We discuss and test these alternative explanations in Section 4. A second caveat is related to the endogeneity assumption of IPOs and relationship lending. In practice both variables could be endogenous due to omitted variables. We address these issues using recursive bivariate probit models in Section 5. We discuss some further robustness tests in Section 6.

2.1.4. Informational rent extraction and firm risk

Rajan (1992) suggests that inside banks can charge informational rents more easily from riskier borrowers than from safer ones, because outside banks will be less inclined to lend once the borrower is revealed as risky. This view suggests that when information asymmetry between inside and outside banks is alleviated, rent extraction will decline for safer firms but not for risky ones. We test to see if this prediction applies to collateral as well (see Hale and Santos (2009) for similar tests on lending rates).

We propose a novel measure of firm risk: whether the first IPO application of a firm was rejected by the CSRC ($Multiapp_{il}$). A firm can be rejected for an IPO by the CSRC for many reasons, such as cash-flow problems, uncertain or weak profitability perspectives, unclear corporate governance structures or suspicious earnings, all of which suggest potential risk factors that do not meet CSRC listing requirements. In a way, this measure is similar to a credit rating (see an application in Hale and Santos, 2009), but now the firm is rated by a government body instead of private sector rating companies. To test this hypothesis, we expand the baseline *Equation (3)* with three-way interaction terms between informational rent variables ($Sizeconcen_{il}$ and $ACR4_{il}$), IPO_{il}, and firm risk proxy $Multiapp_{il}$.

2.2. Data

Unlike most studies on informational rent extraction which focus on advanced economies, we concentrate on China, which is an ideal testing ground for our purpose for several reasons. First, collateral is particularly important in markets where banks lack sufficient tools or expertise to price credit risk, or are inhibited to do so due to price regulations. This has been the case in Chinese bank lending markets for many years. An additional incentive to request collateral in these markets is to reduce the personal risks faced by loan officers, as the "loan officer responsibility system" introduced in 2002 holds individual loan officers accountable for bad loans (Qian et al., 2015). Second, Chinese banking has been characterized by strict interest rate controls, which only very recently have been lifted completely. This suggests that banks have had less discretion in setting prices compared to their counterparts in advanced economies, making rent extraction through collateral an attractive alternative. Third, the protection of creditor rights in China was strengthened after the approval of the Property Law of the People's Republic of China in 2006 (Berkowitz et al., 2015), which increased the value of collateral. As our sample starts in 2007, informational rent extraction though collateral may have become more valuable since then, given the enhanced credit rights protection embedded in the new law. Fourth, bank lending markets in China are relatively segmented and offer significant variation across regions and time. This feature allows us to test if collateral requirements vary with the information configurations embedded in regional bank market structures. Finally, the particular features of equity IPO regulations and procedures in China make IPOs a valid choice as an exogenous informational shock for Chinese credit markets. Firms might expect to go public at some point, but the exact timing of an IPO depends on the approval by the China Securities Regulatory Commission (hereafter the CSRC), which is unpredictable and exogenous to both banks and firms, suggesting that adjustments of loan contract terms prior to an IPO are hardly economically viable.

We manually collect *loan-level* data from listed firms' financial reports, published by Wind Finance Co., Ltd. Hence, our analysis departs importantly from most studies on Chinese loan markets, which either use yearly aggregate *firm-level* data from the China Securities Markets and Accounting

Research Database (CSMAR) (e.g. Firth et al., 2012; Chen et al., 2013) or rely on loan-level datasets provided by *few* state-owned banks (Chang et al., 2014; Qian et al., 2015).

Our dataset consists of 10,654 loans made to 676 firms listed at the Shenzhen Stock Exchange (SZSE) between 2007 and 2013.[14] The size of the sample is reduced by some recording errors, incomplete loan contract information and questionable financial data. In particular, loans issued at rates below the lending rate floor (i.e. below 90% of the baseline lending rate) are removed, because these loans are likely to have been issued at non-commercial terms. We further remove loans to financial institutions and loans made in foreign currencies. This reduces our database to 9,288 loans provided to 649 listed non-financial firms. Our database provides information on multiple borrowings by each firm (on average, each firm has 20 loans in our sample) and from multiple banks (on average 4 banks per firm), including almost all types of Chinese banks.

Summary statistics for all variables are provided in Table I. 66% of the loans in our database are collateralized, which is comparable to figures recorded for other emerging market economies, such as 53% for Mexico (La Porta et al., 2003) and 72% for Thailand (Menkhoff et al., 2006). Our main relationship variable $Sizeconcen_{il}$ has an average value of 0.33, suggesting that on average around one third of loans are obtained from a firm's current lender. The concentration ratio $ACR4_{il}$, which is our proxy for market structure, has an average of 0.55, indicating that the four largest banks in each province on average hold 55% of total provincial banking assets.

The summary statistics for IPO_{il} show that 83% of the loans in our sample were issued after an IPO. Among the 649 firms in our sample, 111 firms reported at least one loan before their IPO and at least one after; in total these firms account for 2,181 loans, representing 23% of all loans. The remaining firms only had loans either before their IPO (142 firms with 660 loans) or after (396 firms with 6,447 loans). Furthermore, our sample consists of relatively old (on average 13 years) and large firms (average total assets of RMB 2,139.5 million). Regarding firm ownership (FT_{il}), firms with state

[14] We concentrate on listed firms from Shenzhen Stock Exchange because firms listing at this stock exchange market are more diverse in terms of size and industry when compared with those listed at the Shanghai Stock Exchange. Our sample starts from 2007 because listed firms were required to comprehensively report their loan records from 2007.

majority ownership represent 33% of all firms in our sample and account for 40% of all loans.

Regarding the controls for loan characteristics, the average maturity of the loans in our sample ($Maturity_{il}$) is around two years (25.9 months), while the average size ($Loansize_{il}$) in real terms is RMB 62.6 million. The average spread between loan lending rates and corresponding deposit rates ($Spread_{il}$) is 2.85%.

Of the other controls, we provide further detail only on the variable that we use to investigate rent extraction and firm risk, i.e. $Multiapp_i$, which measures whether the firm is rejected in its first IPO application. 40 firms, or around 7% of all firms, were rejected for an IPO when they applied for the first time (but were eventually listed, after multiple applications). The definition and summary statistics for each instrumental variable and additional variables are discussed in their respective sections, but are all reported in Table I, panel F, G.

3. Main results

3.1. Univariate tests

This subsection investigates whether the mean values of the key variables differ across relationship intensity, market structure and for pre- and post-IPOs loans. Results are reported in Table II.

Relationship loans, defined as the ones with $Sizeconcen_{il}$ above the sample median, on average enjoy better loan terms such as longer maturity and lower lending spreads. At the same time, these loans are smaller; however collateral requirements do not differ significantly between relationship and non-relationship loans.

Collateral requirements are significantly more severe in concentrated markets, where concentrated markets are defined as the ones with $ACR4_{il}$ above the sample median. Loan maturity does not differ across markets, while loan size and the average lending spread are significantly larger in less concentrated markets. Lastly, loan contract terms such as collateral (-), maturity (+) and loan size (+) change significantly after listing (in brackets change after IPO compared to before), while the average lending spread does not differ for loans issued before and after IPOs.

Firm characteristics do not depict a clear pattern between groups. For instance, firms that borrow from relationship lenders are on average more liquid, less leveraged and have higher tangibility ratios. However, they are also younger and smaller than firms borrowing from non-relationship banks. Firms that borrow in concentrated markets are on average less liquid, smaller, younger and more leveraged, and have higher tangibility ratios. Lastly, firms that borrow after an IPO are less liquid and less profitable, but the leverage ratio of borrowing firms does not differ before and after the IPO.

3.2. Multivariate tests

3.2.1. Do relationship lending and market structure determine collateral incidence?

In this section, we first test the impact of relationship lending and market structure on collateral incidence in a cross-sectional setting by estimating *Equation (1)* in Section 2.1.1. The results are reported in Panel A of Table III. Marginal effects (M.E.) are calculated based on the results in Column (1). To account for the possibility that some loan contract terms such as *Maturity* and *Spread* are endogenous, we follow Berger and Udell (1995) and estimate the model with and without these terms (Columns (1) and (2), respectively). We conduct additional robustness tests for endogeneity issues of loan contract terms in Section 6.2.

Our results show that relationship intensity is positively related to the incidence of collateral and is highly significant. The marginal effects show that a one standard deviation increase in *Sizeconcen* from its sample mean increases the probability of collateralization by 1.4%. This result does not support the "information accumulation" view that relationship lending and collateral are substitutes in mitigating borrower risks (e.g. Berger and Udell, 1995). In contrast, our finding is in line with the other hypotheses discussed in section 2.1.1. (e.g. "hold-up" problem (Sharpe, 1990;

Rajan, 1992), “soft budget constraint” (Dewatripont and Maskin, 1995; Boot, 2000), “bank seniority” (Longhofer and Santos, 2000) and “cost minimization incentive” (Menkhoff et al., 2006)). Results similar to ours have been reported in e.g. Elsas and Krahnen (2000) and Ono and Uesugi (2009).

Banking market structure, measured as the concentration ratio *ACR4*, is positive and highly significant at the 1% level across all specifications. A one standard deviation increase in this ratio increases the likelihood of collateral incidence by 4.45%. This result confirms Hypothesis *H.2* (Section 2.1.1) that concentrated markets are associated with a higher likelihood of collateralization. Our finding is in line with Hainz et al. (2013), but contrasts with Jiménez et al. (2006). As discussed, both the “informational rent extraction” and “market power” hypotheses can explain this positive coefficient.

The coefficient of *Numlender* is significant and positive as well. A one standard deviation increase in the number of lenders of the firm from its mean increases the incidence of collateral by 2.13%.[15] Other relationship control variables such as *First* and *Switch* are not statistically significant; we shall discuss these results in more detail later on.

Loans obtained after an IPO are significantly less likely to be collateralized (marginal effect is -10.39%). This result lends some support to the notion that IPOs are beneficial to firms with respect to the non-price terms of lending. This adds to the empirical findings in Santos and Winton (2008), Hale and Santos (2009) and Schenone (2010) that loan terms improve after bond or equity IPOs, with these studies presenting evidence of a decline in lending rates.

Before moving forward, we discuss briefly other determinants of collateral, which has merit in itself, as the existing literature on Chinese lending markets has investigated this issue only using firm-year data (e.g. Firth et al., 2012; Chen et al., 2013). As expected, the coefficients of *Age* and *Size* are negative and significant, indicating that older and larger firms are less likely to pledge collateral, possibly because these firms are less prone to moral hazard problems. Firms that are more profitable,

[15] This result is in line with Chakraborty and Hu (2006) and Jiménez et al. (2006), but in contrast to Menkhoff et al. (2006).

more liquid, have a higher tangible assets ratio and are less leveraged are less likely to pledge collateral. Similar to Berger and Udell (1990), we find that *Loanconcen* is significantly positive at the 1% level across all specifications.[16] Among all factors, the most important determinant of collateral is firm ownership. Private firms in China have on average a 16.7% higher probability of pledging collateral than state-owned firms, presumably because the latter enjoys the implicit guarantee from the State. This results adds to the previous empirical studies that private firms in China have been financially discriminated in a state-dominant banking system (Cull and Xu, 2003; Allen et al., 2005).

Other loan contract terms affect the incidence of collateral as well. Loans with a longer maturity are more likely to be collateralized. A one standard deviation increase in loan maturity from its sample mean increases the incidence of collateral by 3.39%. This result is in line with the theoretical prediction that banks use shorter loan maturities to solve adverse selection or moral hazard problems (e.g. Berlin and Mester, 1992; Flannery, 1986; Barclay et al., 1995; Degryse and Van Cayseele, 2000). Larger loans (*Loansize*) are less likely to be collateralized. A one standard deviation increase of loan size reduces the incidence of collateral by 3.37%.[17] Finally, loans with a higher interest rate spread (*Spread*) are more likely to be collateralized (marginal effect of 1%) giving some support to the notion that collateral is associated with risky loans. Nevertheless, the results for contract terms on collateral should be treated with caution, as these variables are potentially endogenous. Excluding potentially endogenous loan contract terms such as *Maturity* and *Spread* does not alter our results for other determinants, as shown in Column (2).

In contrast, the monetary policy stance has a limited impact on the incidence of collateral, with only the *7-day Repo* rate being positively related to collateral at the 10% significance level.[18] Regional macroeconomic variables (*CPI*, *NPLratio* and *Realgdpindex*) generally do not affect

---

[16] See for instance Boot et al. (1991), Dennis et al. (2000) and Bharath et al. (2011) for similar results.

[17] This result is consistent with Leeth and Scott (1989), Jiménez and Saurina (2004) and Menkhoff et al. (2006), but in contrast to the findings of Boot et al. (1991).

[18] Jiménez et al. (2006) find that collateral incidence is lower during episodes of monetary tightening. They resort to credit rationing to explain their results, since during tightening periods banks prefer high-quality borrowers (hence less collateral). Bernanke and Gertler (1995) suggest that higher interest rates raise a firm’s default probability, resulting in a higher likelihood of collateral incidence during monetary policy tightening cycles. Our insignificant result could be due to the combined effect of competing theories, which we leave to future research.

collateral decisions. It is likely that the impact of business cycles is captured by time fixed effects. As a further robustness check, we include regional legal and institutional variables (results available upon request).[19] Our results do not materially change when these additional controls are added.

3.2.2. Does rent extraction vary with firm information transparency?

We test in this section if informational rent extraction is less pronounced for transparent firms. To this end, we estimate *Equation (2)* in Section 2.1.2 using various informational transparency proxies. Results are reported in Table III, Panel B and C, where Panel B uses firm characteristics as transparency measures, and Panel C employs stock market information production as transparency measures.

Firms that are not listed at the main board, privately owned, or small, are more likely to pledge collateral when relationship intensity increases, as suggested by the significantly positive coefficients of $Sizeconcen_{il}$ in all specifications of Panel B. For transparent firms, the impact of $Sizeconcen_{il}$ on collateral vanishes, as the null-hypothesis H_0: $Sizeconcen_{il}+Infor_{il}*Sizeconcen_{il} = 0$ is not rejected for all three informational transparency measures. As for the impact of market structure on collateral, a similar pattern prevails. The concentration ratio $ACR4_{il}$ is statistically positive in all specifications, and its interaction term with information transparency measures is significantly negative for all three cases. Unlike for relationship lending, the null hypothesis that market structure has no impact on collateral for transparent firm (e.g. firms listed at the main board or state-owned firms), i.e. $ACR4_{il}+Infor_{il}*ACR4_{il}=0$, is rejected. Both results suggest the inside banks' ability to charge rent decreases with firms' information transparency.

[19] Empirical studies have identified that banks are better able to control for credit risk if legal frameworks allow lenders to seize collateralized assets in times of default (Qian and Strahan, 2007). We employ the indices of legal infrastructure developed by Fan et al. (2011). These indices have been widely applied for China (e.g. Li et al., 2009), with Li et al. (2009) providing a detailed description. As data for these indices end in 2009 (while our sample ends in 2013), we interpolate the missing values by assuming that the indices grow at the average growth rate of 2006-2009. Our results show that collateral is more likely to be pledged in provinces with better legal infrastructure, a result that is similar to Qian and Strahan (2007). These authors suggest that a better protection of credit rights increases the incidence of collateral for firms with more tangible assets. The results that we present in the rest of the paper are not sensitive to the inclusion of these legal and institutional variables. Results are available upon request.

Next we employ stock market information production variables ($Numalst_{il}$ and $Instishare_{il}$) as proxies of firm transparency. Results are reported in Panel C, Columns (6) and (7). All interaction terms are significantly negative, indicating a moderated effect on rent extraction when more information is produced by stock market, a result similar to Panel B. Moreover, the magnitude of the coefficients further suggests a boundary effect of information production on rent extraction. In other words, rent extraction becomes infeasible when sufficient information is produced by stock market. Specifically, in Column (6), when a borrower is followed by more than 11 analysts (65^{th} percentile), the positive impact of *Sizeconcen* vanishes. Similarly, higher market concentration does not increase collateral incidence for borrowers followed by more than 22 analysts (88^{th} percentile). Column (7) reports similar results where *Instishare* serves as a measure of information production.[20] The results in this subsection are in line with the informational rent hypothesis. However, as discussed in section 2.1.1, alternative theories can also support these finding as information transparency measures are often correlated with firm quality or likelihood of financial stress. We proceed in the next subsection using IPO as an identification strategy.

3.2.3. Do equity IPOs reduce informational rents?

In this subsection, we provide a direct test of informational rent extraction, i.e. we compare the impact of $Sizeconcen_{il}$ and $ACR4_{il}$ on collateral incidence for pre-IPO and post-IPO loans where information asymmetry among lenders is significantly lower for the latter group than the former. Estimations are based on *Equation (3)*.

Results are reported in Table IV. Column (1) includes only the interaction term $Sizeconcen_{il}$* IPO_{il}; Column (2) includes only the interaction term $ACR4_{il}$ * IPO_{il}; Column (3) includes both, while Column (4) re-estimates Column (3) excluding possible endogenous loan contract terms (*Maturity*

[20] Arguably, institutional investors not only bring on board more information disclosure, but also active monitoring and better alignment of management incentives, such as reducing tunneling behavior (e.g. Lin et al., 2011). We control for these effects by incorporating corporate governance variables that directly affect firms' tunneling incentives: the "control and cash flow rights wedge" and cash-flow rights. Our results remain intact, and they are available upon request.

and *Spread*). The results show that $Sizeconcen_{il}$ is significantly positive across all models. The coefficient of the interaction term $Sizeconcen_{il}*IPO_{il}$ is negative and significant for the broader specification (Column (3)), while it is marginally insignificant (p-value 0.102) in Column (1). The coefficient of $ACR4_{il}$ is significantly positive while the interaction term with IPO_{il} is significantly negative across all specifications. As the results of these three specifications are quantitatively similar, we provide a detailed explanation of the results presented in Column (3) only, which is our baseline model.

The likelihood of pledging collateral is increasing with relationship intensity for pre-IPO loans (coefficient 0.596***), while for post-IPO loans this positive impact is greatly moderated (coefficient 0.124*, and H_0: $Sizeconcen_{il}+Sizeconcen_{il}*IPO_{il}=0$ is rejected at the 10% level). In terms of marginal effects, a one standard deviation increase in $Sizeconcen_{il}$ increases the probability of pledging collateral by 4.78% for pre-IPO loans, compared to 1.17% for post-IPO loans. This pattern is consistent with Hypothesis *H.3* (Section 2.1.3) that a reduction in informational asymmetry among lenders makes it harder to establish "hold-ups" through relationship lending, therefore lowering the likelihood of rent extraction through collateral.

A similar pattern is observed for banking market structure. The pre-IPO coefficient of the concentration ratio $ACR4_{il}$ is 5.94***, indicating that pre-IPO loans obtained in concentrated markets are significantly more likely to be collateralized. The post-IPO impact of $ACR4_{il}$ is moderated, but remains statistically positive (coefficient 2.43***, H_0: $ACR4_{il}+ACR4_{il}*IPO_{il}=0$ rejected at 1%). Alternatively, looking at the marginal effects, a one standard deviation increase in the concentration ratio increases the probability of collateral incidence by 8.51% for pre-IPO loans, while for post-IPO loans this effect is reduced to 4.15%. Hence, the contribution of concentrated markets in facilitating the extraction of information, or preventing its spillover to competitors, is greatly eroded, since more information about borrowing firms has been disseminated due to the IPO. This more equal distribution of information further reduces *de novo* banks' adverse selection problems and lowers barriers to entry, which is another reason why informational rent extraction is more difficult for post-

IPO loans. This result confirms Hypothesis *H.4* (Section 2.1.3).

We find the positive impact of market concentration on collateral is both statistically and economically significant even for post-IPO loans. The presence of a certain degree of information asymmetry among lenders even post-IPO could explain this results. This result could also lend some support to the view that information asymmetries are not the only channel leading to higher collateral incidence in concentrated markets. The "market power channel", discussed in section 2.1.1, suggests that monopolistic or oligopolistic banks can extract rents by using their market power, increasing collateral requirements even in an environment where all lenders are equally informed. This channel could be particularly important for banking markets characterized by geographic restrictions in branch expansion or restrictions in business scope.

It is likely that firms gain bargaining power vis-à-vis lenders after their IPO, for example because the listing improves their access to capital markets or increase their attractiveness as clients for other lenders. This reduces the positive impact of relationship lending or bank market structure on collateral incidence. Nevertheless, at least part of the bargaining power gain is due to the higher visibility of post-IPO information dissemination, which makes it extremely hard to differentiate information and bargaining power effects. We control for possible shifts in a borrowing firms' bargaining power by introducing an interaction term $Numlender_{il}*IPO_{il}$. Firms that can borrow from different lenders might be expected to benefit from higher intra-bank competition and therefore have more bargaining power vis-à-vis their current lender(s) (Yasuda, 2007). In our univariate tests, we found that an average firm borrows from two banks before an IPO, while this number increases to four after the IPO, suggesting increasing bargaining power. However, the coefficients on $Numlender_{il}$ and $Numlender_{il}*IPO_{il}$ are both insignificant.

Next, we briefly discuss the other control variables. $First_{il}$ is significantly positive for pre-IPO loans, indicating that borrowing for the first time from a certain lender before an IPO is associated with a higher likelihood of collateral pledging. For post-IPO loans, collateral incidence is not affected by whether the loan is the first one from a certain lender or not (H_0: $First_{il}+First_{il}*IPO_{il}=0$ cannot be rejected). This pattern is fairly persistent throughout all our regressions, which further supports the role of IPOs in disseminating information. Before an IPO, the first loan is associated with higher collateral incidence due to limited knowledge of the borrower. However, this significant relationship disappears after the IPO, given that the IPO process and post-IPO information disclosure increases the transparency of the borrowing firm to all potential lenders. Switching lenders ($Switch_{il}$), however, does not affect collateral incidence before or after the IPO. The coefficients on other control variables are similar to those reported in Table III, which are available upon request.

To conclude, using IPOs as an informational shock, the results in this section provide evidence of informational rent extraction, whether the informational advantage is driven by relationship lending or concentrated markets. As discussed in section 2.1.1, the results of this section are subject to caveats related to alternative explanations and endogeneity issues of key variables, which we examine in Section 4 and 5.

3.2.4. Do informational rents vary with firm risk?

Finally, we test whether following an IPO, informational rents reduce for safe firms, but not, or to a lesser extent, for risky firms. We introduce a three-way interaction term between our informational rent variables ($Sizeconcen_{il}$ or $ACR4_{il}$), IPO_{il} and the firm risk proxy $Multiapp_{il}$. Results are reported in Table V.

In the first column, we examine the main effect of $Multiapp_{il}$. A firm with multiple applications is 7% more likely to pledge collateral than first-time approved firms, which is consistent with our belief that being rejected for IPO is associated with higher firm risk. Three-way interaction terms are introduced in Column (2). Our results show that the marginal effects of the informational rents variables ($Sizeconcen_{il}$ and $ACR4_{il}$) on collateral are all *positive* both before and after IPOs. However, whether these marginal effects are moderated after an IPO depends on the riskiness of firms. To see this, we calculate the *change* in the marginal effects of the informational rent variables after and before IPO, for safe ($Multiapp_{il}$=0) and risky firms ($Multiapp_{il}$=1). For safe firms, the marginal effects of $Sizeconcen_{il}$ on collateral drops by 4% after the IPO, while for risky firms, it increases by 3.2%. Similar results are found for market structure. The marginal effect of $ACR4_{il}$ drops by 6% for safe firms after the IPO, but for risky firms it increases by 5.5%.

These results show that the ability of inside banks to charge informational rents after an IPO falls for safer firms, but increases for risky ones. This is because once the borrower is identified as safe, outside banks bid aggressively for lending business, reducing the inside bank's monopoly power. In contrast, outside banks will be less interested in lending to risky firms when the latter's poor creditworthiness is revealed, strengthening the ability of inside banks to extract rents. We test the robustness of these results by removing loan contract terms (Column (3)) and monetary policy and regional macroeconomic variables (Column (4)). In all cases, our results remain the same.

4. Alternative explanations

As noted earlier, the moderated effect of relationship lending on collateral incidence for post-IPO loans could be explained by alternative theories, which we discuss in this section.[21] One possible alternative is that credit quality is significantly higher for listed firms compared to unlisted ones. In other words, it is higher credit quality instead of lower information asymmetry that explains this moderated effect. The second possible explanation is related to potential heterogeneous risk dynamics around the IPO for relationship dependent and non-dependent firms. The final alternative explanation that we explore is that relationship banks reduce their collateral requirements in exchange for corporate bond underwriting business. We do not find supporting evidence for the first two alternative explanations and the last alternative explanation cannot dismiss the informational rent extraction hypothesis.

[21] We can discard one alternative explanation of the positive correlation between collateral incidence and relationship lending intensity that we find. This is the "cost minimization incentive" view (Menkhoff et al., 2006), which we discussed in section 2.1.1. This interpretation is not able to explain our results, as this incentive is unlikely to change depending on whether the borrower is listed or unlisted. Hence, the observed significant and negative coefficient of the interaction term $Sizeconcen_{il}*IPO_{il}$ is not supported by this theory.

4.1. Higher credit quality of listed firms

Boot (2000) and Longhofer and Santos (2000) (see section 2.1.1) predict a weaker positive correlation between relationship lending and collateral incidence for financially sound firms relative to distressed firms. If listed firms are financially healthier than unlisted ones, it would reduce the need to post collateral from the relationship lender's perspective, as the risk of financial distress and the likelihood of engaging in a future rescue is lowered. However, various studies have shown that the operating performance of listed Chinese firms drops markedly after an IPO. For example, Allen et al. (2014) compare the operating performance of listed and non-listed firms in China for the years around an IPO and find that the average return on assets of listed firms drops significantly from 0.12 to 0.07 within a [-3, 3] years window. This sudden drop is not observed for the unlisted firms over the same time horizon. These authors attribute the deterioration in performance to the extremely strict listing requirements of the CSRC,[22] which induce firms to improve earnings in the years prior to an IPO, adjusting operations to generate short-term profits at the possible cost of long-term growth. Similar evidence is also found in our sample where the average return on assets for pre-IPO firms is around 10% higher than post-IPO firms (e.g. from 15% prior to the IPO to 5% after, see Table II).

To further address selection bias in listing status caused by observables, we employ a propensity score matching method. The propensity score of loans being borrowed by listed firms is estimated based on a set of variables determining an IPO. Using nearest neighbor matching, loans borrowed by listed firms are then matched to the ones borrowed by unlisted firms. We drop loans that are outside of the common support to minimize the potential bias introduced by these loans. This process generates a matched sample of loans that are "identical" in every aspect, except for the borrower's listing status. We re-estimate the baseline model in Table IV, Column (3) on this matched sample. Our results do not materially change (available upon request) and so we conclude that higher

[22] To be approved for listing, firms need to report positive earnings in the three consecutive years prior to the IPO or have accumulated at least 30 million in net income. In addition, firms are required to have accumulated net cash flows of more than 50 billion or revenues in excess of 300 million in the three years prior to the IPO.

observed credit quality of listed firms is unlikely to drive our results.

