MEASURING BANK COMPETITION IN CHINA: A COMPARISON OF NEW VERSUS CONVENTIONAL APPROACHES APPLIED TO LOAN MARKETS

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Abstract

Since the 1980s, important and progressive reforms have profoundly reshaped the structure of the Chinese banking system. 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 concluded that bank competition in China declined during the past decade, despite these reforms. In this paper, we show both empirically and theoretically 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 relatively new Profit Elasticity (PE) approach that was introduced in Boone (2008) as Relative Profit Differences (RPD) does not evidence 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. We provide additional empirical evidence that supports our results. We find that these, firstly, are in line with the process of financial reform, as measured by several indices, and secondly 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.

JEL classification: D4, G21, L1.

Keywords: competition, banking industry, China, lending markets, marginal costs, regulation, deregulation.

Resumen

Desde los años ochenta, la estructura del sistema bancario chino se ha visto reconfigurada a través de importantes reformas progresivas. Muchos estudios empíricos sugieren que la reforma financiera ha fomentado la competencia bancaria en la mayor parte de las economías avanzadas y emergentes. No obstante, algunos estudios previos que adoptaron enfoques convencionales para medir la competencia concluyeron que la competitividad bancaria en China descendió durante la década pasada, a pesar de estas reformas. En el presente análisis se muestra de forma teórica y empírica que esta aparente contradicción se debe a una medición errónea. Diversos indicadores convencionales —como el índice Lerner y el estadístico-H de Panzar-Rosse— no logran medir apropiadamente la competencia en los mercados de préstamo chinos, debido al sistema de regulación de los tipos de interés. Por el contrario, el enfoque relativamente novedoso de la elasticidad del beneficio *(Profit Elasticity,* PE) introducido por Boone (2008) a través de diferencias relativas en los *beneficios (Relative Profit Differences,* RPD) no sufre estas deficiencias. Mediante la utilización de información procedente de los balances de una gran muestra de bancos operativos en China durante 1996-2008 se muestra que, en realidad, la competencia se ha visto incrementada durante la pasada década, cuando se emplea el indicador PE. Se proporciona una evidencia empírica adicional que apoya nuestros resultados. En primer lugar, se confirma que estos se encuentran alineados con el proceso de reforma financiera, medido según diversos índices; y, en segundo lugar, que son robustos en un gran número de especificaciones alternativas y métodos de estimación. En términos generales, nuestro análisis sugiere que los mercados de préstamos bancarios en China han sido más competitivos de lo que se asumía en la literatura.

Códigos JEL: D4, G21, L1.

Palabras clave: competencia, sector bancario, China, mercados de crédito, costes marginales, regulación, desregulación.

1 Introduction

In recent years, the Chinese banking sector has become the focus of a growing body of empirical literature. Studies have tried to measure developments in bank efficiency and bank productivity, the contribution of bank intermediation to China's economic growth and particular characteristics of Chinese banks' corporate governance structures. Notwithstanding this progress in empirical research, relatively few econometrical analyses have concentrated explicitly on bank competition in China.

The apparent lack of these studies is somewhat surprising, for a number of reasons. First, although different views persist, several established observers have identified competition as one of the key factors in China's economic reform success. They argue that the expansion of incentives, mobility and markets has created unprecedented business opportunities, exemplified in astonishingly high scales of entry and exit. This process has pushed China's economy towards "extraordinarily high levels of competition", leading to a situation where "intense competition now pervades everyday economic life" (Brandt and Rawski, 2008).[1] These claims make one wonder whether this general picture can be extended to a specific sector such as banking as well.

Second, with China becoming one of the major global economic powers, overtaking Japan as the world's second-biggest economy at the end of 2010, also Chinese banks have become global powerhouses. The "Big Four" state-owned commercial banks (SOCBs) are now among the ten largest banks in the world according to market capitalisation, and China's banks accounted for 20% of global banking profits in 2010 (KPMG, 2011; Löchel and Li, 2011). They have embarked upon a global expansion strategy, opening new branches and subsidiaries abroad and forming cross-border alliances in bank services, while also developing internationally more diversified business lines in insurance and asset management. These global business advances originate from domestic market conditions, including certain degrees of efficiency, contestability and competition, providing additional arguments for investigating the latter.

Third, the process of financial reform has reshaped China's banking sector profoundly. The prevailing view is that after 30 years of financial reform, China's banks have begun to behave more like commercial banks in the developed world (Firth et al., 2009). During this period, the banking landscape in China has changed dramatically. In essence, it moved from a socialist monolithic structure to a pluralistic system comprising various groups of market-oriented banks. The latter development started with the creation of a "two-tier" banking sector in 1978 with the establishment of the People's Bank of China (PBC) as the central bank and four large state-owned specialised banks serving specific sectors of the economy. It took 16 years before the Chinese government started a new major round of reforms, when in 1994 three specialised policy banks were established to take over the policy loans from the "Big Four" banks and the status of the latter was changed to state-owned commercial banks. The conditions for China's WTO accession and its actual entry into this organisation in 2001 triggered a major third round of reforms, including the establishment of an independent bank supervisor (China Banking Regulatory Commission or CBRC) and

1. This despite China being a clear counter-example with respect to the conclusions from the existing literature on law, institutions, finance and growth on the advocated institutional framework conducive to economic development (Allen et al., 2005).

further measures regarding the liberalisation of interest rates, business scope and market entry. Overall, while acknowledging remaining deficiencies, the Chinese banking sector has moved gradually but unmistakably towards a commercially oriented market-type system, in which competitive forces should take on much greater significance.

Last but certainly not least, investigation of bank competition in China contributes to the continuing debate on which empirical approaches may be the most suitable for measuring competition in specific banking systems. This argument 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 and/or regions. 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). These findings suggest that it may be preferable to consider different measures when assessing bank competition.

This paper contributes to the literature on both Chinese banking and bank competition by arguing that conventional measures of competition like the Lerner index and the Panzar-Rosse H-statistic may not assess bank competition in China correctly, mainly due to the existence of interest rate regulations. The former uses the profit or price cost margin (PCM), ie the mark-up in output prices above marginal cost, as an indicator of market power. The latter measures to what extent input and output prices move in step (as they would under perfect competition) or out of step (indicating monopoly or a perfect cartel) (Bikker, 2010). The shortcomings of these approaches in the context of Chinese banking markets can be summarised as follows. First, the Lerner index and Panzar-Rosse H-statistic may bias the results due to the system of interest rate regulation in China. Second, several other characteristics of Chinese banking make these conventional measures less appropriate as well, in our view.

Instead, we argue that the relatively new Profit Elasticity (PE) indicator may be better suited to investigate competitive conditions in Chinese loan markets, given the particularities of the banking industry in China. This indicator, whose theoretical base is the Relative Profit Differences (RPD) concept, is based on the idea that competition rewards efficiency (Boone et al., 2007; Boone, 2008; Van Leuvensteijn et al., 2011 and 2013). In general, an efficient firm will realise higher profits than a less efficient one. Crucial for the PE indicator is that this effect will be stronger the more competitive the market is.[2] This can be explained as follows. In the theoretical setup of RPD, competition increases due to increased interaction between banks or due to lower entry costs. Boone (2008) shows that RPD is an increasing function of the degree of interaction between firms and a decreasing function of entry costs. Hence, RPD increases when competition intensifies, ie stronger competition decreases (increases) profits of more efficient firms by smaller (larger) amounts than for less efficient firms. The underlying intuition is that in a more competitive market, firms are punished more harshly (in terms of profits) for being inefficient. All in all, the PE indicator is a new measure of competition that is more robust from both a theoretical and an empirical point of view when compared with more conventional measures (Boone et al., 2007). Early empirical applications of the PE indicator are Van Leuvensteijn et al. (2007), Van Leuvensteijn (2008) and Bikker and Van Leuvensteijn (2008), while a more explicit empirical validation has been provided by Boone and Van Leuvensteijn (2010).

2. As with all measures of competition, the PE indicator is also based on certain assumptions, such as that of product homogeneity (product innovation does not matter). These assumptions will be discussed in more detail in Section 4.2.

The paper contributes to the existing literature in a number of ways. It combines, according to our knowledge, for the first time the Lerner, Panzar-Rosse and PE approaches in one paper.[3] We also demonstrate theoretically that RPD, from which the PE indicator is derived, is much better suited to assess competition in banking markets where interest rates are regulated, whereas in that case the Lerner index yields biased results. Furthermore, we assess the relationship of the various approaches with the process of financial reform. All in all, the paper provides an extensive discussion of the appropriateness of competition indicators for a specific country, in this case China, and argues that banking system specifics may significantly affect the results of different indicators of competition. Hence, we warn against "auto-pilot" applications of empirical measures to assess banking competition across countries. Moreover, the paper yields important new results regarding the development of competition in Chinese loan markets.

These results can be summarised as follows. First, the conventional measures – both the Lerner index and the Panzar-Rosse H-statistic – show declining competition over time in Chinese loan markets. This despite the comprehensive process of financial reform, which according to a large body of empirical research promoted banking competition in many mature and other emerging market economies. The results for both indicators hold for alternative specifications, suggesting that they in themselves are estimated correctly. Second, we find that the long-run equilibrium assumption on which the H-statistic depends is violated, also for alternative specifications. This may be related to the financial reform process which continues to reshape the banking landscape in China profoundly. Hence, inferring competitive conditions from it for China is likely to give rise to bias (apart from in our view fundamental biases inherent in the Panzar-Rosse model with respect to China). Third, in contrast, the findings for the PE indicator show 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. Fourth, 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 Lerner index and Panzar-Rosse H-statistic as appropriate measures to assess Chinese banking competition validated by the empirical results. The results support our belief that the PE indicator is better suited for this purpose. Overall, we contribute to both the ongoing discussion in the banking competition literature on the appropriateness of different competition measures and the growing empirical body of research on Chinese banking.

The remainder of this paper is organised as follows. Section 2 provides a brief overview of the research literature on empirical measures of bank competition. Section 3 first gives background information on the structure of Chinese banking (3.1), followed by a review of China-specific studies on bank efficiency and competition (3.2). Section 4 presents the methodological framework of the standard and elasticity-adjusted Lerner indices (4.1) and the PE indicator (4.2). Moreover, it demonstrates that the Lerner index yields biased results under binding interest rate regulation and RPD not (4.3). Section 5 shows our data and sample characteristics. Section 6 presents the empirical results for the (elasticity-adjusted) Lerner index (6.1) and the PE indicator (6.2). It then compares the results of the various empirical measures (6.3) and presents a detailed interpretation of the results from the PE indicator, including their relationship with various financial reform indicators (6.4). Section 7 concludes.

