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For additional information, please contact Irene Li, Department of Accountancy, City University of Hong Kong, Tat Chee Avenue, Kowloon Tong, Hong Kong. Telephone: +852 3442 7932. Fax: +852 3442 0349. E-mail: acwoo@cityu.edu.hk.

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Can media exposure improve stock price efficiency in China and why?

Jeong-Bon Kim [a,*], Zhongbo Yu [b], Hao Zhang [c]

[a] *School of Accounting and Finance, University of Waterloo, Canada*
[b] *Shenzhen Stock Exchange Research Institute, China*
[c] *College of Business, City University of Hong Kong, Hong Kong, China*

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ABSTRACT

The media in China has undergone extensive commercialization to become more market-driven over the last 35 years. Based on a sample of over two million newspaper articles, this study investigates whether the media in China has an incremental impact on stock price efficiency. We find that: as media coverage of a firm increases, (1) its stock price synchronicity decreases; (2) the probability of informed trading of its stock increases; and (3) the extent to which its stock price deviates from random walk decreases. Our inter-regional analysis over thirty-one provinces/regions within China reveals that the effects of the media on decreasing stock price synchronicity, increasing the probability of informed trading, and reducing stock price deviation from random walk are stronger in regions of weaker institutional development. Our findings suggest that a market-driven media can play the role of compensating for the underdeveloped governance institutions in transitional economies such as China.

1. Introduction

In transitional economies such as China or Russia, agency problems are severe but the standard governance institutions to protect investors and maintain transparency are limited and weak (e.g., Allen et al., 2005). With investor protection weakened and transparency reduced, stock prices are likely to become less informative because they are less able to incorporate firm-specific information (Jin and Myers, 2006). The resulting

* Corresponding author. Tel.: +1 519 888 4567x38013.
E-mail address: jb6kim@uwaterloo.ca (J.-B. Kim).

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equilibrium is that stock prices are less efficient because they are less able to channel scarce capital into best investments (Tobin, 1982).

In such an institutional environment, can the media play the role of compensating for the weak standard institutions to improve stock price efficiency? Our study investigates this hitherto unexplored question in the context of listed Chinese firms. Our analysis focuses on the impact of firm-level media coverage on three dimensions of stock price efficiency: (1) the ability of stock price to incorporate firm-specific information; (2) the collection of private firm-specific information in stock trading (i.e., the extent of informed trading); and (3) the extent to which stock prices deviate from random walk (i.e., pricing error).[1] In an inter-regional analysis across 31 regional/provincial jurisdictions within China, we also directly test our compensating hypothesis that the effect of media coverage in increasing stock price efficiency is stronger in regions (or countries) with weaker standard governance institutions.

We are motivated to study the effect of firm-level media coverage in China on stock pricing efficiency for at least two reasons. First, La Porta et al. (1998) suggest that in countries where standard institutions are weak, substitutive or compensating mechanisms (such as concentrated ownership) may develop to sustain financial activities.[2] As the largest transitional economy in which standard institutions for investor protection and disclosure transparency are either underdeveloped or still in the developing stage, China provides arguably the best laboratory setting in which to identify compensating mechanisms.[3] Our investigation in effect assesses whether the media can act as a compensating mechanism to increase stock price efficiency.

Second, Morck et al. (2000) find that Chinese stock prices are less able to incorporate firm-specific information than most other stocks in a 40-country sample.[4] Jin and Myers (2006) as well as Hutton et al. (2009) show that the ability of stock prices to incorporate firm-specific information is lower in countries with poor investor protection and low firm-level transparency. In the case of China and other transitional economies, these two related factors stem from the underdevelopment of standard governance institutions. This institutional underdevelopment results in weak legal enforcement, lax financial regulations, poor disclosure quality, insufficient information intermediation, poor external monitoring by financial analysts, independent auditors, institutional investors, or credit rating agencies. Since the development of standard governance institutions is necessarily a gradual process, the potential compensating role of the well-established and increasingly market-driven institution of the Chinese media warrants research attention.[5]

Our empirical strategy involves developing two measures of firm-specific media coverage based on over two millions newspaper articles written about all listed firms in China and published in 14 national financial newspapers and the 10 largest comprehensive newspapers in China in the period of 2000–2009 and identifying three empirical proxies for stock price efficiency, which are the inverse of stock price synchronicity, the probability of informed trading (PIN), and stock pricing error. Our results reveal the following.

First, we find that as the media coverage of a firm increases, the ability of the firm's stock to incorporate firm-specific information improves. Second, we find that as the media coverage of a firm increases, both the probability of informed trading and the probability that a private information event occurs in its stock increase. Third, we find that as the media coverage of a firm increases, the extent to which the firm's stock

[1] To the best of our knowledge, this study is the first to examine the relation between firm-level media coverage (in China or elsewhere) and stock price efficiency. There is, however, a growing literature which looks at the relation between non-media institutions and stock price efficiency (e.g., Kim and Shi, 2012; Haw et al., 2012; Ferreira et al., 2011; Gul et al., 2010; Boehmer and Kelley, 2009; Fernandes and Ferreira, 2009; Ferreira and Laux, 2007; Jin and Myers, 2006; Chan and Hameed, 2006; Morck et al., 2000).

[2] La Porta et al. (1998) use both the words "substitute" (p. 1116) and "compensate" (p. 1140) to describe these mechanisms relative to "complement" mechanisms. We prefer the word "compensate" because these mechanisms cannot *completely* substitute for legal institutions, but they can "compensate, at least in part" (p. 1140).

[3] China ranks poorly in the ranking of countries in terms of rule of law, corruption, and the protection of property rights (La Porta et al., 2004; Pistor and Xu, 2005; Allen et al., 2005; Jiang et al., 2010).

[4] Given that China is the second-largest and fastest-growing economy and that it attracts the largest amount of foreign capital, the *inability* of Chinese stocks to incorporate firm-specific information into their prices is an important concern, particularly from the perspective of foreign equity investors. For example, the fraction of U.S. pension money invested in emerging markets with underdeveloped legal systems such as China is growing rapidly (Dyck et al., 2008, p. 1094).

[5] See next section and Appendix C for detailed arguments.

price deviates from random walk (i.e., stock pricing error) decreases. These findings, taken together, suggest that firm-level media coverage can increase stock price efficiency.

Fourth, to alleviate concerns about potential endogeneity, we apply the *propensity score matching* (PSM) procedure (Rosenbaum and Rubin, 1984). We find that our results from the PSM analysis remain qualitatively identical with our main results using the full sample. Fifth, to alleviate concerns about the potential impact of liquidity on our results, we conduct a further sub-sample analysis by dividing the main sample into sub-samples of liquid and illiquid firms. We find no evidence that variations in liquidity drive our results.

Finally, we test our compensating hypothesis directly by investigating whether the observed effect of media coverage on stock price efficiency differs systematically across 31 different provincial/regional jurisdictions of differing institutional development within China. Our results show that: (1) stocks of firms located in regions of weaker institutional development are less price efficient as expected; and (2) the effect of media coverage in increasing stock price efficiency is stronger (weaker) in regions within China with weaker (stronger) institutions. These results are in support of the hypothesis that institution of the media is able to play the role of compensating (or substituting) for the underdeveloped standard institutions normally expected to provide investor protection, maintain transparency, and thus increase stock price efficiency.

Our study adds to the extant literature in two ways. First, this paper contributes to the stock price efficiency literature by identifying the effect of a hitherto unexamined factor, namely firm-level media coverage, on the efficiency of stock prices. In so doing, we provide useful insights into the channels through which the media can impact the efficiency of stock markets. Second, our paper contributes to the investor protection literature by showing that the media can play the compensating role of protecting investors and maintaining transparency in transitional economies such as China where standard institutions normally expected to perform these functions are relatively weak and underdeveloped.

The remainder of this paper is organized as follows. The next section provides a background discussion of the media in China. Section 2 develops our empirical predictions. Section 3 describes the data and measurement of key variables. Section 4 presents the empirical analysis and results. Section 5 sets forth our conclusions.

2. Empirical predictions

2.1. Does media coverage enhance stock price efficiency?

Research on the role of media coverage in financial economics is relatively recent. There are two strands of the literature relevant to this paper. The first deals with how media coverage affects stock prices. Klibanoff et al. (1998) find that country-specific news reported on the front page of *The New York Times* makes the price of closed-end country funds moving to fundamentals more closely and quickly. Tetlock et al. (2008) show that firm-specific qualitative information in news articles can help predict earnings and stock prices. Fang and Peress (2009) find that investors demand a premium in stocks with no media coverage: They attribute this finding to a lower level of "investor recognition" (Merton, 1987) of stocks with low media coverage due to their lower level of relative information completeness.[6] Hillert et al. (2012) find that firms with media coverage exhibit stronger momentum.