Obviously, the credit quality of listed and unlisted firms may also differ in an immeasurable way. We conduct further analysis in Section 5 to account for these unobserved risk factors.

4.2. Heterogeneous risk dynamics for relationship dependent and non-dependent firms

Suppose "relationship dependent" firms improve credit quality more (or deteriorate less) than "relationship non-dependent" firms after listing. This heterogeneous change in firm riskiness could explain the moderated effect of relationship lending on collateral incidence for post-IPO loans. To address this concern, we perform difference-in-difference tests for observed risk proxies broken down by whether a firm is relationship dependent and whether the loan is borrowed after an IPO. In a fashion similar to Presbitero and Zazzaro (2011), a relationship dependency dummy is defined as equal to 1 if *Sizeconcen* is above or equal to the sample median (0.20). We construct difference-in-differences tests by regressing key financial risk proxies (*ROA*, *Leverage*, *Tangibility*, *Liquidity*, *Size*, *Maturity*, *Spread* and *Loansize*) on IPO_{il}, the relationship dependency dummy and the interaction terms of these two variables. The coefficient on the interaction term and its statistical significance indicate whether changes in risk proxies around the IPO differ according to relationship intensity. Results are reported in the Internet Appendix, Table IA.I. In all these difference-in-differences tests, the interaction terms are statistically insignificant except for *Liquidity*. Hence, heterogeneous risk dynamics are unlikely to be a key driver of our results.

Finally, we conduct matched sample analysis within pre- and post-IPO samples and compare the impact of relationship lending on collateral pledging across samples. This way we remove the possibility that firm-risk dynamics around IPOs could be driving our results. If relationship banks charge informational rents and if IPOs reduce information asymmetry among lenders, the average treatment effect of relationship lending should be positive for pre-IPO loans and be moderated or insignificant for post-IPO loans. We find that relationship dependent firms are on average 10% to 12%

more likely to pledge collateral relative to matched non-dependent firms for pre-IPO loans, while the difference between these two groups vanishes for post-IPO loans. Technical details, estimation results and sensitivity tests (including balancing property of covariates and sensitivity to unobservables) are reported in the Internet Appendix, Section A and Tables IA.II-III.

4.3. Corporate bond underwriting and concurrent lending

Banks may exchange better loan conditions for corporate bond underwriting business.[23] As most firms have a bond IPO after an equity IPO, and many firms choose their relationship banks as underwriters, the moderated effect of relationship lending for post-IPO loans could be the result of exchanging better loan conditions for bond underwriting fees, instead of an informational equalization effect. Our sample includes 1,287 loans that were originated after the firms' bond IPO, which is a sizeable sample. To address this issue, we construct various samples that only incorporate loans granted before a firms' bond IPOs. If our results are driven by concurrent lending and corporate bond underwriting, once we exclude loans borrowed after the bond IPO, the significant results for the interaction term $Sizeconcen_{il}*IPO_{il}$ should vanish. We find that this is not the case. Results are reported in the Internet Appendix, Table IA.IV.

5. Endogeneity of IPOs and relationship lending

In the previous sections, we treated the IPO or relationship lending variables as exogenous. However, they could be endogenous due to unobserved risk factors. We apply recursive bivariate probit models to address the potential endogeneity issue of IPOs in Section 5.1, and that of relationship lending in Section 5.2. Our results are robust after controlling for these endogeneity concerns.

[23] For instance, Yasuda (2007) documents that firms in Japan obtain a fee discount when employing relationship banks as corporate bond underwriters.

5.1. Endogeneity of IPOs

The fact that all of the firms in our sample have eventually completed their IPOs alleviates the endogeneity concern of IPOs to some extent. However, selection bias could still be present due to unobserved factors. As discussed in section 2, the exact timing of an IPO is to a large extent unpredictable for firms, but it is possible that there are uncontrolled factors that could affect both the timing of an IPO and collateral. For instance, firms' political connections (unobserved to econometricians) can speed up the listing process and at the same time lower collateral requirement as banks may consider politically connected firms less risky. This omitted variable problem makes the IPO variable and its interaction terms with other covariates in *Equation (3)* correlated with the error term in the equations, leading to biased estimates. To address this issue, we follow Wooldridge (2010) and implement a recursive bivariate probit model with instrumental variables[24]. The model is estimated with Maximum Likelihood Estimation (MLE). Besides consistency and efficiency of the MLE, a crucial benefit of this approach is that we can easily estimate the interactions of binary endogenous variable with exogenous variables in the structural equation (Wooldridge 2010).[25] One simply needs to specify that the only source of endogeneity comes from the binary treatment variable, treating the interaction terms in the structural equation as if they were exogenous. Specifically, we estimate the following model:

$$Collateral = 1[Z_1\alpha_1 + IPOX_1\beta_1 + \varepsilon_1] > 0$$

$$IPO = 1[Z_2\gamma + \varepsilon_2] > 0 \quad (4)$$

[24] Since IPO is a binary variable, the traditional two-stage least squares models will produce inconsistent estimators (Greene, 2008).

[25] The existence of endogenous interaction terms in the structural equation causes no problem for MLE estimation of the bivariate probit model because the density function of the outcome variable is conditional on all exogenous variables and endogenous binary variable (or function of endogenous binary variable), therefore the conditional density function is the same whether or not endogenous binary variable (or function of endogenous binary variable) enters the structural equation.

where Z_1 is a vector of collateral determinants and X_1 contains unity and variables that are allowed to be interacted with IPO. This *Collateral Equation* is the same as *Equation (3)*. In the *IPO Equation*, Z_2 contains all variables in Z_1 and at least one additional instrumental variable, i.e. it contains some exogenous variable that affects listing status, but does not explain collateral except through firm's listing status[26]. The error terms are assumed to be bivariate normal distributed with correlation ρ, i.e. $\varepsilon_1, \varepsilon_2 \sim \emptyset(0, 0, 1, 1, \rho)$.

We derive our instrumental variables from CSRC IPO suspensions. By the end of 2013, the CSRC has unexpectedly suspended the IPO reviewing and approval process on eight occasions[27]. These suspensions were unforeseeable by banks or borrowers, and therefore can serve as exogenous shocks. During these suspension periods, no new IPOs were approved, while firms that had already started their IPO applications were forced to stop it. These suspensions affect listing status for at least two reasons: firstly, listings will be delayed as the amount of reviewing work for the CSRC to complete piles up; and, secondly, some applicants need to prepare their application documents again as previous documents expire after the IPO suspension; this is costly and sometimes infeasible for firms that have exhausted their resources to boost up their accounting performance.

Naturally, it is unrealistic to assume that IPO applications are affected by all past CSRC suspensions. Only the ones that occur during firms' preparation period should affect their IPOs. The actual dates when firm started their preparation process are unknown, but the preparation and completion of IPO usually takes at least 1 to 3 years. We take the middle value of 2 years prior to actual listing dates as our cut-off point, which ensures that most of the applicants have started their preparation process[28]. Our first instrument is a dummy variable, *Affected_Firms*, which equals 1 if firms experienced at least one CSRC IPO suspension during the two-year window prior to their actual

[26] Wilde (2000) shows that exclusion restrictions are not generally needed in a multi-equation probit system and that identification is achieved if varying exogenous regressors appear in both equations of the bivarate probit model. Wooldridge (2010) however recommends not relying on nonlinearities solely to identify parameters in bivariate probit models.
[27] By the end of 2013, the CSRC IPO suspension periods are: 1) 1994/7/21-1994/12/7; 2) 1995/1/19-1995/6/9; 3) 1995/7/5-1996/1/3; 4) 2001/7/31-2001/11/2; 5) 2004/8/26-2005/1/23; 6) 2005/5/25-2006/6/2; 7) 2008/9/16-2009/7/10; 8) 2012/11/16-2013/12/31.
[28] Defining a 3-year window does not materially change our results. Results are available upon request.

listings. 442 (68% of all firms) firms satisfy this condition, and in total these firms borrowed 6351 loans (68% of all loans) throughout our sample period. We further calculate the number of IPO suspension days within this 2-year window as our second instrument, denoted it as *dd_lag2*. The average suspension days for *Affected_Firms* are 258 days. For unaffected firms, the number of suspension days is zero. To address skewness, we use *log(1+dd_lag2)* in the estimation.

The results of the recursive bivariate probit model are reported in Table VI. For comparison purpose, Column (1) reproduces the baseline mode of Table IV, Column (3). Column (2) and (3) estimate the recursive bivariate probit model using *Affected_Firms* and *log(1+dd_lag2)* as instruments, respectively. For brevity we report the key results only. Looking at the instrumental variables in the *IPO Equation*, we find the coefficients of *Affected_Firms* and *log(1+dd_lag2)* and are negative and statistically significant at the 1% level, consistent with our projection that IPO suspensions affect listing status. More importantly, after controlling for the endogeneity of IPO, the coefficients of the key variables in the structural equation (*Collateral Equation*) are very similar to the single Probit estimation results in Column (1). This result should not come as surprise since the MLE estimates of the correlation coefficient ρ are statistically insignificant in both Column (2) and (3), indicating that the exogeneity assumption of IPO cannot be rejected, which further justifies our estimations in previous sections using a single equation Probit model. [29]

[29] The validity of instruments hinges on the assumption that the CSRC IPO suspensions did not influence collateral incidence directly. Unfortunately, this assumption is not testable. An informal test of the exclusion restriction can be derived by including instrumental variables in the structural equation and test if their coefficients are statistically significant. The coefficients of *log(1+dd_lag2)* and *Affected_Firms* are -0.009 (p-value 0.22) and -0.03 (p-value 0.53), both of which are statistically insignificant. Another caveat is that banks may consider the IPO suspensions as negative shocks to the firms involved. Consequently, banks may raise collateral requirements for loans obtained during the suspension periods. This could relate IPO suspensions directly to the incidence of collateral, and therefore violate the exclusion restriction. To test this, we define a dummy variable *Affected_Loans* which equals one if loans are obtained by *Affected_Firms* during the suspension periods and zero otherwise. We find that 1,410 loans (or 15% of our sample) satisfy this condition. We re-estimate the baseline model (Table IV, Column (3)) including the *Affected_Loans* dummy. If banks consider the IPO suspensions as negative shocks to firms, *Affected_Loans* should be significantly positive. The coefficient of *Affected_Loans* is indeed positive (0.04, with p-value 0.48), but statistically insignificant. Results of these validity tests are available upon request.

5.2. Endogeneity of relationship lending

Relationship lending could also be endogenous due to omitted variables affecting both relationship formation and collateral[30]. For instance, firms with poor credit quality (unobserved to econometricians but known to competing banks) could only borrow repeatedly from their incumbent banks due to limited outside options. Therefore the positive correlation between relationship lending and collateral could be the result of unobserved poor credit quality instead of informational rent. We employ a recursive bivariate probit model with instrumental variables to address this concern. To implement this approach, firstly, we need to transform our continuous measure of relationship lending into a binary variable. In a fashion similar to Presbitero and Zazzaro (2011), a relationship dependency dummy (Rel) is defined to equal 1 if the firm obtains at least 20% (the sample median of the *Sizeconcen*) of bank loans from the lender prior to the current loan, and 0 otherwise. Secondly, at least one exclusion restriction must be imposed, i.e. there exists at least one exogenous variable that determines Rel, but does not affect $\mathrm{Collateral}$ except through relationship lending. We use past regional average lending rates (*Localavrate*) as instruments (definition and summary statistics are in Table I). A similar approach has been applied in Bharath et al. (2011).[31] *Localavrate* is expected to affect relationship lending positively as firms might prefer to borrow from their relationship lenders when past conditions in regional (local) credit markets are tight. It is unlikely that past regional average lending rates will affect the collateral pledged for current individual loans.[32]

Similar to *Equation (4)*, the recursive bivariate probit model is defined by a two-equation

[30] The self-selection issue of borrowing in concentrated or non-concentrated banking markets is not modeled. This self-selection issue is unlikely to be present because cross-regional loans are rare, due to the segmentation of Chinese banking markets. Regional banks such as city commercial banks and rural commercial (co-operative) banks mainly serve clients located in their own region. It is only recently that some city commercial banks have been allowed to establish branches outside their home province to better serve local customers. Banks that operate at the national level such as state-owned commercial banks (SOCBs) and joint-stock commercial banks (JSCBs) have a wide distribution of branch networks, which allows their local branches to provide loans to local firms. It is unlikely that firms will self-select themselves to borrow from banks (branches) outside their home province or in regional markets characterized by specific market structures in order to avoid collateral requirements.

[31] Bharath et al. (2011) invests joint estimations of loan contract terms, employing lagged average lending spread over the last six month as instrument for collateral. They argue lagged average lending spread do not necessary affect non-price terms such as collateral, based on their conversation with bankers.

[32] Unreported results show *Localavrate* is statistically insignificant as a determinant of collateral incidence. Results are available upon request.

system: a *Collateral Equation* and a *Relationship Equation*, where both relationship dependency dummy Rel and its interaction term with IPO ($Rel * IPO$) enter *Collateral Equation*. Other covariates in the *Collateral Equation* correspond to the ones used in Table IV, Column (3). The model is identified once the exclusion restriction *Localavrate* is added to the *Relationship Equation*, together with other determinants of relationship lending[33]. Results are reported in Table VI, Column (4). The estimated correlation between the error terms of the two equations, i.e. ρ, is significantly negative (-0.508***, p-value is 0.002), rejecting the exogeneity assumption of relationship lending and supporting the recursive bivariate probit estimation approach. The coefficient of the instrumental variable (*Localavrate*) in the *Relationship Equation* is 0.115, significant at 1%, indicating firms in provinces with higher past average lending rates are also more likely to borrow from relationship lenders. Turning to the *Collateral Equation*, the estimates controlling endogeneity of relationship lending are consistent with the baseline results of Column (1).

6. Further robustness tests

This section presents further robustness tests accounting for the unobserved firm specific time-invariant risks with fixed effect logit model (6.1); the endogeneity of other loan contract terms using instrumental (IV) probit model (6.2); and the sensitivity of the results to alternative samples (6.3). Our main results are robust to all these tests.

6.1. Firm fixed effects

Including firm fixed effects alleviates the concern that unobserved time-invariant risk factors can drive our results. As the Probit model is not suitable for fixed effects regressions, we use a fixed

[33] Covariates in the *Relationship Equation* include firm and loan characteristics, monetary policy and regional macroeconomic variables, and fixed effects dummies. Excluding potentially endogenous loan characteristics do not change our results. Estimation of the R*elationship Equation* show firms are more likely to borrow from relationship lenders if they are located in concentrated markets, are liquid, smaller, more leveraged, less profitable, have better loan contract terms such as longer loan maturities and lower spreads, and if the loan represents a relatively large portion of the firm's existing debt (*Loanconcen*). Full results of the recursive bivariate probit model are available upon request.

effects Logit model. Table VII reports the full sample results for specifications without potentially endogenous loan contract terms (Column (1)) and with those terms (Column (2)). Column (3) and (4) replicate these regressions for a sample excluding loans originated after a firm's bond IPOs. After controlling for firm fixed effects, the impact of relationship intensity on collateral incidence is significantly positive for pre-IPO loans, but is statistically insignificant across all specifications for post-IPO loans (H_0: $Sizeconcen_{il}+Sizeconcen_{il}*IPO_{il}=0$ cannot be rejected). This result is even stronger than that of the baseline model (Column (3) of Table IV), supporting the hypothesis that IPOs as an informational shock eliminates rent extraction opportunities. The results for market concentration are similar to previous findings, i.e. increasing market concentration increases the likelihood of collateral, and this effect is stronger for pre-IPO loans.

6.2. Endogeneity of loan contract terms

In this subsection we apply instrumental variable (IV) Probit regressions to address the endogeneity issue of loan contract terms. We examine two possibilities: exclude *Spread* from the determinants of collateral and treat *Maturity* as the sole endogenous variable; and treat both *Spread* and *Maturity* as endogenous variables.[34] The instruments chosen for *Maturity* are asset maturity (*Amaturity*, Barclay et al., 1995) and term spread (*Termspread*, Dennis et al., 2000 and Brick and Ravid, 1985). For the lending spread (*Spread*), we use as an instrument the benchmark loan spread (*Benchsprd* = benchmark lending rate minus the benchmark deposit rate), and lagged regional average lending rates (*Localavrate*). *Benchsprd* and *Localavrate* should be correlated with the lending spread but are not likely to be related to whether or not a particular loan is collateralized.[35] Summary statistics and

[34] The existing literature differs in treating which of the loan contract terms should be endogenous in determining collateral. Dennis et al. (2000) and Bharath et al. (2011) consider *Maturity* as the only endogenous contract term that affects collateral. The underlining assumption is that the lending spread is determined after the decision on collateral pledging. On the other hand, Brick and Paila (2007) and Ono and Uesugi (2009) model the spread as an endogenous determinant of collateral. As empirical validations are provided for both assumptions and theoretical advantages of either assumption are unknown a priori, we examine both.

[35] *Benchsprd* and *Localavrate* may reflect changes in the monetary policy stance or business cycle, which in turn might affect the incidence of collateral. See Jiménez et al. (2006). If this were true, these variables cannot serve as valid instruments. However, our estimations show that monetary conditions measured by the reserve requirement ratio or 7-day

definitions of these instrumental variables are in Panel F of Table I. Technical details, results and the relevance and validity of instrumental variables are reported in the Internet Appendix, Section B and Table IA.V. We find loan contract terms are indeed endogenous. Nevertheless, the IV probit results are largely consistent with previous findings, except that $Sizeconcen_{il}$ loses its explanatory power for post-IPO loans ($H_0{:}Sizeconcen_{il}+Sizeconcen_{il}*IPO_{il}=0$ cannot be rejected, p-value=0.99 or 0.86 depending on specifications), which is a even stronger result than for the baseline model. Results for banking market structure are also similar to previous findings.

6.3. Alternative samples

Lastly, we investigate in this section if results from the baseline model are sensitive to alternative samples. First, we focus on a sample of firms that borrowed at least once before its equity IPO and at least once after, which allows us to compare more precisely changes in collateral incidence around IPOs. Second, we restrict the sample to loans that were originated right before and after the IPO (e.g. one loan before and one loan after); four loans closest to IPO dates (e.g. two before and two after); and six loans closest to IPO dates (e.g. three before and three after). These short event windows minimize the possibility that significant events other than IPOs affect our results. Results for these samples are reported in the Internet Appendix, Table IA.VI. Finally, we investigate if our results are driven by non-commercial basis loans. We re-estimate *Equation (3)* by removing progressively loans from policy banks, state-owned banks, trust and investment companies and other financial institutions, on the basis that loans from these institutions could be based on policy preferences, political pressure, or other non-standard credit criteria. Results are reported in the Internet Appendix, Table IA.VII. Our main findings are solid in almost all of these samples.

---

repo rate, or the business cycle measured by regional GDP growth rates, do not impact significantly on collateral incidence, as reported in most of our tables.

7. Conclusions

In this paper, we investigate whether proprietary information obtained from both lending relationship and banking market concentration allow for informational rent through collateral. We find collateral incidence increases with both relationship lending and market concentration, and these effects are less pronounced for transparent firms. Using equity IPOs as informational shocks, we find that collateral incidence increases with both relationship intensity and market concentration for pre-IPO loans, while these effects are greatly moderated for post-IPO loans. Furthermore, we demonstrate that following an IPO, rent extraction through collateral is moderated for safe firms but intensified for risky firms, a result in line with the prediction of Rajan (1992). Further robustness tests suggest that our results are not caused by differences in credit risks, the possible endogeneity of IPOs and relationship lending, concurrent lending and underwriting, or non-commercial basis loans. Our results complement the finding that banks extract informational rents by charging higher lending rates (Hale and Santos, 2009; Schenone, 2009), and in part validate the theoretical predictions that concentrated market structure facilitates accumulation of inside information (Dell'Ariccia et al., 1999; Dell'Ariccia, 2001). Finally, we provide the first loan-level analysis on collateral for China, which has received little attention so far.

Our study opens up a few avenues for future research. A cross-country investigation of rent extraction through collateral could be fruitful. Rent extraction through collateral may be more likely to be observed in less developed markets where banks lack sufficient tools to price credit risks. Another possibility is to check if banks choose methods to charge rents (either through lending rates or collateral) depending on price regulation or monetary policy. A third avenue is to investigate how rent extraction through collateral could vary with the legal and institutional environment, as these aspects crucially determine how valuable collateral is to banks. We leave these issues for future research.

References

Allen., F., Qian, J., and Qian, M.J., 2005. Law, finance and economic growth in China. *Journal of Financial Economics* 77 (1): 57-116.

Allen, F., Qian, J., Shan, C.Y., Zhu, L., 2014. The best performing economy with the worst performing market: Explaining the poor performance of the Chinese stock market. *Unpublished manuscript.*

Barclay, M., Marx, L., and Smith, C., 1995. The maturity structure of corporate debt, *Journal of Finance*, 50(2), 609-631.

Berger A., Espinosa-Vega, M., Frame, W., and Miller, N., 2011. Why do borrowers pledge collateral? New empirical evidence on the role of asymmetric information, *Journal of Financial Intermediation*, 20(1), 55-70.

Berger, A., and Udell, G., 1990. Collateral, loan quality, and bank risk, *Journal of Monetary Economics*, 25(1), 21-24.

Berger, A., and Udell, G., 1995. Relationship lending and lines of credit in small firm finance, *Journal of Business*, 68(3), 351-381.

Berkowitz, D., Lin, C., and Ma, Y., 2015. Do property rights matter? Evidence from a property law enactment. *Journal of Financial Economics* 116: 583-593.

Bernanke, B., and Gertler, M., 1995. Inside the black box: The credit channel of monetary policy transmission, *Journal of Economic Perspectives*, 9(4), 27-48.

Berlin, M., and Butler, A., 2002. Collateral and competition, Federal Reserve Bank of Philadelphia, Working Papers 02-22.

Berlin, M., and Mester, L., 1992. Debt covenants and renegotiation, *Journal of Financial Intermediation*, 2(2), 95-133.

Berlin, M., and Mester, L., 1999. Deposits and relationship lending, *Review of Financial Studies,* 12(3), 579-607.

Besanko, D., and Thakor, A., 1987. Collateral and rationing: Sorting equilibria in monopolistic and competitive credit markets, *International Economic Review*, 28(3), 671-689.

Bester, H., 1985. Screening vs. rationing in credit markets with imperfect information, *American Economic Review*, 75(4), 850-855.

Bharath, S., Dahiya, S., Saunders, A., and Srinivasan, A., 2011. Lending relationship and loan contract terms, *Review of Financial Studies*, 24(4), 1141-1203.

Boot, A., 2000. Relationship banking: What do we know?, *Journal of Financial Intermediation*, 9(1), 7–25.

Boot, A., and Thakor, A., 1994. Moral hazard and secured lending in an infinitely repeated credit market game, *International Economic Review*, 35(4), 899-920.

Boot, A., and Thakor, A., 2000. Can relationship banking survive competition?, *Journal of Finance*, 55(2), 679-713.

Boot, A., and Thakor, A., 2010. The accelerating integration of banks and markets and its implications for regulation. In: A. Berger, P. Molyneux and J. Wilson (eds.), *The Oxford Handbook of Banking*, 58-89, Oxford University Press.

Boot, A., Thakor, A., and Udell, G., 1991. Secured lending and default risk: Equilibrium analysis, policy implications and empirical results, *Economic Journal*, 101(406), 458-472.

Brick, I., and Palia, D., 2007. Evidence of jointness in the terms of relationship lending, *Journal of Financial Intermediation*, 16(3), 452-476.

Brick, I., and Ravid, A., 1985. On the relevance of debt maturity structure, *Journal of Finance*, 40(5), 1423-1437.

Cerqueiro, G., Ongena, S., and Roszbach, K., 2016. Collateralization, bank loan rates, and monitoring, *Journal of Finance*, 71(3), 1295-1322.

Chang, C., Liao, G,. Yu, X,. and Ni, Z., 2014. Information from relationship lending: Evidence from loan defaults in China, *Journal of Money, Credit and Banking*, 46(6), 1225-1257.

Chakraborty, A., and Hu, C., 2006. Lending relationships in line-of-credit and nonline-of-credit loans: Evidence from collateral use in small business, *Journal of Financial Intermediation*, 15(1), 86-107.

Chan, Y., Greenbaum, S., and Thakor, A., 1986. Information reusability, competition and bank asset quality, *Journal of Banking and Finance*, 10(2), 243-253.

Chen, J., Lobo, G., Wang, Y., and Yu, L., 2013. Loan collateral and financial reporting conservatism: Chinese evidence, *Journal of Banking and Finance*, 37(12), 4989-5006.

Degryse, H., and Van Cayseele, P., 2000. Relationship lending within a bank-based system: Evidence from European small business data, *Journal of Financial Intermediation*, 9(1), 90-109.

Dell'Ariccia, G., 2001. Asymmetric information and the structure of banking industry, *European Economic Review*, 45(10), 1957-1980.

Dell'Ariccia, G., and Marquez, R., 2006. Lending booms and lending standards, *Journal of Finance*,61(5), 2511-2546.

Dell'Ariccia, G., Friedman, E., and Marquez, R., 1999. Adverse selection as a barrier to entry in the banking industry, *Rand Journal of Economics*, 30(3), 515-534.

Dennis, S., Nandy, D., and Sharpe, L., 2000. The determinants of contract terms in bank revolving credit agreements, *Journal of Financial and Quantitative Analysis*, 35(1), 87-110.

Dewatripont, M., and Maskin, E., (1995). Credit and efficiency in centralized and decentralized economies, *Review of Economic Studies*, 62(4), 541-555.

Elsas, R., and Krahnen, J., 2000. Collateral, default risk, and relationship lending: An empirical study on financial contracting, Center for Financial Studies, University of Frankfurt, CFS Working Paper No.1999/13, revised version, 25 November.

Fan, G., Wang, X., and Zhu, H., 2011. NERI Index of marketization of China's provinces: 2011 Report, Economic Sciences Press.

Firth, M., Malatesta, P., Xin, Q., and Xu, L., 2012. Collateral, leverage and corporate investment, unpublished manuscript.

Flannery, M., 1986. Asymmetric information and risky debt maturity choice, *Journal of Finance*, 41(1), 19-37.

Greene, W.H., 2008. *Econometric Analysis*, 6th edition, Prentice Hall.

Hainz, C., 2003. Bank competition and credit markets in transition economies, *Journal of Comparative Economics*, 31(2), 223-245.