3. The Panzar-Rosse H-statistic is presented in Appendix A.

2 Review of empirical literature on bank competition

Competition in the banking sector has generally been analysed on the basis of two concepts, ie market power and efficiency, which have sometimes been tested jointly. Here, market power reflects the ability of specific banks to control the market for bank products, whereas efficiency relates to the ability of specific banks to produce output (such as loans) at minimal cost.

With respect to the first concept for investigating bank competition, a well-known approach to measuring market power is suggested by Bresnahan (1982) and Lau (1982). These authors analyse bank behaviour on an aggregate level and estimate the market power of an average bank following a specific short-run model. Empirical studies based on this approach are rather scarce though, as it is very data-intensive. Another approach based on market power has been proposed by Panzar and Rosse (1987), which is the so-called "H-statistic" and which will be discussed in greater detail in Appendix A. The specific value of the "H-statistic" measure indicates how competitive the market is, ranging from a monopoly (or perfect collusion) to a situation of perfect competition. A third indicator for market power is the "Hirschman-Herfindahl Index" (HHI), which measures the degree of market concentration. This indicator is often used in the context of the "Structure Conduct Performance" (SCP) model, which assumes that market structure affects banks' behaviour, which in turn determines their performance. The idea is that a highly concentrated banking sector (with a few banks occupying significant market shares) can impair competition, in the sense that concentration translates into greater market power, resulting in collusive behaviour and excess profits for banks. Finally, market power may also be related to profits, in the sense that extremely high profits may be indicative of a lack of competition. A traditional measure of profitability is the price-cost margin (PCM), which is equal to the output price minus the marginal costs. The PCM is frequently used in the empirical industrial organisation literature as an empirical approximation of the Lerner index (see Section 4.1) and in this context applied to the banking sector as well.

With respect to the second concept for measuring bank competition, ie efficiency, this indicator is often seen as a proxy for competition, in the sense that the most efficient banks (and therefore the most competitive ones) will gain market share at the cost of the less efficient banks. A relatively recent method, which has been named the profit elasticity (PE) model or the "Boone" indicator, can be seen as an elaboration on this efficiency hypothesis. This measure has gained considerable support more recently (Van Leuvensteijn et al., 2007, 2011 and 2013; Van Leuvensteijn, 2008; Schaeck and Cihák, 2010; Delis, 2012; Tabak et al., 2012). The underlying model will be explained in more detail in Section 4.2.

The actual literature on the measurement of competition is generally categorised into two major streams (Bikker, 2004; Tabak et al., 2012). So-called structural approaches are based on the SCP model and use concentration indicators as proxies for competition, such as the HHI and the CRn which measures the market shares of the n largest banks. In contrast, non-structural approaches have been promoted within the so-called New Empirical Industrial Organisation (NEIO) literature. They estimate parameters that reflect the degree of competition in specific markets based on bank-level data and specific assumptions on the behaviour of banks. The Bresnahan, Lau and Panzar-Rosse approaches mentioned above, as well as the Lerner index and PE elasticity, fall into this part of the literature. While these measures have been broadly accepted, there is no consensus regarding which is the “best” indicator for gauging bank competition (Carbó Valverde et al., 2009). As a matter of fact, they often reach different conclusions for banking systems of the same countries and groups of countries. Consequently, leading experts in the field acknowledge that the ability of empirical research to capture the degree of bank competition is still imperfect and that developing proper competitiveness tests and methodologies will remain an important area of research (Claessens, 2009, p. 95; Claessens and Laeven, 2004, p. 581).

At the same time, a view that has been gaining support is that concentration may not be the most appropriate indicator for measuring bank competition (Bikker, 2004; Casu and Girardone, 2006 and 2009; Schaeck et al., 2009). Moreover, concentration does not measure the competitive conduct of banks at the margin. In addition, concentration indices such as the HHI do not distinguish between small and large countries and may incorrectly suggest that competition is declining, while in fact concentration and competition are increasing simultaneously as a result of bank consolidation (Van Leuvensteijn, 2008; Van Leuvensteijn et al., 2013).

One of the main issues that has been at the core of empirical studies on bank competition, and which may be especially relevant for China, is the impact of financial reform on the degree of competitiveness of banking systems and subsequent efficiency and welfare gains. There is considerable empirical evidence that deregulation fostered competitive conditions in banking markets and led to greater product differentiation, lower cost of and improved access to financial services and greater financial stability (Claessens, 2009). These effects have been validated empirically especially for the United States and emerging market economies, while results have been less conclusive for Europe.

3 Background on Chinese banking

3.1 The structure of the Chinese banking industry[4]

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.[5] In this section, we provide an overview of developments which are of particular relevance for the investigation of competition in Chinese loan markets. These are the start of commercial banking in China, the entrance of new players, which was promoted by China joining the WTO in 2001, and the deregulation of the credit control system and of interest rates.

The Commercial Bank Law of the People's Republic of China was promulgated in May 1995, which paved the way for the development of a commercial banking system in China and the entrance of important new players (Fu and Heffernan, 2009). In this context, 12 so-called joint-stock commercial banks (JSCBs) and more than 100 city commercial banks (CCBs) were established. The former initially offered banking services only regionally, but later they were allowed to operate freely nationwide, competing with the state-owned commercial banks (SOCBs) for the large firms and with the CCBs for small and medium-sized enterprises. 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 (Sun and Yamori, 2011).

The commercialisation of Chinese banking was triggered by mounting problems at the four state-owned specialised banks which concentrated on providing loans to state-owned enterprises in specific sectors. These "Big Four" (Bank of China, Agricultural Bank of China, China Construction Bank, and Industrial and Commercial Bank of China) experienced a significant deterioration of their asset quality in the early 1990s. This was reflected in a rapid increase of their non-performing loans, which consisted predominantly of loans to state-owned enterprises, granted mainly for political instead of business reasons. Three specialised policy banks were established to take over the policy-lending business of the "Big Four" banks. The latter were converted into state-owned commercial banks (SOCBs), and four asset management companies were given the task of absorbing their non-performing loans. 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 1, 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,

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4. This section will only pay attention to the major banks which are covered in our study and ignore smaller banks such as rural credit co-operatives. In the same vein, only those aspects of financial reform that are related to the banks in our sample will be discussed.

5. According to the China Banking Regulatory Commission (CBRC), financial reform in China can be classified in three major stages (1978-1993, 1994-2002, 2003-present), which are discussed extensively in Liu (2009). Clear overviews in English are presented in, for example, Matthews and Zhang (2010) and Chang et al. (2012).

during 2002-2008.[6] These declines in market shares have been mirrored in considerable increases in those of especially the JSCBs and of the CCBs as well.

Overview of the Chinese banking sector 1996–2008

Table 1

	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
	Share of total securities (%)				Share of pre-tax profits (%)				Return on assets (%)			
	SOCB	JSCB	CCB	FOREIGN	SOCB	JSCB	CCB	FOREIGN	SOCB	JSCB	CCB	FOREIGN
Average 1996–2001	81.16	15.62	3.15	0.06	71.21	23.83	4.52	0.43	0.16	0.70	1.03	0.99
Average 2002–2008	76.45	17.20	5.95	0.41	71.85	21.19	5.90	1.05	0.72	0.45	0.63	0.93
Average 1996–2008	78.62	16.47	4.66	0.25	71.77	21.54	5.72	0.97	0.46	0.57	0.82	0.96

Source: BankScope, authors' own calculations. The data are presented for the full sample of 1996–2008 and the two subsamples that we use in the estimations (1996–2001 and 2002–2008), with 2001 being the year of China's entry in 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. Under the conditions of WTO membership, the activities of foreign banks were liberalised profoundly. For example, these 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 take equity stakes of up to 25%. Chinese banking markets have been opened to foreign competition rather comprehensively since the end of 2006, with foreign banks receiving in principle the same regulatory treatment as their Chinese counterparts and expanding their business further into the rest of the country (Yao et al., 2008; Matthews and Zhang, 2010; Xu, 2011). The foreign liberalisation of Chinese banking resulted in a sharp increase in the number of foreign players, whose number increased in our sample from four in 1996 to 26 in 2008 (see Section 5). Table 1 shows that the market shares of foreign banks in total assets and loans increased as well during 1996-2008, but remained below 1%. Their share in pre-tax profits rose to an annual average of above 1% in 2001-2008, reflecting a return on assets that has been the highest of all banking groups.

Despite the relatively small market shares of foreign banks in China, in our view the importance of their role in Chinese banking should not be underestimated. Namely, it has been argued rather widely that the impact of foreign banks on bank efficiency and competition may have significantly exceeded what their modest presence in China may suggest. First, various studies suggest that both the threat of foreign entry and its actual realisation forced Chinese banks to respond in terms of improving their business models, efficiency, market practices and hence their degree of competitiveness (He and Fan, 2004; Leung and Chan, 2006). Second, foreign banks have obtained minority equity stakes in various Chinese banks, which overall seem to have had positive effects on the efficiency and/or performance of the latter and hence likely on competition in Chinese banking markets

6. 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.

(Berger et al., 2010). Third, foreign banks have tried to penetrate into other parts of China beyond the main cities through equity partnerships or less institutionalised forms of co-operation with Chinese banks with either geographical significance or appreciable levels of national coverage (He and Fan, 2004; Leung and Chan, 2006). Fourth, some observers claim that foreign banks have demonstrated that they are capable of snatching significant market shares away from Chinese banks, such as in RMB loan and deposit markets in key cities such as Beijing, Shanghai and Guangzhou (Xu and Lin, 2007). Finally, Xu (2011) provides empirical evidence that foreign bank entry into China has been supportive of developing a more competitive banking industry. All in all, we believe that these arguments support the inclusion of foreign banks in empirical investigations of bank competition in China.

Another important reform affecting Chinese loan markets was the replacement of the PBC's 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 (Mo, 1999). 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 emphasise the use by the PBC of quantitative instruments aimed at controlling credit growth, despite the discontinuation of the binding credit plan system (Goodfriend and Prasad, 2006; Liu and Xie, 2006; Delatte, 2007; Geiger, 2008; Du, 2010; Fukumoto et al., 2010; Huang et al., 2010; He and Wang, 2011 and 2012; Ma et al., 2011; Chen et al., 2012; Martin, 2012; World Bank, 2012). 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 in principle the sectoral direction of lending, although it is suspected that in practice the guidance has also affected the amount of lending (Green, 2005; Okazaki, 2007).

The reform of the credit control system has been followed – in terms of degree of deregulation – by interest rate liberalisation, resulting in relatively liberalised bank interest rates in 2004, when the deposit rate floor and the lending rate ceiling were eliminated for the major banks (Figure 1, left-hand panel).[7] The liberalisation of the ceiling on the lending rate and the floor on the deposit rate in October 2004 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). The lending rate floor should inhibit competition to some extent, although probably not in a binding manner.[8] In fact, during December 2004 and

7. The PBC started to widen the floating band on banks' interest rates from 1998 onwards, after it liberalised interbank interest rates. Hence, it gave commercial banks more discretion in setting loan rates (People's Bank of China, 2005; Feyzioğlu et al., 2009).