The second strand of the literature deals with the watchdog function of the media. The question typically asked is whether more "sunshine," resulting from more media coverage translates into more investor protection. In a sample of U.S. firms whose accounting practices were challenged and sanctioned by the Securities and Exchange Commission (SEC), Miller (2006) finds that media coverage played a role in the early identification of accounting fraud. Dyck et al. (2008) show that coverage in the Anglo-American press increases the probability that a corporate governance violation is reversed in Russian firms. Dyck et al. (2010) show that the media was responsible for "blowing the whistle" on between 17% and 24% of corporate frauds in the U.S. between 1996 and 2004. Two important and related implications arise from these studies. First, more firm-specific "sunshine" resulting from more firm-level media coverage can improve investor protection

[6] In other words, media coverage of a firm increases the relative "investor recognition" of its stock because it increases the stock's relative information completeness.

through facilitating law/regulatory enforcements (the enforcement effect) and/or creating public pressure on the firms covered (the reputation effect). Second, better investor protection improves the information environment of the firm by improving governance. With better governance, firms become more transparent because they have less to hide (Leuz et al., 2003).

Combining these two strands of research and extending them, we conjecture that stocks of firms with more media coverage are more price efficient because there is more firm-specific information available in the market about these stocks. The underlining rationale is that these firms are more transparent and investors of these firms enjoy better investor protection (Jin and Myers, 2006). First, firm-level media coverage generates more firm-relevant *public* information. The resultant enforcement and reputation effects improve firm-level governance and thus firm-level transparency accordingly (Leuz et al., 2003). Second, media coverage provides more investor protection through the watchdog function of the media. Better investor protection is critically responsible for encouraging investors to collect more firm-specific *private* information before they trade on a stock. With more information and higher transparency available, any deviation of a transaction stock price from the efficient price (i.e. stock pricing error) would become smaller and disappears more quickly.

2.2. The media in China as a watchdog?

Because of its ability to monitor and expose corporate wrongdoing, the media in the U.S. have been viewed as a watchdog for small investors (e.g., Miller, 2006; Dyck et al., 2010). That is, the watchdog function, if carried out effectively, should increase firm-level transparency and improve investor protection, resulting in improvements in the firm-level information environment and stock price efficiency. Moreover, the watchdog effect of the media is expected to be comparatively greater when other standard investor-protecting institutions (such as the courts) are underdeveloped (e.g., La Porta et al., 1998).

In the Chinese context, standard investor-protecting institutions are weak (La Porta et al., 2004; Allen et al., 2005; Jiang et al., 2010) and the media institution is much better developed (Liebman, 2005; Lin, 2008). Thus, it is of interest to find out whether, and to what extent, the media in China can perform the watchdog function. At the outset, it is not apparent whether the state-controlled (i.e., non-independent) Chinese media can ever be a watchdog for small investors. However, the discussion in Appendix C about the media reform in China over the last 35 years suggests that the situation is more nuanced and reveals that a strict dichotomy between media independence and non-independence is not informative of the evolving Chinese media. Specifically, two factors are worth noting.

First, there has been a change in the media information paradigm as the media model in China changed from a propagandist model to a commercialized model, resulting in not only more information, but also more diversified information, conveyed to the public by the media.[7] The change in the media model reflects the fact that the legitimacy basis of the ruling Party has changed from one of communist ideology to one of economic performance; it also explains the Party's permanent retreat since the early 1990s on news reporting in the lifestyle-related *white zone* (see Appendix C, Footnote 21).

Second, the near-complete financial autonomy, the demand for credible and reliable information from the public, and competition among news media force most newspapers in China to become responsive to the information demands of the subscribing public. In particular, strong market competition and the increasing dissonance with a dying ideology have caused Chinese journalists to focus more on the battleground of *gray zone* news reporting, covering wrong-doing, fraud, corruption, and other unfair actions at the regional and corporate levels (see Appendix C, Footnote 22).

Based on the discussions above and in Appendix C, we conjecture that the media in China are able to perform a limited watchdog role at the regional and corporate levels for two main reasons. The first (arising from *market* logic) has to do with the strong market incentives unleashed by commercialization that motivate the media to push the limits of what is allowed. The second reason (arising from "*Party*" logic) is related to the congruence between the Party-state's overall interests in sustaining economic development as well as maintaining social stability at the national level and curbing corruption as well as fraud at the regional and corporate levels.

[7] By a "commercialized" media, it is meant that the media receives little or no state subsidies and must rely on private subscription and advertising revenue to survive in the marketplace. In short, a commercialized media needs to be market-driven.

In sum, we predict that stocks of firms with more media coverage in China are more price-efficient because these firms are more transparent and there is more informed trading of these firms' stocks. As a result, there is: (1) more firm-specific information being incorporated into stock prices; (2) more private information being transmitted to the public through informed trading; and (3) the deviation of a transaction stock price from the efficient price being reduced. We therefore hypothesize in alternative form:

H1a. Media coverage of a firm is positively associated with the amount of firm-specific information incorporated into its stock price, measured by the inverse of stock price synchronicity, all else being equal.

H1b. Media coverage of a firm is positively associated with the amount of firm-specific private information available, measured by the probability of informed trading in the stock market, all else being equal.

H1c. Media coverage of a firm is negatively associated with the deviation of a stock price from the efficient price, measured by stock pricing error, all else being equal.

2.3. Does institutional development matter?

Because investors normally rely on legal and other governance institutions for protection, these institutions are important to both the development and efficiency of the stock market (La Porta et al., 2000).[8] Evidence from Morck et al. (2000) suggests that poor investor protection increases the cost and risk of collecting firm-specific private information on the part of investors, and thus makes stock prices less informative. La Porta et al. (1998) present evidence to suggest that in countries with a weak legal system, compensating institutions (such as concentrated ownership) may develop to sustain financial activities.

Similar to La Porta et al. (1998), we argue that the incremental impact of the media as a compensating institution in providing investor protection is likely to be greater in countries (or regions within a country) where extant institutions that are normally expected to perform such functions are less developed. For example, the incremental effects of media coverage on investor protection and transparency are likely to be weaker in the U.S., the most developed market economy with mature institutional infrastructures, than in China, where institutional infrastructures are much less mature. Similarly, the incremental impact of media coverage is likely to vary in accordance with the extent of institutional development across 31 provincial/regional jurisdictions within China itself. Thus, the regional variation in institutional development within China provides an opportunity to test our hypothesis directly by examining the interplay between the role of the media and that of non-media institutions in influencing stock price efficiency. The compensating hypothesis on the role of the media suggests that the association between media coverage and stock price efficiency is likely to be stronger in regions of relatively weaker institutional development. We hypothesize in alternative form:

H2. The associations of media coverage with (a) stock price synchronicity, (b) firm-specific private information available, and (c) deviation of a transaction stock price from the efficient price, are more pronounced in regions (or provinces) with relatively weaker institutional development, all else being equal.

3. Data and variables

3.1. Empirical measures of media coverage

To test hypotheses H1 through H2, we need to construct empirical measures of firm-level media coverage in the Chinese language setting. This poses challenges, particularly because no previous research has addressed

[8] Acemoglu and Johnson (2005), for example, distinguish "contracting institutions" (which support economic activities among firms and/or individuals) and "property rights institutions" (which protect individuals and investors from expropriation by corporate insiders and politically powerful persons or organizations). However, these authors find that both sets of institutions are required to support the development of stock markets.

this issue in the Chinese language setting.[9] Unlike U.S.-based studies, where linguistic classification dictionaries (e.g., the commonly used General Inquirer's Harvard-IV-4 classification dictionary) are available to measure different types of media coverage in English (e.g., Tetlock et al., 2008; Fang and Peress, 2009), in this study we must develop our own for the Chinese language media.

To avoid subjective interpretations and to include as large a coverage range as possible, we choose to focus on *general* media coverage based on newspaper articles written about every listed firm (excluding financial firms) in China and published in all 14 national financial newspapers and the 10 largest comprehensive newspapers by circulation in China from 2000 to 2009.[10] To minimize measurement errors, we use a task-specific computer search program to download these articles. This approach has several advantages. By focusing on general media coverage, we need not decide subjectively whether coverage is positive or negative and how positive or negative it is. In addition, the ability to include a wide range of media coverage enables us to capture both the information enlargement effect and the information dissemination effect (Engelberg and Parsons, 2011), both of which support the watchdog/oversight function of the media. Moreover, this approach is justified because we are interested in how the watchdog function of media coverage (as opposed to how a specific type of coverage) facilitates firm-specific information flow or improves information-based trading in the stock market.