Hainz, C., Weill, L., and Godlewski, C., 2013. Bank competition and collateral: Theory and evidence, *Journal of Financial Services Research*, 44(2), 131-148.

Hale, G., and Santos, J., 2009. Do banks price their informational monopoly?, *Journal of Financial Economics*, 93(2), 185-206.

Hauswald, R., and Marquez, R., 2006. Competition and strategic information acquisition in credit markets, *Review of Financial Studies*, 19(3), 967-1000.

Inderst, R., and Mueller, H., 2007. A lender-based theory of collateral, *Journal of Financial Economics*, 84(3), 826-859.

Jiménez, G., and Saurina, J., 2004. Collateral, type of lender and relationship banking as determinants of credit risk, *Journal of Banking and Finance*, 28(9), 2191-2212.

Jiménez, G., Salas, V., and Saurina, J., 2006. Determinants of collateral, *Journal of Financial Economics*, 81(2), 255-281.

Kiyotaki, N., and Moore, J., 1997. Credit cycles, *Journal of Political Economy*, 105(2), 211-248.

La Porta, R., López-de-Silanes, F. and Zamarripa, G., 2003. Related lending, *Quarterly Journal of Economics*, 118(1), 231-268.

Leeth, J., and Scott, J., 1989. The incidence of secured debt: Evidence from the small business community, *Journal of Financial and Quantitative Analysis*, 24(3), 379-394.

Li, K., Yue, H., and Zhao, L., 2009. Ownership, institutions, and capital structure: Evidence from China, *Journal of Comparative Economics*, 37(3), 471-490.

Lin, C., Ma, Y., Malatesta, P., and Xuan, Y., 2011. Ownership structure and the cost of corporate borrowing, *Journal of Financial Economics*, 100(1), 1-23.

Longhofer, D., and Santos, J., 2000. The importance of bank seniority for relationship lending, *Journal of Financial Intermediation*, 9(1), 57-89.

Manove, M., Padilla, A., and Pagano, M., 2001. Collateral versus project screening: A model of lazy banks, *Rand Journal of Economics*, 32(4), 726-744.

Marquez, R., 2002. Competition, adverse selection, and information dispersion in the banking industry, *Review of Financial Studies*, 15(3), 901-926.

Menkhoff, L., Neuberger, D., and Suwanaporn, C., 2006. Collateral-based lending in emerging markets: Evidence from Thailand, *Journal of Banking and Finance*, 30(1), 1-21.

Ono, A., and Uesugi, I., 2009. Role of collateral and personal guarantees in relationship lending: Evidence from Japan's SME loan market, *Journal of Money, Credit and Banking*, 41(5), 935-960.

Petersen, M., and Rajan, R., 1995. The effect of credit market competition on lending relationships, *Quarterly Journal of Economics*, 110(2), 407-444.

Presbitero, A., and Zazzaro, A., 2011. Competition and relationship lending: Friends or foes?, *Journal of Financial Intermediation*, 20(3), 387-413.

Qian, J., and Strahan, P., 2007. How laws and institutions shape financial contracts: The case of bank loans, *Journal of Finance*, 62(6), 2803-2834.

Qian, J., Strahan, P., and Yang, Z., 2015. The impact of incentives and communication costs on information production and use: Evidence from bank lending, *Journal of Finance*, 70(4), 1457–1493.

Rajan, R., 1992. Insiders and outsiders: The choice between informed and arm’s-length debt, *Journal of Finance*, 47(4), 1367-400.

Santos, J. and Winton, A., 2008. Bank loans, bonds, and information monopolies across the business cycle, *Journal of Finance*, 63(3), 1315–1359.

Schenone, C., 2010. Lending relationships and information rents: Do banks exploit their information advantages?, *The Review of Financial Studies*, 23(3), 1149-1199.

Sharpe, S., 1990. Asymmetric information, bank lending, and implicit contracts: A stylized model of customer relationships, *Journal of Finance*, 45(4), 1069-87.

Smith, J.K., 1987. Trade credit and informational asymmetry. *Journal of Finance*, 41 (4), 863-872.

Wilde, J., 2000. Identification of multiple equation probit models with endogenous dummy regressors. *Economics Letters*, 69 (3), 309-312.

Wooldridge, J.M., (2010). *Econometric Analysis of Cross Section and Panel Data*, Second Edition, The MIT Press.

Yasuda, A., 2007. Bank relationships and underwriter competition: Evidence from Japan, *Journal of Financial Economics*, 86(2), 369-404.

Xu, B., van Rixtel, A., and van Leuvensteijn, M., 2013. Measuring bank competition in China: A comparison of new versus conventional approaches applied to loan markets, Bank for International Settlements, Working Paper No.422.

Table I: Summary statistics and variable definition

Variable	Definition	N	Mean	S.D	Min	Max
Panel A: Market structure						
ACR4	The market share (in terms of assets) of the top four banks in the province. Measured at one semi-accounting year prior to current loan.	9288	0.55	0.06	0.35	0.97
Panel B: Firm characteristics						
Size	Natural logarithm of total assets in millions of RMB deflated to year 2006 value. Measured at one semi-accounting year prior to current loan.	8779	7.67	1.16	4.01	12.72
Leverage	Outstanding debt/total assets, measured at one semi-accounting year prior to current loan.	8779	0.56	0.19	0.02	2.37
ROA	Return on assets, measured at one semi-accounting year prior to current loan.	8779	0.06	0.07	-0.44	1.71
Age	Natural log of firm age. Firm age is the difference in months between the firm's establishment date and the loan initiation date.	9288	5.03	0.40	2.77	6.62
Tangibility	(Net property, plants and equipment)/total assets, measured at one semi-accounting year prior to current loan.	8779	0.27	0.19	0.00	0.92
FT	= 1 if majority stake is owned by the State, and 0 otherwise.	9288	0.40	0.49	0	1
Liquidity	Current assets/total assets, measured at one semi-accounting year prior to current loan.	8779	0.55	0.23	0.01	1
Loanconcen	Loan concentration ratio. Defined as Loansize / (Loansize and debt outstanding).	8779	0.04	0.07	0.00	0.93
IPO	= 1 if loan is issued after the IPO, and 0 otherwise.	9288	0.83	0.37	0	1
Panel C: Loan characteristics						
Collateral	= 1 if loan is secured by collateral, and 0 otherwise.	9288	0.66	0.47	0	1
Maturity	Natural log of loan maturity. Measured in months.	9288	3.25	0.79	0.00	5.70
Spread	Difference between lending rate and benchmark deposit rate of corresponding maturity. Measured in percentage.	9288	2.85	1.21	0.71	13.60
Loansize	Natural log of loan size. Measured in millions of RMB deflated to year 2006 value.	9288	3.13	1.41	-3.70	8.97
Panel D: Relationship variables						
Numlender	Number of different lenders the firm has borrowed from prior to origination of current loan.	9288	3.93	3.45	0	28
Sizeconcen	The amount of loans that a firm has borrowed from its current lender as a proportion of the total amount of loans it obtained prior to the current loan.	9288	0.33	0.35	0	1
Numconcen	The number of loans that a firm has borrowed from its current lender as a proportion of the total number of loans it borrowed prior to the current loan.	9288	0.34	0.34	0	1
First	= 1if the current loan is the first loan borrowed from this lender, and 0 otherwise.	9288	0.24	0.43	0	1
Switch	= 1 if the current loan is borrowed from the same lender as the previous loan, and 0 otherwise.	9288	0.40	0.49	0	1
Panel E: Monetary and regional macroeconomic variables						
RRR	Reserve Requirement Ratio for the month when the loan is issued.	9288	0.17	0.03	0.10	0.21
Repo	7-day repo rate for the month when the loan is issued, in percentage.	9288	2.55	1.21	0.94	6.92
CPI	Provincial consumer price index, measured at one semi-account year prior to current loan.	9288	1.03	0.03	0.98	1.10
NPLratio	Provincial non-Performing loan ratio, measured at one semi-account year prior to current loan.	9288	0.03	0.03	0.00	0.21
Realgdpindex	Provincial real GDP growth rate, measured at one semi-account year prior to	9288	0.09	0.03	0.01	0.18

	current loan					
Panel F: Instrumental variables						
Amaturity	((current assets/total assets)*(current assets/cost of goods sold)+(fixed assets/total assets)*(fixed assets/depreciation))/1000	9288	10.68	6.64	0.18	55.33
dd_lag2	The number of CSRC IPO suspension days during the 2-year window prior to listing date.	9288	188.6	168.8	0	523
Affected_Firms	Dummy variable equals 1 if firm experienced at least one CSRC IPO suspension during the 2-year window prior to listing date.	9288	0.68	0.47	0	1
Termspread	Yield difference between 5-year Treasury bond and 1-year Treasury bond, for the month when the loan is issued, in percentage.	9288	0.86	0.44	-0.19	1.54
Localavrate	People's Bank of China reports on a yearly basis the percentage of loans that are issued below/at/above the corresponding benchmark rate. The actual lending rate to benchmark rate ratio is classified in seven groups: [0.9,1], [1], [1.0-1.1], [1.1-1.3],[1.3-1.5],[1.5-2.0] and [above 2.0]. We take the middle value of each group and calculate the weighted average ratio using the percentage of loans within each group as weight. This weighted average is then multiplied with the one-year reference rate to calculate the regional average lending rates. Measured at one semi-account year prior to the current loan. In percentage.	9288	6.79	0.94	5.14	9.88
Benchsprd	Benchmark lending rate minus benchmark deposit rate of corresponding maturity, for the month the loan is issued. In percentage.	9288	2.42	0.55	1.4	3.78
Panel G: Additional variables						
Numalst	Number of analysts following the firms measured at one semi-accounting year before loan origination.	7719	11.01	10.90	0	66
Instishare	Percentage of shares held by institutional investors measured at one semi-accounting year before loan origination, in percentage.	7367	29.07	22.03	0	96.33
Multiapp	Dummy variable that equals 1 if firm applied for its IPO multiple times before eventually listed, and 0 if succeeded in the first IPO application.	9288	0.05	0.22	0	1
Affected_Loans	Dummy variable equals 1 if the loan is borrowed by firms that experienced CSRC IPO suspension during the suspension periods.	9288	0.15	0.36	0	1

Table II: Univariate tests

	Panel A: Sizeconcen			Panel B: ACR4			Panel C: IPO		
	<Median	>=Median	Mean diff	<Median	>=Median	Mean diff	Pre-IPO	Post-IPO	Mean diff
Relationship variables									
Sizeconcen	--	--	--	0.32	0.35	-0.02***	0.40	0.32	0.08***
Numconcen	0.22	0.73	-0.51***	0.33	0.35	-0.02***	0.41	0.33	0.08***
Numlender	4.65	3.21	1.44***	4.41	3.46	0.96***	2.17	4.29	-2.11***
Market structure									
ACR4	0.55	0.55	-0.00*	-	-	-	0.56	0.55	0.01***
Loan characteristics									
Collateral	0.66	0.66	-0.00	0.62	0.70	-0.08***	0.86	0.62	0.24***
Maturity	3.19	3.32	-0.13***	3.26	3.25	0.00	3.12	3.28	-0.16***
Spread	2.99	2.70	0.30***	2.87	2.82	0.04*	2.85	2.85	0.01
Loansize	3.19	3.07	0.12***	3.17	3.10	0.08**	2.32	3.30	-0.97***
Firm characteristics									
FT	0.42	0.39	-0.03**	0.42	0.39	0.03***	0.11	0.46	-0.35***
Liquidity	0.55	0.54	0.01*	0.60	0.50	0.10***	0.58	0.54	0.04***
Total Assets	7.76	7.58	0.18***	7.81	7.53	0.28***	6.32	7.85	-1.53***
Leverage	0.57	0.55	0.02***	0.55	0.57	-0.02***	0.55	0.56	-0.00
ROA	0.07	0.06	0.00	0.06	0.07	-0.00	0.15	0.05	0.09***
Age	5.04	5.02	0.02***	5.06	5.00	0.06***	4.70	5.10	-0.40***
Tangibility	0.27	0.27	-0.01*	0.24	0.31	-0.07***	0.27	0.27	-0.01

*** p<0.01, ** p<0.05, * p<0.1.

Table III: Collateral determinants and borrower information transparency

Panel A shows the results for the estimation of *Equation (1)*. M.E are the marginal effects calculated on the basis of the results in Column (1). Panel B estimates *Equation (2)*. It reports the impact of $Sizeconcen_{il}$ and $ACR4_{il}$ on collateral incidence differentiated by the informational transparency of borrowers ($Infor_{il}$), which is defined by three proxies: *Borrower ownership* (*FT*=1 if state owned and 0 otherwise); *Listed Board* (*Listmain*=1 if listed in the main board and 0 otherwise); and *Firm Size* (*Medianta*=1if *log(total assets)* is above the provincial median and 0 otherwise). Panel C estimates *Equation (2)* using stock market information production (*Numalst* and *Instishare*) as measures of informational transparency of borrowers. The sample is restricted to post-IPO loans for Column (6) and (7). In all panels, the control variables include firm characteristics, loan contract terms, monetary policy variables, regional macroeconomic variables and a set of fixed effects, including *Industry*, *Province*, *Banktype* and *Loan-year* dummies. In column (2), *Maturity* and *Spread* are excluded for endogeneity concerns. Removing these terms in Panel B and C do not affect our results, which are available upon request. Results for fixed effects dummies are not reported to save space. The equations are estimated with the Probit model. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	Panel A: Main Effects			Panel B: Borrower Information Transparency			Panel C: Stock Market Infor Production	
	With contract terms	Without contract terms	M.E of model (1) (%)	Board of listing	Ownership	Firm size	Numalst	Instishare
VARIABLES	(1)	(2)		(3)	(4)	(5)	(6)	(7)
Sizeconcen	0.153**	0.170**	1.40	0.231***	0.256***	0.287***	0.209**	0.277***
	(0.068)	(0.068)		(0.085)	(0.082)	(0.076)	(0.088)	(0.097)
ACR4	2.685***	2.623***	4.45	3.826***	3.463***	3.482***	4.912***	4.897***
	(0.805)	(0.802)		(0.895)	(0.858)	(0.832)	(0.901)	(0.924)
Listmain*Sizeconcen				-0.129				
				(0.098)				
FT*Sizeconcen					-0.203**			
					(0.098)			
Medianta*Sizeconcen						-0.390***		
						(0.102)		
Numalst*Sizeconcen							-0.010**	
							(0.005)	
Instishare*Sizeconcen								-0.770***
								(0.240)
Listmain*ACR4				-1.664***				
				(0.616)				
FT*ACR4					-1.603***			
					(0.619)			
Medianta*ACR4						-2.051***		
						(0.571)		
Numalst*ACR4							-0.149***	
							(0.032)	
Instishare*ACR4								-4.924***
								(1.318)
Listmain				0.705**				
				(0.346)				
Medianta						1.334***		
						(0.316)		
Numalst							0.074***	
							(0.017)	
Instishare								2.574***
								(0.722)
FT	-0.606***	-0.594***	-16.7	-0.565***	0.335	-0.618***	-0.597***	-0.568***

	(0.047)	(0.046)		(0.048)	(0.340)	(0.047)	(0.050)	(0.050)
First	0.036	0.049	0.94	0.048	0.044	0.019	-0.030	-0.042
	(0.056)	(0.055)		(0.056)	(0.056)	(0.056)	(0.059)	(0.059)
Switch	-0.028	-0.064	-0.75	-0.033	-0.028	-0.023	-0.020	-0.023
	(0.039)	(0.039)		(0.040)	(0.039)	(0.039)	(0.042)	(0.042)
IPO	-0.412***	-0.387***	-10.39	-0.322***	-0.391***	-0.405***		
	(0.071)	(0.071)		(0.073)	(0.071)	(0.071)		
Numlender	0.024***	0.018**	2.13	0.021***	0.024***	0.020***	0.027***	0.024***
	(0.007)	(0.007)		(0.007)	(0.007)	(0.008)	(0.008)	(0.008)
Liquidity	-0.458***	-0.545***	-2.76	-0.504***	-0.447***	-0.375**	-0.558***	-0.689***
	(0.155)	(0.153)		(0.156)	(0.155)	(0.155)	(0.168)	(0.167)
Size	-0.221***	-0.215***	-7.29	-0.191***	-0.222***	-0.233***	-0.163***	-0.217***
	(0.027)	(0.027)		(0.028)	(0.027)	(0.030)	(0.033)	(0.030)
Leverage	0.941***	1.049***	4.53	1.040***	0.926***	0.951***	0.891***	0.963***
	(0.127)	(0.126)		(0.129)	(0.127)	(0.127)	(0.138)	(0.137)
ROA	-1.134***	-1.084***	-2.22	-1.124***	-1.102***	-1.160***	-0.583*	-0.704**
	(0.277)	(0.282)		(0.279)	(0.278)	(0.276)	(0.330)	(0.325)
Age	-0.415***	-0.432***	-4.50	-0.331***	-0.419***	-0.409***	-0.385***	-0.422***
	(0.058)	(0.057)		(0.060)	(0.058)	(0.058)	(0.064)	(0.064)
Tangibility	-0.852***	-0.891***	-4.43	-0.893***	-0.855***	-0.782***	-1.028***	-1.021***
	(0.179)	(0.178)		(0.180)	(0.179)	(0.179)	(0.189)	(0.188)
Maturity	0.169***		3.39	0.169***	0.169***	0.171***	0.187***	0.200***
	(0.028)			(0.028)	(0.028)	(0.028)	(0.030)	(0.030)
Spread	0.031*		1.00	0.036**	0.031*	0.035**	0.021	0.023
	(0.017)			(0.017)	(0.017)	(0.017)	(0.018)	(0.018)
Loansize	-0.089***	-0.070***	-3.37	-0.090***	-0.090***	-0.090***	-0.095***	-0.095***
	(0.020)	(0.020)		(0.020)	(0.020)	(0.020)	(0.021)	(0.021)
Loanconcen	1.830***	1.921***	3.37	1.956***	1.804***	1.866***	1.779***	1.672***
	(0.413)	(0.408)		(0.410)	(0.414)	(0.415)	(0.440)	(0.434)
RRR	-0.071	-0.021	-0.05	0.050	-0.202	-0.188	0.645	0.422
	(2.902)	(2.884)		(2.909)	(2.904)	(2.907)	(3.068)	(3.068)
Repo	0.048*	0.045*	1.51	0.044	0.048*	0.050*	0.054*	0.047*
	(0.027)	(0.027)		(0.027)	(0.027)	(0.027)	(0.029)	(0.029)
CPI	1.475	2.003	1.04	1.241	1.320	1.518	2.608	2.614
	(1.510)	(1.501)		(1.514)	(1.513)	(1.513)	(1.601)	(1.597)
NPLratio	-0.535	-0.647	-0.42	-0.305	-0.526	-0.685	-0.414	-0.121
	(1.135)	(1.132)		(1.137)	(1.135)	(1.140)	(1.183)	(1.179)
Realgdpindex	1.097	1.548	1.00	0.763	0.787	0.975	1.606	1.198
	(1.435)	(1.429)		(1.441)	(1.442)	(1.439)	(1.500)	(1.496)
Constant	-0.566	-0.644		-1.577	-0.850	-1.123	-7.478	-6.924
	(1.874)	(1.869)		(1.888)	(1.879)	(1.884)	(106.776)	(106.273)
Observations	8,741	8,753		8,741	8,741	8,741	7,620	7,620
Pseudo R2	0.287	0.283		0.289	0.288	0.290	0.291	0.291
H_0:Sizeconcen+Infor*Sizeconcen=0				0.102	0.052	-0.103		
H_0: ACR4+Infor*ACR4=0				2.162***	1.860**	1.431		

Table IV: Identify informational rents through IPOs

This table reports estimates based on various versions of *Equation (3)*. Column (1) to Column (3) add the interaction terms $Sizeconcen_{il}*IPO_{il}$ and $ACR4_{il}*IPO_{il}$ progressively. Column (4) excludes the potentially endogenous contract terms *Spread* and *Maturity* and re-estimates Column (3). M.E. are marginal effects based on Column (3). For variables interacting with IPO_{il}, we report marginal effects of said variable from before and after the IPO. Results for control variables and fixed effects dummies are not reported to save space. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

VARIABLES	(1)	(2)	(3)	(4)	M.E. of Model (3)
Sizeconcen	0.493**	0.169**	0.596***	0.604***	4.78
	(0.215)	(0.069)	(0.218)	(0.218)	
ACR4	2.806***	5.617***	5.935***	5.931***	8.51
	(0.807)	(1.201)	(1.216)	(1.211)	
Sizeconcen*IPO	-0.369		-0.471**	-0.463**	1.17
	(0.226)		(0.229)	(0.228)	
ACR4*IPO		-3.218***	-3.503***	-3.574***	4.15
		(1.000)	(1.016)	(1.012)	
First	0.423**	0.203	0.478**	0.462**	10.78
	(0.194)	(0.143)	(0.195)	(0.195)	
First*IPO	-0.430**	-0.190	-0.485**	-0.454**	-0.19
	(0.201)	(0.144)	(0.203)	(0.203)	
Switch	0.177	0.153	0.175	0.133	4.14
	(0.126)	(0.126)	(0.126)	(0.126)	
Switch*IPO	-0.218*	-0.189	-0.215	-0.207	-1.06
	(0.132)	(0.132)	(0.132)	(0.132)	
Numlender	-0.000	-0.023	0.009	-0.002	0.78
	(0.033)	(0.028)	(0.033)	(0.033)	
Numlender*IPO	0.025	0.051*	0.016	0.021	2.34
	(0.034)	(0.029)	(0.034)	(0.034)	
IPO	-0.132	1.396**	1.914***	1.951***	-7.13
	(0.206)	(0.572)	(0.627)	(0.626)	
Constant	-1.063	-2.417	-2.936	-3.025	
	(1.886)	(1.946)	(1.964)	(1.959)	
Fixed effects dummies	Industry, Province, Bank Type, Time				
Other loan contract terms	Yes	Yes	Yes	No	
Controls variables	firm characteristics, monetary policy and regional macro variables				
Observations	8,741	8,741	8,741	8,753	
Pseudo R2	0.288	0.289	0.289	0.285	
H_0:Sizeconcen+Sizeconcen*IPO=0	0.124*		0.124*	0.141**	
H_0: ACR4+ACR4*IPO=0		2.399***	2.431***	2.357***	
H_0:First+First*IPO=0	-0.007	0.013	-0.007	0.008	
H_0:Switch+Switch*IPO=0	-0.041	-0.036	-0.039	-0.074*	

Table V: Informational rents and firm risk

This table investigates how informational rents vary with firm risk. Firm risk is proxied by a dummy variable *Multiapp* that equals one if the firm applied multiple times before eventually being listed, and zero if being listed in its first IPO application. Column (1) tests the main effect of *Multiapp*. Column (2) introduces three-way interaction terms among informational rent variables (*Sizeconcen* and *ACR4*), listing status (*IPO*) and *Multiapp*. For these two columns, other control variables are the same as in Table III (Column (1)). Column (3) and (4) removes progressively loan contract terms and monetary and regional macroeconomic variables. Results of control variables and fixed effects dummies are not reported to save space. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

VARIABLES	(1)	(2)	(3)	(4)
Sizeconcen	0.600***	0.634***	0.648***	0.646***
	(0.219)	(0.225)	(0.225)	(0.225)
ACR4	5.979***	6.073***	6.081***	5.741***
	(1.217)	(1.254)	(1.249)	(1.226)
Sizeconcen*IPO	-0.476**	-0.532**	-0.526**	-0.526**
	(0.229)	(0.236)	(0.235)	(0.235)
ACR4*IPO	-3.558***	-4.368***	-4.441***	-4.419***
	(1.016)	(1.060)	(1.055)	(1.054)
Multiapp	0.286***	0.730	0.925	0.820
	(0.094)	(2.131)	(2.093)	(2.098)
Sizeconcen*Multiapp		-0.462	-0.497	-0.510
		(0.471)	(0.465)	(0.465)
ACR4*Multiapp		-1.493	-1.856	-1.647
		(3.676)	(3.608)	(3.617)
Multiapp*IPO		-4.872**	-4.873**	-4.791**
		(2.364)	(2.327)	(2.331)
Sizeconcen*Multiapp*IPO		0.944*	0.959*	0.974*
		(0.552)	(0.546)	(0.546)
ACR4*Multapp*IPO		9.315**	9.305**	9.143**
		(4.085)	(4.019)	(4.026)
IPO	1.962***	2.347***	2.384***	2.379***
	(0.627)	(0.650)	(0.647)	(0.647)
Constant	-2.854	-2.794	-2.904	-0.632
	(1.963)	(1.972)	(1.967)	(0.925)
Fixed effects dummies	Industry, Province, Bank Type, Time			
Firm characteristics	Yes	Yes	Yes	Yes
Other loan contract terms	Yes	Yes	No	No
Monetary policy variables	Yes	Yes	Yes	No
Regional macro variables	Yes	Yes	Yes	No
Observations	8,741	8,741	8,753	8,753
Pseudo R2	0.290	0.293	0.289	0.289

Table VI: Bivariate Probit Models

This table reports the results of recursive Bivariate Probit models with instrumental variables. Column (1) replicates the Probit model results of Table IV, column (3) for comparison purposes. Column (2) and (3) treat *IPO* as endogenous variable. Column (4) treats relationship lending dummy *Rel* as endogenous variable, where *Rel* is a dummy variable equals 1 if the firm obtains at least 20% (i.e. the sample median of the *Sizeconcen*) of bank loans from the lender prior to the current loan, and 0 otherwise. In all specifications, the variables in the *Collateral Equation* correspond to the ones used in Table IV, column (3), except that in Column (4) where *Sizeconcen* and *Sizeconcen*IPO* are replaced by *Rel* and *Rel*IPO*, respectively. Variables in the *IPO Equation* include one instrument (*Affected_Firms* or *Log(1+dd_lag2)*) and all variables in the *Collateral Equation*, except *IPO* and its interaction terms with other covariates. Variables in the *Relationship Equation* include one instrument (*Localavrate*) and all variables in the *Collateral Equation*, except *Rel*, *Rel*IPO*, relationship control variables (*Relcontrols* defined in section 2.1.1), and their interactions with *IPO*. The instrumental variables are defined as following: *Affected_Firms* is a dummy variable equals 1 if the firm has experienced at least one CSRC IPO suspension within the 2-year window prior to the firm's actual listing; *Log(1+dd_lag2)* is the logarithm of 1 plus the number of CSRC IPO suspension days within the 2-year window prior to the firm's actual listing; *Localavrate* is the regional average lending rate one semi-accounting year before the current loan. Full results of Bivariate Probit models are available upon request. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	Probit	Bivariate probit		Bivariate Probit
		IPO as endogenous		*Rel* as endogenous
		IV: Affected_Firms	IV: Log(1+dd_lag2)	IV: Localavrate
VARIABLES	(1)	(2)	(3)	(4)
Collateral Equation				
Sizeconcen (Rel)	0.596***	0.589***	0.589***	1.314***
	(0.218)	(0.217)	(0.217)	(0.247)
ACR4	5.935***	5.873***	5.848***	4.999***
	(1.216)	(1.214)	(1.214)	(1.178)
Sizeconcen*IPO (Rel*IPO)	-0.471**	-0.460**	-0.460**	-0.521***
	(0.229)	(0.228)	(0.228)	(0.148)
ACR4*IPO	-3.503***	-3.487***	-3.469***	-3.198***
	(1.016)	(1.013)	(1.012)	(0.935)
IPO Equation				
Affected_Firms		-0.681***		
		(0.094)		
Log(1+dd_lag2)			-0.080***	
			(0.016)	
Relationship Equation				
Localavrate				0.115***
				(0.040)
ρ		-0.129 (p=0.12)	-0.114 (p=0.17)	-0.508***(p=0.002)
Observations	8741	8,765	8,765	8765

Table VII: Firm fixed effects

This table reports the results for the fixed effects Logit model for alternative samples, and for specifications with and without loan contract terms. Results for firm characteristics and fixed effects dummies are not reported to save space. Monetary policy variables and regional macro variables are not included in this estimation. Including them does not change our results. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	Fixed effects Logit model			
	All loans		*Loans originated before corporate bond IPOs*	
	Without loan contract terms	With loan contract terms	Without loan contract terms	With loan contract terms
VARIABLES	(1)	(2)	(3)	(4)
Sizeconcen	1.645***	1.634***	1.750***	1.713***
	(0.543)	(0.544)	(0.542)	(0.543)
ACR4	23.247***	24.007***	23.356***	24.055***
	(5.305)	(5.284)	(5.337)	(5.309)
Sizeconcen*IPO	-1.472***	-1.453**	-1.774***	-1.722***
	(0.564)	(0.565)	(0.567)	(0.568)
ACR4*IPO	-17.824***	-18.051***	-19.251***	-19.548***
	(5.210)	(5.177)	(5.209)	(5.169)
First	1.074***	1.080***	1.292***	1.287***
	(0.389)	(0.388)	(0.397)	(0.395)
First*IPO	-1.209***	-1.199***	-1.547***	-1.527***
	(0.400)	(0.399)	(0.410)	(0.408)
Switch	0.407	0.448	0.325	0.374
	(0.300)	(0.299)	(0.303)	(0.302)
Switch*IPO	-0.472	-0.476	-0.365	-0.368
	(0.311)	(0.310)	(0.316)	(0.315)
Numlender	0.023	0.033	0.063**	0.075**
	(0.028)	(0.029)	(0.030)	(0.030)
IPO	10.171***	10.272***	10.954***	11.097***
	(2.978)	(2.959)	(2.978)	(2.954)
Observations	5,856	5,851	4,816	4,811
Number of firms	291	291	255	255
Pseudo R2	0.137	0.142	0.138	0.144
H_0:Sizeconcen+Sizeconcen*IPO=0	0.173	0.181	-0.024	-0.009
H_0: ACR4+ACR4*IPO=0	5.423***	5.967***	4.105*	4.506*

Internet Appendix for "Do banks extract informational rents through collateral?"