8. Fu and Heffernan (2009) mention that Chinese banks have always had some discretion over certain loan rates. For example, from 1999, small business loans could carry a premium of up to 30% (50% for the rural credit cooperatives) over the central bank rate.

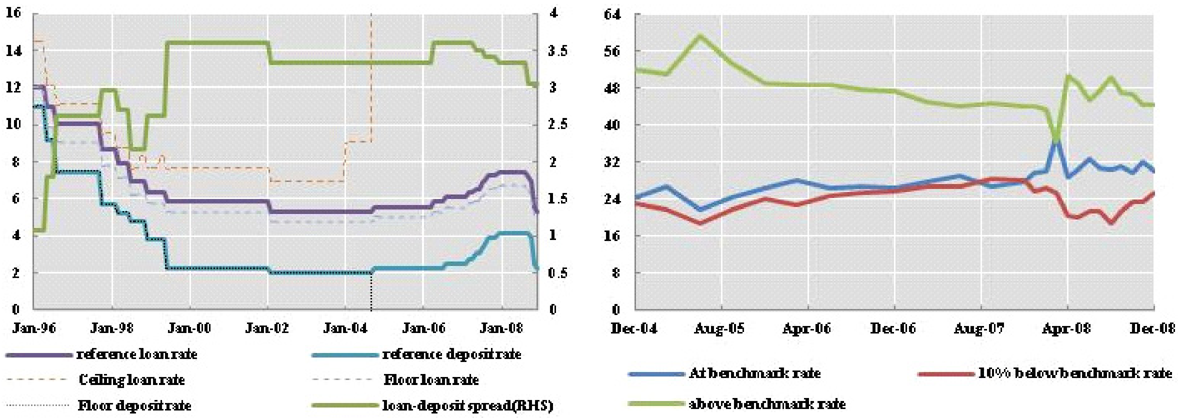
December 2008 only between 19% and 29% of all loans were made at the floor lending rate, suggesting that most loan rates were higher (Figure 1, right-hand panel). In contrast, empirical research has suggested that the ceiling on deposit rates has been binding and put them at a level below equilibrium (Feyzioğlu et al., 2009; He and Wang, 2011; Ma et al., 2011).

Interest rates in China

Figure 1

Panel A: One-year interest rates (in %)

Panel B: Share of loans issued at different rates (in %)



Sources: CEIC database, Authors' own calculations. All rates are one-year rates.

The liberalisation of both credit controls and interest rates has had important positive repercussions for Chinese banks' lending policies. Many observers adhere to the view that these policies have become more market-oriented, characterised by increasing attention to risk management and loan monitoring and by greater diversification into less traditional areas such as consumer lending (Allen et al., 2005; Leung and Chan, 2006; Abiad et al., 2010; Herd et al., 2010; Qian et al., 2011; IMF, 2011). In principle, this development should have been supportive of enhancing competitive conditions in Chinese loan markets.

This development has been aided by the establishment of the China Banking Regulatory Commission (CBRC) in 2003, which helped bring lending policy more into line with market-conform assessment and approval criteria (Yeung, 2009). Moreover, since October 2004, when Chinese banks were permitted to use their own judgement in setting lending rates, credit risk has been much better reflected in the lending rate setting process (People's Bank of China, 2005). Another important improvement is that lending decisions have become more centralised through a reduction of the autonomy of regional offices and branches, and of any possible interference by local and regional governments and Party officials (He and Fan, 2004; Dobson and Kashyap, 2006). All in all, it has been suggested that banks in China have been considerably freed from political pressure, for example to make loans to support state-owned enterprises (Lardy, 2006).[9] This should foster a more competitive environment.

To conclude, this brief summary of financial reform and banking structure in China yields a number of important insights for our analysis of competitive conditions in Chinese loan markets. First, the process of financial reform has led to a fundamental change in the orientation of Chinese banks from mere instruments of economic and industrial policy

9. Notwithstanding proofs of considerable government interference, Cull and Xu (2000) provide evidence that lending by Chinese banks to state-owned companies was guided by credit risk considerations as well.

towards in essence commercially oriented providers of a broad range of financial services to non-financial corporations, central, regional and local governments and households (Cheng and Degryse, 2010).[10] Second, more specifically, the lending policies of Chinese banks have been founded much more on market-based criteria. This has been aided by the liberalisation of interest rates and strict credit controls. Loan rates have been liberalised gradually since 1998, first by increasing their ceilings and later by removing them altogether. Moreover, the abolishment of binding credit plans has provided banks in China with greater opportunities to diversify and optimise their lending business. As demonstrated by Abiad et al. (2010), the process of financial reform has made the biggest advances regarding credit controls. We concur with Huang (2010) that credit allocation has increasingly become more market-oriented. Third, notwithstanding the growing adherence to market principles in Chinese banks' lending policies, the fact that the PBC has continued, at least during certain episodes, to provide loan guidelines is important for our investigation. While the effectiveness of these quantitative instruments is not clear and may actually have been limited, their continued use makes us rather sceptical about employing banks' market shares to assess competitive conditions in Chinese loan markets.[11] Finally, the international liberalisation of China has promoted the presence and most likely the influence of foreign banks, arguably most importantly through their impact on the strategies and operations of Chinese banks. Overall, the broad range of financial reform measures implemented may have promoted over time a priori competitive conditions in Chinese banking markets.

3.2 Empirical literature on bank efficiency and competition in China

One of the most frequently investigated aspects of the Chinese banking sector is its efficiency.[12] Most studies have characterised Chinese banks as still being relatively inefficient, despite improvements since the highly regulated and inefficient banking system of the past (Fu and Heffernan, 2007; Yao et al., 2007; Berger et al., 2009a; García-Herrero et al., 2009). This also applies when they are compared with their international peers operating outside China (Singh Pritam Singh and Munisamy, 2008; Löchel and Li, 2011; Allen et al., 2012).[13] To what extent bank efficiency in China has improved over time is still subject to considerable debate, although some studies have argued that considerable efficiency gains have been realised in more recent years (see for example: Matthews et al., 2009; Barros et al., 2011).

The discussion on efficiency improvements has concentrated on two issues, which in empirical studies often yield mixed results (Berger et al., 2009a; Chang et al., 2012): the impact of bank deregulation on bank efficiency (Kumbhakar and Wang, 2007; Fu and Heffernan, 2007) and the relationship between the efficiency and performance of banks in China. On the former, some support the view that efficiency has improved in parallel with the process of financial reform (Feyzioğlu, 2009), while others are more sceptical and tend to

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10. It needs to be acknowledged that some observers remain rather sceptical about the degree of progress achieved in Chinese financial reform. For example, World Bank (2012) concludes that "[d]espite the many reforms introduced so far, the Chinese financial system remains oppressed, unbalanced, costly to maintain and potentially unstable. […] Continued protection and intervention in the business decisions of financial institutions make them convenient policy instruments, the use of which prolongs the bureaucratic culture and distorted incentives that have prevented banks from full commercialization and from allocation of financial resources to the most productive uses".

11. Van Leuvensteijn et al. (2011 and 2013) use market shares in loan markets to investigate competition in a number of mature economies.

12. Most of the empirical studies focus on cost and profit efficiency using the stochastic frontier approach; examples include: Fu and Heffernan (2007), Feyzioğlu (2009), Berger et al. (2009a). Some researchers focus on technology efficiency and sources of total factor productivity growth by constructing various productivity indices (Chang et al., 2012; Matthews et al., 2009). Chang et al. (2012) provide a comprehensive review of various efficiency measures applied to Chinese banking.

13. In contrast, Matthews et al. (2009) argue that Chinese banks have achieved significant improvements in bank efficiency and hence may not be out of line with international bank efficiency benchmarks.

highlight that, despite tangible liberalisation efforts, China's banking sector is still relatively inefficient or not sufficiently commercially oriented (Allen et al., 2005; Dobson and Kashyap, 2006; Podpiera, 2006). A few studies document that reforms during the early 1990s did not seem to have affected performance (Park and Sehrt, 2001). On the latter, some studies demonstrate that Chinese bank performance has been affected positively by efficiency improvements, and that more profitable banks tend to be more efficient than less profitable banks (Heffernan and Fu, 2010; García-Herrero et al., 2009). In contrast, alternative ones emphasise that the high profitability of Chinese banks which has been observed in recent years, for example in terms of both return on average assets and equity in comparison with those of international peers, is not correlated with their efficiency, but is largely driven by wide interest rate spreads between loan and deposit rates and low personnel costs (Feyzioğlu, 2009; Löchel and Li, 2011).

Results of empirical studies on the efficiency of Chinese banks are the most conclusive when conducted for specific banking groups. The SOCBs seem to have been the least efficient commercial banks in China, especially in comparison with the JSCBs and/or foreign banks (Kumbhakar and Wang, 2007; Fu and Heffernan, 2007; Yao et al., 2007; Ariff and Can, 2008; Berger et al., 2009a; Feyzioğlu, 2009; Fu and Heffernan, 2009; Lin and Zhang, 2009; Matthews et al., 2009). Some argue that this may be related to their relative lack of shareholder diversification, compared with JSCBs, CCBs and foreign banks, whose plurality of shareholders may reduce political interference (Ferri, 2009). Shih et al. (2007) and Feyzioğlu (2009) find that the JSCBs perform significantly better than both the SOCBs and CCBs. Moreover, the impact of Chinese WTO membership since 2001 on bank efficiency has been investigated. Some studies show that the efficiency of domestic Chinese banks weakened after China's accession to the WTO, while that of foreign banks increased (Rezvanian et al., 2011). Regarding specific aspects of international liberalisation, a recent study shows that reforms which promoted foreign (and reduced state) ownership of Chinese banks had strong favourable effects on their efficiency (Berger et al., 2009a). García-Herrero and Santabárbara (2008) demonstrate that this applies especially to foreign participation through minority strategic partnerships (see also Hasan and Xie, 2012). Berger et al. (2010) show that foreign ownership has some positive effects on performance, as it mitigates certain adverse effects of business diversification. On the other hand, Heffernan and Fu (2010) argue that foreign equity investment did not have a significant influence on performance (measured by different indicators).