Specifically, we take the following three steps to construct a data set containing media coverage of listed firms using the WiseNews database.[11] First, for each firm in our sample, we download articles that mention the name of the firm for each firm-year after filtering out articles generated by the firm in question (e.g., various firm-generated announcements and reports). Here, our objective is to exclude self-generated media coverage.

Second, for each firm-year, we count the number of articles that mention the name of the firm and denote this total as *MediaCov1*. Although articles that mention the name of a particular firm in question may or may not be about the firm, the number does capture the publicity dimension of media coverage of the firm and enhances its "investor recognition." To ensure that an article is about the firm, researchers in U.S.-based studies normally rely on a "relevance score" provided by a classification dictionary in English. Fang and Peress (2009), for example, retain an article containing a firm's name if the article's "relevance score" is 90% or above. Since there is no classification dictionary in Chinese to rely on, we use a different empirical strategy in the next step.

Third, in each article, we count the number of times the name of the firm in question is mentioned and the number of times that the name of each of *all* listed firms in China is mentioned in the same article (if any).[12] For example, for given firm in a given year, we may find that the name of the firm in question (Firm A) in an article is mentioned 3 times, but the name of another firm (Firm B) is mentioned 5 times in the article, and the name of yet another firm (Firm C) is mentioned 6 times in the same article. We keep an article only if the name of the firm in question is mentioned more times than any other names on the assumption that this article is therefore likely to be about this firm. In other words, we would drop the article about Firm A in our example even though the name of Firm A is mentioned 3 times, because Firm B (Firm C) is mentioned 5 (6) times. We then re-count the number of articles in each firm-year and denote this total as *MediaCov2*.

By construction, *MediaCov1* measures media coverage of a firm by all articles that mention the name of the firm in question, whereas *MediaCov2* measures media coverage of a firm by articles in which the name of the firm is mentioned most often. While the *MediaCov1* captures the publicity or "investor recognition" dimension of media coverage, it is very likely that not all these articles are written about the firm. In comparison, *MediaCov2* is more likely to capture articles actually written about the firm, though this measure may still contain measurement errors. To the extent that both measures are likely to contain articles not actually written about the firm, they are biased against researchers finding significant results.[13]

[9] To the best of our knowledge, our study is the first attempt to empirically measure firm-level media coverage in China.

[10] Miller (2006) finds that, as far as financial frauds, the financial press tends to produce more original information, whereas the non-financial press tends to repeat information already reported elsewhere.

[11] WiseNews is a Hong Kong-based electronic news clipping service that provides access to news from about 550 newspapers and magazines published in Mainland China, Hong Kong, Taiwan, and Macau.

[12] In our sample of 2,144,850 articles, the average number of listed firms mentioned by name in an article is 6.

[13] Manually reading through the 2,114,850 articles in our sample is not feasible and would be subject to the uncertainty of subjective and inconsistent interpretations.

3.2. Empirical measures of stock price efficiency

To enhance the validity of our finding on the impact of media coverage on stock price efficiency, we assess stock price efficiency from three different but inter-related dimensions using three measures: (1) the ability of a stock price to incorporate firm-specific information as measured by the inverse of stock price synchronicity developed by Morck et al. (2000), (2) the amount of private firm-specific information about a stock available in the stock market as measured by the probability of informed trading (PIN) developed by Easley et al. (1996), and (3) the deviation of a transaction stock price from the efficient price as measured stock pricing error developed by Hasbrouck (1993).

Stock price synchronicity, denoted by *SYNCH*, reflects the amount of firm-specific information incorporated into stock prices through informed trading. High synchronicity of a stock indicates that its price reflects more common information relative to firm-specific information, or a less informative stock price.[14] The probability of informed trading denoted by *PIN*, is used to capture the amount of firm-specific *private* information available in the stock market. Previous empirical work generally supports the use of *PIN* as a valid measure of the probability of informed trading (Vega, 2006; Chen et al., 2007; Ferreira and Laux, 2007; Ferreira et al., 2011).[15] Following Hasbrouck (1993) and Boehmer and Kelley (2009), we measure a stock pricing error ($V(s)/V(p)$) as the dispersion or standard deviation of the extent to which a stock price deviates from random walk ($V(s)$), scaled by the standard deviation of the stock price itself ($V(p)$). Appendix B provides detailed explanations on procedures used to compute all measures mentioned above.

3.3. Empirical measures of institutional development

H2 concerns whether and how non-media institutional development across regions affects the impact of media coverage on stock price informativeness. To test H2, we need to obtain empirical proxies for the strength of institutional development across the economically fragmented 31 regions/provinces within China.[16] Given that administrative regions or provinces in China typically exhibit large variations in economic and institutional development, Fan et al. (2009) construct institutional development indices for each province to capture the efficacy or strength of institutions at the provincial level within China.[17] These indices, ranging from small (bad) to large (good), include: (1) overall market-institution development (denoted *Inst_General*); (2) government intervention in markets (denoted *Inst_Gov*); (3) private enterprise development (denoted *Inst_Private*); (4) regional protectionism (denoted *Inst_Protectionism*); (5) financial market development (*Inst_Financial*); and (6) legal environment (denoted *Inst_Law*). Since the above indices capture different aspects of the strength of regional institutions, we use them all, one by one, to test H2.

3.4. Other variables

Exact definitions and sources of the key variables above and other control variables used in the paper are provided in Appendix A. Financial information needed to construct non-media and non-institutional variables is collected from the China Stock Market and Accounting Research (CSMAR) database.

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[14] This measure has been used in a large number of studies to measure firm-specific information, including Piotroski and Roulstone (2004), Chan and Hameed (2006), Ferreira and Laux (2007), Fernandes and Ferreira (2009), Gul et al. (2010), and Ferreira et al. (2011).

[15] To avoid the effects of order flows, we also discompose *PIN* into the probability of a private information event occurring (*ALPHA*) and order flows and focus on *ALPHA* alone (see Appendix B for details).

[16] The problem of regional economic fragmentation within China has been described as "one country, thirty-one economies" (Huang, 2003, p. 141).

[17] Periodic reports based on these indices have been compiled by the National Economic Research Institute of the China Reform Foundation since 2001. These indices are used to gauge the varying degree of "marketization" across regions within China. Academics are beginning to use these indices to conduct China-based research (e.g., Srinidhi et al., 2012).

4. Empirical analysis

4.1. Descriptive statistics

Panels A and B of Table 1 show the extent of media coverage about our sample firms by year and industry, respectively. Overall, there are a total of 2,114,860 (499,876) articles when media coverage is measured as *MediaCov1* (*MediaCov2*) written about our sample of 9613 firm-years over the period 2000–2009. On average, there are 220 (52) articles per firm in a year when media coverage is measured as *MediaCov1* (*MediaCov2*). Given that there are, on average, 220 articles per firm in a year, a listed firm on average gets its name mentioned in an article at least once in a few days in any given year. Because all firms in our sample are listed and the range of our media coverage is quite extensive, there is no firm in the sample that received no media coverage in any given year. As shown in Panel A, the extent of media coverage per year (*MediaCov1*) appears to increase over time, with a low of 108 per firm in 2000 and a high of 307 per firm in 2008, though there is no clear trend. As shown in Panel B, more than half the firms in our sample are in the manufacturing industry. Firms in the mining industry tend to have the greatest media coverage, though there appears to be no significant difference in the distribution of coverage per firm by industry, irrespective of whether the coverage is measured by *MediaCov1* or *MediaCov2*.

Table 2 presents descriptive statistics. The mean and median of the R^2 for the market model specified in Appendix B are 0.4082 and 0.4027, respectively, with a standard deviation of 0.1635, suggesting that the R^2 is reasonably distributed. The R^2 for our Chinese sample is relatively high, reflecting that Chinese stocks tend to co-move more closely with common factors rather than firm-specific factors.[18] Consistent with the R^2 statistics, the mean and median of *SYNCH* are −0.4193 and −0.3939 with a standard deviation of 0.7035, much higher than for the U.S. sample of Piotroski and Roulstone (2004), who report a mean and median synchronicity of −1.742 and −1.754, respectively. This again suggests that Chinese stocks are more synchronous than U.S. stocks. The mean and median of *PIN* (*ALPHA*) in our sample are 0.1383 (0.3304) and 0.1398 (0.3340), respectively, with a standard deviation of 0.0600 (0.1259), suggesting *PIN* (*ALPHA*) is reasonably distributed. The mean (median) of $V(s)/V(p)$ in the sample is 0.1156 (0.1079), with a standard deviation of 0.0531, similarly suggesting a reasonable distribution.