This appendix provides technical details and results of the propensity score matching analysis (Section 4.2) and the instrumental variable Probit model (Section 6.2). Results of propensity score matching analysis are reported in Table IA.II and Table IA.III. Results controlling for endogeneity of loan contract terms are reported in Table IA.V. Moreover, details and results are reported of several additional tests discussed in Section 4.2 ("difference-in-difference" tests, Table IA.I.) and Section 4.3 (corporate bond underwriting and concurrent lending, Table IA.IV), and for alternative samples such as conducted in Section 6.3 (Table IA.VI-VII).

A. Propensity score matching

This section presents the technical details of propensity score matching (e.g. Heckman et al., 1998). We divide our sample into two subsamples: pre-IPO loans and post-IPO loans, with the former presumably subjected to a higher degree of information asymmetries for non-relationship banks. Within each subsample, we estimate the propensity score of loans borrowed from relationship lenders using a Logit model. Specifically, for each sample, we regress the relationship dummy on the following covariates: *ACR4*, *FT*, *Liquidity*, *Size*, *Leverage*, *ROA*, *Age* and *Tangibility.*[36] For the sake of robustness, we further expand the covariates list by introducing their square terms.[37] Relationship dummies equal one if *Sizeconcen* is greater or equal to the sample median of the respective samples (0.25 for the pre-IPO sample and 0.19 for the post-IPO sample, respectively). Then we match each relationship loan (treatment group) with a (set) of non-relationship loans (control group) that have the

[36] Estimates on propensity scores are available upon request.
[37] The main purpose of propensity score estimation is not to predict selection into treatment as good as possible, but to balance all covariates (Augurzky and Schmidt, 2000).

closest propensity scores to that specific relationship loan. The average treatment effects of relationship intensity on collateral incidence are expected to be significantly positive for the pre-IPO loans, and moderated or insignificant for the post-IPO loans.

To compute the average treatment effects, two alternative matching methods are used, i.e. "nearest neighbor" matching and "kernel" matching. We drop all loans that are outside of the common support to minimize the potential bias introduced by these loans. Bootstrap standard errors based on 50 replications are reported.

Next, we test the balancing property of covariates. The estimated average treatment effects are biased if the covariates determining participation in the treatment group are not sufficiently balanced. The standardized bias of Rosenbaum and Rubin (1985) is a common statistic to test the balance of the distribution of the covariates in both the control and treatment groups. For brevity, we only report the mean bias of the matched sample.[38] Several other overall balancing tests including the pseudo-R^2, Rubin's B and Rubin's R are also reported. All of these diagnoses confirm that the covariates of the matched sample are balanced. In more detail: the mean bias for the matched sample is below the 5% threshold; the pseudo-R^2 for the matched sample is fairly low; Rubin's B is below 25 thresholds for most of the cases, and Rubin's R is within [0.5, 2].[39] Results are reported in the Internet Appendix Table IA.II.

Finally, we test the sensitivity of our results to unobserved variables that affect both relationship lending and collateral incidence. Rosenbaum (2002) developed a bounding approach to address whether or not inference about treatment effects may be affected by unobserved factors. We focus on pre-IPO loans, because as noted by Hujer et al. (2004), sensitivity analysis for insignificant treatment effects is not meaningful. Results are reported in Internet Appendix Table IA.III. Taking into account that the estimated treatment effect is positive for pre-IPO loans, the lower bounds (*Q_mh-*) – under the assumption that the true treatment effect has been underestimated – are less interesting

[38] The standardized biases of individual covariates are available upon request.

[39] Sianesi (2004) suggests that a low pseudo-R^2 for the post matching sample is an indicator of balanced matching. Rubin's B is the absolute standardized difference of the means of the linear index of the propensity score in the treated and matched sample. Rubin's R is the ratio of treated to matched variances of the propensity score index. Rubin (2001) recommends that Rubin's B is less than 25 and Rubin's R lies between 0.5 and 2 for the samples to be sufficiently balanced.

(Becker and Caliendo, 2007). Therefore, we focus on the upper bounds (*Q_mh+*). We report the Rosenbaum bounds for propensity score model II with the nearest neighbor matching (*NN(20)*). The results for the bounds are similar for propensity score model I and other matching methods. The critical level e^{γ}, at which one would question the positive effect of relationship lending on collateral incidence, is 1.85, a fairly large value by normal standards (see e.g. Bharath et al., 2011, for further discussion). Note that a critical value of 1.85 does not mean that relationship lending has no effect on collateral incidence and that unobserved heterogeneity exists. It only states that the confidence interval for the treatment effect would include zero if unobserved variables caused the odds ratio of relationship lending to differ between relationship borrowers and non-relationship borrowers by a factor 1.85. We conclude that it is unlikely that our causal inference of the positive effect of relationship lending on collateral incidence for pre-IPO loans could be challenged by powerful unobserved variables.

B. Endogeneity of loan contract terms: IV Probit model

This section addresses the endogeneity issue of loan contract terms using IV Probit estimations. We examine two possibilities: exclude *Spread* from the determinants of collateral and treat *Maturity* as the sole endogenous variable; and treat both *Spread* and *Maturity* as endogenous variables.[40] Our choices of instruments are guided by the existing literature and the specific characteristics of Chinese banking regulation. For *Maturity*, we follow Barclay et al. (1995) and employ asset maturity (*Amaturity*) as instrument, as firms may match their debt maturity with that of their assets to mitigate agency costs.[41] In addition, as proposed in Dennis et al. (2000) and Brick and Ravid (1985), loan maturity is expected to be positively related to the slope of the yield curve, proxied by the term spread (*Termspread*). This spread is defined as the yield difference between the 5- and 1-year government bonds for the month when the loan was originated. Regarding the lending spread, we use as instrument the benchmark loan spread (*Benchsprd*) for maturities that correspond with that of loan *l* in the month of the loan origination (*Benchsprd* = benchmark lending rate minus the benchmark deposit rate). Another instrument we introduce is the lagged regional average lending rate (*Localavrate*), measured at one semi-accounting year before the current loan. *Benchsprd* and *Localavrate* should be correlated with the actual lending spread, but they are not likely to be related to whether a particular loan is collateralized or not. Summary statistics and definitions of these instrumental variables are in Panel F of Table I.

Results of the IV Probit model are reported in Internet Appendix, Table IA.V. Column (1) excludes *Spread* from the determinants of collateral and treats *Maturity* as the sole endogenous variable, whereas Column (2) treats both *Spread* and *Maturity* as endogenous variables. Newey's

[40] The existing literature differs in treating which of the loan contract terms should be endogenous in determining collateral. Dennis et al. (2000) and Bharath et al. (2011) consider *Maturity* as the only endogenous contract term that affects collateral incidence. The underlining assumption is that the lending spread is determined after the decision on collateral pledging. On the other hand, Brick and Paila (2007) and Ono and Uesugi (2009) model the spread as an endogenous determinant of collateral. As empirical validations are provided for both assumptions and theoretical advantages of either assumption are unknown a priori, we examine both.

[41] Bharath et al. (2011) and Barclay et al. (2003) provide in-depth discussions of the validity of using asset maturity as an instrument for debt maturity. We follow Li et al. (2009) in defining asset maturity. See Table I, Panel F for definitions. Missing data for asset maturity is replaced by the industry median.

efficient two-step estimator is employed to obtain coefficient estimates for both specifications. The relevance and validity of our instruments in the IV Probit model are reported at the bottom rows.[42] In both Column (1) and (2), the null hypotheses that *Maturity* alone or *Maturity* and *Spread* together are exogenous are strongly rejected (Wald-test p-value=0.0192 and 0.0000, respectively), validating the IV Probit approach. Nevertheless, the IV Probit results are largely consistent with our previous findings, except that $Sizeconcen_{il}$ loses its explanatory power for post-IPO loans ($H_0: Sizeconcen_{il} + Sizeconcen_{il} * IPO_{il} = 0$ cannot be rejected, p-value=0.99 or 0.86 depending on specifications), which is an even stronger result than the one obtained in our baseline model. Results for market structure are also similar to our previous findings. The results of the conditional likelihood-ratio (*CLR*) test, *K* test and Anderson-Rubin Chi square test (*AR*) all reject the null hypothesis that the coefficients of the endogenous regressors in the structural equation are (jointly) zero. We also conduct the *J* statistics test, which assesses the validity of the instruments, i.e. the null hypothesis is that the instruments are uncorrelated with the error term. In both Column (1) and (2), the *J* statistics are statistically insignificant, confirming the validity of our instruments for the endogenous loan contract term *Maturity*, or for both *Maturity* and *Spread*.

[42] See Finlay and Magnusson (2009) for details on weak instrument robustness tests for limited dependent variable models.

References:

Augurzky, B. and Schmidt, C., 2000. The propensity score: A means to an end, Working Paper, University of Heidelberg.

Barclay, M., Marx, L., and Smith, C., 1995. The maturity structure of corporate debt, *Journal of Finance*, 50(2), 609-631.

Barclay, M., Marx, L., and Smith, C., 2003. The joint determination of leverage and maturity, *Journal of Corporate Finance*, 9(2), 149-157.

Becker, S., and Calendo, M., 2007. Sensitivity analysis for average treatment effects, *The Stata Journal*, 7(1), 77-83.

Bharath, S., Dahiya, S., Saunders, A., and Srinivasan, A., 2011. Lending relationship and loan contract terms, *Review of Financial Studies*, 24(4), 1141-1203.

Brick, I., and Palia, D., 2007. Evidence of jointness in the terms of relationship lending, *Journal of Financial Intermediation*, 16(3), 452-476.

Brick, I., and Ravid, A., 1985. On the relevance of debt maturity structure, *Journal of Finance*, 40(5), 1423-1437.

Dennis, S., Nandy, D., and Sharpe, L., 2000. The determinants of contract terms in bank revolving credit agreements, *Journal of Financial and Quantitative Analysis*, 35(1), 87-110.

Finlay, K., and Magnusson, L., 2009. Implementing weak-instrument robust tests for a general class of instrumental-variables models, *The Stata Journal*, 9(3), 398-421.

Heckman, J., Ichimura, H., and Todd, P., 1998. Matching as an econometric evaluation estimator, *Review of Economic Studies*, 65(2), 261-294.

Hujer, R., Caliendo, M., and Thomsen, S., 2004. New evidence on the effects of job creation schemes in Germany: A matching approach with threefold heterogeneity, *Research in Economics*, 58(4), 257-302.

Li, K., Yue, H., and Zhao, L., 2009. Ownership, institutions, and capital structure: Evidence from China, *Journal of Comparative Economics*, 37(3), 471-490.

Ono, A., and Uesugi, I., 2009. Role of collateral and personal guarantees in relationship lending: Evidence from Japan's SME loan market, *Journal of Money, Credit and Banking*, 41(5), 935-960.

Rosenbaum, P., 2002. *Observational studies*, 2nd edition, Springer-Verlag, New York.

Rosenbaum, P., and Rubin, D., 1985. Constructing a control group using multivariate matched sampling methods that incorporate the propensity score, *The American Statistician*, 29(1), 33-38.

Rubin, D., 2001. Using propensity score to help design observational studies: Application to the tobacco litigation. *Health Services & Outcomes Research Methodology*, 2, 169-188.

Siansei, B., 2004. An evaluation of the active labour market programmes in Sweden, *The Review of Economics and Statistics*, 86(1), 133-155.

Table IA.I Difference-in-Difference

This table reports the difference-in-difference tests in key risk factors for post- and pre-IPO samples (*post-IPO-pre-IPO*) and for both relationship dependent and non-dependent firms. Relationship dependent firms are the ones with *Sizeconcen* greater or equal to the sample median, while the rest are non-dependent firms. *** p<0.01, ** p<0.05, * p<0.1.

	Mean differences (*post-IPO-pre-IPO*)							
	ROA	Leverage	Tangibility	Liquidity	Size	Maturity	Spread	Loansize
Relationship dependent Firms	-0.03***	0.21***	0.08***	0.23***	4.04***	0.14**	0.25**	1.06***
Relationship non-dependent firms	-0.05***	0.20***	0.11***	0.15***	3.94***	0.19***	-0.06	0.83***
Difference-in-Differences	-0.01	-0.00	0.03	-0.08***	-0.11	0.04	0.19	-0.23

Table IA.II: Selection of observables – Propensity score matching on relationship lending.

This table reports average treatment effects of relationship lending on collateral incidence for pre-IPO and post-IPO loans. Propensity Score Model I in Panel A employs the following variables: *ACR4*, *FT*, *Liquidity*, *Size*, *Leverage*, *ROA*, *Age* and *Tangibility*. The Propensity Score Model II in Panel B includes all variables used in Panel A and the square terms of these variables (except the square term of *FT*). Logit regression is adopted in both panels. Bootstrap standard errors based on 50 replications are reported. NN(20) and NN(50) are the nearest neighbor matching estimators with 20 and 50 nearest neighbors. Epanechnikov kernel with bandwidth 0.06 is applied for the kernel matching estimator. Observations of common support are discarded. All balancing tests are based on matched samples. *** p<0.01, ** p<0.05, * p<0.1.

Panel A: Propensity Score Model I						
	Pre-IPO loans			Post-IPO loans		
	NN(20)	NN(50)	Kernel	NN(20)	NN(50)	Kernel
ATE	0.126***	0.116***	0.115***	-0.005	-0.005	-0.008
Std.Err.	(0.033)	(0.032)	(0.029)	(0.012)	(0.014)	(0.012)
Pseudo R2	0.006	0.004	0.006	0.002	0.002	0.001
Mean Bias	4.7	3.2	4.3	3.1	2.5	2.0
Rubin's B	17.6	15.0	18.3	10.7	9.5	8.0
Rubin's R	0.99	1.16	1.01	1.28	1.46	1.36
Panel B: Propensity Score Model II						
ATE	0.103***	0.102***	0.108***	0.007	-0.002	0.002
Std.Err.	(0.033)	(0.036)	(0.037)	(0.015)	(0.014)	(0.011)
Pseudo R2	0.013	0.013	0.007	0.002	0.002	0.002
Mean Bias	3.3	4.4	3.3	1.8	1.4	1.9
Rubin's B	27.0*	27.4*	20.2	11.0	11.2	9.8
Rubin's R	1.16	1.23	1.04	1.42	1.41	1.44

Table IA.III: Sensitivity test-Rosenbaum bounds.

This table reports results for the Rosenbaum bounds test for Propensity Score Model II with nearest neighbor matching (NN(20)). e^{γ} is the odds of differential assignment due to unobserved factors. *Q_mh+* and *Q_mh-* are the upper and lower bounds of the Mantel-Haenszel statistic. With increasing e^{γ}, the bounds move apart, reflecting uncertainty about the test-statistics in the presence of hidden bias. *p_mh+* and *p_mh-* are significance levels for upper and lower bounds.

e^{γ}	Q_mh+	Q_mh-	p_mh+	p_mh-
1	4.51	4.51	0.00	0.00
1.05	4.24	4.78	0.00	0.00
1.1	3.98	5.04	0.00	0.00
1.15	3.74	5.29	0.00	0.00
1.2	3.51	5.53	0.00	0.00
1.25	3.29	5.77	0.00	0.00
1.3	3.07	6.00	0.00	0.00
1.35	2.87	6.22	0.00	0.00
1.4	2.68	6.43	0.00	0.00
1.45	2.49	6.64	0.01	0.00
1.5	2.31	6.84	0.01	0.00
1.55	2.13	7.04	0.02	0.00
1.6	1.97	7.23	0.02	0.00
1.65	1.80	7.42	0.04	0.00
1.7	1.64	7.60	0.05	0.00
1.75	1.49	7.78	0.07	0.00
1.8	1.34	7.95	0.09	0.00
1.85	1.20	8.13	0.12	0.00

Table IA.IV: Corporate bond underwriting and concurrent lending

This table reports the results for samples of loans issued before corporate bond IPOs using the Probit model. Column (1) reports results for the full sample. Column (2) report results for a sample of firms that borrowed both before and after their equity IPOs. In both columns, loans borrowed after corporate bond IPOs are excluded. Results for firm characteristics and fixed effects dummies are not reported to save space. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	Loans before corporate bond IPOs	
	All firms	Firms that borrowed both before and after equity IPO
VARIABLES	(1)	(2)
Sizeconcen	0.642***	1.531***
	(0.190)	(0.326)
ACR4	4.651***	12.911***
	(1.228)	(2.637)
Sizeconcen*IPO	-0.511**	-0.813**
	(0.201)	(0.398)
ACR4*IPO	-3.777***	-4.129*
	(1.022)	(2.460)
First	0.542***	1.083***
	(0.154)	(0.252)
First*IPO	-0.562***	-1.079***
	(0.160)	(0.288)
Switch	0.106	0.500***
	(0.121)	(0.188)
Switch*IPO	-0.182	-0.592***
	(0.128)	(0.222)
Numlender	0.027***	0.111***
	(0.008)	(0.028)
IPO	2.086***	3.371**
	(0.601)	(1.425)
FT	-0.631***	-0.731***
	(0.052)	(0.255)
Constant	-0.341	-7.682
	(0.920)	(182.973)
Fixed effects dummies	Industry, Province, Bank Type, Time	
Firm characteristics	Yes	Yes
Monetary policy variables	No	No
Regional macro variables	No	No
Other contract terms	No	No
Observations	7,453	1,606
Pseudo R2	0.270	0.401
H_0:Sizeconcen+Sizeconcen*IPO=0	0.131*	0.719***
H_0: ACR4+ACR4*IPO=0	0.875	8.781***

Table IA.V: Endogeneity of loan contract terms

This table reports IV Probit regression results, treating other loan contract terms as endogenous variables. Column (1) treats *Maturity* as the sole endogenous variable, assuming that *Spread* does not affect collateral incidence. Column (2) treats both *Spread* and *Maturity* as endogenous variables. The instruments for *Maturity* are asset maturity (*Amaturity*) and term spread (*Termsprd*). Instruments for *Spread* are the lagged local average lending rate (*Localavrate*) and benchmark loan spread (*Benchsprd*). Definitions and summary statistics for these instrumental variables are reported in Table I, Panel F. Results for fixed effects dummies and first stage estimations of IV Probit regression are not reported to save space. They are available upon request. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	IV Probit	
VARIABLES	(1)	(2)
$\widehat{Maturity}$		0.597**
		(0.273)
$\widehat{Spread}$	0.996**	0.746***
	(0.426)	(0.271)
Sizeconcen	0.503**	0.591**
	(0.250)	(0.242)
ACR4	4.972***	5.279***
	(1.314)	(1.320)
Sizeconcen*IPO	-0.501**	-0.608**
	(0.251)	(0.253)
ACR4*IPO	-3.013***	-3.364***
	(1.099)	(1.098)
First	0.394*	0.345
	(0.217)	(0.223)
First*IPO	-0.446**	-0.480**
	(0.225)	(0.228)
Switch	0.530***	0.358**
	(0.191)	(0.146)
Switch*IPO	-0.392***	-0.368**
	(0.149)	(0.148)
Numlender	0.076	-0.016
	(0.049)	(0.039)
Numlender*IPO	-0.018	0.021
	(0.040)	(0.037)
IPO	1.648**	1.920***
	(0.684)	(0.683)
FT	-0.671***	-0.534***
	(0.067)	(0.056)
Liquidity	0.090	-0.242
	(0.329)	(0.201)
Size	-0.260***	-0.172***
	(0.038)	(0.036)
Leverage	0.372	0.667***
	(0.262)	(0.155)
ROA	-1.460***	-1.077***
	(0.335)	(0.351)
Age	-0.452***	-0.521***
	(0.071)	(0.064)
Tangibility	-0.587**	-0.788***
	(0.284)	(0.222)
Loansize	-0.200***	-0.107***

	(0.060)	(0.024)
Loanconcen	1.471***	1.665***
	(0.523)	(0.475)
RRR	-3.191	-0.083
	(3.273)	(3.755)
Repo	0.045	0.068**
	(0.030)	(0.031)
CPI	-1.791	-1.389
	(1.949)	(1.839)
NPLratio	0.891	-0.905
	(1.309)	(1.382)
Realgdpindex	-1.625	-0.290
	(1.858)	(1.647)
Constant	0.385	-2.186
	(2.193)	(2.468)
Observations	8,159	8,159
Fixed effects dummies	Industry, Province, Bank Type, Time	
H_0:Sizeconcen+Sizeconcen*IPO=0	0.002 (p=0.99)	-0.017 (p=0.86)
H_0: ACR4+ACR4*IPO=0	1.959**(p=0.03)	1.914** (p=0.03)
H_0:First+First*IPO=0	-0.052 (p=0.46)	-0.136 (p=0.14)
H_0:Switch+Switch*IPO=0	0.138 (p=0.21)	-0.009 (p=0.86)
Wald test (p-value)	Chi2(1)=5.48 (0.0192)	Chi2(2)=20.36 (0.0000)
CLR (p-value)	6.12 (0.0146)	23.94 (0.0000)
K (p-value)	Chi2(1)=6.12 (0.0134)	Chi2(2)=23.23 (0.0000)
J (p-value)	Chi2(1)=0.00 (0.9488)	Chi2(2)=1.81 (0.4041)
AR (p-value)	Chi2(2)=6.12 (0.0469)	Chi2(4)=25.04 (0.0000)

Table IA.VI: Alternative samples – Firms which borrowed both before and after IPO

This table reports the results for a sample of firms that borrowed both before and after their equity IPOs. Panel A reports results for all loans. Panel B further restricts this sample to loans around IPO dates. Results for firm characteristics and fixed effects dummies are not reported to save space. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	Firms borrowed both before and after IPO			
	Panel A: All loans	Panel B: Loans around IPOs dates		
		One loan before and one after equity IPO	Two loans before and two after equity IPO	Three loans before and three after equity IPO
VARIABLES	(1)	(2)	(3)	(4)
Sizeconcen	1.532***	2.293**	1.099**	1.173***
	(0.324)	(0.921)	(0.534)	(0.441)
ACR4	12.211***	14.652	11.515**	7.357*
	(2.543)	(9.731)	(5.416)	(4.284)
Sizeconcen*IPO	-0.713*	-1.108	-1.165*	-1.076*
	(0.394)	(1.208)	(0.683)	(0.552)
ACR4*IPO	-4.224*	-0.766	-8.850*	-8.722**
	(2.405)	(8.008)	(4.901)	(4.031)
First	1.121***	2.497***	1.439***	0.854**
	(0.251)	(0.842)	(0.499)	(0.378)
First*IPO	-1.069***	-1.086	-1.351**	-0.860*
	(0.286)	(0.873)	(0.564)	(0.447)
Switch	0.491***	-0.815	-0.049	0.277
	(0.188)	(0.623)	(0.381)	(0.283)
Switch*IPO	-0.588***	-0.138	-0.423	-0.465
	(0.220)	(0.831)	(0.494)	(0.376)
Numlender	0.114***	0.367**	0.176**	0.131**
	(0.027)	(0.158)	(0.084)	(0.058)
IPO	3.353**	1.439	6.374**	6.165***
	(1.394)	(4.649)	(2.837)	(2.332)
FT	-0.683***	-5.019***	-2.392***	-1.880***
	(0.244)	(1.291)	(0.555)	(0.410)
Constant	-7.514	-14.636	-12.967	-8.227
	(159.820)	(326.925)	(326.330)	(242.200)
Fixed effects dummies	Industry FE, Province FE, Bank Type FE, Time FE			
Firm characteristics	Yes	Yes	Yes	Yes
Firm fixed effects	No	No	No	No
Monetary policy variables	No	No	No	No
Regional macro variables	No	No	No	No
Other loan contract terms	No	No	No	No
Observations	1,663	215	421	564
Pseudo R2	0.403	0.553	0.452	0.364
H_0:Sizeconcen+Sizeconcen*IPO=0	0.819***	1.184	-0.066	0.096
H_0: ACR4+ACR4*IPO=0	7.987***	13.886	2.665	-1.365