Only a few papers investigate explicitly and in-depth competition in the Chinese banking sector by using econometric tools, while others discuss it on the sidelines and often adopt more descriptive approaches. One of the first to address this issue adopts the structural approach (see Section 2) and calculates concentration indicators (HHI) which show high degrees of concentration that may inhibit competition (Wong and Wong, 2001). Furthermore, this study concentrates qualitatively on institutional characteristics which inhibited bank competition in China during the 1990s, such as government interference, information deficiencies and a weak legal infrastructure. At the same time, this study acknowledges that reforms stipulated under the conditions for China's WTO accession helped to create a more competitive and efficient 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 and Heffernan (2009) look at the effect of financial reform on China's banking structure and performance during 1985-2002 and draw the conclusion that the estimation of structure-performance models lends some support to the existence of relative market power in the 1980s and early 1990s. Overall, they conclude that to improve competition, new policies should be adopted to encourage market entry and to increase the market share of the most efficient banks. 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. Moreover, Fu (2009) shows that the H-statistic for core commercial banking activities was higher before WTO entry than after, suggesting that competition in Chinese loan markets was higher during the period before joining the WTO than in the period after. Bikker et al. (2007), 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 (ie lowest Lerner index). 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). Both papers report Lerner indices for China that lie predominantly between 0.3 and 0.4 for 2002-2008.

All in all, the results of empirical studies, both for China and more generally (see Section 2), point to several lessons for the analysis of bank competition in China. First, it may be worthwhile to compare the results of various measures of competition, such as the Lerner index, Panzer-Rosse and PE approaches, since a priori the literature does not suggest a superior method. Second, concentration indicators do not seem to be appropriate competition measures. Hence, we do not employ them in our analysis. Finally, banking deregulation and foreign bank entry generally have affected bank competition favourably. This may be a finding especially relevant for China, taking into account its elaborate process of financial reform.

4 Competition measures: methodology and theory.

In this section, we discuss the conventional Lerner and elasticity-adjusted Lerner measures (4.1) and the relatively new Profit Elasticity (PE) indicator (4.2). We theoretically show that the former fail to measure competition correctly when interest rates are regulated, while the latter does not suffer from this shortcoming (4.3).

Our discussion of the conventional Panzar-Rosse H-statistic is conducted in Appendix A, as its flaws in measuring competition under regulated interest rates are rather clear and do not require a theoretical assessment. Namely, if due to binding regulation deposit and lending rates move in step, the H-statistic will be biased upwards and measure incorrectly a higher degree of competition. We adopt the H-statistic merely to replicate and check the results of Yuan (2006) and Fu (2009).

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., 2005 and 2007; Berger et al., 2009b).[14] 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:

$$L_{it} = \left(p_{it} - mc_{it}\right)/p_{it} \tag{4.1}$$

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_{it} are marginal costs of loans.

In order to be able to calculate marginal costs of loans, we first estimate a Translog Cost Function (TCF) using individual bank observations (Van Leuvensteijn et al., 2007). This function assumes that the technology of an individual bank can be described by one multiproduct production function. Under proper conditions, a dual cost function can be derived from such a production function, using output levels and factor prices as arguments. A TCF is a second-order Taylor expansion around the mean of a generic dual cost function with all variables appearing as logarithms. It is a flexible functional form that has proven to be an effective tool in explaining multiproduct bank services. Our TCF has the following form:

$$\begin{aligned} \ln c_{it} &= \alpha_0 + \Sigma_{t=1,\dots,(T-1)} \gamma_t d_t + + \Sigma_{h=1,\dots,(H-1)} \eta_h d_h \\ &+ \Sigma_{j=1,\dots,K} \delta_j \ln x_{ijt} + \Sigma_{j=1,\dots,K} \Sigma_{k=1,\dots,K} \epsilon_{jk} \ln x_{ijt} \ln x_{ikt} + \nu_{it} \end{aligned} \tag{4.2}$$

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,*

14. The various measures that we adopt in this paper can be linked theoretically. For example, Shaffer (1983) demonstrated that the Lerner index can be derived in terms of the Panzar-Rosse H-statistic. We consider this as future work.

CCB).[15] 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, eg 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., 2007). v_{it} is the error term.

Two standard properties of cost functions are linear homogeneity in the input prices and cost exhaustion (see eg Beattie and Taylor, 1985; Jorgenson, 1986). They impose the following restrictions on the parameters, assuming – without loss of generality – that the indices j and k of the two sum terms in equation (4.2) are equal to 1, 2 or 3, respectively, for wages, funding rates and prices of other expenses:

$$\delta_1 + \delta_2 + \delta_3 = 1, \epsilon_{1j} + \epsilon_{2jk} + \epsilon_{3j} = 0 \textit{ for } j = 1,2,3, \textit{ and } \epsilon_{k,1} + \epsilon_{k,2} + \epsilon_{k,3} = 0 \textit{ for } k = 4,...,K \quad (4.3)$$

The first restriction stems from cost exhaustion, reflecting the fact that the sum of cost shares is equal to unity. In other words, the value of the three inputs is equal to total costs. Linear homogeneity in the input prices requires that the three linear input price elasticities ($\boldsymbol{\delta}_i$) add up to one, whereas the squared and cross terms of all explanatory variables (ε_{ij}) add up to zero. Again without loss of generality, we also apply symmetry restrictions $\boldsymbol{\varepsilon}_{j,k} = \boldsymbol{\varepsilon}_{jk}$ for j,k = 1,…, K.

The marginal costs of loans are obtained by differentiating the TCF (4.2) with respect to loans, namely:

$$mc_{i1t} = c_{it} / x_{i1t} \left(\delta_1 + 2\epsilon_{1l} \ln x_{ilt} + \Sigma_{k=1,...,K;k \neq l} \epsilon_{1k} \ln x_{ikt} \right) \quad (4.4)$$

Once marginal costs of loans are obtained, an individual bank's Lerner index is calculated according to equation (4.1). The yearly Lerner index L_t is then the average of the 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 pre-WTO (1996-2001) and post-WTO (2002-2008).

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 normalises 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-maximising problem:

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15. 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 E (E.5) as an additional robustness test.

$$Max_{qi}\Pi = p(Q)q_i - C(q_i, w_i) \tag{4.5}$$

where $Q= \Sigma_i q_i$ 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:

$$p_i = C'(q_i, w_i) - \Theta_i/\varepsilon \tag{4.6}$$

where Θ_i is defined as the conjectural elasticity of total industry output with respect to the output of bank i, and $\boldsymbol{\varepsilon}$ is the market semi-price elasticity of demand, namely $\Theta_i = (dQ/dq_i)/(Q/q_i)$ and $\boldsymbol{\varepsilon} = (dQ/dp)/Q$. In a perfectly competitive market, Θ_i equals zero for all banks i, 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 $\lambda = \Theta_i/\boldsymbol{\varepsilon}$ if the goal is to evaluate the industry's overall degree of market power.[16] The elasticity-adjusted Lerner index will then be defined as $L = \Lambda/p$, 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 a 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 (4.2) and (4.4), respectively.

Substituting the marginal costs equation (4.4) into the supply equation (4.6), we obtain:

$$p_{it} = \frac{c_{it}}{x_{ilt}}\left(\delta_1 + 2\epsilon_{1l} lnx_{ilt} + \sum_{k=1\dots K;k\neq l}\epsilon_{1k} lnx_{ikt}\right) + \sum_{t=1\dots T-1}\lambda_t d_t + \varepsilon_{it} \tag{4.7}$$

where d_t is a year dummy and $\boldsymbol{\varepsilon}_{it}$ is the error term.

To access the evolution of bank competition measured by the elasticity-adjusted Lerner index, we perform two types of regressions: yearly estimates and subsample estimates. The yearly elasticity-adjusted Lerner index is then derived as λ_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 (4.2) and (4.7), replacing year dummies d_t with subsample dummies in both equations. The subsample elasticity-adjusted Lerner index is then defined as $\lambda_{subsample}/p_{subsample}$, where $p_{subsample}$ is the average loan rate for each subsample. The estimation is carried out with 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.

Unlike the traditional Lerner index, the elasticity-adjusted Lerner index has the advantage of allowing for formal tests whether its estimated values over time are statistically different from zero or one and whether subsample estimates differ significantly. These

16. As a robustness test, we estimate explicitly the conjectural variation parameter as a direct measure of competition in Appendix E (E.3).

advantages are crucial for our research, as we want to investigate if competitive conditions changed significantly over time.

4.2 The Profit Elasticity (PE) model

The PE indicator, or Relative Profit Differences (RPD), 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 heavier the competition in that market is (Van Leuvensteijn et al., 2007). While RPD can be seen as the theoretical model underlying this work, the PE indicator is the empirical operationalisation of this model. Boone (2008) shows that there is a continuous and monotonically increasing relationship between RPD and the level of competition if firms are ranked by decreasing efficiency. In other words, there is a negative relationship between efficiency, measured in terms of marginal costs, and profits; the more intense this negative relationship is, the more competitive markets will be. So, in practice, the PE indicator will have a negative sign when the relationship between marginal costs and profits is estimated, and it will be more negative the higher the level of competition is.

The fact that this relationship is both continuous and monotonic is the main advantage of RPD over more traditional measures of competition such as the HHI and Lerner index (or 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.[17]

The PE indicator (also referred to as the Boone indicator) has been developed in a broad set of theoretical models (Boone, 2000, 2001 and 2008; Boone et al., 2004; Boone et al., 2007; CPB, 2000). Following Boone et al. (2004), 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:

$$p\left(q_i, q_{j\neq i}\right) = a - bq_i - d\Sigma_{j\neq i} q_j \tag{4.8}$$

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:

$$a - 2bq_i - d\Sigma_{i\neq j} q_j - mc_i = 0 \tag{4.9}$$

When *N* banks produce positive output levels, we can solve the *N* first-order conditions (4.9), yielding:

$$q_i\left(mc_i\right) = \frac{\left[\left(2b/d - 1\right)a - \left(2b/d + N - 1\right)mc_i + \Sigma_j mc_j\right]}{\left[\left(2b + d\left(N-1\right)\right)\left(2b/d - 1\right)\right]} \tag{4.10}$$

We define profits π_i as variable profits excluding entry costs $\boldsymbol{\varepsilon}$. Hence, a bank enters the banking industry if, and only if, $\pi_i \geq \boldsymbol{\varepsilon}$ in equilibrium. Note that Equation (4.10) provides a relationship between output and marginal costs.

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17. Cournot competition is an economic model used to describe an industry structure in which companies compete on the amount of output they will produce, which they decide on independently of each other and at the same time. Bertrand competition assumes that firms compete on price and not output quantity (market shares). For more detail, see Tirole (1988).