With respect to descriptive statistics on other variables, the following are noteworthy: First, the mean for *Private Firm* of 0.317 indicates that about 32% of our sample firms are non-state-controlled listed firms, while the rest are state-controlled firms. Second, about 8% of our sample firms issue both A-shares for domestic investors and B-shares for foreign investors. Third, about 2% of our sample firms issue both A-shares and H-shares traded on the Hong Kong stock market. Finally, on average, each of our sample firms is followed by about one analyst.

Fig. 1a–d is graphic illustrations of the relations between *MediaCov1* in deciles and the corresponding *SYNCH*, *PIN*, *ALPHA*, and $V(s)/V(p)$ averages. The illustrations generally suggest negative relations of media coverage with *SYNCH* and $V(s)/V(p)$ and positive relations of media coverage with *PIN* and *ALPHA*. We obtain similar patterns when using *MediaCov2* to measure media coverage. These results are generally consistent with H1a that media coverage reduces stock price synchronicity, H1b that media coverage increases the probability of informed trading, and H1c that the media coverage reduces the deviation of a stock price from random walk. Taken together, they are consistent with the view that more intensive media coverage is associated with higher price efficiency for Chinese stocks.

4.2. The impact of media coverage on stock price efficiency

H1a concerns the impact of media coverage on the ability of stock prices to incorporate firm-specific information, measured by stock price synchronicity (*SYNCH*), while H1b and H1c re-examine H1a, focusing on the impact of media coverage on the amount of private firm-specific available, measured by the probability

[18] For example, Gul et al. (2010) report that the mean R^2 for their Chinese sample is about 45%, whereas Piotroski and Roulstone (2004) report a much lower mean R^2 of 19.3% for their U.S. sample.

Table 1
Media sample characteristics.

Year	No. of firms	No. of articles as measured by *MediaCov1*	No. of articles as measured by *MediaCov2*	Average media coverage	
				MediaCov1 per firm	*MediaCov2* per firm
Panel A: by year					
2000	285	30,780	5700	108	20
2001	707	108,171	21,917	153	31
2002	804	106,128	27,336	132	34
2003	893	175,921	40,185	197	45
2004	1019	106,995	23,437	105	23
2005	1086	280,188	65,160	258	60
2006	1144	323,752	85,800	283	75
2007	1158	279,078	68,322	241	59
2008	1256	385,592	79,128	307	63
2009	1261	254,722	63,050	202	50
Total	9613	2,114,860	499,876	220	52

Industry	No. of firms	Average media coverage	
		MediaCov1 per firm	*MediaCov2* per firm
Panel B: by industry			
Agriculture, Forestry, and Fishing	192	181	40
Mining	144	312	66
Manufacturing	5168	213	50
Electric, Gas, and Sanitary Services	439	202	47
Construction	167	215	60
Transportation	364	265	65
Communication	571	238	63
Wholesale and Retail Trade	729	219	53
Real Estate	705	242	55
Public Utility	335	218	66
Culture, Sport and Entertainment	88	255	43
Conglomerate	711	214	47
Total	9613	220	52

This table presents average media coverage distributions as measured by number of newspaper articles.

of informed trading (*PIN*) and the extent to which stock price deviates from random walk, measured by stock pricing error ($V(s)/V(p)$). To test H1a, H1b, and H1c, we specify the following regression:

$$StockPriceEfficiency = a_0 + a_1 Media\,Coverage + \sum_k a_k Control_k + (error) \tag{1}$$

where the dependent variable, *StockPriceEfficiency*, refers to the pricing efficiency of stocks. When testing H1a, *StockPriceEfficiency* is proxied by *SYNCH*, which is inversely related to the ability of stock price to incorporate firm-specific information. When testing H1b, *StockPriceEfficiency* is proxied by *PIN*, which has a direct relation with the probability of informed trading in a stock. When testing H1c, *StockPriceEfficiency* is proxied by $V(s)/V(p)$, which is positively related to the extent to which a stock price deviates from random walk. In Eq. (1), we measure the extent of media coverage about a firm using two alternative proxies: *Ln*(*MediaCov1*) and *Ln*(*MediaCov2*). Given the positive relation between *Media Coverage* and *StockPriceEfficiency*, H1a translates as $a_1 < 0$ when Eq. (1) is estimated with *StockPriceEfficiency* = *SYNCH*. H1b translates as $a_1 > 0$ when Eq. (1) is estimated with *StockPriceEfficiency* = *PIN*. H1c translates as $a_1 < 0$ when Eq. (1) is estimated with *StockPriceEfficiency* = $V(s)/V(p)$.

Following previous related research (Chan and Hameed, 2006; Ferreira and Laux, 2007; Boehmer and Kelley, 2009; Gul et al., 2010; Ferreira et al., 2011), we include a set of control variables in our regressions. These variables are firm size (*SIZE*), leverage (*LEV*), earnings volatility (*STDROA*), market-to-book ratio (*M*/*B*), the number of firms in the industry to which the firm belongs (*INDNUM*), the size of the industry

Table 2
Descriptive statistics.

	Q1	Mean	Median	Q3	Std dev.	*N*
Panel A: PIN and synchronicity						
PIN	0.1085	0.1383	0.1398	0.1734	0.0600	9613
ALPHA	0.2710	0.3304	0.3340	0.4038	0.1259	9613
SYNCH	−0.8812	−0.4193	−0.3939	0.0789	0.7035	9316
R^2	0.2929	0.4082	0.4027	0.5197	0.1635	9316
$V(s)/V(p)$	0.0705	0.1156	0.1070	0.1486	0.0531	9316
Panel B: Variables of interest used in the main regression						
MediaCov1	95	220	156	276	182.77	9613
MediaCov2	15	52	32	62	56.71	9613
Ln(MediaCov1)	0.0913	0.1013	0.1011	0.1125	0.0181	9613
Ln(MediaCov2)	0.0555	0.0693	0.0699	0.0829	0.0213	9613
Inst_General	6.11	7.78	7.66	9.55	2.163	9613
Inst_Gov	7.52	8.49	8.64	9.80	1.498	9613
Inst_Private	6.39	8.49	8.93	10.29	2.823	9613
Inst_Protectionism	8.90	9.50	10.29	10.77	1.831	9613
Inst_Financial	6.33	7.90	8.07	10.01	2.156	9613
Inst_Law	4.32	7.09	5.96	9.42	3.443	9613
Panel C: Other explanatory variables						
PrivateFirm	0	0.317	0	1	0.465	9613
B_SHARE	0	0.081	0	0	0.272	9613
H_SHARE	0	0.024	0	0	0.153	9613
Big_4	0	0.063	0	0	0.245	9613
ANALYST	0	1.075	0.693	2.079	1.217	9613
AD	0.017	0.059	0.037	0.072	0.051	9613
Panel D: Control variables						
SIZE	20.628	21.323	21.271	21.978	0.9546	9613
LEV	0.3861	0.5213	0.5281	0.6583	0.1851	9613
M/B	1.0467	1.5076	1.2677	1.7522	0.6328	9613
INDNUM	4.2341	4.5641	4.5218	4.9698	0.5567	9613
INDSIZE	25.887	26.477	26.453	27.088	0.8489	9613
STDROA	0.0133	0.0458	0.0261	0.0568	0.0489	9613
VOL	0.0448	0.1113	0.0875	0.1563	0.0787	9613

All variables are defined in Appendix A.

to which the firm belongs (*INDSIZE*), and trade volume (*VOL*). Year and industry dummies are also included to account for year and industry fixed effects.

Table 3 presents the results on the impact of media coverage on the amount of firm-specific information incorporated into stock prices. As shown, where *SYNCH* is used as the dependent variable, we find that the coefficients on both media coverage measures are highly significant with an expected negative sign at less than the 1% level. This is consistent with H1a, suggesting that a greater amount of firm-specific information is incorporated into the stock prices of firms with high media coverage than those with low coverage.

Panel A of Table 4 reports the results using *PIN* as the dependent variable. As shown in Panel A, we find that the coefficients on both media coverage measures are also highly significant with an expected positive sign at less than the 1% level. This is consistent with H1b. One concern with using PIN as a proxy of private firm-specific information is that it reflects both the probability of a private information event occurring (*ALPHA*) and trade order flows. To alleviate this concern, we discompose *PIN* into *ALPHA* and trade order flows. Panel B of Table 4 reports the results using *ALPHA* as the proxy for the amount of private firm-specific information available in the market. As shown in Panel B, we find that the coefficients of both media coverage measures are highly significant with an expected positive sign at less than the 1% level. These findings support the following view: Media coverage encourages information arbitrageurs to process more public information into value-relevant private information, and/or to collect and produce more private information.