Table IA.VII: Alternative samples – Excluding non-commercially viable loans

This table reports results for samples of loans provided by different types of banks. We exclude progressively loans that are less likely to be issued on a commercial basis. The model specification is based on *Equation (4)* excluding: *Maturity*, *Spread*, monetary variables and regional macroeconomic variables. Including these variables does not affect our results. Column (1) excludes loans borrowed from state-owned banks (SOCBS). Column (2) excludes loans from policy banks (PBs). Column (3) excludes loans from both policy banks and state-owned banks. Column (4) further excludes loans borrowed from trust and investment companies (TICs). Column (5) further excludes loans from other financial companies (Other), which leaves loans from joint-stock commercial banks, city commercial banks, rural commercial (cooperative) banks and foreign banks remaining. Results for firm characteristics and fixed effects dummies are not reported to save space. The equation is estimated with the Probit model. Standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

	Excluding SOCBs	Excluding PBs	Excluding SOCBs&PBs	Excluding SOCBs&PBs &TICs	Excluding SOCBs&PBs &TICs&Other
VARIABLES	(1)	(2)	(3)	(4)	(5)
Sizeconcen	1.323***	0.556***	0.792**	0.958**	0.957**
	(0.344)	(0.194)	(0.368)	(0.388)	(0.393)
ACR4	11.231***	5.047***	10.203***	9.115***	9.108***
	(1.972)	(1.212)	(2.076)	(2.173)	(2.195)
IPO*Sizeconcen	-1.229***	-0.409**	-0.680*	-0.722*	-0.706*
	(0.358)	(0.203)	(0.385)	(0.406)	(0.412)
IPO*ACR4	-7.334***	-3.153***	-6.541***	-4.875***	-5.358***
	(1.718)	(1.022)	(1.779)	(1.871)	(1.881)
First	0.703***	0.501***	0.528**	0.682***	0.616**
	(0.227)	(0.157)	(0.246)	(0.257)	(0.262)
IPO*First	-0.673***	-0.474***	-0.446*	-0.605**	-0.550**
	(0.234)	(0.162)	(0.254)	(0.264)	(0.269)
Switch	0.316*	0.030	0.070	0.077	0.110
	(0.190)	(0.123)	(0.207)	(0.217)	(0.221)
IPO*Switch	-0.444**	-0.126	-0.263	-0.277	-0.308
	(0.200)	(0.129)	(0.217)	(0.227)	(0.231)
Numlender	0.024**	0.026***	0.034***	0.027**	0.028**
	(0.009)	(0.007)	(0.010)	(0.011)	(0.012)
IPO	4.511***	1.731***	3.773***	2.811**	3.039***
	(1.017)	(0.604)	(1.065)	(1.120)	(1.127)
FT	-0.520***	-0.565***	-0.440***	-0.477***	-0.476***
	(0.070)	(0.048)	(0.075)	(0.083)	(0.084)
Constant	-9.580	-0.111	-8.429	-8.433	-6.706
	(165.908)	(0.917)	(95.904)	(92.578)	(80.646)
Fixed effects dummies	Industry, Province, Bank Type, Time				
Firm characteristics	Yes	Yes	Yes	Yes	Yes
Firm fixed effects	No	No	No	No	No
Monetary policy variables	No	No	No	No	No
Regional macro variables	No	No	No	No	No
Other loan contract terms	No	No	No	No	No
Observations	4,098	8,273	3,573	3,274	3,132
Pseudo R2	0.313	0.286	0.317	0.322	0.312
H0: ACR4+IPO*ACR4=0	3.897***	1.894**	3.662***	4.239***	3.750***
H0:Sizeconcen+IPO*Sizeconcen=0	0.094	0.147*	0.112	0.238*	0.251*

Collateral and the disruption of firm as non-financial intermediary: Evidence from Chinese Property Law

Bing Xu[1]

Universidad Carlos III de Madrid

Abstract

By allowing large classes of movable assets to be used as collateral, the Chinese Property Law reform transformed firms' role as non-financial intermediaries in China. We find after the legal reform, firms relied on trade credit financing could substitute to more bank credit. Accordingly, the providers of trade credit reduced significantly their provision of trade credit. In particular, we find the Property Law has disrupted the practice in which firms borrow short-term debts and redistribute them via trade credit. After the reform, instead of providing trade credit, firms started to accumulate more fixed asset investment, which in turn allowed for more long-term borrowing from banks. Our findings are not driven by confounding factors such as liquidity drain due to financial crisis. Our results also cannot be explained by other important reforms which were introduced around the same time. Our paper highlights the importance of looking at other financing channels when investigating the effect of collateral laws.

JEL Classification: G21; G28; G32; K22.
Key Words: Collateral; Trade Credit, Movable Assets; Property Law.

[1] The author is indebted to Bank of Finland (BOFIT) for hosting his visit during which part of this research was written. The author thanks Ricardo Mora, Georges Siotis, Adrian van Rixtel, Mauricio Larrain, Mikko Makinen, Honglin Wang, participants of the Conference on China's Financial Markets and the Global Economy (BOFIT and City University of Hong Kong), European Central Bank China Experts Network Annual Meeting, and seminar prticipants of Banco de España and Universidad Carlos III de Madrid for very helpful comments and suggestions. All remaining errors are my own.

1. Introduction

Theories and empirical studies highlight the links between enforceability of secured contracts and access to external finance (e.g. Calomiris et al., 2017; Haselmann, Pistor and Vig, 2010; Lilienfeld-Toal, Mookherjee and Visaria, 2012; Vig, 2013), and the availability of collateral and debt capacity in the presence of contract incompleteness (e.g. Stiglitz and Weiss, 1982; Hart and Moore, 1994). Studies find that the lack of sufficient collateral is a key obstacle to access to external finance across countries. The problem is aggravated in countries with weak collateral laws, because inadequate legal infrastructure excludes important asset types as permissible collateral classes. In general, these studies have found better collateral laws could promote firms' access to bank credit. However, this literature has largely ignored that some firms served as non-financial intermediaries by providing trade credit to their customers, particularly because of weak collateral law prohibits some firms' access to bank credit. In this regard, changing collateral law could affect not only credit transaction between banks and firms, but also credit transaction between firms. This link has largely been left out in the collateral law and finance literature. In this article, we intend to fill this void.

Trade credit is widely used by firms to finance their own purchase of inputs (as accounts payable), as well as provide financial support to their customers (accounts receivable). Previous studies find that trade credit plays crucial roles when bank credit is not available (e.g. Petersen and Rajan, 1997, Nilsen, 2002, Fisman and Love 2003), and it is particularly important in developing countries where legal and financial institutions are less developed (Demirguc-Kunt and Maksimovic, 2001). Linking these literatures, if better legal and financial institutions improve access to bank credit, they could also reduce the firms' reliance on trade credit. In addition, if some firms served as non-financial intermediaries and redistribute their obtained bank credit via trade credit, such redistribution should decline under better collateral law, given that better collateral law could facilitate more direct financing from banks. This is the main hypothesis we take to data.

To this end, we explore the passage of the Property Law in China in 2006. The Property Law allowed a broad class of movable assets as permissible collateral for bank credit (see section 2), including in particular accounts receivable, which is registered as firms' provision of trade credit. Therefore, by allowing accounts receivable to be pledged as collateral, the Property Law potentially can affect both trade credit and firms' access to bank credit.

The crucial element that differentiates our work from others is that we take into account the role of trade credit when discussing the impact of the collateral law. We argue a legal reform such as the Property Law, which could plausibly change the access to bank credit, should in turn influence the provision or demand of trade credit. In particular, we consider firms as receivers or providers of trade credit, and both types of firms make their decisions on trade credit depending on the availability of short-term bank credit. Specifically, for the receivers of trade credit, bank credit serves as a plausibly less expensive substitute, and therefore if more short-term bank credit is available, they might reduce their reliance on trade credit and substitute to more bank credit. For suppliers of trade credit, on the one hand, having easier access to bank credit may allow them to extend more trade credit. On the other hand, if their clients could also access more bank credit directly, the suppliers of trade credit may face negative demand shock, resulting less provision of trade credit. Therefore, when accessing the effect of collateral law, it is crucial to take into account the fact that credit transactions occur not only between banks and firms, but also among firms.

As the Property Law reform only pertains to movable assets, it should affect more the firms with intensive use of these assets. This policy wrinkle allows for an investigation in a difference-in-differences framework in which the effect of the Property Law is evaluated as a function of firms' *pre-reform* movable assets, controlling for other firm characteristics, firm fixed effects and year fixed effects. Specifically, we compare various outcome of interest before and after the passage of the Property Law as a function of firms' pre-reform position of movable assets. The difference-in-differences framework allows us to control observed and unobserved factors that could affect equally firms with different level of movable assets. We also control industry specific (province specific) time-variant shocks by introducing interaction between industries and time (provinces and time). As a result, our specification compares firms with different level of moveable assets within the same industry (province).

The crucial assumption for the difference-in-differences strategy is that treated and control group should behave similarly in the absence of the shock, that is, the “parallel trend” assumption. Although cannot be tested, we provide supporting evidence on its validity by investigating placebo (non-exist) reforms took place before the actual passage of the Property Law. To further mitigate concerns that there might be latent unobserved diverging trends in the variable of interest, we also augment our specification with linear treatment trend (*Treated*Trend*). The underlining assumption is that the *changes* in the outcome of interest for treated group would have been the same as that of the control group in the absence of the shock. Inclusion of linear treatment trend usually leads to underestimation of treatment effect, as part of the effects could be absorbed by the linear trend. Crucial for difference-in-difference strategy is that the shock needs to be plausibly exogenous and unanticipated. Otherwise, firms may have adjusted their asset mix in anticipation of the reform, which could compromise the core of our empirical strategy. Berkowitz, Lin, and Ma (2015) provide one of the first empirical evidences that the passage of the Property Law was a complete shock to stock market. As will be discussed in more detail in section 3.2, anticipating the passage of the Property Law could not have been possible. Nevertheless, to control the possibility that some firms may have inside information when the Property Law was heavily debated in 2006 before its final passage, we remove observation from 2006. We further conduct additional robustness test by removing also observations from 2005.

We investigate the effect of the Property Law using a large sample of listed firms during 2001-2011. Focusing on public listed firms could to some extent alleviate an identification challenge, that is, to differentiate demand and supply side arguments. Specifically, for this sample of large firms that are less likely to be financially constrained, their needs to pledge movable assets in order to access bank finance are relatively low. On the contrary, for vast population of unlisted firms, which are smaller and more likely to be financially constrained, the Property Law provides the opportunity to expand access to bank credit and reduce reliance on trade credit. Hence we speculate for our sample of listed firms, the potential increase in the provision of trade credit would be outweighed by the fall in the demand for trade credit from the vast population of unlisted firms. Using data from a sample of Chinese listed firms during 2001-2011, we find that passage of the Property Law significantly changed firms' reliance and provision of trade credit, and the relationship between bank credit and trade credit.

First, consistent with substitution theory (Biais and Gollier, 1997; Burkart and Ellingsen, 2004), we find that firms relied on trade credit financing reduced their usage of trade credit after the Property Law, and could substitute almost entirely their financing needs with short-term bank credit. In addition, the substitution effect is much stronger for the ex-ante financially constrained firms, consistent with the causal explanation that the substitution was caused by relaxation of access to bank credit due to the Property Law.

Second, our findings suggest that providers of trade credit reduce more their provision (accounts receivable) after the Property Law, consistent with the demand side argument that better access to bank credit due to the Property Law reduces the importance of trade credit financing, and therefore forces the suppliers to reduce their provision of trade credit. These findings are in contrast to the supply side argument: suppliers pledge their movable assets to obtain more bank credit after the Property Law, and better access to bank credit could allow the suppliers to provide more trade credit. This finding is mostly driven by our sample of large public traded firms, which are less financially constrained compared to unlisted small and medium sized firms, and therefore are less likely to reply on movable assets to access bank credit.

Third, we find the Property Law disrupted a practice call the "redistribution of bank credit via trade credit". Specifically, before the legal reform, firms that borrowed more short-term bank credit provided more trade credit, suggesting these firms redistribute bank credit to their clients via trade credit. The redistribution was disrupted after the Property Law, consistent with our conjecture that better collateral law could reduce firms' role as financial intermediaries.

Finally, we document implications for corporate asset structure and debt maturity. Our analysis suggests the providers of trade credit shifted away from short-term assets and increased their investments in fixed assets. Likewise, consistent with theory that firms match asset and debt maturity (e.g. Myers, 1977; Milbradt and Oehmke, 2014), they also increased their holdings in long-term bank credit. These findings suggest that when bank credit becomes more available, the providers of trade credit could reduce their provision of trade credit and consequently redirect more resources for investment. In addition, we do not find the providers of trade credit to increase their short-term bank credit after the reform. The finding is consistent with our conjecture that after taking into account trade credit, the effect of the Property Law on short-term bank credit is muted. On the one hand, the providers of trade credit are able to borrow more short-term debt because they can now pledge their movable assets. On the other hand, once their customers can access bank credit more easily after the Property Law, their importance as non-financial intermediary decreases, in the sense that these firms stopped borrowing short-term debt to finance the provision of trade credit to their customers. The aggregate effect on short-term debt therefore becomes insignificant.

Our findings provide supporting evidence that better access to bank credit due to collateral law reform could change provision and usage of trade credit, and in particular, disrupt the redistribution of bank credit via trade credit. We further demonstrate these findings could not be explained by several alternative theories, other contemporary reforms, and these results continue to hold for additional robustness tests. First, we check if global financial crisis could drive our results. The global financial crisis may have drained the liquidity of banking sector, leaving firms with more outstanding short-term debt particularly vulnerable due to the difficulty to roll-over short-term debt, and consequently, they were unable to pass on scarce bank credit to their clients via trade credit. To investigate this possibility, we re-examine our hypotheses on a short-event window that covers only one year before and one year after the Property Law, which excludes the influence of global financial crisis. In addition, we also investigate precisely when the disruption of redistribution occurred during the post-reform period. We find that the disruption of the redistribution was strongest one year after the Property Law, which is before the global financial crisis. This finding implies that the global financial crisis is unlikely to be the main reason of our results. Our results are also not driven by other contemporary reforms, in particular, our results still hold after controlling an adjacent reform that reduced related-party transitions. Related-party transactions are usually inter-corporate loans provided by listed firms to their controlling shareholders or affiliates. Such transactions could confound our results as the providers of trade credit are prone to related party transition. Finally, we verify the robustness of our findings through a battery of additional tests, which ensure that our results could not be attributed to anticipation of the legal reform, changes in firm characteristics other than movable assets, specific definitions of trade credit, and alternative samples.

This article is closely related to the literature that investigates how legal reforms on collateral affect corporate financial policies, in particular firms' access to bank credit (e.g. Campello and Larrain, 2015; Aretz, Campello and Marchica, 2015; Cerqueiro, Ongena and Roszbach, 2015; Love, Martinez Peria and Sandeep, 2016; Calomiris et al., 2017). Our paper contributes to this literature by extending the effect of collateral law on trade credit, and the inter-play between trade credit and bank credit. Our results highlight the importance to analyze alternative sources of finance in order to have a more complete picture of the potential effects of collateral law reform.

Our paper also fits into the trade credit literature. Garcia-Appendini and Montoriol-Garriga (2013) shows during the credit crunch phase of global financial crisis, liquid firms, that is firms with more cash holding, could offer more trade credit to their less liquid customers. Love, Preve and Sarria-Allende (2007) investigate in a cross country sample of firms from emerging markets around financial crises, and find firms that are financially vulnerable to financial crises extend less trade credit to their customers, a result consistent with the "redistribution view" of trade credit. These studies explored the liquidity shock caused by financial crisis. Our paper complements this literature by exploring liquidity shock due to better collateral law. In this regard, our work is similar to Shenoy and Williams (2017), which explores the exogenous liquidity shocks caused by the U.S. interstate bank branching laws. Our paper also contribute to the literature that argues trade credit can offer alternative financial sources in developing countries where formal bank credit are scarce, and financial institutions are less developed (Fisman and Love 2003; Allen et al., 2005). In particular, we complement existing literature focus on trade credit in China (Ge and Qiu, 2007; Cull et al., 2009). These studies investigate the relationship between trade credit and bank credit, while ours explore how such relationship could change when legal institution improves.

The remainder of the paper is organized as follows. Section 2 provides a brief description of the institutional background governing secured transactions in China. Section

3 presents our hypotheses, identification strategy and data. Section 4 presents the main results. Section 5 discusses various alternative explanations and robustness tests. Finally, Section 6 concludes.

2. The institutional background

2.1 The Security Law

Before the approval of the Property Law by end of 2006, secured financing was governed by the 1995 Security Law. This law specifies certain types of *existing* movable assets which can be pledged as collateral. Non-possessory security interests were allowed only for the use of equipment and motor vehicles as collateral (under Article 34 of the Security Law). Other movable assets such as accounts receivables, future acquired properties, properties that cannot be fixed in type, quantity or location, could not serve as permissible collateral. The Security Law did not exclude inventory as permissible collateral; however, it could be used only as collateral by way of possessory security interests. In practice, the amount of inventory had to be fixed at the time of financing and was required to be relocated (or the ownership certificate had to be transferred) to creditors.

Furthermore, a secured interest had to be registered to be enforceable, while no centralized registration system existed. In China, numerous registries dealt with different types of collateral, and had ultimate discretion in rejecting or accepting the registration of secured interests. Moreover, these registries required collateral to be appraised and the legality of security agreements to be certified. As a result, creating and registering secured interests was costly, time consuming and subject to uncertainty. Another problem was that the Security Law did not provide clear rules on the determination of priority among competing claims on the same collateral. Secured lenders might have to compete with other claimants for underlying collateral, which in turn increased the cost of credit.

The limited permissible asset types and prohibitive process in creating and registering secured interest impeded secured transactions using movable assets as collateral. As a result, secured transactions strongly favored real property as security when lending to enterprises. A joint survey by People’s Bank of China and World Bank Group (2007) shows that less than 7% of loans in China were secured purely by movables assets, which were mostly inventories and equipment.[2]

2.2. The Property Law

On December 29, 2006, the 5th Session of the 10th Standing Committee of the National People's Congress (NPC) accepted a draft of the Property Law of the People's Republic of China. The Law was eventually passed on March 16th, 2007 and put into effect on October 1st of that year. The Property Law was supplemented by two additional implementation measures: the Measures for Chattel Mortgage Registration, issued by the State Administration of Industry and Commerce (SAIC); and the Measures for the Registration of Pledged Receivables, issued by the People's Bank of China. The former governs general movable properties while the latter governs receivables. These measures together with the Property Law provide detailed guidance on the scope of permissible collateral and registration systems for security interests.

Under the new law, the range of permissible security was greatly expanded, which now includes accounts receivable, existing and future production equipment, raw materials, semi-finished goods and inventories. The registration of security interests is also simplified: for general movable assets (except accounts receivable), the registration can be done at the local office of the SAIC for the county in which the debtor is domiciled, and it requires only basic information about the parties, the debt and the underlying security.

In addition, specific rules and registration systems are created to guide secured transactions in accounts receivable, which are arguably one of the most important movable assets classes. Account receivables are broadly defined in Chapter 17 of the Property Law as "… the right to require payment from debtors arising out of sales of goods, services or facilities, including existing and future monetary claims and proceeds, but not including those arising from negotiable instruments or other negotiable securities". The Measure for the Registration of Pledged Receivables provides further clarification by listing five types of accounts receivables as permissible collateral, including, but not limited to, the following: 1) claims from sales; 2) claims from leases; 3) claims from rendering services; 4) rights to charge fees from immovable property such as toll roads, bridges, tunnels, ferries, etc.; and 5)

[2] Source: People's Bank of China (PBOC)-FIAS-CPDF survey of financial institutions (The "Lender Survey"), p.56.

claims from granting loans or other credit. To facilitate the creation of secured interests in accounts receivables, the Credit Reference Centre of the People's Bank of China (Centre) is created as a centralized registration authority for the pledging of accounts receivables. The Centre also sets up a search system to publicize registration information of the pledge of accounts receivables, which allows lenders to obtain information about borrowers or other registered security interests. Apart from allowing more permissible collateral and establishing centralized registration systems, the Property Law also provides clearer references to the determination of priorities among competing claims on the same collateral. Specifically, priority is determined by the date of registration of security interests.

As the result of these legal changes, secured transactions against movable assets have expanded greatly. During 2008-2010, the number of loans backed by movable assets increased by 21% per year, while the value of loans increased by 24% per year. Since the creation of the Credit Reference Centre in 2007, more than 1.7 million receivable-backed loans have been recorded by the end of July 2015, or a remarkable annual growth rate of 51%. These loans amounted to 57 trillion RMB, among which 30 trillion was given to 220,000 small and medium-sized enterprises.[3]

3. Hypotheses, identification strategy and data

3.1. Literature and hypotheses

Previous literature have investigated how legal reforms that expanded pledgable asset categories could change firms' access to bank credit, asset composition, resource allocation, and industry composition. Campello and Larrain (2015) investigate the reforms in Eastern Europe that permitted the use of movable assets (e.g. machinery and equipments) as collateral, and find that such reforms promoted access to external finance, and reallocated assets and employments towards firms with more movable assets. Aretz, Campello and Marchica (2015) analyze the reform of the Napolenoic Code in France, and find that increased access to collateral – by expanding it to hard assets – increased firms' debt capacity and prolonged debt maturity. Cerqueiro, Ongena and Roszbach (2015) examine

[3] Source: Independent Evaluation of the IFC Secured Transactions Advisory Project in China (2011) and Credit Reference Center of People's Bank of China

legal reform in Sweden that reduced the value of collateral (e.g. floating liens). They show that such reform reduced debt capacity and shortened debt maturity, and eventually contributed to distortions in corporate investment and asset allocation. Love, Martinez Peria and Sandeep (2016) investigate the effects of the existence of collateral registries on access to finance across a large number of countries. Calomiris et al., (2017) demonstrate that in countries with weak movable collateral laws, lending is biased towards loans backed with immovable assets, and resource allocations across sectors are distorted towards immovable-based production and investment. In general, these studies support the view that the expanded capacity of collateralization and better enforceability of secured contracts facilitate firms' access to bank credit. However, this literature primarily focuses on accessibility to bank credit, while the effects of collateral law on alternative financial channels have largely been left out. In particular, some firms that had better access to bank credit due to weak collateral law served as non-financial intermediaries and redistributed their bank credit to their customers via trade credit. Having better collateral law could promote more direct financing from banks for those who relied on trade credit, which implies less demand for trade credit, and consequently less redistribution of bank credit through firms. These are potentially important effects of collateral law that have largely been ignored in previous studies.

Indeed, firms obtain credit from financial institutions as well as from other firms via trade credit. Burkart and Ellingsen (2004) provide one of the first theoretical models that explain that when banks ration credit, suppliers are often better positioned to provide credit. The suppliers extend trade credit because they have an advantage to overcome moral hazard and asymmetric information frictions with respect to banks. Moreover, suppliers obtain a markup on trade credit over their funding costs, making the extension of trade credit profitable from the suppliers' perspective. Suppliers may also extend trade credit because they have implicit equity stakes in their customers, and therefore willing to help financially constrained customers to overcome financial difficulties. From the borrowers' perspective, Petersen and Rajan (1995) find that, when confronted with bank lending constraints, firms are more inclined to borrow more expensive trade credit provided that investment returns exceed the cost of funding. Recent empirical studies have largely confirmed that accessibility to bank credit determines the supply and demand of trade credit.

Garcia-Appendini and Montoriol-Garriga (2013) investigates how a negative shock to bank credit caused by the global financial crisis affects demand and supply of trade credit. They find that liquidity-rich suppliers can extend more trade credit during the financial crisis when external finance is difficult to access. This finding supports the view that suppliers serve as liquidity providers (Cuñat, 2007), and trade credit is a substitute to bank credit (Biais and Gollier, 1997; Burkart and Ellingsen, 2004). In addition, a few papers investigate firms' financing choices with respect to legal institutions. Demirguc-Kunt and Maksimovic (2001) find that trade credit is more prevalent in countries with worse legal institutions. However, these studies usually rely on cross-country heterogeneity in legal institutions, which makes it more difficult to control country specific factors that could have driven the differences in trade credit. Shenoy and Willams (2017) explore an exogenous shock to banking liquidity due to the implementation of the Interstate Banking and Branching Efficiency Act in the United States. They provide evidence that legal reform which promotes bank competition improve banks' liquidity. Consequently more bank liquidity allows suppliers to provide more trade credit. Likewise, the customers who have access to bank liquidity can rely much less on trade credit. Their findings suggest that although trade credit are credit transactions in between firms, the accessibility of external bank credit plays a crucial role in firms' decisions of trade credit.

Motivated by these studies, our study tries to answer the following questions: Firstly, if collateral law improves access to bank credit, would firms substitute plausibly more expensive trade credit with bank credit? Secondly, how does collateral law affect the provision of trade credit? Thirdly, how does collateral law change the "redistribution of bank credit via trade credit", a practice where firms with easier access to bank credit redistribute their obtained credit to less fortunate firms via trade credit? Lastly, if collateral law changes the provision of trade credit, how do firms change their asset and debt compositions accordingly?

The Chinese financial system is particularly suitable for answering these questions. First of all, the main source of external financing in China is bank loan, while other funding resources such as equity market and corporate bond market represent only very small share of overall financing. Before 2007, the annual funds raised from equity and corporate bond markets represented less than 5% of GDP, while that from banks represented 90% of GDP.

Secondly, the Chinese banking sector is notorious for misallocating credit (Cull and Xu, 2005). State owned, politically connected, and large firms have preferential access to bank credit, while small and medium sized firms have difficulties in accessing bank loans. The lack of alternative financing channels implied that many Chinese firms have to rely on trade credit to finance their short term financing needs (Allen, Qian, and Qian, 2005; Ge and Qiu, 2007; Cull et al., 2009). Indeed, Cull et al., (2009) find for a sample of 100,000 large and medium sized Chinese industrial firms, accounts receivables ranges from 18% of total sale for private firms, and 36.5% for state owned enterprises. On average these figures are relatively higher than the 18.5% for the U.S. Compustat firms (Peterson and Rajan, 1997). Ge and Qiu (2007) investigates a smaller sample of survey data and reports that the accounts receivables and accounts payables represent 13% and 14% of firm assets, respectively. These figures suggest trade credit is important source of financing for Chinese firms. Lastly, as argued in the introduction and in section 3.2 in more detail, the passage of the Property Law provided an ideal exogenous shock on the quality of legal institution, which bears particularly importance on firms' choices of external financing.