It follows from $\pi_i\ (mc_i) = (p_i - mc_i)\ q_i$ that profits depend on marginal costs in a quadratic way, ie

$$\pi_i(mc_i) = \frac{(2b/d-1)a-(2b/d+N-1)mc_i+\sum_j mc_j}{[(2b+d(N-1))(2b/d-1)]}(p_i - mc_i) \tag{4.11}$$

The theoretical concept RPD is then defined as $(\pi(mc^{**}) - \pi(mc))/(\pi\ (mc^{*}) - \pi(mc))$ 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 $\boldsymbol{\varepsilon}$ decline. Boone (2008) proves that RPD is an increasing function of interaction among existing firms ($dRPD/dd>0$) and a decreasing function of entry costs ($dRPD/d\boldsymbol{\varepsilon}<0$). In other words, RPD increases when competition intensifies, ie 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 normalised by calculating its relative profit difference against the profits of the most and the least efficient firms. This procedure yields a normalised RPD curve as a function of normalised relative efficiency differences. The level of competition is then represented by the area under the normalised 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 (ie 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, ie $dln(\pi)/dln(mc)$, as the PE indicator (Boone et al., 2004 and 2007; Van Leuvensteijn et al., 2007; 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) conducted simulations for the PE indicator and found 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 the next section.

We note that the PE indicator model, like every other model, is a simplification of reality (Van Leuvensteijn et al., 2007). 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 3.1). 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 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 Boone indicator approach focuses on one important relationship (that between efficiency and profits), thereby disregarding other aspects (see also Bikker and Bos, 2005). 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. Naturally, annual estimates of the PE indicator are more likely to be impaired by these distortions than the estimates covering the full sample period. Hence, in addition to annual estimates, we provide point estimates for the full 1996-2008 period and the subsample periods 1996-2001 and 2002-2008, 2001 being the year of China's entry into the WTO.

4.3 Competition measures under interest rate regulation

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, eg 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 $mc_i(\boldsymbol{\varepsilon}) = mc_i - \boldsymbol{\varepsilon}$ ($i = 1,\ldots, N$). $\boldsymbol{\varepsilon}$ is a regulation parameter, which measures the extent to which deposit rate regulation is binding. $\boldsymbol{\varepsilon} \in (-\ddot{\varepsilon}, mc)$, where mc is the marginal cost of the most efficient bank and $\ddot{\varepsilon}$ is some positive number that allows the least efficient bank to remain profitable and stay in the market. A positive $\boldsymbol{\varepsilon}$ reflects a binding deposit rate ceiling, while a negative $\boldsymbol{\varepsilon}$ corresponds to a binding deposit rate floor. Higher values of $\boldsymbol{\varepsilon}$ 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 (4.8), (4.9) and (4.10), and imposing $\boldsymbol{\varepsilon}$, we derive the effect of binding deposit rate regulation on optimal output:

$$f(\varepsilon) = q_i(\varepsilon) - q_i = \frac{\left(\frac{2b}{d} - 1\right)\varepsilon}{\left(2b + d(N-1)\right)\left(\frac{2b}{d} - 1\right)} \quad (4.12)$$

where q_i is the optimal output without deposit rate regulation. Given $0<d \leq b$, $f(\boldsymbol{\varepsilon})$ is increasing in $\boldsymbol{\varepsilon}$ and takes the same sign as $\boldsymbol{\varepsilon}$. 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 $\boldsymbol{\varepsilon}$:

$$L_i(\varepsilon) = \frac{bq_i(\varepsilon)}{\left(bq_i(\varepsilon) + mc_i(\varepsilon)\right)} = \frac{b\left(q_i + f(\varepsilon)\right)}{\left(b\left(q_i + f(\varepsilon)\right) + mc_i - \varepsilon\right)} \quad (4.13)$$

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

$$sign\left(\frac{dL_i(\varepsilon)}{d\varepsilon}\right) = sign\left(b\left(q_i + \frac{f(\varepsilon)}{\varepsilon} mc_i\right)\right) > 0 \quad (4.14)$$

Hence, a higher value of $\boldsymbol{\varepsilon}$ increases an individual bank's Lerner index, indicating less competition, as theory would suggest. However, the aggregate Lerner index – ie 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 $s_i(\boldsymbol{\varepsilon}) = q_i(\boldsymbol{\varepsilon})/\Sigma_j q_j(\boldsymbol{\varepsilon})$, 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:

$$s_i(\varepsilon) = \frac{1}{N} \frac{\left(\frac{2b}{d} - 1\right)a - \left(\frac{2b}{d} + N - 1\right)mc_i + \Sigma_j mc_j + \left(\frac{2b}{d} - 1\right)\varepsilon}{\left(\frac{2b}{d} - 1\right)a - \left(\frac{2b}{d} + N - 1\right)mc_k + \Sigma_j mc_j + \left(\frac{2b}{d} - 1\right)\varepsilon} \quad (4.15)$$

Taking the derivative with respect to $\boldsymbol{\varepsilon}$ yields:

$$sign\left(\frac{ds_i}{d\varepsilon}\right) = sign\left(\left(\frac{2b}{d} - 1\right)\left(\frac{2b}{d} + N - 1\right)\left(mc_i - mc_k\right)\right) \quad (4.16)$$

It is immediately clear that the market share of bank *i* increases with a higher $\boldsymbol{\varepsilon}$ 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:

$$\frac{dL}{d\varepsilon} = \sum_{i=1}^{k} \frac{ds_i}{d\varepsilon} L_i + \sum_{i=k+1}^{N} \frac{ds_i}{d\varepsilon} L_i + \sum_{i=1}^{N} s_i \frac{dL_i}{d\varepsilon} \quad (4.17)$$

Denote banks i = 1,…, k as low-efficiency banks, which will see their market shares increase. In contrast, the market share of high-efficiency banks $i = k+1,…, N$ 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.[18]

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:

$$RPD(\varepsilon) = \frac{\left(q_A + f(\varepsilon)\right)^2 - \left(q_C + f(\varepsilon)\right)^2}{\left(q_B + f(\varepsilon)\right)^2 - \left(q_C + f(\varepsilon)\right)^2} \tag{4.18}$$

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

$$sign\left(\frac{dRPD}{d\varepsilon}\right) = sign\left(\frac{2f'(\varepsilon)(q_B - q_A)}{\left(q_B + q_c + 2f(\varepsilon)\right)^2}\right) \tag{4.19}$$
$$< 0$$

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 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.[19] 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* ($da > 0$),[20] 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

18. 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.

19. 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.

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

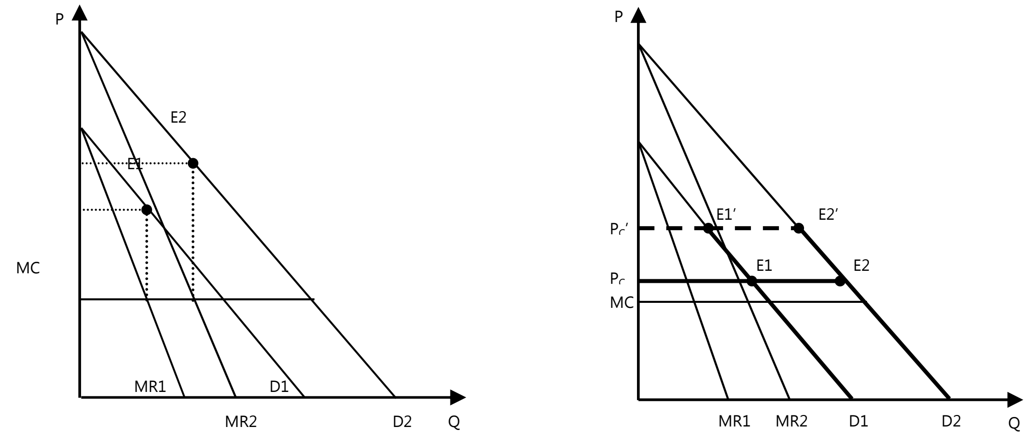
exogenous shifts in the demand curve can be correctly picked up by the Lerner index. However, this is not the case if the lending rate ceiling is binding, as demonstrated in Panel B. This ceiling (*P_c*) prevents the bank from choosing the optimal price-output combination according to *MR = MC*. In contrast, 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 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.

Lerner index and price ceiling — Figure 2

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 favour 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:

$$q_i^* = q_i + s_i\left(Q^* - Q\right) = s_i Q^*, \text{ where } s_i = \frac{1}{N}\frac{a-\left(N+1\right)mc_i + \Sigma_j mc_j}{a-\left(N+1\right)mc_k + \Sigma_j mc_j}. \tag{4.20}$$

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:

$$RPD\left(a\right) = \frac{\left(\bar{P} - mc_A\right)s_A - \left(\bar{P} - mc_C\right)S_c}{\left(\bar{P} - mc_B\right)s_B - \left(\bar{P} - mc_C\right)S_c} \tag{4.21}$$

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

$$sign\left(\frac{dRPD}{da}\right) = sign\left(\left(N+1\right)\left(mc_C - mc_B\right)\left(mc_A - mc_C\right)\left(mc_B - mc_A\right)\right) \tag{4.22}$$
$$< 0$$

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.

5 Data and sample

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 under consideration of data availability and to capture various banking sector reforms, including those related to WTO accession. 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). 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. First, in the 1990s, trust and investment corporations were important financial institutions that operated similarly to commercial banks, but with less restrictions and regulations (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. Second, 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. Third, there are significant data limitations for especially the large number of small banks that are excluded from the sample.

In order to exclude irrelevant and unreliable observations, banks are incorporated in our sample only if they fulfilled the condition that total assets, loans, deposits, equity and other non-interest income should be positive. We lost 43 observations after applying this criterion, mainly due to negative non-interest income (34 observations). 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[21] and 82 CCBs. Table 2 summarises the distribution of the observations. Table 3 gives a short description of the variables used in the 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.

21. Banks with more than 50% foreign ownership are classified as foreign banks. We only include foreign banks that provide separate balance sheet data for the People's Republic of China. 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.