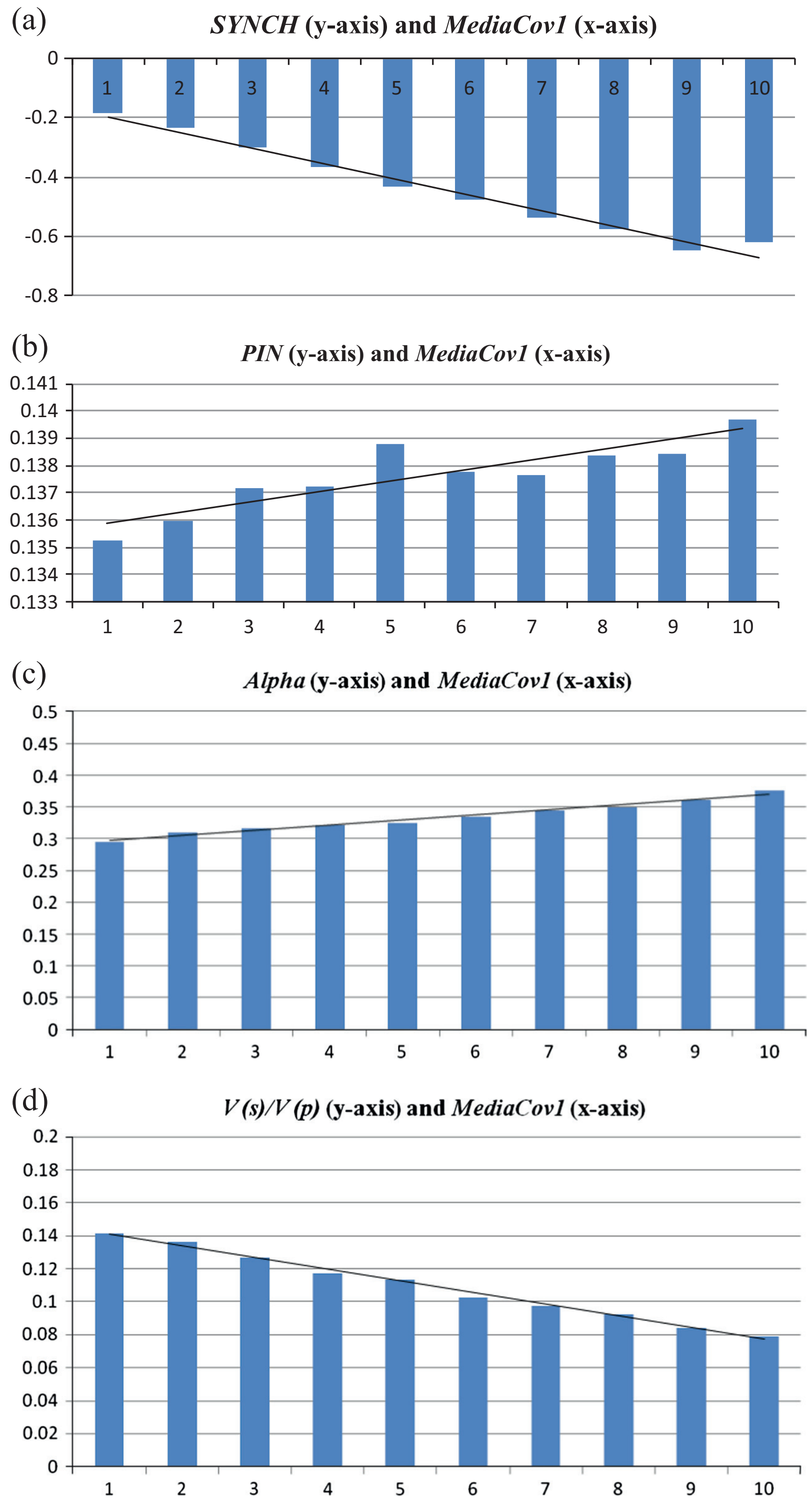

Figure 1. The figures illustrate the relation between media coverage in deciles and the corresponding synchronicity averages, the relation between media coverage in deciles and the corresponding *PIN* (*ALPHA*) averages, and the relation between media coverage in deciles and the corresponding $V(s)/V(p)$. All variables are defined in Appendix A.

Table 3
Media coverage and synchronicity.

Dependent variable = *SYNCH*	(1)	(2)
Constant	−2.8220	-3.1383
	(4.82)***	(5.32)***
Ln(MediaCov1)	−6.479	
	(9.32)***	
Ln(MediaCov2)		−3.432
		(6.79)***
SIZE	0.1833	0.1719
	(13.54)***	(12.51)***
LEV	0.0101	0.0122
	(1.80)*	(2.32)***
M/B	−0.0087	−0.0091
	(1.36)	(1.39)
INDNUM	0.0350	−0.0252
	(0.86)	(0.61)
INDSIZE	−0.0200	−0.0135
	(0.69)	(0.46)
STDROA	−0.2541	−0.2727
	(3.25)***	(3.34)***
VOL	−1.4290	−1.440
	(9.41)***	(9.38)***
Industry & year dummies	Yes	Yes
Observations	9316	9316
Adjusted R^2	0.398	0.389

This table presents regression results of the effects of media coverage on stock price efficiency as measured by synchronicity. The variables are as defined in Appendix A. Absolute values of z are reported in parentheses. Standard error is clustered by firm.
* Statistical significance level at 10%.
** Statistical significance level at 5%.
*** Statistical significance level at 1%.

Table 5 presents the results on the impact of media coverage on stock pricing error. As shown, where $V(s)/V(p)$ is used as the dependent variable, we find that the coefficients on both media coverage measures are highly significant with an expected negative sign at less than the 1% level. This is consistent with H1c, suggesting a smaller deviation from random walk for stock prices of firms with high media coverage than those with low coverage.

In short, the results from Tables 3–5 suggest the media coverage is able to increase stock price efficiency by increasing the ability of stock price to incorporate firm-specific information and the probability of having private firm-specific information available as well as reducing the deviation of a stock price from the efficient price.

4.3. Endogeneity

In the regressions specified above, we implicitly assume that stock price synchronicity, the probability of informed trading, and stock pricing error are influenced by media coverage. Intuitively, it seems unlikely that a stock (or the firm behind it) is able to affect media coverage by newspapers across the entire country. The firm behind a stock cannot decide the amount of media coverage it receives at any point in time. For example, there is no compelling reason to believe that firms with low stock price synchronicity are more or less able to attract media attention. Therefore, the results reported in Tables 3–5 are unlikely to be driven by reverse causality.[19]

[19] Engelberg and Parsons (2011), for example, find that media coverage strongly predicts financial market activities rather than the reverse.

Table 4
Media coverage, *PIN*, and *ALPHA*.

Dependent variable = *PIN*	(1)	(2)
Panel A.: Media coverage and PIN		
Constant	0.3665	0.3787
	(8.08)***	(8.35)***
Ln(*MediaCov1*)	0.4850	
	(5.25)***	
Ln(*MediaCov2*)		0.1492
		(4.26)***
Size	−0.0118	−0.0116
	(16.10)***	(15.58)***
LEV	−0.00030	−0.0002
	(0.51)	(0.48)
M/B	−0.0021	−0.0001
	(1.70)*	(1.67)*
INDNUM	0.0022	0.0025
	(0.85)	(0.93)
INDSIZE	−0.0002	−0.0010
	(16.10)***	(0.12)
STDROA	0.0002	0.0021
	(1.88)*	(1.87)*
VOL	−0.1687	−0.168
	(10.50)***	(10.45)***
Industry & year dummies	Yes	Yes
Observations	9613	9613
Adjusted R^2	0.156	0.155

Dependent variable = *ALPHA*	(1)	(2)
Panel B: Media coverage and ALPHA		
Constant	−0.4542***	−0.5149***
	(−4.19)	(−4.71)
Ln(*MediaCov1*)	0.0110***	
	(4.85)	
Ln(*MediaCov2*)		0.0073***
		(4.19)
SIZE	0.0238***	0.0247***
	(10.41)	(11.08)
LEV	−0.0099	−0.0110
	(−1.24)	(−1.38)
M/B	0.0220***	0.0234***
	(5.73)	(6.28)
INDNUM	−0.0033	−0.0032
	(−0.57)	(−0.55)
INDSIZE	0.0098**	0.0100**
	(2.10)	(2.13)
STDROA	0.1522***	0.1573***
	(4.65)	(4.80)
VOL	0.1209***	0.1404***
	(3.39)	(4.00)
Industry & year dummies	Yes	Yes
N	9613	9613
Adjusted R^2	0.131	0.131

Panel A presents regression results of the effects of media coverage on stock price efficiency as measured by *PIN*. The variables are as defined in Appendix A. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm. Panel B presents regression results of the effects of media coverage on stock price efficiency as measured by *ALPHA*. The variables are as defined in Appendix A. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

* Statistical significance level at 10%.

** Statistical significance level at 5%.

*** Statistical significance level at 1%.

Table 5
Media coverage and stock pricing error.