We develop our arguments as below. First we look at the usage of trade credit. Following previous literature, collateral law that expanded pledgable asset type could improve firms' access to bank credit. In addition, according to the substitution view of trade credit and bank credit, improved access to bank credit reduces demand for trade credit, given that the latter is usually a more expensive substitute of bank credit (Biais and Gollier, 1997; Burkart and Ellingsen, 2004). Therefore, if the Property Law expands access to bank credit, the usage of trade credit should decrease disproportionally more for those who relied on trade credit financing before the reform, and consequently, those firms could substitute more their financing needs with short-term bank credit. To further establish the causal relationship between accessibility of external financial resources and the substitution between trade credit and bank credit, we explore firm heterogeneity in financial constraints. If as argued that the Property Law increases the availability of bank credit, it is expected to affect more financially constrained firms, in the sense that these firms should reduce more their reliance on trade credit and substitute more to short-term debt. In summary, we derive our first hypothesis:

Hypothesis 1: *The Property Law should allow trade credit dependent firms to substitute their usage of trade credit with more short-term bank credit; and the substitution effect should be more pronounced for ex-ante financially constrained firms.*

Next we look at the provision of trade credit. The impact of the Property Law on the provision of trade credit is determined by a net outcome of supply and demand. On the one hand, because the Property Law allows for movable assets to be pledged as collateral, the providers of trade credit may have an incentive to extend trade credit, given that the provision of trade credit (i.e. accounts receivable, which is a type of movable assets) can now be pledged against bank credit. On the other hand, as long as the customers of trade credit substitute trade credit with short-term bank credit, the providers of trade credit face a negative demand shock, and consequently, they may have to reduce their provision of trade credit. The aggregate impact of the Property Law on the provision of trade credit depends on both supply and demand side factors. These two forces are negatively correlated, and therefore, if we observe a decrease in the equilibrium quantity of trade credit provision, the demand side effect must outweigh the positive supply side effect. Without a matching sample of supplier and customer of trade credit, it is extremely hard to fully differentiate the demand and supply effects. Our sample of listed firms could alleviate this problem to some extent. Because these firms are generally large and have less difficulties in accessing external finance comparing to smaller unlisted firms, their needs to pledge movable assets in order to access bank finance are relatively low. On the contrary, for most unlisted firms, which are smaller and more likely to be financially constrained, the Property Law provides the opportunity to expand access to bank credit and reduce their reliance on trade credit. For these firms, substituting trade credit with bank credit could be cost saving as bank loans are usually cheaper, and banks are specialized in providing loan services. Hence we speculate that for our sample of listed firms, the effect of the Property Law on trade credit will predominantly go through the demand side: broadened access to finance reduces demand for trade credit, causing the suppliers to reduce their provision of trade credit. We proxy firms' capacity to provide trade credit using their pre-reform median level of movable assets to total asset ratio, with higher value indicating more capacity to provide trade credit.

Hypothesis 2: *Facing negative demand shock, firms with more pre-reform movable assets should reduce disproportionally more the provision of trade credit.*

Thirdly we investigate whether the Property Law could disrupt or moderate the practice called the "redistribution of bank credit via trade credit". Previous studies have found firms finance their provision of trade credit with short-term bank credit, in the sense that the suppliers borrowed short-term bank credit and redistribute it to their clients in the form of trade credit (Calomiris, Himmelberg and Wachtel, 1995; Petersen and Rajan, 1997; Nilsen, 2002). This practice emerges because the suppliers might have better access to bank credit, due to an advantage in overcoming information asymmetries with respect to banks, or the presence of credit rationings. The suppliers often are incentivized to provide support to their customers because they could obtain a markup on trade credit over their own funding costs (Ng et al., 1999; Klapper et al., 2012), or have implicit stakes in the survival of their clients (Wilner, 2000; Cuñat, 2007).

If the Property Law reduces the demand for trade credit, the suppliers' incentives to borrow short-term debt to finance the provision of trade credit must decrease, causing the disruption (moderation) of redistribution of bank credit via trade credit. On the other hand, as the Property Law also allowed the supplier of trade credit to access bank credit more easily, these firms could now redistribute more bank credit via trade credit. As argued before, for our sample of listed firms, demand should dominate supply. To this end, we first investigate whether the provision of trade credit as a function of lagged short-term debt (or pre-reform short-term debt) experiences a structural break around the Property Law. Following the "redistribution view", short-term bank credit should be positively related to the provision of trade credit before the reform, indicating these firms borrowed bank credit to finance the provision of trade credit. After the reform, if the clients of trade credit substitute their demand for trade credit with short-term bank loans, the suppliers' role as non-financial intermediary must decrease, which will lead to a disruption (moderation) of the redistribution. In other words, the correlation between lagged short-term debt (pre-reform short-term debt) and provision of trade credit should be less pronounced after the reform. Next, to reinforce our argument, we explore the pre-reform heterogeneity in movable assets. Specifically, since the redistribution of bank credit should mostly be conducted by the providers of trade credit, we expect the disruption (moderation) effect should be more pronounced for these firms. To this end, we employ the same proxy of suppliers of trade credit as before, that is the firms with high pre-reform median movable assets ratio, and investigate if the disruption of the redistribution is driven by these firms. We derive our third hypothesis:

Hypothesis 3: *The redistribution of bank credit via trade credit should be moderated or disrupted by the Property Law; and the moderation or disruption effect should be driven by the traditional providers of trade credit.*

Finally, we investigate other implications of the Property Law, particularly on firms' asset compositions and debt maturity. If the providers reduce their provision of trade credit after the reform, they could shift their asset composition towards other types of short-term assets, for instance cash holding, or prolong their asset maturity by investing more in fixed assets. Changes in asset composition could subsequently lead to shifts in capital structure. As we argue that the Property Law should affect directly firms' liability on the short-term end, we do not expect to see a direct impact on their long-term leverage. However, if firms match their asset and debt maturities, long-term leverage could also change accordingly.

The impact on short-term leverage is ex-ante ambiguous: On the one hand, firms with more movable assets could obtain short-term bank credit more easily after the Property Law, suggesting an increase in short-term leverage for these firms. On the other hand, these firms also redistributed bank credit via trade credit before reform. If these firms face a negative demand shock on trade credit, their incentive to borrow short-term bank credit to finance the provision of trade credit must decrease. The aggregate impact on short-term debt should be determined by the two opposite forces.

3.2.Identification strategy

We investigate these hypotheses with a difference-in-differences setting where we compare various outcomes of interests before and after the enactment of the Property Law as a function of firms' pre-reform level of movable assets. The identification hinges on the fact that the Property Law pertains to movable assets only, and therefore firms with ex-ante higher reliance on movable assets in their operations are more affected by this legal change. The advantage of the difference-in-differences approach is that it allows us to control observed and unobserved factors that could affect treaded and control firms alike. Like any method, it relies on crucially a few assumptions. Firstly, the pre-reform trends for the treated and control group must be similar, i.e. the parallel trend assumption. We provide supporting evidence on this assumption by investigating placebo (non-exist) reforms that took place before the actual passage of the Property Law. Finding non-zero effects on the outcome of interest would suggest violation of parallel trend assumption. To further mitigate concerns that there might be latent unobserved diverging trends in the variable of interest, we also augment our specification with linear treatment trend. The underlining assumption is that the *changes* in the outcome of interest for treated group would have been the same as that of the control group in the absence of the shock. Inclusion of linear treatment trend usually leads to underestimation of treatment effect, as part of the effects could be absorbed by the linear trend. Second, crucially for the difference-in-differences framework is that external shock needs to be plausibly exogenous and unanticipated. If firms could anticipate the legal reform, they could have adjusted their asset mix accordingly before the actual reform, and therefore compromise the core of our empirical methodology which relies on the pre-reform measure of movable assets. We argue it is very unlikely that firms could have anticipated the passage of the Property Law beforehand, as well as predicting precisely the content of the Property Law. According to Zhang (2008), various versions of the Property Law were discussed before the final version, and each version faced strong oppositions from the conservatives and was blocked. The Property Law therefore had to be redrafted many times, making it impossible for firms to plan their response accordingly, given that the actual content of the Property Law was unknown beforehand. In addition, the actual timing of the passage of the law was also difficult to predict. As stated in Berkowitz, Lin, and Ma (2015), the enactment of the Property Law constitutes an exogenous shock, because the exact timing of the passage of the law was a product of complex political decision-making. They provide solid empirical evidence that the announcement of the Property Law on December 29, 2006 was unexpected by comparing the stock market reaction on the announcement date and the rest of the trading days of 2006. Indeed, the Property Law came as quite a surprise: before its final passage, the law was actually withdrawn from the People's Congress in March of 2006 due to strong opposition from conservatives, which at that time suggested the new law was very unlikely to be accepted. On December 24, 2006, the standing committee of People's Congress conducted an unprecedented seventh reading of the law to discuss its suitability, suggesting that even 5 days before its approval, it was still uncertain whether or not the law would pass. And finally, when the Property Law was approved on Dec 29th, 2006, it shocked the stock market. These facts suggest that the passage of the Property Law was exogenous. Nevertheless, it is possible that some politically connected firms may have inside information about the potential passage of the law, and could have changed their asset mix in advance. To further mitigate such anticipation effect, we remove observations from 2006 when the Property Law was approved by the National People's Congress. Therefore in our main analysis we use observations before 2005, which is plausibly immune to any anticipation effect. We conduct further robustness test by further removing observations from 2005, and use only information from pre-2004.

We define movable assets as the sum of *Inventory* and *Accounts Receivable*. This definition captures the main groups of the assets that are allowed to be pledged as collateral after the reform. These firms were also more likely (or have more capacity or incentive) to provide trade credit. We argue that following the Property Law, these firms are better positioned to extend more trade credit using their pledgeable movable assets, or face a large negative demand side shock on trade credit and reduce their provisions instead. The movable ratio (*Movratio*) is defined as $(Inventory_t+Account\ Receivable_t)/Asset_t$. For each firm we calculate its pre-reform median movable ratio (calculated over 2001-2005), and interact it with a dummy indicating whether the year in question is before or after the passage of the law. Firms with higher pre-reform median movable ratio are supposed to be affected more by the Property Law. In addition to this continuous measure, we also provide nonparametric results where we divide pre-reform movable ratios into three equal sized bins (bottom 33%, mid 33% and top 33%), and interact each size bin dummy with the post-reform dummy. We expect stronger results for firms located in higher bin of pre-reform movable assets distribution. The generalized difference-in-differences specification is as follows:[4]

$$\mathrm{Y_{it}} = \alpha_\mathrm{i} + \gamma_\mathrm{t} + \beta\mathrm{Fmov_i} * \mathrm{After_t} + \rho\mathrm{Fmov_i} * \mathrm{Trend_t} + \delta\mathrm{X} + \varepsilon_\mathrm{it} \qquad (1)$$

where *i* indexes for firm and *t* for accounting year. Y_{it} represents various outcomes of interest. Firm fixed effects α_i control for time-invariant differences in firm behavior, while the time fixed effects γ_t control for aggregate time-varying shocks. $Fmov_i$ is a continuous measure of pre-reform movable ratio, with higher value indicating higher percentage of movable assets in the firm's asset mix. Since $Fmov_i$ is time invariant, the level of $Fmov_i$ is subsumed into fixed effects. $After_t$ is a binary variable that takes the value one for the years after the Property Law reform (2007-2011), and zero otherwise (2001-2005*).* $Trend_t$ is a linear trend. *X* denotes control variables including: $Size_{it-1}$, $Tangibility_{it-1}$, $Liquidity_{it-1}$, $Profitability_{it-1}$, $Sale_{it-1}$, Age_{it-1}, $List_{it}$, $Split_{it}$ and $State_{it}$. To mitigate endogeneity concerns, most of these controls enter the model with lagged values. ε_{it} is the error term. Following Bertrand, Duflo and Mullainathan (2004), standard errors are clustered at the firm level. The difference-in-differences estimator is β, which measures the effect of the Property Law as a function of pre-reform movable assets ratio.

One potential concern with this specification is that some industry specific shocks occurred around the passage of the law, and these industries have higher movable assets ratio. To address this issue, the baseline specification is augmented with *Industry-Year* fixed effects to control for time-varying industry specific shocks. Similarly, *Province-Year* fixed effects are included to control for time-varying regional economic shocks. We also conduct tests on placebo (non-exist) reforms to validate parallel trend assumption and rule out possibility of anticipation effect. Additional robustness tests to rule out confounding factors are discussed in section 5.

[4] $After_t$ and $Fmov_i$ do not enter the regression as stand-along variables because they are absorbed by year fixed effects and firm fixed effects.

3.3.Data

Our primary data source is the WIND Information database. This firm-level annual database contains for a sample of listed firms detailed balance sheets, income statements information and general information such as industry classification, location, established year, listed year, and ownership type. It also provides a detailed breakdown of firms' liabilities, including information on total debt, long-term debt, and short-term debt. WIND Information also provides detailed breakdowns of asset categories, which is crucial for the construction of movable assets ratio (discuss shortly).

Firms from financial industries, with missing values in total assets and with unclear industry classification are excluded from the sample. To be qualified in the sample, firms are also required to have annual reports both before and after 2006 (the year of the reform). The final sample includes firms listed in either the Shanghai Stock Exchange or the Shenzhen Stock Exchange from 2001 to 2011. As discussed before, we remove observations from the reform year 2006. In total, our database contains more than 12,000 firm-year observations from around 1,200 firms, covering firms located in all provinces of mainland China and from 58 industries.

We winsorized all continuous variables at 1% and 99%. Table 1 (Panel A) shows average movable assets represent 26% of total assets. A decomposition of the movable assets shows that inventories represent the lion share of movable assets (17% of total assets), while accounts receivable account for 9% of total assets. Firms' usage of trade credit is represented as accounts payable, which account for 9% of total assets. Panel B provides the summary statistics on the debt side. Total leverage is defined as the ratio of bank credit over lagged assets ($Debt_t/Asset_{t-1}$), where $Debt_t$ is the sum of long-term bank credit ($LongDebt_t$) and short-term bank credit ($ShortDebt_t$).[5] The average $Debt_t/Asset_{t-1}$ is 0.33, with standard deviation of 0.22. Total leverage is further decomposed into long-term leverage ($LongDebt_t/Asset_{t-1}$) and short-term leverage ($ShortDebt_t/Asset_{t-1}$). Average long-term leverage is 0.13 while that of short-term leverage is 0.21. These figures suggest that

[5] Following standard accounting standards, short-term debt is debt matures within one year, while long-term debt is debt with maturity longer than one year.

the majority of corporate debts are in short term.

Panel C reports summary statistics on the asset side. The mean value of *Log(1+Asset$_t$)* is 21.25, which translates to a book value of total assets of 1,693 million RMB. Finally, Panel D describes briefly the control variables employed in the analysis. *Tangibility* is defined as the ratio of fixed assets to total assets (*FixedAsset$_t$/Asset$_t$*). Average tangibility is 0.29. *Liquidity* is defined as cash divided by total assets (*Cash$_t$/Asset$_t$*). The average liquidity is 0.16. *Profitability* is defined as the ratio of net profits over total assets (*Netprofit$_t$/Asset$_t$*). *Sale* is the logarithm of one plus total sale. *Age* is defined as the logarithm of one plus the number of years since incorporation. The average age is 11 years. *List* is a dummy variable that equals one for firm-year observation after firm's IPO, and zero otherwise. *Split* is a dummy variable that equals one for firm-year observation after firm's completion of the split share reform, and zero otherwise. *State* is a dummy variable that equals one if the controlling shareholder is government and zero if the controlling shareholder is a private entity. 69% of the firms in the sample are state owned firms.

4. **Results**

4.1.Usage of trade credit and substitution to bank credit

We first investigate whether firms relied more on trade credit could reduce their usage of trade credit and substitute to more short-term bank credit. Specifically, we expect to see a higher decrease in the usage of trade credit for firms that relied more on such form of financing, and a higher increase in their usage of short-term bank credit. We proxy firms' dependence on trade credit using the pre-reform median level of accounts payable to total asset ratio, *Fpay*, with higher value indicating more dependence on trade credit. We regress *AP/TA* (*Accounts Payable$_t$/Total Asset$_{t-1}$*) on the interaction term *Fpay*After* and a set of firm controls. The coefficient on the interaction term examines how differently firms change their reliance on trade credit as a function of their pre-reform level of accounts payable. A negative coefficient would suggest trade credit dependent firms to decrease disproportionally more their reliance on trade credit after the reform. Column (1) of Table 2 finds the coefficient on *Fpay*After* to be negative and highly significant. In terms of economic significance, this result suggests a one-standard-deviation increase in the pre-reform reliance on trade credit reduces the payable ratio by 1.5%, which is economically significant given the average accounts payable ratio for our sample is around 10%. In monetary terms, it implies that an average firm would reduce the usage of trade credit by around RMB 79 million. Column (2) provides nonparametric results in which we sort firms into three equal sized bins based on their pre-reform payable ratio, and interact the size bin dummies with *After*. As expected, the results suggest that firms in the third size bin reduced more reliance on trade credit after the Property Law.

One might suspect the observed negative coefficient could be explained by the reversal of the usage on trade credit: that is, high dependence of trade credit in the past is followed naturally by low usage of trade credit in later periods. If this argument was driving our results, the negative correlation should be observed at any given point of time. To investigate this possibility, we repeat our analysis in columns (3) and (4) for several placebo reforms that happen in years before the actual reform. Since we use lagged control variables and we need at least one year of observation before and after the placebo reforms, we could design two placebo reforms occurring in 2003 and 2004, respectively. In addition to test the alternative explanation above, the placebo reforms further check whether the parallel trend assumption hold, which is crucial for our DID framework. In all these tests, *After* is an indicator variable takes value one for years after the placebo reform, and *Fpay* is the median payable ratio measured before each placebo reform. For none of these placebo reforms do we observe statistically significant correlation between pre-reform payable ratio and the usage of trade credit. These results rule out the alternative explanation explained above, and provide evidence supporting the parallel trend assumption.

Secondly, we investigate if trade-credit dependent firms could expand their access to bank credit after the reform. Given the short-term nature of trade credit, we expect to see broadened access to short-term bank credit, but not to long-term. Table 3 provides the results. Columns (1)-(3) employ $ShortDebt_t/Asset_{t-1}$ as dependent variable. In column (1), the coefficient on the *Fpay*After* is positive and statistically significant at 10%, confirming firms that relied more on trade credit before the reform could expand disproportionally more access to short-term bank credit after the reform. The estimate suggests that a one-standard-deviation increase in pre-reform payable ratio increases short-term leverage by

1.2%. For an average firm, it is equivalent to an increase of RMB 66 million in short-term debt. Combining with the results in column (1) of Table 2, these estimates suggest these firms could substitute almost entirely their usage of trade credit with short-term bank loans. These findings provide supporting evidence that the Property Law could allow trade credit dependent firms to shift away from plausibly expensive trade credit to bank credit, and validate the presence of a negative demand shock of trade credit.

Columns (2) and (3) show estimates for two placebo reforms took place in 2003 and 2004. In both cases, the coefficients on the interaction terms are negative and statistically insignificant, supporting the parallel trend assumption. More importantly, these estimates rule out an important alternative explanation, i.e., firms that had higher level of accounts payable in the past needed to borrow more short-term debt in the future. If this argument was driving our findings, we would expect to see positive interaction terms at any pre-reform period, however, this is not what we found in columns (2) and (3).

Next we investigate if firms could substitute their usage of trade credit with long-term bank credit and total bank credit, using $LongDebt_t/Asset_{t-1}$ (column (4)) and $Debt_t/Asset_{t-1}$ (column (5)) as dependent variable, respectively. In both cases, the coefficients on the interaction terms *Fpay*After* are statistically insignificant, suggesting firms could not substitute accounts payable, which is short-term in nature, to long-term bank credit.

These findings provide supporting evidence that after the Property Law, firms substitute their usage of trade credit with short-term bank credit. Arguably, access to external finance is not the only factor that could change firms' usage of trade credit. For instance, the usage of trade credit may change because of changes in market power (Wilner, 2000; Fabbri and Klapper, 2016), information opacity (Bias and Gollier, 1997), transaction costs (Ferris, 1981), or relationships between suppliers and customers (Summers and Wilson, 2002). Inclusion of firm fixed effects and the interaction term between *Treated*Trend* could only alleviate some of these concerns. To further validate the causal relationship between access to finance and the usage of trade credit, we explore cross-sectional variations in firms' financial constraints. Existing literature provides both theoretical arguments and empirical evidence that financially constrained firms use more trade credit (Burkart and Ellingsen, 2004; Bias and Gollier, 1997; Pertersen and Rajan, 1997). Following this line, if the availability of external financial recourses is the main reason behind the substitution effect, the reduction in the usage of trade credit and the increase in short-term debt should be stronger for the ex-ante financially constrained firms.

Table 4 shows results for subsamples divided by pre-reform firm size, a standard measure for financial constraints. Smaller firms are more financially constrained, and therefore if the Property Law promotes access to bank credit, the reduction in the usage of trade credit as well as the improvements in the access to short-term debt should be larger. We re-estimate for firms that belong to the highest 33% percentile in the pre-reform size distribution, and the lowest 33% percentile. Columns (1) and (2) of Table 4 compare the results on the usage of trade credit. The coefficient in column (1) suggests that firms belonging to the lowest 33% percentile in pre-reform size distribution experienced the largest decrease in the usage of trade credit, while that for the firms in the top 33% percentile in column (2) reduced much less. Specifically, a one-standard-deviation increase in the payable ratio decreases the dependence on trade credit by 1.8% for small firms, while for large firms, dependence only decreases by 0.5%. The test on the equality of the interaction terms across two subsamples however could not reject the null, although the p-value would be significant on conventional level for one-sided test. Likewise, column (3) shows that, the pre-reform trade-credit dependent small firms increase their access to short-term bank credit. Compared with the coefficient in column (1), small firms could substitute all their demand for trade credit with short-term bank credit. For large firms, in column (4), the increase in short-term debt is not significant. We test again the equality of the interaction terms across two subsamples. The result shows that the p-value would be significant on conventional level for one-sided test.

Although we cannot rule out other factors that could to have changed firms' usage of trade credit, these results exploring cross-sectional variations in financial constraints could further strengthen the causal relationship between access to bank credit and the dependence on trade credit. In addition, since our sample is composed of large public traded companies which are less financially constrained compared to small and medium unlisted companies, the substitution effects we find here may underestimate the average substitution effect for the entire population of Chinese companies. Nevertheless, even for our sample, these findings provide suggestive evidence that the Property Law may have caused a negative demand shock on trade credit.

4.2.Provision of trade credit

This section investigates the *provision of trade credit*, which is registered as accounts receivable. As argued before, the effect of the Property Law on accounts receivable is ambiguous: on the one hand, firms with more pre-reform movable assets may accumulate more accounts receivable, because such asset type is pledgable after the Property Law reform. Therefore, the extension of trade credit could increase more for firms with more pre-reform movable assets. On the other hand, if the Property Law allowed for better access to bank credit, the customers of trade credit could shift to potentially less costly bank credit after the reform, leaving the providers of trade credit to face less demand. The aggregate impact of the Property Law on the provision of trade credit should depend on both demand and supply factors. We proxy the suppliers using the pre-reform median level of movable assets ratio, *Fmov*. To recall, *Fmov* is defined as the median pre-reform movable assets ratio, where movable assets are composed of inventory and accounts receivable (which is the provision of trade credit). The assumption is that firms provided more trade credit or had more inventories in the past were more likely to be suppliers of trade credit, and consequently, should be more affected by the legal change.[6]

If the demand side effect dominates, we expect firms with higher level of pre-reform movable assets to reduce more their provision of trade credit. Table 5 provides results using *AR/TA* ($Accounts\ Receivable_t/Total\ Asset_{t-1}$) as dependent variable. Column (1) include only *Fmov*After* and $Fmov*Trend_t$. The negative coefficient of -0.063 implies that firms with high pre-reform movable assets ratio decreases more their provision of trade credit. The economical impact is sizeable: a one-standard-deviation increase in *Fmov* reduces the provision of trade credit (as share of total assets) by 1%. Given the average *AR/TA* ratio is 9.7%, a 1% decrease is economically significant. Quantitatively, it is equivalent to reduction of RMB 48 million in the supply of trade credit for an average firm. Column (2) adds firm control variables to reduce residual variation. The results remain similar.

[6] Unreported tests show similar results if the providers are proxied by pre-reform accounts receivable to asset ratio.

Columns (3) and (4) provide nonparametric results in which we sort firms into three equal sized bins s based on their pre-reform movable ratio, and interact the size bin dummies with *After*, for specifications with and without firm controls, respectively. In both specifications, the firms with more pre-reform movable assets experienced larger drops in the provision of trade credit: the reduction in the third size bin is twice as large as the one in the middle size bin.

These results seem to suggest the demand side effect dominates: the Property Law provided more access to bank credit which reduced the importance to of trade credit financing, and therefore the provision of trade credit dropped. There could be an alternative explanation: accounts receivable may be cyclical in nature, which implies that a high level of accounts receivable in the past would naturally be followed by less provision of trade credit. To address this possibility, we employ two placebo reforms defined in the same way as in Table 2. If the cyclical nature of accounts receivable is driving the results, we should expect to see a disproportional decrease in accounts receivable for high movable firms at any given point of time. The results in columns (5) and (6) show both interactions are negative, but statistically insignificant and economically small (2 to 3 times smaller than column (2)), which allows us to disregard the alternative explanation. Moreover, as argued before, the placebo tests also provide supporting evidence to the parallel trend assumption.

4.3. Redistribution of bank credit via trade credit

This section tests *Hypothesis 3*. Following the "redistribution view" of trade credit, firms borrow short-term loans from banks, and redistribute this short-term liquidity to their customers via trade credit. After the Property Law, because alternative bank credit became available, the redistribution of bank credit could be disrupted.

We employ an approach similar to Cull, Xu, and Zhu (2009) by regressing provision of trade credit (*AR/TA*) on lagged short-term leverage (SD_{it-1}) and its interaction term with the post reform dummy ($SD_{it-1}*After$). A positive coefficient on SD_{it-1} implies that firms which borrowed more short-term debt could redistribute more trade credit before the Property Law, and a negative coefficient on the interaction term $SD_{it-1}*After$ would imply a moderation or disruption of the redistribution effect after the Property Law.

Importantly, firms could also fund the provision of trade credit with long-term debt or internal resources such as cash holding or retained earnings. Because the exogenous shock of the Property Law primarily works through the availability of short-term bank credit, it should not have a direct impact on the provisions of trade credit via long-term bank credit or internal resources. Finding evidence that the provision of trade credit does not experience structure changes as function of long-term bank credit or internal resources would further validate our argument that the Property Law mainly provided a shock on the access to short-term bank credit.