Number and bank distribution of observations

Table 2

	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

Mean values of key variables by bank group

Table 3

Panel A

	Total costs as a share of total assets					Loans as a share of total assets					Securities as a share of total assets					Other services as a share of total income					Interest expenses as a share of total funding				
	SOCB	JSCB	CCB	FOREIGN	WHOLE	SOCB	JSCB	CCB	FOREIGN	WHOLE	SOCB	JSCB	CCB	FOREIGN	WHOLE	SOCB	JSCB	CCB	FOREIGN	WHOLE	SOCB	JSCB	CCB	FOREIGN	WHOLE
Average 1996-2001	6.57	4.4	3.99	4.05	4.52	60.28	51.19	48.28	47.77	50.91	12.38	15.29	21.78	6.32	16.06	3.96	5.92	13.43	7.46	8.32	5.6	3.08	2.94	4.85	3.75
Average 2002-2008	2.92	3.22	3.51	2.83	3.31	55.43	58.76	54.11	57.85	55.55	28.78	21.77	22.96	10.55	21.4	9.26	6.08	9.47	18.15	10.3	1.52	1.83	2.15	2.12	2.06
Average 1996-2008	4.61	3.71	3.57	3.16	3.6	57.66	55.61	53.38	55.12	54.46	21.21	19.08	22.81	9.81	20.22	7.26	6.01	9.98	15.4	9.85	3.4	2.35	2.25	2.78	2.45

Panel B

	Other expenses as a share of fixed assets					Average market share of lending[1]					Interest income as a share of total assets					Personnel expenses as a share of total assets[2]					Interest income as a share of total loans				
Average 1996-2001	33.8	69.12	58.18	108.77	68.04	22.01	1.07	0.16	0.02	3.57	6.95	5.23	4.4	6.14	5.41	0.5	0.4	0.63	0.57	0.52	11.64	10.44	9.54	16.36	11.47
Average 2002-2008	40.04	68.95	56.22	261.46	93.66	17.83	1.89	0.11	0.06	1.28	3.69	3.85	4.35	3.51	4.09	0.57	0.41	0.49	0.59	0.5	6.86	6.65	8.07	8.46	7.86
Average 1996-2008	37.16	69.02	56.47	224.22	87.68	19.76	1.55	0.12	0.05	1.82	5.19	4.43	4.36	4.14	4.39	0.54	0.41	0.51	0.59	0.5	9.07	8.22	8.26	10.36	8.69

Panel C

	Other non-earning assets to total assets ratio[3]					Funding mix[4]					Equity to total assets ratio					Other income to interest income ratio				
Average 1996-2001	6.87	4.83	6.91	4.03	5.55	92.03	91.14	90.41	97.14	91.75	4.54	5.64	5.78	36.43	12.17	2.04	6.77	17.36	7.30	9.32
Average 2002-2008	0.24	1.15	2.63	2.55	2.27	96.08	91.59	89.05	58.76	84.93	3.04	4.10	5.02	26.93	8.67	10.65	6.47	11.00	16.66	11.28
Average 1996-2008	3.23	2.68	3.18	2.95	3.04	94.30	91.41	89.22	64.96	86.36	3.72	4.74	5.11	29.50	9.49	6.77	6.60	11.80	14.41	10.83

All data are in percentages. "WHOLE" represents the figure for all banks in the sample.

1 This represents the average market share of total loans (by all banks) per individual SOCB, JSCB, CCB or foreign bank. 2 Personnel expenses to assets ratio serves as a proxy of wage rate. 3 Other non-earning assets to total assets ratio is defined as: (total assets minus loans minus other earning assets)/total assets. 4 The funding mix is defined as: customer deposits/(total funding minus deposits from banks).

Source: BankScope, authors' own calculations.

We divide the whole sample into two subsamples, ie the pre-WTO era (1996-2001) and post-WTO era (2002-2008). We choose 2001 as the break year because China's WTO accession at the end of 2001 could potentially have affected the operation of Chinese banks, as it was accompanied by important financial reforms, and therefore may fundamentally have changed the competitive structure of Chinese banking markets. On average, total costs as a share of total assets were significantly lower for 2002-2008 for all types of banks (but the most for SOCBs), which is mostly due to lower interest expenses driven by lower deposit rates in the post-WTO era.

The wage rate, which is proxied by the ratio of personnel expenses to total assets, did not change significantly. The ratio of other expenses to fixed assets stayed relatively stable for Chinese banks as well, while it increased significantly in the post-WTO era for foreign banks.

Turning to revenues, the ratio of interest income to total assets decreased on average significantly for all types of banks, which is most likely the result of lower lending rates in the post-WTO era and increasing diversification of business activities. The latter is reflected in the development of the ratio of other services to total income, which increased significantly for all types of banks except the CCBs. However, it should be noted that income from other services only represented around 10% of total income, suggesting that income predominantly was generated through lending activity.

To assess the relative importance of individual banks in Chinese loan markets, we have calculated the average market share of total loans (provided by all banks) per individual SOCB, JSCB, CCB and foreign bank. SOCBs had by far the largest individual share of total lending at around 20% of the full sample period, while those of individual CCBs and foreign banks were below 1%. The average market share of individual SOCBs dropped by nearly 5% in the post-WTO era compared with the pre-WTO era, which was the result of the increasing entry of new banks, as more CCBs were established and more foreign banks were allowed to start and expand business in China. Notwithstanding these new entries, individual JSCBs increased on average their market shares in loan markets post-WTO, reflecting their growing importance in lending activity. The average market share of total lending by individual CCBs and foreign banks decreased, which mainly reflects the fact that the number of these banks increased significantly in the post-WTO era.

The ratio of other non-earning assets to total assets decreased on average significantly for all types of banks, possibly reflecting improving profitability. The funding mix, defined as the ratio of customer deposits to total funding (excluding bank deposits), stayed relatively stable for all bank types except foreign banks. The equity to assets ratio generally decreased mildly. The other income to interest income ratio increased significantly for SOCBs and foreign banks, reflecting the increasing importance of non-traditional banking activities for these groups.

As in other empirical investigations of Chinese banking markets, the most troublesome data are those on wages. Ideally, the wage rate is the ratio of personnel expenses to the number of staff. However, many banks do not provide information on the number of their staff members, and for some banks personnel expenses data are missing as well. Therefore, we need to find an appropriate proxy for wages. 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; Rezvanian and

Mehdian, 2002; Vander Vennet, 2002). This approach might not be appropriate for our sample, as very few CCBs report the number of employees, so its 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 have complete data on total assets; so, to generate this proxy, we only need to have relatively complete data on personnel expenses. We adopted the following procedure to approximate missing data. 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.

6 Results[22]

6.1 Empirical results: Lerner and elasticity-adjusted Lerner indices

The estimation results for the cost equation (4.2) and supply equation (4.7) that form the basis for estimating the elasticity-adjusted Lerner index are reported in Appendix B, Table B.1. We summarise the results for the traditional Lerner index and elasticity-adjusted Lerner index in Table 4.

The results for both the traditional Lerner index and the elasticity-adjusted Lerner index suggest a general increasing level of bank competition up to around 2002 and a decreasing level of bank competition afterwards. 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 and one for all years, rejecting the null hypothesis that Chinese loan markets are in a state of either perfect competition or monopoly.

Turning to subsample estimates, the estimated values of the elasticity-adjusted Lerner index are significantly different from zero and one for each subsample as well. The results for both the traditional and elasticity-adjusted Lerner indices suggest that bank competition in the post-WTO period was lower than in the pre-WTO period, with the lowest level of competition registered for both indices in 2007. To test this conclusion formally, we conducted Chi-squared distributed Wald tests with one degree of freedom to determine whether the results of the elasticity-adjusted Lerner index are significantly different across the subsamples. Test statistics confirm that the elasticity-adjusted Lerner index for the pre-WTO period is lower than that for the post-WTO period, indicating strong evidence that competition worsened after WTO accession.

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

	λ_t	Average loan rate	Average loan deposit spread	Elasticity-adjusted Lerner index	Lerner index	Mc_e	MC_l
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	0.050	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	0.020	0.068	0.047	0.288	0.340	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	0.380	0.052	0.048
2007	0.036	0.079	0.052	0.452	0.434	0.050	0.046
2008	0.041	0.094	0.061	0.439	0.402	0.060	0.055
1996–2001	0.029	0.115	0.059	0.249	0.320	0.071	0.066
2002–2008	0.027	0.079	0.052	0.342	0.375	0.051	0.047

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

λ_t are statistically different from zero for all years at 1% significance level; *prewto* represents 1996-2001 period; *postwto* represents 2002–2008 period;

MC_e and MC_l are average marginal costs used to calculate elasticity-adjusted Lerner index and traditional Lerner index, respectively.

22. The results for the Panzar-Rosse H-statistic are reported in Appendix A.

Our results are reinforced by Soedarmono et al. (2013) and Fungáčová et al. (2012),[23] which also document a general decreasing trend of bank competition in China during 2002-2008. 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. (2009b) found an average Lerner index of 0.22 for 23 industrial countries calculated over the period 1999-2005, while Carbó Valverde et al. (2009) obtained 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.

Therefore, regardless of whether the elasticity-adjusted Lerner index or traditional Lerner index is adopted as the indicator to measure competition, we conclude that Chinese loan markets were relatively competitive in the pre-WTO period, while competitive conditions worsened later.

6.2 Empirical results: Price 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. More specifically, we estimate one separate TCF for the SOCBs, JSCBs, CCBs and the foreign banks, which should improve the accuracy of the estimation of marginal costs. The estimation of the TCF is reported in Appendix C, Table C.1. 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:

$$\ln \pi_{ilt} = \alpha + \Sigma_{t=1,\dots,(T-1)} \gamma_t d_t + \Sigma_{t=1,\dots,T} \beta_t d_t \ln mc_{ilt} + u_{ilt} \tag{6.1}$$

where π_{ilt} stands for profits, 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 term. This provides us with the coefficient $\boldsymbol{\beta}_t$, ie the PE indicator (as is given by PE = $d(\ln\pi_i)/d(\ln mc_i)$). $\boldsymbol{\beta}_t$ is negative in theory, reflecting that higher marginal costs reduce profits for all banks.[24] Moreover, the more competitive the market is, the lower the value of $\boldsymbol{\beta}_t$. In other words, banks are punished more harshly for being inefficient in more competitive markets. Note that the indicator $\boldsymbol{\beta}_t$ is time-dependent.

Profits are defined as:

$$\pi_{ilt} = x_{ilt}\left(p_{ilt} - mc_{ilt}\right) \tag{6.2}$$

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

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23. Soedarmono et al. (2013) used the elasticity-adjusted Lerner index, while Fungáčová et al. (2012) estimated the traditional Lerner index.

24. In practice, a positive β_t is possible (Van Leuvensteijn et al., 2007), which could be the result of extreme collusion, market regulation or banks competing on quality (Tabak et al., 2012).

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.[25] 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 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.[26]

To assess the evolution of bank competition, we first estimate the yearly PE indicator, which is based on equation (6.1). Table 5 reports the results. 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. 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.

25. The Stock-Yogo weak ID test critical value at 10% (maximal LIML size) is 16.38 (Stock and Yogo, 2005).
26. Results with 2SLS are available upon request.

Yearly PE Indicator

Dependent variable: ln(Profits) Table 5

	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	13.10
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		

z-values in parenthesis; ** represents significance level of 5%, *** represent significance level of 1%. 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-Yogo (2002) and (2005) 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.