Dependent variable = $V(s)/V(p)$	(1)	(2)
Constant	0.5116***	0.5003***
	(17.31)	(16.68)
Ln(MediaCov1)	−0.0033***	
	(−4.06)	
Ln(MediaCov2)		−0.0032***
		(−5.04)
SIZE	−0.0151***	−0.0147***
	(−21.48)	(−20.93)
LEV	0.0198***	0.0199***
	(6.77)	(6.89)
M/B	−0.0024*	−0.0026**
	(−1.93)	(−2.07)
INDNUM	0.0024	0.0025
	(1.27)	(1.33)
INDSIZE	−0.0027**	−0.0028**
	(−2.16)	(−2.24)
STDROA	−0.0179	−0.0129
	(−1.44)	(−1.03)
VOL	−0.1754***	−0.1778***
	(−21.15)	(−21.73)
Industry & year dummies	Yes	Yes
N	9316	9316
Adjusted R^2	0.672	0.675

This table presents regression results of the effects of media coverage on stock price efficiency as measured by stock pricing error. The variables are as defined in Appendix A. Absolute values of z are reported in parentheses. Standard error is clustered by firm.
* Statistical significance level at 10%.
** Statistical significance level at 5%.
*** Statistical significance level at 1%.

Nonetheless, one cannot completely rule out the possibility of reverse causality or endogeneity. To alleviate this concern, we address the possible endogeneity problem using a *propensity score matching* (PSM) analysis (Rosenbaum and Rubin, 1984). To conduct a PSM analysis, we construct a reduced sample that consists of the sub-group of stocks with the highest 20% media coverage (high media coverage stocks, coded as 1) and a sub-group of stocks with the lowest 50% media coverage (low media coverage stocks, coded as 0). Using this reduced sample, we first estimate a probit model in which the likelihood of a stock getting high media coverage is linked to firm-specific variables that are deemed to affect the media coverage. Dyck et al. (2008) suggest that the level of media interest in a firm is determined by firm-specific attributes. Since these attributes are likely to correlate with dimensional characteristics that distinguish one firm from another, we include the following variables in the probit model: size (*SIZE*), risk (*Lev*), growth prospects (*M/B*), earnings volatility (*STDROA*), trading volume of its stock (*VOL*), the number of firms in the industry to which the firm belongs (*INDNUM*), and the size of the industry to which the firm belongs (*INDSIZE*). Year and industry dummies are also included to account for year and industry fixed effects.

Using this probit model, we obtain the *predicted* likelihood of a stock getting high media coverage, i.e., the propensity score, for each stock. Based on this score and using the one-to-one nearest neighbor matching method, we match stocks in the high media coverage group with stocks from the low media coverage group. When media coverage is measured as *MediaCov1* (*MediaCov2*), the mean allowable range of likelihood between a high media coverage stock and its PSM low media coverage match is 0.13% (0.15%), with the maximum at 3.2% (8.8%). By construction, in our PSM sample, a high media coverage stock and its PSM-matched, low media coverage stock are identical with respect to the predicted likelihood of media coverage, and thus, are equally likely ex ante to receive the same amount of media coverage (although they

Table 6
Propensity Score Method (PSM) analysis.

	(1) $\Delta SYNCH$	(2) ΔPIN	(3) $\Delta ALPHA$	(4) $\Delta V(s)/V(p)$
Panel A: Media coverage = MediaCov1				
Constant	0.1026	−0.0224***	0.0583**	0.0029
	(1.26)	(2.75)	(2.22)	(0.66)
Δ *Ln*(*MediaCov1*)	−0.0546*	0.0050**	0.0142**	−0.0081***
	(1.88)	(2.37)	(2.21)	(4.62)
$\Delta SIZE$	0.1570***	−0.0237***	−0.0034	−0.0249***
	(4.43)	(7.92)	(0.40)	(13.81)
ΔLEV	−0.5523***	0.0135***	0.0082	0.0196***
	(6.67)	(2.67)	(0.63)	(6.87)
$\Delta M/B$	−0.2346***	−0.0072***	0.0077	−0.0105***
	(8.63)	(3.51)	(1.35)	(8.36)
$\Delta INDNUM$	−0.0868***	0.0070**	−0.0185**	0.0106***
	(2.59)	(2.01)	(2.14)	(4.55)
$\Delta INDSIZE$	0.0943***	−0.0049*	0.0162**	−0.0115***
	(4.26)	(1.91)	(2.50)	(7.88)
$\Delta STDROA$	0.2070	−0.0272***	−0.0473**	−0.0428***
	(0.54)	(3.33)	(2.25)	(8.39)
ΔVOL	−0.8805***	−0.1954***	0.0921*	−−0.2820***
	(5.04)	(11.00)	(1.84)	(25.56)
Industry & year dummies	Yes	Yes	Yes	Yes
Observations	1863	1922	1922	1863
Adjusted R^2	0.391	0.216	0.038	0.550
Panel B: Media coverage = MediaCov2				
Constant	0.0477	−0.0376***	0.0221	−0.0089**
	(0.52)	(4.45)	(0.82)	(2.45)
Δ *Ln*(*MediaCov2*)	−0.0618***	0.0064***	0.0109**	−0.0036**
	(2.71)	(3.27)	(1.98)	(2.06)
$\Delta SIZE$	0.1536***	−0.0231***	0.0181*	−0.0243***
	(4.51)	(6.11)	(1.76)	(11.18)
ΔLEV	−0.4052***	0.0066	−0.0272*	0.0034
	(6.79)	(1.14)	(1.79)	(1.10)
$\Delta M/B$	−0.1334***	−0.0095***	0.0253***	−0.0108***
	(6.09)	(4.25)	(3.93)	(8.92)
$\Delta INDNUM$	−0.1227***	0.0006	0.0014	0.0064***
	(3.68)	(0.20)	(0.14)	(2.77)
$\Delta INDSIZE$	0.1525***	0.0012	0.0091	−0.0081***
	(5.68)	(0.49)	(1.12)	(4.56)
$\Delta STDROA$	0.1880	−0.0169	0.0126	−0.0186**
	(1.47)	(1.51)	(0.41)	(2.51)
ΔVOL	−1.7743***	−0.2003***	0.3070***	−0.3565***
	(8.42)	(9.68)	(5.45)	(27.84)
Industry & year dummies	Yes	Yes	Yes	Yes
Observations	1863	1922	1922	1863
Adjusted R^2	0.400	0.183	0.075	0.490

The matches for the sub-group of stocks with the highest 20% media coverage (high media coverage stocks, coded as 1) are selected from the sub-group of stocks with the lowest 50% media coverage (low media coverage stocks, coded as 0) using PSM. The panels below present the regression results of the impact of the ex post difference between matching pairs in actual media coverage on the difference in $SYNCH$, PIN, $ALPHA$, and $V(s)/V(p)$. Note that a high media coverage stock and its low media coverage match are PSM identical and thus *equally* likely ex ante to receive the same media coverage. The variables are as defined in Appendix A. Absolute values of z are reported in parentheses.

* Statistical significance level at 10%.
** Statistical significance level at 5%.
*** Statistical significance level at 1%.

in fact did not, ex post). Therefore, for the firms in this PSM sample, the difference in actual ex post media coverage, if any, is likely to be exogenous.

Using this reduced PSM sample of matching pairs, we repeat our regression analyses reported in Tables 3–5 by regressing the differences between pairs in *SYNCH*, *PIN*, *ALPHA*, and $V(s)/V(p)$, respectively on the differences between pairs in actual ex post media coverage (both measures) and in the controls. The results are reported in Panels A (for *MediaCov1*) and B (for *MediaCov2*) of Table 6. As shown in Panels A and B, the PSM-estimated coefficients of both media coverage variables are significant for the four dependent variables and with the expected signs. This is consistent with our full sample results reported in Tables 3–5. Our PSM results reported in Panels A and B of Table 6, taken together, suggest that our main results in Tables 3–5 are unlikely to be driven by potential endogeneity bias. The PSM results lend further support to H1a, H1b, and H1c that media coverage improves the pricing efficiency of stocks.

4.4. Variations in liquidity

Although all our regressions have been adjusted for trade volume (*VOL*), we carry out additional analysis to assess whether variations in liquidity drive our results. We do this by dividing the sample into liquid and illiquid sub-samples. The lack of liquidity or illiquidity is measured by the ratio of the daily absolute return to the RMB trade volume, averaged for each year. This illiquidity ratio gives the absolute (percentage) price change per RMB of daily trading volume, or the daily price impact of the order flow. A stock is classified as liquid (illiquid) if its illiquidity ratio is below (above) the median ratio in the whole sample. We then re-estimate the regressions of Tables 3–5 in these two sub-samples. The estimated coefficients of media coverage (both measures) with *SYNCH*, *PIN*, *ALPHA*, and $V(s)/V(p)$, respectively, as dependent variables for liquid and illiquid sub-samples are reported in Table 7. As shown, the media coverage coefficients of all four dependent variables remain significant and of the correct signs. In other words, out results are unlikely to be driven by variations in liquidity.