Table 6 shows whether the provision of trade credit as function of short-term debt (column (1)), or other financial resources (columns (2)-(4)) experienced structural shifts around the Property Law. The dependent variable is *AR/TA*. In column (1), the coefficient on SD_{t-1} is positive and statistically significant, which suggests that in the pre-reform era, firms that borrowed more short-term bank credit provided more trade credit (accounts receivable), consistent with the view that firms redistribute their bank credit via trade credit. This finding is similar to Cull, Xu, and Zhu (2009) who investigate a large sample of Chinese industrial firms for 1998-2003. The redistribution effect however was disrupted after the Property Law reform, which is reflected in the negative coefficient on the interaction term $SD_{it-1}*After$. The sum of the two coefficients is 0.017 and statistically indifferent from zero (p-value=0.17). This finding suggests that after the Property Law reform, increases in the short-term bank credit do not increase the provision of trade credit, consistent with the view that the Property Law has disrupted the redistribution channel.

Column (2) investigates whether firms could rely on long-term loans to fund their provision of trade credit, and if so, whether the pattern experienced structural changes around the Property Law reform. The coefficients on LD_{it-1}, $LD_{it-1}*After$, and their sum are all statistically insignificant, indicating that in our sample firms do not finance accounts receivable with long-term debt either before or after the reform. Columns (3) and (4) investigate if firms fund accounts receivable with internal resources such as cash holding or retained earnings. For the pre-reform era, internal resources such as cash holding (column (3)) and retained earnings (column (4)) strongly determine the extension of trade credit, as suggested by the significant coefficient on $FinResource_{it-1}$. However, the interaction term $After*FinResource_{it-1}$ is insignificant regardless how $FinResource_{it-1}$ is proxied, which implies that their impact on the provision of trade credit did not experience structural break after the Property Law. This evidence further strengthens our argument that the Property Law mainly provided exogenous shock on short-term external liquidity, but not on internal resources. but as suggested by the insignificant interaction terms, their patterns remain unchanged after the Property Law.

The results above suggest that before the Property Law some short-term debt was used to fund the provision of trade credit. Such practice was disrupted by the Property Law. To further validate this hypothesis, we introduce a triple interaction term to test whether the disruption in the redistribution is larger for the suppliers of trade credit. Table 7 reports results. The dependent variable is *AR/TA*. Column (1) includes triple interaction terms SD_{it-1}**After*Fmov* to investigate if firms with more pre-reform movable assets experienced more disruption of redistribution after the Property Law reform. The negative coefficient on the triple interaction term indicates that firms with higher level of pre-reform moveable assets experienced more pronounced disruption of the redistribution of bank credit. For ease of interpretation, columns (2) and (3) show results for subsample of firms divided by the level of pre-reform movable assets. Specifically, column (2) investigates the relationship between lagged short-term bank credit and the provision of trade credit for firms in the lowest 33% percentile of pre-reform movable assets, and column (3) inspects for firms located in the highest 33% percentile of pre-reform movable assets. In column (2), we find the correlation between lagged short-term bank credit and provision of trade credit is insignificant either before or after the Property Law, which suggest these firms did not redistribute bank credit via trade credit. In contrast, for firms with high level of pre-reform movable assets, column (3), lagged short-term bank credit was positively related to provision of trade credit before the Property Law, consistent with the idea that these firms borrowed short-term bank credit to finance their provision of trade credit. This redistribution was disrupted after the reform, as indicated by the negative and statistically significant coefficient on the interaction term SD_{it-1}**After*. These findings are consistent with the hypothesis that the Property Law caused a negative demand shock on trade credit, which in turn disrupted the redistribution of bank credit via trade credit, particularly for firms with more movable assets.

Although lagged by one period, short-term bank credit could still be endogenous.

We investigate the robustness of our results using (only) short-term leverage measured in 2005 (SD_{05}), one year before the Property Law, to proxy firms' access to short-term debt. This is similar to an instrumental variables approach in which the identification is that pre-reform short-term leverage is not correlated with unobserved with-firm *changes* in trade credit following the Property Law reform (see Duchin, Ozbas and Sensoy (2010) for a similar application). Columns (4)-(6) re-estimate previous columns using SD_{05} to proxy firms' access to short-term leverage and find similar results.

To conclude, the findings in this section provide supporting evidence that the Property Law effectively disrupted the redistribution of bank credit via trade credit. We argue this is because the Property Law promoted access to bank financing, and consequently redistributing bank credit via trade credit became less important.

4.4. Asset structure and debt maturity

This section investigates how firms adjust their asset composition after the Property Law. Table 8 provide results, using *Log(TA)*, *Fixed Asset/Total Asset*, and *Cash/Total Asset* as dependent variables, respectively. We first look at changes in total assets. Column (1) uses the logarithm of total assets (Log(TA)) as dependent variable. The interaction term on *Fmov*After* is 0.315 and statistically significant at 5%, suggesting that firms with higher pre-reform movable assets ratio increase relatively more in total assets. However, the magnitude of the coefficient suggests a small economic effect: a one-standard-deviation increase in the movable assets ratio (*s.d.*=0.148) increases the logarithm of total assets by around 0.05 (0.315*0.148). Given the median value of *Log(TA)* is 21.3, this increase in size is economically small. Columns (2) and (3) test if the previous findings hold for placebo reforms happened in 2003 and 2004. In both cases, we do not find statistically significant results and therefore validating the parallel trend assumption.

Columns (4)-(6) investigate the changes in fixed asset investment (*Fixed Asset/Total Asset*). The coefficient on the interaction term of *Fmov*After* in column (4) is 0.095 and statistically significant. This implies that a one-standard-deviation increase in movable ratio increase tangibility by 1.4%, which is equivalent to RMB 68 million for an average firm. This result suggests that firms with high pre-reform movable assets shift towards more fixed asset investment. As before, we test for the placebo reforms that took place in 2003 and 2004 to check the parallel trend assumption. The insignificant interaction terms in columns (5) and (6) provide supporting evidence that the parallel trend assumption hold. Columns (7)-(9) investigate if firms change their cash holding after the Property Law, and for two placebo reforms defined as before. None of the columns shows differential changes in cash holding. Taken together, these results suggest that firms mostly adjust their asset composition towards longer maturity in terms of more fixed asset investments.

Next we examine the effect of the Property Law on corporate leverage and debt maturity. Table 9 presents the results. The positive and statistically significant coefficient on the interaction term *Fmov*After* in column (1) suggests firms with more pre-reform movable assets are able to obtain more bank credit after the reform. The increase is economically significant, as a one-standard-deviation increase in movable assets ratio increases total leverage by 1.6%, which is equivalent to a RMB 77 million increase in total debt for an average firm of our sample. Column (2) examines the impact of the law on long-term leverage. The coefficient on *Fmov*After* indicates that firms with more pre-reform movable assets increase disproportionally more long-term leverage after the reform. Quantitatively, a one-standard-deviation increase in the movable assets ratio implies a 1.4% increase in long-term leverage. This figure suggests that the increase in total leverage found in column (1) is almost driven entirely by the increase of long-term leverage. Column (3) further confirms that short-term leverage did not change depending on the pre-reform level of movable assets ratio, given the statistically insignificant and economically small coefficient on *Fmov*After*.

Columns (4) to (6) provide nonparametric results in which we sort firms into three equal sized bins based on their pre-reform movable ratio, and interact the size bin dummies with *After*. We expect to see larger effects for firms located in the higher size bin of pre-reform movable ratio. Columns (4) and (5) show that firms in the third size bin experienced 4.2% relative increase in leverage and 3.7% in long-term leverage, comparing to firms with lowest level of movable assets. In column (6), we do not find differential changes in short-term debt for all size bins of movable assets.

The findings on short-term debt is consistent with our conjecture that after taking into account trade credit, the effect of the Property Law on short-term bank credit is muted. On the one hand, firms with more movable assets may be able to borrow more short-term debt because they can now pledge their movable assets. On the other hand, once their customers can access bank credit more easily after the Property Law, their importance as non-financial intermediary decreases, in the sense that these firms stopped borrowing short-term debt to finance the provision of trade credit to their customers. The aggregate effect on short-term debt therefore becomes insignificant.

Combining the findings from Tables 8 and 9, we find that firms adjusted both their asset and debt composition following the Property Law. In both cases, the composition shifts towards longer end of maturities, that is more fixed asset investment and more long-term debt.[7] This is in line with the notion that firms match the maturity of assets and liabilities (e.g. Myers, 1977; Milbradt and Oehmke, 2014). Our findings also lend some support to "credit multiplier" effects (e.g. Bernanke, Gertler and Gilchrist, 2000; Campello and Hackbarth, 2012): higher external finance promotes more investments in fixed assets, which in turn could be used as collateral to further increase the debt capacity of these firms.

One might suspect that movable assets could be positively related to corporate leverage for other reasons. For instance, it is possible that firm extended more trade credit in the past or had more inventory needed to borrow more to keep the company afloat. If this argument is driving our results, it would imply a positive correlation between the movable assets ratio and leverage at any given point of time. We therefore repeat our analysis in Table 10 for two placebo reforms occurred in 2003 and 2004. For none of these placebo reforms do we observe statistically significant correlation between pre-reform movable ratio and corporate leverage (debt maturity). These results rule out the alternative explanation stated above, and provide evidence in support of the parallel trend assumption. In unreported tests we show that previous results still hold using debt levels instead of leverage as dependent variables, confirming our results are not driven by the variations in the denominator (i.e. total assets).

[7] See Gopalan, Mukherjee and Singh (2015) for similar findings. They find that better contract enforcement allows firms to better match debt and asset maturity.

5. **Robustness**

This section checks the robustness of previous results. For the sake of brevity, we focus on the disruption of the redistribution of bank credit. We first show the main findings could not be explained by the global financial crisis which hit China since 2008. Secondly, we check if other contemporary reform could explain the main finding, specifically the China Securities Regulatory Commission's (CSRC) regulations on related party transaction and illegal loan guarantee (2005-2006) (hereafter denoted as tunneling reform). We also investigate if changes in some innate firm characteristics other than movable assets are driving our results. We finally repeat other robustness tests relating to alternative definitions and alternative samples, which are available upon request.

5.1. Property Law or financial crisis

The disruption of redistribution could be explained by credit crunch due to the global financial crisis: the shortage in bank credit limited firms' ability to redistribute bank credit, simply because external financial resources were shut down and there was no bank credit to redistribute (Love, Preve and Sarria-Allende, 2007). In addition, firms with more outstanding short-term debt were particularly vulnerable due to the difficulty to roll-over short-term debt during financial crisis, and consequently, they were unable to pass on scarce bank credit to their clients via trade credit. Therefore, these firms shifted away from redistribution of bank credit because of liquidity drain and the needs to mitigate their own liquidity risks during crisis.

We argue that if the global financial crisis in 2008 was driving previous results, one would expect the disruption of the redistribution to occur predominately during the crisis period, but not *before* the crisis. To this end, we design two empirical strategies: Firstly we investigate a short event window: one year before and one year after the Property Law. Since the Property Law was formally approved at the end of 2006, firms could not have anticipated the financial crisis which hit China two years later. Therefore, this short event window could effectively mitigate the possible confounding factors caused by the financial crisis. In addition, the short event window could also mitigate confounding factors caused by other reforms (policies) that took place in the post-reform period. For instance, the RMB

4 trillion stimulus package introduced over 2009-2010 may have differential impacts on firms with different level of pre-reform movable assets. By focusing on the short event window, we remove this possibility. Our second strategy estimates precisely when the disruption occurred during the post-reform period. Specifically, we break down the post-reform periods into the pre-crisis year of 2007 (*Post1*), the crisis year of 2008 (*Crisis*), the years of stimulus package 2009-2010 (*Post2*), and the final year of the sample (*Post3*). By adding interaction terms between each post-reform dummy, access to short-term debt *SD*, and pre-reform movable assets *Fmov*, we could identify the moment when the disruption occurred. If financial crisis was driving our results, the coefficient on interaction term *Crisis*SD*Fmov* should be significantly negative, while that of $Post1*SD*Fmov_i$ should not be significant. As before we proxy access to short-term bank credit *SD* with lagged short-term leverage $(ShortDebt/TA)_{t-1}$ or short-term leverage measured in year 2005, $(ShortDebt/TA)_{2005}$.

Table 11 reports results. First, looking at the results from the short event window (columns (1) and (2)), we find the triple interaction terms in both columns are statistically and economically more significant than columns (1) and (4) of Table 7. This result suggests the disruption of the redistribution is particularly pronounced one year after the Property Law reform, which cannot be due to the financial crisis. In columns (3) and (4), we find the coefficients on the interaction terms *SD*Post1*Fmov* to be statistically significant and negative, consistent with previous columns that the first year after the Property Law observed the stronger disruption of redistribution. More importantly, the coefficients on *SD*Crisis*Fmov* are statistically insignificant, regardless how we proxy the access to short-term debt. This finding invalidates the alternative theory that credit crunch due to financial crisis could drive our findings.

5.2 Property Law or anti-related party transaction

Firms provide another type of inter-corporate loans which are called related party transaction. These types of transactions usually occur in between listed firms and their controlling holders and affiliates. Unlike trade credit, these transactions are not based on selling or purchasing of goods. Instead, as demonstrated in Jiang, Lee, and Yue (2010), this type of inter-corporate lending is one of the main channels for controlling share holders to "tunnel" corporate assets.[8] During 2005 and 2006, the State Council and China Securities Regulatory Commission issued several statements to tackle tunneling activities, including joint statements by eight ministries threatening to take personal actions against top managers of controlling entities if the tunneling problem was not resolved by the end of 2006 (hereafter tunneling reform).[9] According to many observers, these strict rules have successfully reduced the tunneling activities (e.g. Jiang, Lee, and Yue, 2010; Li et al., 2015). The tunneling reform could confound the previous results because firms with more movable assets are also more likely to engage in tunneling activities, as these firms have strong incentives to borrow short-term bank credit to lend to their controlling share holders. After the effective tunneling reform which prohibited such activities, the redistribution of short-term debt was disrupted. In other words, the tunnel reform may offer an alternative explanation to our findings.

We tackle this issue by controlling explicitly for firms' tunneling risks. Related party transactions are registered under the entry *"Other Receivables"* (*OREC*), and according to Jiang, Lee and Yue (2010), firms with more other receivables are prone to tunneling risks. We therefore use other receivables ratio ($Other\ Receivable_t/Asset_t$) measured as the pre-reform median to proxy the tunneling risk, denoted as *Forec*.[10] Because the tunneling reform started from 2005, we define a dummy variable *Tunnel* equals zero for the pre-2005 period, and one otherwise. We introduce the interaction term *SD*Forec*Tunnel* together with interaction term *SD*After*Fmov*. If the tunneling reform was causing the disruption of the redistribution, we expect the coefficient on *SD*Forec*Tunnel* to be negative, and that of $SD*After*Fmov_i$ to lose its significance. Again, we proxy access to short-term debt (*SD*) with either lagged short-term leverage $(ShortDebt/TA)_{t-1}$ or short-term leverage measured in

[8] Tunneling is an activity by which controlling shareholders or managers extract firm values. (see e.g. Johnson, La Porta, Lopez-de-Silanes, and Shleifer, 2000; La Porta and Lopez-de-Silanes, and Zamarripa, 2003).

[9] Several failed attempts in regulating tunneling activities took effect in the early 2000s. Eventually, in November 2005, the State Council issued a Directive on behalf of CSRC, titled " On improving the Quality of Listed Companies", which states the top management of controlling shareholders or colluding firms will be personally punished for tunneling activities. In November 2006, eight government ministries issued a joint announcement, making it clear that the top management of controlling entities would be fired and face disciplinary punishment if tunneling activities remained by December 31, 2006. See Jiang, Lee, and Yue (2010) and Li et al. (2015) for more details.

[10] An alternative proxy of tunneling risk is the wedge between control rights and cash flow rights (*wedge*). See for instance Lin et al., (2012). Additional tests are conducted using such proxy and our results remain similar. Data of control rights and cash flow rights are collected from China Stock Market & Accounting Research (CSMAR). Results are available upon request.

year 2005, *(ShortDebt/TA)*$_{2005}$.The results in Table 12 column (1) and (2) show that after controlling the tunneling reform, the coefficient on the interaction term *SD*After*Fmov* is very similar to our previous results. This implies that our previous findings are not driven by the tunneling reform.

5.3. Other robustness tests

In previous sections, we use lagged firm control variables. In this section, we control firm characteristics with pre-reform median values. Specifically, we calculate the pre-reform median value of each control variable, except *Age*, *List* and *State*, and denote them as *FX*. For *Age*, *List* and *State*, we use (only) the 2005 value, which is one year before the reform. We replace all lagged control variables with interaction terms between *FX*After*, where the coefficients on *FX*After* indicate if firms react to the Property Law differently depending on these firms' innate features. Table 12 columns (3) and (4) report results using either lagged short-term leverage or short-term leverage measured at 2005 as proxy of access to short-term debt. The coefficients on the triple interaction terms are very similar as before, suggesting our findings could not be driven by changes in firm characteristics other than movable assets.

Finally, some additional robustness tests are available upon request, including 1) Re-estimation after removing observations from 2005. Even though we have removed observations from 2006 to control for the possibility that some firms may have anticipated the passage of the law in 2006, we go one step further by removing data from 2005. With this sample, it is safe to argue that anticipating the reform to take place back in 2004 is impossible for any firm. 2) Re-define the pre-reform movable ratio. Instead of using pre-reform median movable ratio, we explore if the results still hold using only the movable ratio measured in 2005. 3) Redefine trade credit using the definition of Petersen and Rajan (1997). Specifically, instead of scaling with total assets, we scale accounts receivable with net sale, and accounts payable with cost of goods sold. 4) Re-estimation over alternative samples: a) sample of firms that have been listed throughout the entire sample period, to alleviate the concern that differential reputational improvements due to listing or accessibility to capital markets might drive the results;[11] b) a sample of firms that never changed their ownership type throughout the sample period, which removes possible confounding effects due to privatization. Our main results hold for all these tests.

6. Conclusions

By allowing movable assets to be pledged as collateral, the Chinese Property Law may have extended access to bank credit for vast population of small and medium firms. This exogenous shock in short-term bank liquidity reduced the demand for trade credit, a possibly more expensive substitute to bank credit. As a result, the providers of trade credit reduced their provision of trade credit. More importantly, the Property Law disrupted a practice in which firms borrow bank credit to finance their provision of trade credit. Therefore, even though the Property Law could also allow the providers of trade credit to access short-term bank credit more easily, the aggregate effect on short-term borrowing was muted. Instead, we find the providers of trade credit started to accumulate more fixed asset investment, which in turn led to more borrowing in long-term debt. Our findings highlight the importance of taking into account other financing channels when investigating the effect of collateral law. Our study also has important policy implications: by removing collateral constraint, legal reform could re-direct funding flows back to the banking sector, which to some extent could facilitate a better implementation of monetary policy.

[11] All firms in the sample are eventually listed at stock exchanges, but their annual reports started to be published several years before the actual listings.

References:

Allen., F., J., Qian, and M.J. Qian, 2005. Law, finance and economic growth in China. *Journal of Financial Economics* 77 (1): 57-116.

Alvarez de la Campa, A., 2011. Increasing access to credit through reforming secured transactions in the MENA region. Working Paper, *World Bank Policy Research*.

Aretz, K., M. Campello, and M.T.Marchica., 2015. Conflicting security laws and the democratization of credit: France's reform of the Napoleonic code. Working Paper

Berkowitz, D., C. Lin, and Y., Ma., 2015. Do property rights matter? Evidence from a property law enactment. *Journal of Financial Economics* 116: 583-593.

Bernanke, B., M. Gertler, and S.Gilchrist., 2000. The financial accelerator in a quantitative business cycle framework. In *Handbook of macroeconomics*, Vol. 1C, 1341-1393. Eds J. Taylor and M. Woodford, M. North-Holland, Amsterdam: Elsevier.

Bertrand, M., E. Duflo, and S.Mullainathan., 2004. How much should we trust differences-in-differences estimates? *Quarterly Journal of Economics* 119: 249-275.

Biais, B., and C., Gollier., 1997. Trade credit and credit rationing. *Review of Financial Studies* 10, 903-937.

Burkart, M., and T., Ellingsen., 2004. In-kind finance: a theory of trade credit. *American Economic Review* 94, 569-590.

Calomiris, C., C., Himmelberg, and P., Wachtel., 1995. Commercial paper, corporate finance, and the business cycle: a microeconomic perspective. *Carnegie-Rochester Conference Series on Public Policy* (42), 203-250.

Calomiris, C., M.Larrain, J.Sturgess., 2017. How collateral laws shape lending and sectoral activity. *Journal of Financial Economics* 123 (1), 163-188.

Campello, M., and D. Hackbarth, 2012. The firm-level credit multiplier. *Journal of Financial Intermediation* 21: 446-472.

Campello, M., and M. Larrain, 2015. Enlarging the contracting space: collateral menus, access to credit and economic activity. *Review of Financial Studies*. Advance Access published online November 14, 2015. doi:10.1093/rfs/hhv069

Cerqueiro, G., S. Ongena, and K. Roszbch, 2015. Collateral damage? On collateral, corporate financing and performance. Working Paper

Cull, R., L.C. Xu., 2005. Institution, ownership and finance: the determinants of profit reinvestment among Chinese firms. *Journal of Financial Economics* 77 (1), 117-146.

Cull, R., L.C. Xu, and T. Zhu, 2009. Formal finance and trade credit during China's transition. *Journal of Financial Intermediation* 18 (2): 173-192.

Cuñat, V., 2007. Trade credit: suppliers as debt collectors and insurance providers. Review of Financial Studies 20 (2), 491-527.

Demirguc-Kunt, A., V. Maksimovic., 2001. Firms as financial intermediaries: evidence from trade credit data. World Bank Working Paper.

Duchin, R., O. Ozbas, and B., Sensoy., 2010. Costly external finance, corporate investment, and the subprime mortgage credit crisis. Journal of Financial Economics 97, 413-432.

Fabbri, D., and L., Klapper, 2016. Bargaining power and trade credit. *Journal of Corporate Finance* 41, 66-80.

Ferris, J.S., 1981. A transactions theory of trade credit use. *The Quarterly Journal of Economics* 96 (2), 243-270.

Fisman, R., and I. Love., 2003. Trade Credit, Financial Intermediary Development, and Industry Growth. *The Journal of Finance* 58 (1), 353-374.

Garcia-Appendini, E. and J. Montoriol-Garriga, 2013. Firms as liquidity providers: evidence from the 2007-2008 financial crisis. *Journal of Financial Economics* 109: 272-291.

Ge, Y., and J. Qiu., 2007. Financial development, bank discrimination and trade credit. *Journal of Banking and Finance* 31, 513-530.

Gopalan, R., A., Mukherjee, and M., Singh. 2015. Do debt contract enforcement costs affect financing and asset structure. Accepted manuscript by *Review of Financial Studies*.

Haselmann, R., K. Pistor, and V. Vig., 2010. How law affects lending. *Review of Financial Studies* 23: 549-580.

Jiang, G.H., C.M.C., Lee, and H. Yue, 2010. Tunneling through intercorporate loans: The China experience. *Journal of Financial Economics* 98 (1): 1-20.

Johnson, S., R. La Porta, and F., Lopez-de-Silanes, 2000. Tunneling. American Economic Review 90 (2): 22-27.

Klapper, L., L., Laeven and R., Rajan., 2012. Trade credit contracts. *Review of Financial Studies* 25, 838-867.

La Porta, R., F. Lopez-de-Silanes, and G. Zamarripa, 2003. Related lending. The Quarterly Journal of Economics 118 (1): 231-268.

Li, K., L. Lu, J. Qian., and J. L. Zhu., 2015. Enforceability and the effectiveness of laws and regulations. Working Paper. Shanghai Advanced Institute of Finance.

Lin, C., Y., Ma, P., Malatesta, and Y.H., Xuan. 2012. Corporate ownership structure and bank loan syndicate structure. *Journal of Financial Economics* 104 (1): 1-22.

Lilienfeld-Toal, U., D. Mookherjee, and S. Visaria. 2012. The distributive impact of reforms in credit enforcement: Evidence from Indian debt recovery tribunals. *Econometrica* 80: 497-558.

Love, I., S. Martinez Peria, and S. Sandeep., 2016. Collateral registries for movable assets: Does their introduction spur firms' access to bank finance? *Journal of Financial Services Research* 49: 1-37.

Love, I., L.A. Preve, and V. Sarria-Allende., 2007. Trade credit and bank credit: evidence from recent financial crises. *Journal of Financial Economics* 83, 453-469.

Milbradt, K. and M. Oehmke, 2014. Maturity rationing and collective short-termism. *Journal of Financial Economics* 118 (3): 553-570.

Myers, S., 1977. Determinants of corporate borrowing. *Journal of Financial Economics* 5: 147-175.

Nilsen, J.H., 2002. Trade credit and the bank lending channel. *Journal of Money, Credit, and Banking* 34 (1), 226-253.

Ng, C., J.K., Smith and R.L., Smith. Evidence on the determinants of credit terms used in interfirm trade credit. *The Journal of Finance* 54, 1109-1124.

Petersen, M.A. and R.G., Rajan, 1995. The effect of credit market competition on lending relationships. *Quarterly Journal of Economics* 110, 407-443.

Petersen, M.A. and R.G., Rajan. 1997. Trade credit: theories and evidence. *Review of Financial Studies* 10 (3): 661-691.

Shenoy, J. and R., Willams, 2017. Trade credit and the joint effects of supplier and customer financial characteristics. *Journal of Financial Intermediation* 29, 68-80.

Wilner, B., 2000. The exploitation of relationships in financial distress: the case of trade credit. *The Journal of Finance* 55, 153-178.

Wilson, N. and B. Summers, 2002. Trade credit terms offered by small firms: survey evidence and empirical analysis. *Journal of Business Finance & Accounting* 29, 317-351.

Vig, V. 2013. Access to collateral and corporate debt structure: Evidence from a natural experiment. *Journal of Finance* 68: 881-928.

Zhang, M., 2008. From public to private: the newly enacted Chinese property law and the protection of property rights in China. *Berkeley Business Law Journal* 5 (2), 317.