The subsamples are defined in the same way as in the previous sections, eg pre-WTO (1996-2001) and post-WTO (2002-2008). We estimate one PE indicator for each subsample and test whether competition changed significantly after WTO accession. 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. A point estimate of the PE indicator for the whole sample 1996-2008 is provided as well. Estimations are based on the following equation:

$$\ln \pi_{ilt} = \alpha + \gamma Trend + \beta \ln mc_{ilt} + u_{ilt} \tag{6.3}$$

where *Trend* is a time trend.[27]

Table 6 reports the results for the subsample estimations. The Kleibergen-Paap rk LM statistics are significant at the 1% level for the whole sample and each subsample, rejecting the null hypothesis that the model is unidentified. The Kleibergen-Paap Wald rk F-statistic for each sample is larger than 16.38, 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. To test whether competitive conditions in Chinese loan markets experienced significant structural change after WTO accession, we performed a Chi-squared distributed Wald test with one degree of freedom to determine whether the PE indicators are significantly different across various subsamples. The rejection of the null hypothesis (H0: pre-WTO PE indicator <= post-WTO PE indicator) indicates that the level of bank competition was significantly higher in the post-WTO period.

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

Point estimates PE indicator: Whole sample and subsamples

Dependent variable: ln(Profits) Table 6

	1996–2008	1996–2001	2002–2008
PE Indicator	−2.388***	−1.514	−3.570***
	(–5.78)	(–1.43)	(–7.74)
Time Trend	−0.0332	−0.519**	0.345***
	(–0.82)	(–2.37)	–4.71
Constant	−0.24	4.966**	−8.050***
	(–0.19)	(2.18)	(–4.51)
H_0:prewto -postwto <=0(p-value)		3.61** (0.0288)	
Nr. Obs	457	87	370
F	16.78	2.97	33.67
Centered R^2	0.089	0.141	0.18
Kleibergen-Paap rk Wald F	211.4	30.98	130.8
Kleibergen-Paap rk LM (p-value)	62.00(0.0000)	13.94(0.0001)	44.22(0.0000)

z-values in parenthesis; ** represents significance level of 5%, *** represents significance level of 1%. Since there is only one endogenous variable, we use Kleibergen-Paap rk Wald F and Kleibergen-Paap rk 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.

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. These robustness tests are presented in Appendix E, for an alternative definition of the PE indicator (E.4) and for an alternative calculation of marginal costs (E.5). The latter analysis shows that the slightly different calculation of marginal costs for the (elasticity-adjusted) Lerner index and the PE indicator does not drive the divergence between their results.

6.3 Comparison of the various empirical measures

The results for the various measures that we presented in the previous sections revealed significant differences in the evolution of competition and its level, for both the yearly and subsample estimates. The fact that different competition measures yield inconsistent results for the same banking market and country is well documented in the empirical banking literature, as discussed in Sections 1 and 2. At the same time, this finding may be especially relevant for relatively regulated markets such as Chinese loan markets. Hence, following Carbó Valverde et al. (2009), we formally test for the consistence of different competition measures by calculating pair-wise correlation coefficients. Since all competition measures except the H-statistic imply higher competition with a lower value, we multiply the H-statistic results by –1 to make comparison easier, so that now a higher value implies lower competition for all measures. The results of the pair-wise correlations are reported in Table 7.

Pair-wise correlation coefficients				Table 7
	PE	H	Lerner	Elasticity-adjusted Lerner
PE	1			
H	–0.1884*	1		
Lerner	–0.0413	0.4154*	1	
Elasticity-adjusted Lerner	0.1433*	0.3522*	0.9321*	1
TIME	–0.4152*	0.2779*	0.7602*	0.5869*

* represents significance level of 1%.

Testing the pair-wise correlations of these measures of competition first reveals that the PE indicator is negatively correlated with the H-statistic and the traditional Lerner index (the latter not significant at 1%). This finding confirms that the former indicator yields diametrically opposed results to those from the latter two traditional measures. At the same time, the PE indicator is positively correlated with the elasticity-adjusted Lerner index. Second, in order to test whether the competition measures produce similar conclusions on the evolution of competition over time, we provide pair-wise correlation coefficients with time. A negative (positive) value indicates improved (worsened) competition over time. The PE indicator suggests improving competition across the sample years, while the other measures suggest the opposite, a result that is consistent with our results in the previous sections.

6.4 Interpretation results PE indicator

As we have argued above, the PE indicator is the most appropriate measure to assess competitive conditions in Chinese loan markets. This section provides an in-depth analysis of its results. Overall, we are generally well able to explain the specific development of this indicator over time, which strengthens our belief that it is superior to the Panzar-Rosse and Lerner approaches in the context of China. In fact, we find it rather difficult to offer plausible explanations for the results obtained with the latter two measures.

The key to understanding why the results for the PE indicator are so different from those of more conventional competition indicators like the Panzar-Rosse H-statistic and the (elasticity-adjusted) Lerner index lies in the system of interest rate regulation in China. If interest regulation is binding, it can substantially bias the traditional measures but not RPD. We have proved this theoretically for the Lerner index in Section 4.3. In the case of the H-statistic, we discuss this bias in Appendix A (A.3).

Whether and to what extent interest 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 is considered to be non-binding in practice (He and Wang, 2012; see Section 3.1); b) the deposit rate floor and ceiling are binding (Feyzioğlu, 2009; Ma et al., 2011; He and Wang, 2012;[28] PBC, 2009; Yi, 2009); and c) the lending rate ceiling was most likely binding during the pre-WTO period (Yi, 2009). In this section, we argue that the bias in the H-statistic and Lerner index predominantly results

28. 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."

from binding interest rate regulation. The interest rate regulation discussed here is mainly the ceilings on deposit and lending rates, but similar analysis can be extended to the floors applied to these rates.

Turning to the specific results we obtained in 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 Section 6.2, Table 5). We are encouraged by this result, as one would expect that, during the years when Chinese banking markets were still heavily regulated, more efficient banks would not be more profitable, with competition being suppressed. In other words, there was no reward for being more competitive than one's competitors. Actually, this finding reminds us of the results for Japanese loan markets during the 1990s in Van Leuvensteijn et al. (2007 and 2011), when the PE indicator was positive (and significant). This could be related to the regulated "convoy system" in Japan where market shares were more or less guaranteed 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 efficient banks. The PE indicator improved further until 2003, when it reached its lowest value of –6.3 (eg highest level of competition). From an international perspective, this value is comparable to the most competitive yearly results we obtained for several mature economies (Van Leuvensteijn et al., 2007 and 2011).

Then, after 2003, we find a gradual decline in competition in Chinese loan markets (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 re-regulation may be responsible for this pattern of slightly declining competition. In general, there is evidence that, for both mature and emerging market economies, financial deregulation has often been intertwined with concomitant prudential re-regulation (Zhao et al., 2010). This seems also to have occurred to a certain extent in China. In 2004, the 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 PBC, 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. As formulated by Fukumoto et al. (2010, p. 3): "The newly introduced credit limits were similar to the credit plan which had existed until 1998 in the sense that both of these measures set rigorous constraints on the growth of bank lending. The growth of bank lending started to slow down once credit limits were implemented." These lending restrictions were kept in place until the fall of 2008 and can be characterised as a major step of re-regulation, as they re-instated elements of the old credit plan system. It may be regarded that this policy move had a detrimental impact on bank competition in China, as it frustrated commercially oriented lending decisions and disincentivised competition.

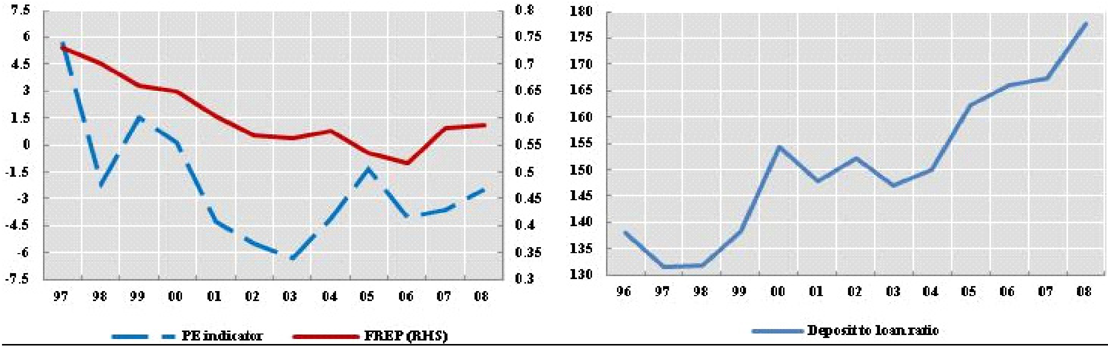
The element of re-regulation is picked up nicely by the financial repression index developed for China in Huang and Wang (2011) (see Figure 3, left-hand panel). It is based on six financial repression variables, including two interest rates, two loan market share variables, reserve requirements and capital 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 are depicted for illustrative purposes in Figure 3 (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 3, 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.

Interpretation results PE indicator

Figure 3

Panel A: PE indicator and FREP

Panel B: Deposit-to-loan ratio (in %)



FREP = Financial repression index.

Sources: Authors' own calculations. We are grateful to Yiping Huang and Xun Wang for sharing the values of FREP index with us.

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 liberalisation indices. To this end, we employ two additional indicators of financial reform: the overall financial liberalisation and interest rate liberalisation indices developed by Abiad et al. (2010). Their values are shown in Appendix D, Table D.1. The former index measures the overall degree of financial liberalisation, with values ranging from 0 to 1, with a higher value indicating a more liberalised financial system. The latter index, which takes the values 0, 1, 2 or 3, indicates fully repressed, partially repressed, partially liberalised and fully liberalised 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 four measures of competition that we used in Section 6.3.

Again, we multiply the H-statistic results by –1 to make comparison of the correlations easier. The results are reported in Table 8. Should financial reform promote competition, one would expect positive correlations between the financial repression index and the competition measures. This is because both the index and the measures show improved conditions with lower values and vice versa. In contrast, one would expect negative correlations between the two other financial liberalisation indices and the competition measures if financial reform promotes competition. Namely, the two indices indicate a more liberalised financial system with higher values, while the indicators of competition suggest more competition with lower values. Since financial reform may affect banking behaviour with a time lag, we use one-period lagged values of the three indices.[29]

Pair-wise correlation coefficients with financial reform indices — Table 8

	PE	H	Lerner	Elasticity-adjusted Lerner
FREP	0.6560*	–0.2159*	–0.5794*	–0.3104*
Fin_Lib Index	–0.4223*	0.2689*	0.5015*	–0.055
Int_Lib Index	–0.2206*	0.4175*	0.7917*	0.3285*

* represents significance level of 1%. 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. The values of the indices are reported in Appendix D, Table D.1.

In case a more liberalised 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 (ie the H-statistic and the Lerner indices) that are significant all have the opposite sign, suggesting that increased liberalisation is associated with weaker competition.