4.5. Robustness

Our regression analyses thus far include a set of seven control variables (*SIZE*, *LEV*, *M/B*, *INDNUM*, *INDSIZE*, *STDROA*, and *VOL*). We now include six additional control variables to alleviate concern about omitted correlated variables. Using a sample of Chinese listed firms, Gul et al. (2010) find that stock price syn-

Table 7
High liquidity versus low liquidity.

Dependent variables	Low liquidity		High liquidity	
	Ln(MediaCov1)	*Ln(MediaCov2)*	*Ln(MediaCov1)*	*Ln(MediaCov2)*
SYNCH	−0.1479***	−0.0717***	−0.1646***	−0.0630***
	(6.24)	(3.95)	(5.78)	(3.32)
PIN	0.0054***	0.0028*	0.0078***	0.0030**
	(2.81)	(1.90)	(3.89)	(2.00)
ALPHA	0.0164***	0.0082*	0.0144***	0.0068**
	(2.67)	(1.86)	(3.35)	(2.36)
$V(s)/V(p)$	−0.0060***	−0.0047***	−0.0061***	−0.0043***
	(3.33)	(3.46)	(4.47)	(3.54)

This table presents the estimated coefficients of media coverage from regressions using *SYNCH*, *PIN*, *ALPHA*, and $V(s)/V(p)$, respectively, as dependent variables in sub-samples divided according to illiquidity. The lack of liquidity or illiquidity of a stock is measured by the ratio of the daily absolute return to the RMB trade volume, averaged for each year. This illiquidity ratio gives the absolute (percentage) price change per RMB of daily trading volume, or the daily price impact of the order flow. A stock is classified as a high liquidity (low liquidity) stock if its illiquidity ratio is below (above) the median ratio in the whole sample. The variables are as defined in Appendix A. Absolute values of z are reported in parentheses.

* Statistical significance level at 10%.

** Statistical significance level at 5%.

*** Statistical significance level at 1%.

Table 8
Media coverage, synchronicity, PIN, stock price error and other explanatory variables.

Dependent variable = *SYNCH*	(1)	(2)
Panel A: Synchronicity		
Constant	−2.2851	−2.5945
	(3.79)***	(4.29)***
Ln(MediaCov1)	−6.0599	
	(8.74)***	
Ln(MediaCov2)		−2.9879
		(5.95)***
PrivateFirm	−0.0439	−0.0481
	(1.73)*	(1.87)*
B_SHARE	−0.0468	−0.0389
	(1.30)	(1.05)
H_SHARE	−0.1167	−0.1287
	(2.04)**	(2.14)**
Big_4	−0.02153	−0.0331
	(0.65)	(0.97)
Ln (ANALYST)	−0.0046	−0.0119
	(0.39)	(0.98)
AD	−0.5792	−0.6056
	(3.70)***	(3.87)***
SIZE	0.1808	0.1721
	(11.64)***	(10.88)***
LEV	0.0091	0.0106
	(01.56)	(1.93)*
M/B	−0.0082	−0.0084
	(1.29)	(1.30)
INDNUM	0.0494	0.0413
	(1.22)	(1.01)
INDSIZE	−0.0393	−0.0343
	(1.36)	(1.19)
STDROA	−0.2445	−0.2608
	(3.20)*	(3.27)*
VOL	−1.456	−1.484
	(9.80)***	(9.89)***
Industry & year dummies	Yes	Yes
Observations	9316	9316
Adjusted R^2	0.402	0.392

Dependent variable = *PIN*	(1)	(2)
Panel B: PIN		
Constant	0.3326	0.3473
	(7.05)***	(7.37)***
Ln(MediaCov1)	0.2622	
	(5.42)***	
Ln(MediaCov2)		0.1538
		(4.20)***
PrivateFirm	0.0027	0.0028
	(1.58)	(1.67)*
B_SHARE	0.0070	0.0070
	(2.92)***	(2.86)***
H_SHARE	0.001	0.0013
	(0.15)	(0.31)
Big_4	−0.000	−0.000
	(−.0.17)	(0.10)
Ln (ANALYST)	−0.0051	−0.0048
	(5.00)***	(4.73)***

(*continued on next page*)

Table 8 (*continued*)

Dependent variable = *PIN*	(1)	(2)
AD	−0.0028	−0.0031
	(0.28)	(0.29)
SIZE	−0.0093	−0.0091
	(10.63)***	(10.33)***
LEV	−0.0000	−0.0000
	(0.25)	(0.26)
M/B	−0.0002	−0.0001
	(1.66)*	(1.66)*
INDNUM	0.0031	0.0033
	(1.17)	(1.26)
INDSIZE	−0.0006	−0.0007
	(0.31)	(0.36)
STDROA	0.0002	0.0002
	(1.88)*	(1.87)*
VOL	−0.1776	−0.168
	(10.92)***	(10.45)***
Industry & year dummies	Yes	Yes
Observations	9613	9613
Adjusted R^2	0.161	0.159

Dependent variable = *ALPHA*	(1)	(2)
Panel C: ALPHA		
Constant	−0.4199***	−0.3273***
	(−3.69)	(−2.84)
Ln(*MediaCov1*)	0.0079***	
	(3.40)	
Ln(*MediaCov2*)		0.0058***
		(3.28)
PrivateFirm	0.0021	0.0016
	(0.57)	(0.43)
B_SHARE	−0.0014	−0.0017
	(−0.89)	(−0.83)
H_SHARE	−0.0105	−0.0103
	(−0.80)	(−0.90)
Big_4	0.0028	0.0027
	(0.42)	(0.40)
Ln (*ANALYST*)	−0.0102***	−0.0093***
	(−4.09)	(−3.76)
AD	−0.0001***	−0.0001***
	(−8.96)	(−8.32)
SIZE	0.0274***	0.0267***
	(10.02)	(9.66)
LEV	−0.0121	−0.0107
	(−1.33)	(−1.17)
M/B	0.0167***	0.0175***
	(3.57)	(3.72)
INDNUM	0.0078	0.0081
	(1.58)	(1.62)
INDSIZE	0.0039	0.0045
	(0.84)	(0.97)
STDROA	0.1423***	0.1412***
	(3.80)	(3.72)
VOL	−0.0604	−0.0483
	(−1.38)	(−1.11)
Industry & year dummies	Yes	Yes
Observations	9613	9613
Adjusted R^2	0.128	0.128

Table 8 (*continued*)

Dependent variable = $V(s)/V(p)$	(1)	(2)
Panel D: V(s)/V(p)		
Constant	0.5819***	0.5756***
	(16.67)	(16.42)
Ln(MediaCov1)	−0.0020***	
	(2.61)	
Ln(MediaCov2)		−0.0021***
		(3.40)
PrivateFirm	0.0001	0.0001
	(0.04)	(0.12)
B_SHARE	0.0056	0.0058
	(1.27)	(1.30)
H_SHARE	0.0057	0.0055
	(1.44)	(1.41)
Big_4	0.0012	0.0013
	(0.53)	(0.58)
Ln (ANALYST)	0.0028***	0.0028***
	(4.56)	(4.55)
AD	0.0001	0.0001
	(0.71)	(0.11)
SIZE	−0.0163***	−0.0161***
	(17.89)	(17.59)
LEV	0.0216***	0.0216***
	(6.95)	(6.92)
M/B	−0.0052***	−0.0052***
	(3.68)	(3.73)
INDNUM	0.0029	0.0029
	(1.40)	(1.44)
INDSIZE	−0.0028*	−0.0029**
	(1.94)	(1.99)
STDROA	0.0035	0.0044
	(0.49)	(0.62)
VOL	−0.0770***	−0.0779***
	(8.48)	(8.63)
Industry & year dummies	Yes	Yes
Observations	9316	9316
Adjusted R^2	0.663	0.663

Panel A presents regression results of the effects of media coverage on stock price efficiency as measured by *SYNCH* with the inclusion of other explanatory variables. The variables are as defined in Appendix A. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

Panel B presents regression results of the effects of media coverage on stock price efficiency as measured by *PIN* with the inclusion of other explanatory variables. The variables are as defined in Appendix A. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

Panel C presents regression results of the effects of media coverage on stock price efficiency as measured by *ALPHA* with the inclusion of other explanatory variables. The variables are as defined in Appendix A. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

Panel D presents regression results of the effects of media coverage on stock price efficiency as measured by $V(s)/V(p)$ with the inclusion of other explanatory variables. The variables are as defined in Appendix A. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

* Statistical significance level at 10%.