Table 1: Summary statistics

This table reports summary statistics for selected variables. Sample period is 2001 to 2011, excluding observations from the reform year 2006. The sample contains firms listed in either Shanghai Stock Exchange or Shenzhen Stock Exchange. Data is obtained from Wind Information. $Movratio_t$ is defined as $(Inventory_t+Account\ Receivable_t)/Asset_t$. *AR/TA* is the accounts receivable scaled by total assets. *AP/TA* is accounts payable scaled by total assets. *OREC/TA* is other receivable scaled by total assets. $Tangibility_t$ is the ratio of fixed assets to total assets ($FixedAsset_t/Asset_t$). $Liquidity_t$ is cash divided by total assets ($Cash_t/Asset_t$). $Profitability_t$ is the ratio of net profits over total assets ($Netprofit_t/Asset_t$). $Sale_t$ is the logarithm of one plus total sale. Age_t is defined as the logarithm of one plus the number of years since incorporation. $List_t$ is a dummy variable equals one for firm-year observation after firm's IPO, and zero otherwise. $Split_t$ is a dummy variable equals one for firm-year observation after firm's completion of the split share reform, and zero otherwise. $State_t$ is a dummy variable equals one if the controlling shareholder is government and zero if the controlling shareholder is private entity.

VARIABLES	Observations	Mean	Ste.Dev.	Min	Max
Panel A:Movable assets and trade credit					
$Movratio_t$	12445	0.26	0.17	0.01	0.76
AR/TA	12518	0.09	0.08	0.00	0.37
AP/TA	12630	0.09	0.07	0.01	0.36
$Inventory_t/Asset_t$	12570	0.17	0.15	0.00	0.75
OREC/TA	12692	0.04	0.07	0.00	0.38
Panel B:Debt					
$Debt_t/Asset_{t-1}$	6884	0.33	0.22	0.02	1.44
$LongDebt_t/Asset_{t-1}$	7239	0.13	0.15	0.00	0.85
$ShortDebt_t/Asset_{t-1}$	10218	0.21	0.16	0.00	0.88
Panel C:Assets					
$Log(1+Asset_t)$	12720	21.25	1.17	18.42	24.70
$FixedAsset_t/Asset_{t-1}$	11415	0.32	0.22	0.00	1.07
Panel D: Controls					
$Tangibility_t$	12680	0.29	0.19	0.00	0.79
$Liquidity_t$	12707	0.16	0.12	0.00	0.56
$Profitability_t$	12717	0.03	0.08	-0.41	0.22
$Sale_t$	12701	20.56	1.50	15.94	24.50
Age_t	12720	2.48	0.48	0.00	4.12
$List_t$	12720	0.96	0.20	0	1
$Split_t$	12720	0.50	0.50	0	1
$State_t$	12720	0.69	0.46	0	1

Table 2: Usage of trade credit

This table report results on the usage of trade credit. The dependent variable is *Accounts Payable*$_t$*/Total Assets*$_{t-1}$ (*AP/TA*). In column (1), *Fpay*$_i$ is the median payable ratio measured over 2001-2005. *After*$_t$ is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. In column (2), D^{Fpay_mid} and D^{Fpay_high} are indicator variables that take value one if the pre-reform median payable ratio belongs to the middle and top tertile, and zero otherwise. Columns (3) and (4) report results for placebo reforms take place in 2003 and 2004, respectively. In the placebo regression, *After* is an indicator variable that takes value one for years after the placebo reform, and *Fpay*$_i$ is the median movable ratio measured over years before the placebo reform. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep. Var	AP/TA			
			Placebo 2003	Placebo 2004
	(1)	(2)	(3)	(4)
$Fpay_i$*After	-0.155***		-0.037	-0.076
	(0.060)		(0.059)	(0.061)
$D^{Fpay\_mid}{}_i$*After		-0.008		
		(0.006)		
$D^{Fpay\_high}{}_i$*After		-0.012*		
		(0.007)		
Log(TA_{t-1})	-0.050***	-0.052***	-0.066***	-0.066***
	(0.003)	(0.003)	(0.006)	(0.006)
$Tangibility_{t-1}$	-0.027***	-0.031***	-0.039***	-0.039***
	(0.010)	(0.010)	(0.013)	(0.013)
$Liquidity_{t-1}$	-0.027**	-0.032***	-0.026*	-0.026*
	(0.010)	(0.011)	(0.013)	(0.013)
$Profitability_{t-1}$	0.024	0.021	0.034*	0.035*
	(0.015)	(0.016)	(0.018)	(0.018)
$Sale_{t-1}$	0.017***	0.018***	0.008**	0.009**
	(0.003)	(0.003)	(0.004)	(0.004)
Age_{t-1}	0.010	0.014*	0.004	0.004
	(0.008)	(0.008)	(0.010)	(0.010)
$List_t$	-0.001	-0.003	0.014***	0.014***
	(0.006)	(0.006)	(0.005)	(0.005)
$Split_t$	0.013***	0.013***	0.008*	0.008*
	(0.005)	(0.005)	(0.004)	(0.004)
$State_t$	0.004	0.005	-0.001	-0.001
	(0.004)	(0.005)	(0.006)	(0.006)
Fpay*Trend	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes
Observations	11,343	11,343	5,052	5,052
R-squared	0.229	0.212	0.215	0.216
Number of firms	1,271	1,271	1,269	1269

Table 3: Substitution between bank credit and trade credit

This table estimates the substitution between trade credit and bank credit. The dependent variable is *ShortDebt/TA*, *LongDebt/TA*, and *Debt/TA*, respectively. In column (1), $Fpay_i$ is the median payable ratio measured over 2001-2005. $After_t$ is a dummy variable that equals one for the years 2007-2011, and zero otherwise. Columns (2) and (3) report results for two placebo reforms took place in 2003 and 2004, respectively. In both of these columns, *After* is an indicator variable that takes value one for years after the placebo reform, and $Fpay_i$ is the median movable ratio measured over years before the placebo reform. The sample ends by year-end of 2005 for these columns. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	ShortDebt/TA			LongDebt/TA	Debt/TA
		Placebo 2003	Placebo 2004		
	(1)	(2)	(3)	(4)	(5)
$Fpay_i$*After	0.203*	-0.025	-0.073	0.010	-0.041
	(0.110)	(0.099)	(0.090)	(0.132)	(0.255)
$Log(TA_{t-1})$	-0.042***	-0.088***	-0.087***	-0.032***	-0.102***
	(0.006)	(0.012)	(0.012)	(0.007)	(0.013)
$Tangibility_{t-1}$	-0.007	-0.024	-0.023	-0.068***	-0.098***
	(0.019)	(0.029)	(0.029)	(0.022)	(0.036)
$Liquidity_{t-1}$	-0.034	-0.049	-0.048	-0.059**	-0.053
	(0.024)	(0.030)	(0.030)	(0.027)	(0.046)
$Profitability_{t-1}$	-0.121***	-0.067*	-0.066*	0.076**	-0.069
	(0.032)	(0.037)	(0.037)	(0.035)	(0.061)
$Sale_{t-1}$	0.012**	0.007	0.007	-0.020***	-0.006
	(0.005)	(0.006)	(0.006)	(0.006)	(0.010)
Age_{t-1}	0.058***	0.063***	0.061**	0.035*	0.132***
	(0.018)	(0.024)	(0.024)	(0.020)	(0.032)
$List_t$	0.008	0.010	0.015	-0.071***	-0.041*
	(0.011)	(0.012)	(0.013)	(0.016)	(0.024)
$Split_t$	0.008	0.010	0.011	0.027***	0.036**
	(0.009)	(0.009)	(0.009)	(0.009)	(0.016)
$State_t$	-0.003	-0.001	-0.001	-0.011	-0.011
	(0.010)	(0.013)	(0.013)	(0.011)	(0.017)
Fpay*Trend	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes	Yes
Observations	10,202	4,678	4,684	7,226	6,875
R-squared	0.206	0.162	0.160	0.192	0.217
Number of firms	1,261	1,242	1,244	1,159	1,144

Table 4: Financial constraints and substitution between bank credit and trade credit

This table estimates the substitution between trade credit and bank credit for samples of small and large firms. Small firms are the ones with lowest 33% percentile of pre-reform total assets, while large firms are the ones with highest 33% percentile of pre-reform total assets. The dependent variable is AP/TA in columns (1) and (2), and *ShortDebt/TA* in columns (3) and (4). $Fpay_i$ is the median payable ratio measured over 2001-2005. $After_t$ is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	AP/TA		ShortDebt/TA	
	Small	Large	Small	Large
	(1)	(2)	(3)	(4)
$Fpay_i$*After	-0.298**	-0.094	0.305*	0.010
	(0.123)	(0.097)	(0.175)	(0.120)
F (p-value)	1.72 (0.19)		2.26 (0.13)	
Firm controls	Yes	Yes	Yes	Yes
Fpay*Trend	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes
Observations	3,759	3,786	3,256	3,505
R-squared	0.362	0.364	0.334	0.338
Number of firms	423	424	420	420

Table 5: Provision of trade credit

This table report results on the provision of trade credit. The dependent variable is *Accounts Receivable$_t$/Total Assets$_{t-1}$* (*AR/TA*). In columns (1)-(2), *Fmov$_i$* is the median movable ratio measured over 2001-2005. *After$_t$* is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. In columns (3)-(4), D^{Fmov_mid} and D^{Fmov_high} are indicator variables that take value one if the pre-reform median movable ratio belongs to the middle and top tertile, and zero otherwise. *After$_t$* is defined the same way as in columns (1)-(2). Columns (5) and (6) report results for placebo reforms took place in 2003 and 2004, respectively. The sample ends by year-end of 2005 for these two columns. In these placebo regressions, *After* is an indicator variable that takes value one for years after the placebo reform, and *Fmov$_i$* is the median movable ratio measured over years before the each placebo reform. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	AR/TA					
					Placebo 2003	Placebo 2004
	(1)	(2)	(3)	(4)	(5)	(6)
$Fmov_i$*After	-0.063***	-0.057***			-0.022	-0.030
	(0.020)	(0.019)			(0.015)	(0.020)
$D^{Fmov\_mid}_i$*After			-0.011*	-0.010*		
			(0.006)	(0.005)		
$D^{Fmov\_high}_i$*After			-0.023***	-0.022***		
			(0.007)	(0.007)		
Log(TA_{t-1})		-0.039***		-0.040***	-0.062***	-0.062***
		(0.003)		(0.003)	(0.006)	(0.006)
$Tangibility_{t-1}$		-0.033***		-0.037***	-0.034**	-0.038***
		(0.010)		(0.011)	(0.014)	(0.014)
$Liquidity_{t-1}$		-0.039***		-0.043***	-0.053***	-0.060***
		(0.010)		(0.010)	(0.014)	(0.014)
$Profitability_{t-1}$		0.060***		0.061***	0.082***	0.081***
		(0.013)		(0.013)	(0.018)	(0.018)
$Sale_{t-1}$		0.021***		0.022***	0.007**	0.008**
		(0.003)		(0.003)	(0.003)	(0.003)
Age_{t-1}		0.006		0.006	0.002	0.001
		(0.010)		(0.010)	(0.012)	(0.013)
$List_t$		-0.003		-0.002	0.023***	0.022***
		(0.006)		(0.006)	(0.006)	(0.006)
$Split_t$		0.005		0.005	0.004	0.003
		(0.004)		(0.004)	(0.004)	(0.004)
$State_t$		-0.004		-0.004	-0.006	-0.005
		(0.005)		(0.005)	(0.006)	(0.007)
Firm controls	No	Yes	No	Yes	Yes	Yes
$Fmov_i$*Trend	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes	Yes	Yes
Observations	11,229	11,207	11,229	11,207	5,013	5,015
R-squared	0.159	0.217	0.156	0.217	0.270	0.266
Number of firms	1,266	1,266	1,266	1,266	1,263	1,264

Table 6: Redistribution of bank credit via trade credit

This table report results on the redistribution of bank credit via trade credit. The dependent variable is *Accounts Receivable$_t$/Total Assets$_{t-1}$* (*AR/TA*). *After$_t$* is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. The lagged financial resource variables that are interacted with *After* dummy include: *ShortDebt/TA* (column (1)), *LongDebt/TA* (column (2)), *Cash/TA* (column (3)) and *Retained Earnings/TA* (column (4)), all lagged by one year. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep. Var	AR/TA			
FinResource$_{it-1}$=	SD/TA	LD/TA	Cash/TA	RetainEearning/TA
	(1)	(2)	(3)	(4)
FinResource$_{it-1}$	0.047***	-0.027	-0.062***	0.038***
	(0.015)	(0.020)	(0.014)	(0.009)
After*FinResource$_{it-1}$	-0.030*	0.028	0.025	-0.009
	(0.017)	(0.027)	(0.016)	(0.009)
Log(TA$_{t-1}$)	-0.042***	-0.034***	-0.042***	-0.047***
	(0.004)	(0.004)	(0.003)	(0.004)
Tangibility$_{t-1}$	-0.047***	-0.041***	-0.048***	-0.044***
	(0.012)	(0.010)	(0.011)	(0.011)
Liquidity$_{t-1}$	-0.043***	-0.053***		-0.048***
	(0.012)	(0.013)		(0.011)
Profitability$_{t-1}$	0.071***	0.044***	0.060***	
	(0.014)	(0.015)	(0.013)	
Sale$_{t-1}$	0.022***	0.019***	0.023***	0.021***
	(0.003)	(0.002)	(0.003)	(0.002)
Age$_{t-1}$	0.007	0.020*	0.007	0.007
	(0.011)	(0.010)	(0.010)	(0.010)
List$_t$	-0.001	0.002	-0.004	-0.003
	(0.007)	(0.007)	(0.006)	(0.007)
Split$_t$	0.002	-0.000	0.004	0.001
	(0.004)	(0.005)	(0.003)	(0.004)
State$_t$	-0.004	0.001	-0.002	-0.003
	(0.005)	(0.005)	(0.005)	(0.005)
FinResource+After*FinResource	0.017	0.001	-0.037***	0.029***
	(0.012)	(0.016)	(0.013)	(0.007)
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes
Observations	10,204	7,152	11,235	11,222
R-squared	0.207	0.231	0.197	0.206
Number of firms	1,262	1,164	1,271	1,271

Table 7: Redistribution of bank credit and movable assets

This table report results on redistribution of bank credit and movable assets. The dependent variable is *Accounts Receivable*$_t$*/Total Assets*$_{t-1}$ (*AR/TA*). Columns (1)-(3) use lagged short-term bank credit scaled by total asset, and columns (4)-(6) use short-term bank credit to total asset ratio measured at 2005. In column (1) and (4), *Fmov*$_i$ is the median movable ratio measured over 2001-2005. *After*$_t$ is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. Columns (2) and (5) estimate for a sample of firms that locate in the lowest 33% tertile of pre-reform movable assets, while columns (3) and (6) estimate for firms with the highest 33% tertile of pre-reform movable assets. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	AR/TA					
	SD=(ShortDebt/TA)$_{it-1}$			SD=(ShortDebt/TA)$_{i2005}$		
	All firms	Low Movable	High Movable	All firms	Low Movable	High Movable
VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)
SD	0.044*	-0.012	0.059**			
	(0.027)	(0.022)	(0.028)			
SD*After	0.024	0.035	-0.074**	0.021	0.027	-0.096***
	(0.033)	(0.026)	(0.030)	(0.031)	(0.023)	(0.027)
SD*After*Fmov$_i$	-0.196*			-0.196*		
	(0.114)			(0.117)		
SD*Fmov$_i$	-0.012					
	(0.094)					
Fmov$_i$*After	-0.015			-0.020		
	(0.031)			(0.031)		
H_0: SD+After*SD=0		0.023	-0.015			
		(0.018)	(0.022)			
Firm controls	Yes	Yes	Yes	Yes	Yes	Yes
Treated*Trend	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes	Yes	Yes
Observations	10,179	3,229	3,440	10,315	3,276	3,514
R-squared	0.232	0.383	0.401	0.231	0.354	0.399
Number of firms	1,257	416	420	1,163	371	395

Table 8: Asset structure

This table estimates the effect on asset structure. The dependent variables Log(TA), $FixedAsset_t/Asset_{t-1}$ (FA/TA), and $Cash_t/Asset_{t-1}$ (CA/TA). For each dependent variable, we report results for the actual reform and two placebo reforms take place in 2003 and 2004. For actual reform, *After* is a dummy variable that equals one for the years 2007-2011, and zero otherwise; $Fmov_i$ is the median movable ratio measured over 2001-2005. In these placebo regressions, *After* is an indicator variable that takes value one for years after the placebo reform, and $Fmov_i$ is the median movable ratio measured over years before the each placebo reform. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	LogTA	LogTA	LogTA	FA/TA	FA/TA	FA/TA	CA/TA	CA/TA	CA/TA
	Actual Reform	Placebo 2003	Placebo 2004	Actual Reform	Placebo 2003	Placebo 2004	Actual Reform	Placebo 2003	Placebo 2004
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
$Fmov_i$*After	0.315**	-0.105	0.009	0.095**	-0.021	0.032	-0.015	0.103	0.028
	(0.146)	(0.064)	(0.067)	(0.047)	(0.030)	(0.037)	(0.112)	(0.091)	(0.108)
Firm controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Treated*Trend	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	11,376	5,046	5,048	11,358	5,045	5,047	11,375	5,046	5,048
R-squared	0.581	0.519	0.518	0.203	0.149	0.146	0.179	0.267	0.264
Number of firm	1,266	1,263	1,264	1,266	1,263	1,264	1,266	1,263	1,264

Table 9: Debt structure

This table estimates the effect on debt maturity. The dependent variables are $Debt_t/Asset_{t-1}$ (*DT/TA*), $LongDebt_t/Asset_{t-1}$ (*LD/TA*), and $ShortDebt_t/Asset_{t-1}$ (*SD/TA*), respectively. *After* is a dummy variable that equals one for the years 2007-2011, and zero otherwise. $Fmov_i$ is the median movable ratio measured over 2001-2005. D^{Fmov_mid} and D^{Fmov_high} are indicator variables that take value one if the pre-reform median movable ratio belongs to the middle and top tertile, and zero otherwise. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	DT/TA (1)	LD/TA (2)	SD/TA (3)	DT/TA (4)	LD/TA (5)	SD/TA (6)
$Fmov_i$*After	0.110*	0.088**	-0.006			
	(0.060)	(0.039)	(0.048)			
$D^{Fmov\_mid}_i$*After				-0.002	0.001	-0.013
				(0.019)	(0.011)	(0.019)
$D^{Fmov\_high}_i$ *After				0.042*	0.037***	-0.009
				(0.022)	(0.014)	(0.016)
$Log(TA_{t-1})$	-0.102***	-0.030***	-0.045***	-0.103***	-0.031***	-0.046***
	(0.013)	(0.008)	(0.006)	(0.013)	(0.008)	(0.006)
$Tangibility_{t-1}$	-0.098***	-0.076***	-0.002	-0.097***	-0.075***	-0.002
	(0.036)	(0.024)	(0.021)	(0.036)	(0.023)	(0.020)
$Liquidity_{t-1}$	-0.051	-0.067**	-0.032	-0.049	-0.066**	-0.033
	(0.046)	(0.029)	(0.025)	(0.046)	(0.029)	(0.025)
$Profitability_{t-1}$	-0.063	0.086**	-0.133***	-0.069	0.082**	-0.134***
	(0.061)	(0.038)	(0.034)	(0.061)	(0.038)	(0.034)
$Sale_{t-1}$	-0.006	-0.022***	0.012**	-0.006	-0.022***	0.012**
	(0.010)	(0.006)	(0.006)	(0.010)	(0.006)	(0.006)
Age_{t-1}	0.131***	0.039*	0.056***	0.131***	0.040*	0.055***
	(0.032)	(0.021)	(0.019)	(0.032)	(0.021)	(0.020)
$List_t$	-0.043*	-0.074***	0.010	-0.044*	-0.075***	0.009
	(0.024)	(0.017)	(0.011)	(0.024)	(0.016)	(0.011)
$Split_t$	0.034**	0.026***	0.006	0.034**	0.027***	0.006
	(0.016)	(0.010)	(0.009)	(0.016)	(0.010)	(0.009)
$State_t$	-0.010	-0.010	-0.004	-0.010	-0.010	-0.004
	(0.017)	(0.012)	(0.010)	(0.017)	(0.012)	(0.010)
Treated*Trend	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes	Yes	Yes
Observations	6,853	7,202	10,162	6,853	7,202	10,162
R-squared	0.216	0.197	0.157	0.217	0.199	0.158
Number of firms	1,140	1,155	1,256	1,140	1,155	1,256

Table 10: Placebo reforms

This table report results on debt maturity for placebo reforms that took place in 2003 (columns (1)-(3)) and 2004 columns (4)-(6)). The sample ends by year-end of 2005. The dependent variables are $Debt_t/Asset_{t-1}$ (*DT/TA*), $LongDebt_t/Asset_{t-1}$ (*LD/TA*), and $ShortDebt_t/Asset_{t-1}$ (*SD/TA*), respectively. In these placebo regressions, *After* is an indicator variable that takes value one for years after the placebo reform, and $Fmov_i$ is the median movable ratio measured over years before the each placebo reform. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var	DT/TA	LD/TA	SD/TA	DT/TA	LD/TA	SD/TA
	Placebo 2003	Placebo 2003	Placebo 2003	Placebo 2004	Placebo 2004	Placebo 2004
	(1)	(2)	(3)	(4)	(5)	(6)
$Fmov_i$*After	-0.033	-0.002	-0.019	0.035	0.049	0.010
	(0.056)	(0.035)	(0.044)	(0.052)	(0.032)	(0.040)
Firm controls	Yes	Yes	Yes	Yes	Yes	Yes
Treated*Trend	Yes	Yes	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3,067	3,171	4,658	3,069	3,173	4,660
R-squared	0.247	0.173	0.158	0.247	0.175	0.158
Number of firms	1,003	1,020	1,236	1,004	1,021	1,237

Table 11: Property Law or global financial crisis?

This table tests if the disruption of redistribution is driven by global financial crisis. Columns (1) and (2) estimate for short event window covering one year before the Property Law and one year after. Columns (3) and (4) breakdown the post-reform period into sub-periods: pre-crisis year of 2007 (*Post1*), the crisis year of 2008 (*Crisis*), the years of stimulus package 2009-2010 (*Post2*), and the final year of the sample 2011(*Post3*). The dependent variable in this table is $Accounts\ Receivable_t/Total\ Assets_{t-1}$ (*AR/TA*). To proxy short-term bank credit, columns (1) and (3) use lagged short-term bank credit scaled by total asset, and columns (2) and (4) use short-term bank credit to total asset ratio measured at 2005. $Fmov_i$ is the median movable ratio measured over 2001-2005. $After_t$ is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var.	AR/TA			
SD=	$(ShortDebt/TA)_{it-1}$	$(ShortDebt/TA)_{i2005}$	$(ShortDebt/TA)_{it-1}$	$(ShortDebt/TA)_{i2005}$
	(1)	(2)	(3)	(4)
SD	-0.023		0.043*	
	(0.056)		(0.026)	
After*SD	0.080	0.072		
	(0.052)	(0.045)		
After*$Fmov_i$	0.076*	0.068*		
	(0.041)	(0.035)		
After*SD*$Fmov_i$	-0.515***	-0.487***		
	(0.167)	(0.145)		
Post1*SD			0.051	0.061
			(0.050)	(0.047)
Post1* $Fmov_i$			0.078**	0.079**
			(0.039)	(0.038)
Post1* SD*$Fmov_i$			-0.434***	-0.456***
			(0.163)	(0.154)
Crisis* SD			-0.000	0.022
			(0.037)	(0.046)
Crisis* $Fmov_i$			-0.021	0.007
			(0.040)	(0.041)
Crisis* SD*$Fmov_i$			-0.092	-0.207
			(0.125)	(0.154)
Post2* SD			0.032	0.009
			(0.033)	(0.031)
Post2* $Fmov_i$			0.047	0.046
			(0.042)	(0.046)
Post2* SD*$Fmov_i$			-0.143	-0.079
			(0.126)	(0.126)
Post3* SD			0.008	0.001
			(0.037)	(0.034)
Post3* $Fmov_i$			0.099*	0.113**
			(0.052)	(0.057)
Post3* SD*$Fmov_i$			-0.085	-0.125
			(0.132)	(0.137)
SD*$Fmov_i$	0.174		-0.022	
	(0.236)		(0.092)	
Treated*Trend	No	No	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes

Observations	2,301	2,300	10,179	10,315
R-squared	0.168	0.168	0.237	0.235
Number of firms	1,193	1,157	1,257	1,163

Table 12: Other robustness

This table tests the disruption of redistribution after controlling the tunneling reform (columns (1) and (2)), and controlling initial firm conditions (columns (3) and (4)). The dependent variable in this table is *Accounts Receivable$_t$/Total Assets$_{t-1}$* (*AR/TA*). Columns (1) and (3) use lagged short-term bank credit scaled by total asset, and columns (2) and (4) use short-term bank credit to total asset ratio measured at 2005. *Fmov$_i$* is the median movable ratio measured over 2001-2005. *After$_t$* is a dummy variable that equals one for the years 2007-2011, and zero otherwise. The observations from the reform year 2006 are excluded from the sample. *Fmorec$_i$* is the pre-reform median other receivable to asset ratio (*Other Receivable/Asset*). *Tunnel$_t$* is a dummy equals one for years after the tunneling reform in 2005, and zero otherwise. In columns (3) and (4), the initial controls are measured as pre-reform median value of each firm control variable used in Table 2, except *Age*, *List* and *State*. For *Age*, *List* and *State*, we use (only) value measured in 2005. Standard errors clustered at firm level are reported in parentheses. ***, **, and * implies significance at 1%, 5%, and 10% level, respectively.

Dep.Var.	AR/TA			
SD=	$(ShortDebt/TA)_{it-1}$	$(ShortDebt/TA)_{i2005}$	$(ShortDebt/TA)_{it-1}$	$(ShortDebt/TA)_{i2005}$
	(1)	(2)	(3)	(4)
SD	0.079***		0.027	
	(0.024)		(0.025)	
SD*After	0.048	0.048	0.032	0.030
	(0.031)	(0.030)	(0.031)	(0.031)
SD*After*Fmov$_i$	-0.188*	-0.196*	-0.191*	-0.201*
	(0.108)	(0.111)	(0.116)	(0.121)
SD*Fmov$_i$	-0.031		-0.007	
	(0.086)		(0.090)	
After*Fmov$_i$	-0.015	-0.019	-0.001	-0.009
	(0.031)	(0.031)	(0.034)	(0.034)
SD*Tunnel	-0.050**	-0.054***		
	(0.021)	(0.020)		
SD* Tunnel*Fmorec$_i$	0.273	0.290		
	(0.243)	(0.190)		
SD*Fmorec$_i$	-0.353			
	(0.229)			
Tunnel*Fmorec$_i$	-0.046	-0.044		
	(0.077)	(0.060)		
Firm controls	Yes	Yes	-	-
Initial controls*After	-	-	Yes	Yes
Treated*Trend	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Industry*Year	Yes	Yes	Yes	Yes
Province*Year	Yes	Yes	Yes	Yes
Observations	10,179	10,315	10,187	10,331
R-squared	0.234	0.233	0.191	0.182
Number of firms	1,257	1,163	1,257	1,163