Although we did not formally test the impact of financial reform on competition in Chinese loan markets, the pair-wise correlation coefficients that we find for the PE indicator, all associating more reform with more competition, tend to bolster our confidence in it. The empirical literature on the relationship between financial reform and competition is not always conclusive, but generally its results for emerging market economies have shown a beneficial link between the two (see Section 2). That we find this for China is in our view encouraging.

Finally, a number of structural developments in the Chinese financial system should support increasing competition in Chinese banking. First, the market share of the four SOCBs in bank lending has declined steadily over the past decade. This should have contributed to improving competition. Second, China has experienced a steady rise of the shadow banking system, which should have increased competition in bank loan markets, as banks have to compete for a smaller share of total credit intermediation in China. Third, Chinese banks have introduced personal incentive-driven remuneration packages, which link the remuneration of individual loan officers much more to their actual performance. This should have contributed to competition in Chinese loan markets as well.

29. 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.

7 Conclusions

This paper investigates the evolution of competition in Chinese loan markets. We believe that this investigation makes sense, as after 30 years of financial reform, China's banks have begun to behave more like commercial banks in mature economies (Firth et al., 2009; Herd et al., 2010; IMF, 2011). The impact of financial reform on bank competition has been investigated extensively for many countries, and for most mature and emerging economies empirical studies suggest that it promoted bank competition.

However, some earlier studies that adopted conventional approaches to measure competition concluded that bank competition in China declined during the past decade, despite these reforms. In this paper, we compare the results obtained from conventional indicators such as the Lerner index and Panzar-Rosse H-statistic with those estimated using the relatively new Profit Elasticity (PE) approach. We argue that traditional measures of competition fail to measure competition in the Chinese banking sector properly, and we provide arguments – both theoretically and empirically – to support this.

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 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 characterised 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 and the H-statistic. We doubt these findings, as they may be distorted by various factors, including restrictions on market shares and interest rates.

This study yields two major insights. First, the theoretical foundation of the PE indicator, which is the RPD model, is not biased due to interest rate regulation. This makes the PE indicator a much better measure to gauge competition in Chinese loan markets 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. Second, applying this unbiased competition indicator to Chinese loan markets shows that financial reform indeed has contributed to significant improvements in competition. This result is much in line with those obtained for other emerging economies. Again, we find contradictory results for the conventional measures. Moreover, our results for both the PE indicator and the 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. It may also provide an answer to questions raised in other research on this issue. For example, Fungáčová et al. (2012) use the Lerner index and find results similar to ours for the same index, indicating that bank competition in China declined over time. They go on to note that "at first glance, it is somewhat remarkable that China's accession to WTO has not led to greater competition in the banking industry". As we have argued in this paper, this may not be that remarkable after all. It may be the case that competition in the Chinese banking system has not been assessed with the proper method. Of course, further work is definitely needed to substantiate this claim. For example, it would be interesting to see how the results that we report here compare with those for other banking systems of comparable structure and stage of development to China's. Moreover, given the strong differences in regional economic and financial development across China, the PE indicator and conventional measures could be estimated for the regional banking market in China. For those that de facto operate under more liberalised conditions, the differences between the PE indicator and other measures may be much smaller, for example when compared with those for much less developed and de facto more regulated regions.

More generally, our findings for China 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 interest rate and other regulations. Regarding the former, empirical work to assess whether, and to what extent, interest rate regulation is binding is crucial to validating whether conventional measures of competition may obtain unbiased results.

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Appendix: Measuring bank competition in China: a comparison of new versus conventional approaches applied to loan markets

Appendix A. Panzar-Rosse H-statistic

A.1. 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).[30] 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.

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

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

The dependent variable $\ln(II_{it} / TA_{it})$ is the logarithm of the ratio of interest income to total assets.[31] 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 $H=\beta+\gamma+\delta$.

30. 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.

31. 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.1).

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.

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 (A.1):

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

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.

A.2. EMPIRICAL RESULTS PANZAR-ROSSE H-STATISTIC

Estimations are carried out with recursive least squares.[32] 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 estimations (Bikker and Spierdijk, 2008). We estimate Equation (A.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 A.1, Panel A. To ensure standard errors and statistics that are robust to arbitrary heteroskedasticity and autocorrelation, kernel-

32. 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.2). We also tested 3-year rolling-window regressions and found similar results to recursive least squares. Results are available upon request.

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 A.1. 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.[33]

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.[34] The results for the H-statistic are reported in Table A.2, while the long-run market equilibrium condition tests for the whole sample and sub-samples are reported in Table A.3. 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 subsamples for any conventional significance level, suggesting no significant structural change. Table A.3 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.[35]

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.[36]

A.3. BIAS PANZAR-ROSSE H-STATISTIC DUE TO INTEREST RATE REGULATION

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

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33. Table A.2 reports the coefficients of the various control variables. The positive sign for *LNS_TA* and the negative signs for *ONEA_TA* and *OI_II* confirm our prior expectations.

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

35. 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).

36. We demonstrate in Appendix E (E.1) by using the unscaled version of the H-statistic, which is theoretically more sound than the scaled version (see Bikker et al., 2007, and Bikker et al., 2012), that the market structure for the pre-WTO period featured perfect competition and that for the post-WTO period monopolistic competition.

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 A.2, 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.

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

Table A.1

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.

H-statistic point estimates: Whole sample and subsamples						Table A.2
	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%

Long-run equilibrium condition: Whole sample and subsamples

Dependent variable: lnROA

Table A.3

	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	

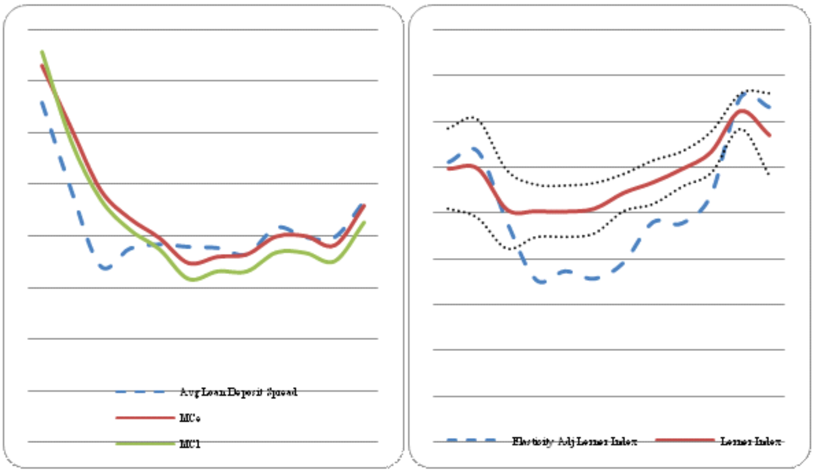
z-values in parenthesis; * represents significance level of 10%, ** represents significance level of 5%, *** represents significance level of 1%.

Appendix B. Underlying estimations 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)
λ_{1997}	0.0616***	(9.17)		
λ_{1998}	0.0449***	(7.89)		
λ_{1999}	0.0294***	(5.89)		
λ_{2000}	0.0208***	(4.84)		
λ_{2001}	0.0191***	(4.66)		
λ_{2002}	0.0151***	(4.19)		
λ_{2003}	0.0167***	(4.84)		
λ_{2004}	0.0196***	(5.96)		
λ_{2005}	0.0227***	(7.39)		
λ_{2006}	0.0255***	(9.18)		
λ_{2007}	0.0359***	(13.99)		
λ_{2008}	0.0415***	(13.18)		
$\lambda_{1996-2001}$			0.0287***	(9.38)
$\lambda_{2002-2008}$			0.0269***	(14.12)
Number of observations:	453		453	

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

Chart B.1: Marginal costs and traditional and elasticity-adjusted Lerner indices



MCe and MCl are the average marginal costs used in the calculation of the elasticity-adjusted Lerner and traditional Lerner indices, respectively. The dotted line in the right panel represents 95% confidence interval of traditional Lerner index.

Appendix C. 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:

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

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 (C.2) holds for each bank type h:

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

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:

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

The term $\partial \ln c^h_{it} / \partial \ln x_{ilt}$ is the first derivative of Equation (C.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:

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

Estimate translog cost function by bank type

Table C.1

	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	
$Adj\text{-}R^2$	0.9997		0.9998		0.9990		0.9987	

Appendix D. Financial reform indices

Financial reform indices — Table D.1

	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.

Appendix E. 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 in Appendix E whether the main results are sensitive to: 1) unscaled version of the Panzar-Rosse H-statistic; 2) parametric approach of Panzar-Rosse; 3) alternative Lerner index (conjectural variation); 4) alternative definition of PE indicator; 5) calculation marginal costs.

E.1. UNSCALED PANZAR-ROSSE H-STATISTIC

In our estimation of the Panzar-Rosse H-statistic (Appendix A), we used the scaled approach, ie 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.[37] 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 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.

E.2. PARAMETRIC APPROACH OF PANZAR-ROSSE

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:

$$\begin{aligned}\ln\left(II_{it}/TA_{it}\right) &= \alpha + \left(\beta\ln\left(AFR_{it}\right) + \gamma\ln\left(PPE_{it}\right) + \delta\ln\left(PCE_{it}\right)\right)^{*}\exp\left(\zeta^{*}t\right) + \eta_1\ln\left(LNS\_TA_{it}\right)\\ &+\eta_2\ln\left(ONEA\_TA_{it}\right) + \eta_3\ln\left(DPS\_F_{it}\right) + \eta_4\ln\left(EQ\_TA_{it}\right) + \eta_5 OI\_II_{it} + \Sigma_{h=1\ldots H-1}\varsigma_h d_i^h + error_{it}\end{aligned} \quad \text{(E.2.1)}$$

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

37 Results are available upon request.

(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.

E.3. ALTERNATIVE LERNER INDEX (CONJECTURAL VARIATION)

In Section 6.1 we calculated the elasticity-adjusted Lerner index L by first estimating λ, ie 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 4.2), the supply equation (Equation 4.6) 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.

E.4. 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.[38] 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 E.1 respectively E.2. 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.

38. 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.

E.5. 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 4.2). 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), ie 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 TCF. We calculate again yearly and subsample values, which are shown in Table E.3. The modification in the TCF turns out to change the elasticity-adjusted Lerner index only very marginally for both the yearly and subsample estimations[39]. 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 E.1 for the yearly results and of Table E.2 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.

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

Yearly estimates of alternative PE indicators

Table E.1

	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.

Whole sample and subsample estimates of alternative PE indicators

Table E.2

	Panel A: Independent variable $\ln(MC_e)$			Panel B: Dependent variable: $\ln(\mathit{variable\ profits})$		
	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(0.00)	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

Lerner indices assuming different TCFs for each bank group

Table E.3

	λ_t	Elasticity adjusted Lerner index	Lerner index	MC_e	MC_l
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_l are average marginal cost estimated from elasticity-adjusted Lerner index and traditional Lerner index, respectively.

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