** Statistical significance level at 5%.

*** Statistical significance level at 1%.

Table 9
Media coverage, stock price synchronicity, and institutional development.

Dependent variable = *SYNCH*	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Synchronicity and MediaCov1						
Constant	−3.3563	−3.0673	−3.4235	−3.0975	−3.3246	−3.4911
	(5.40)***	(4.62)***	(5.64)***	(4.85)***	(5.43)***	(5.83)***
Ln(MediaCov1)	−5.6136	−9.9233	−5.7884	−8.7713	−6.3348	−3.6719
	(3.20)***	(3.61)***	(3.78)***	(3.18)***	(3.72)***	(3.31)***
Inst_General	−0.0483					
	(2.15)***					
Inst_General * *Ln(MediaCov1)*	0.4717					
	(2.25)***					
Inst_Gov		−0.0833				
		(2.50)***				
Inst_Gov * *Ln(MediaCov1)*		0.9497				
		(3.01)***				
Inst_Private			−0.0412			
			(2.37)***			
Inst_Private * *Ln(MediaCov1)*			0.4548			
			(2.74)***			
Inst_Protectionism				−0.0594		
				(2.13)**		
Inst_Protectionism * *Ln(MediaCov1)*				0.7647		
				(2.68)***		
Inst_Financial					−0.0529	
					(2.54)***	
Inst_Financial * *Ln(MediaCov1)*					0.5632	
					(2.75)***	
Inst_Law						−0.0289
						(2.08)**
Inst_Law * *Ln(MediaCov1)*						0.2433
						(1.90)*
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year and industry dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	7972	7972	7972	7972	7972	7972
Adjusted R^2	0.503	0.505	0.502		0.505	
Panel B: Synchronicity and MediaCov2						
Constant	−3.3238	−3.1998	−3.2969	−3.2775	−3.2825	−3.2943
	(5.50)***	(5.32)***	(5.64)***	(5.59)***	(5.62)***	(5.67)***
Ln(MediaCov2)	−2.4322	−5.6343	−2.7662	−4.2813	−2.8758	−0.9435
	(1.89)*	(2.94)***	(2.48)***	(2.36)***	(2.31)***	(1.16)
Inst_General	−0.0253					
	(2.09)**					
Inst_General * *Ln(MediaCov2)*	0.3463					
	(2.19)**					
Inst_Gov		−0.0356				
		(2.08)**				
Inst_Gov * *Ln(MediaCov2)*		0.6970				
		(3.12)***				
Inst_Private			−0.0206			
			(2.20)**			
Inst_Private * *Ln(MediaCov2)*			0.3590			
			(2.87)***			
Inst_Protectionism				−0.0156		
				(1.16)		
Inst_Protectionism * *Ln(MediaCov2)*				0.4988		
				(2.63)***		

Table 9 (*continued*)

Dependent variable = *SYNCH*	(1)	(2)	(3)	(4)	(5)	(6)
Inst_Financial					−0.0239	
					(2.19)***	
*Inst_Financial * Ln(MediaCov2)*					0.3977	
					(2.61)***	
Inst_Law						−0.0160
						(2.10)**
*Inst_Law * Ln(MediaCov2)*						0.1653
						(1.70)*
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year and industry dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	7972	7972	7972	7972	7972	7972
Adjusted R^2	0.502	0.501	0.503	0.506	0.503	0.505

Panel A presents regression results of the effect of *MediaCov1* (lagged one period) on *SYNCH* across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

Panel B presents regression results of the effect of *MediaCov2* (lagged one period) on *SYNCH* across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

* Statistical significance level at 10%.

** Statistical significance level at 5%.

*** Statistical significance level at 1%.

chronicity is higher for government controlled firms, while it is lower for A-share firms that simultaneously issue B-shares for foreign investors (traded on the Shanghai and Shenzhen stock exchanges) and H-shares traded on the Hong Kong stock exchange. Their study also finds that synchronicity is lower for Big 4-audited firms. Following their study, we now additionally include in Eq. (1) four indicator variables: (1) *Private Firm* equal to 1 for non-government controlled firms and 0 for government controlled firms; (2) *B-SHARE* equal to 1 for firms with both A- and B-shares; (3) *H-SHARE* equal to 1 for firms with both A- and H-shares; and (4) *Big 4* equal to 1 for firms with Big 4 auditors and 0 otherwise.

Chan and Hameed (2006) show that analyst following is positively associated with synchronicity in emerging markets. To further check the sensitivity of our results, we also control for the intensity of analyst coverage of a firm by including *Ln* (*ANALYST*). One may argue that a firm's exposure to media (as well as the ability to attract analyst attention) could be positively associated with the intensity of a firm's advertising activities. In an attempt to isolate the media coverage effects on firm-specific information flow from advertising (and analyst coverage), we also include the *AD* variable, which is measured by the ratio of a firm's selling expenses to sales. Panels of A, B, C, and D of Table 8 present the regression results after including the six additional control variables mentioned above.

As shown in Panel A of Table 8, when *SYNCH* is used as the dependent variable, the coefficients on both media coverage variables are highly significant at less than the 1% level with an expected negative sign. This suggests that our main results are robust to the inclusion of these additional control variables. The synchronicity-reducing impact of media coverage holds even after the additional control variables are accounted for. The following is noteworthy with respect to the additional control variables: First, we find that the coefficient on *Private Firm* is negative and marginally significant at less than the 10% level. This suggests that synchronicity is lower for non-government controlled firms than for government controlled firms, a finding consistent with Gul et al. (2010). Second, we find that synchronicity is significantly lower for A-share firms that also issue H-shares for investors in the Hong Kong stock market. This is consistent with the finding of Gul et al. (2010). Third, consistent with our previous findings, we find that synchronicity is lower for firms with higher advertising expenditures. Finally, the coefficients on other control variables are, overall, qualitatively identical with those reported in Panel A of Table 3.

Table 10
Media coverage, PIN, and institutional development.

Dependent variable = *PIN*	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: *PIN* and *MediaCov1*						
Constant	0.2355	0.1973	0.2429	0.1868	0.2442	0.2696
	(4.71)***	(3.73)***	(4.95)***	(3.57)***	(4.92)***	(5.36)***
Ln(*MediaCov1*)	0.4023	0.6493	0.2717	0.7939	0.3247	0.2350
	(2.76)***	(3.01)***	(2.25)***	(4.06)***	(2.35)***	(2.49)***
Inst_General	0.0048					
	(2.59)***					
Inst_General * *Ln*(*MediaCov1*)	−0.0422					
	(2.31)***					
Inst_Gov		0.0083				
		(3.14)***				
Inst_Gov * *Ln*(*MediaCov1*)		−0.0687				
		(2.68)***				
Inst_Private			0.0033			
			(2.31)***			
Inst_Private * *Ln*(*MediaCov1*)			−0.0235			
			(1.70)*			
Inst_Protectionism				0.0092		
				(4.20)***		
Inst_Protectionism * *Ln*(*MediaCov1*)				−0.0780		
				(3.64)***		
Inst_Financial					0.0038	
					(2.16)***	
Inst_Financial * *Ln*(*MediaCov1*)					−0.0318	
					(1.80)*	
Inst_Law						0.0023
						(1.84)*
Inst_Law * *Ln*(*MediaCov1*)						−0.0215
						(1.83)*
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year and industry dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8514	8514	8514	8514	8514	8514
Adjusted R^2	0.239	0.250	0.251	0.252	0.249	0.250
Panel B: *PIN* and *MediaCov2*						
Constant	0.2463	0.2176	0.2470	0.2223	0.2504	0.2646
	(5.07)***	(4.38)***	(5.11)***	(4.50)***	(5.17)***	(5.50)***
Ln(*MediaCov2*)	0.3449	0.5816	0.2405	0.5798	0.3213	0.1985
	(2.86)***	(3.32)***	(2,38)***	(4.02)***	(2.81)***	(2.65)***
Inst_General	0.0033					
	(3.01)***					
Inst_General * *Ln*(*MediaCov2*)	−0.0389					
	(2.57)***					
Inst_Gov		0.0059				
		(3.32)***				
Inst_Gov * *Ln*(*MediaCov2*)		−0.0646				
		(3.12)***				
Inst_Private			0.0025			
			(3.08)***			
Inst_Private * *Ln*(*MediaCov2*)			−0.0242			
			(2.10)***			
Inst_Protectionism				0.0053		
				(4.54)***		
Inst_Protectionism * *Ln*(*MediaCov2*)				−0.0597		
				(3.76)***		

Table 10 (*continued*)

Dependent variable = *PIN*	(1)	(2)	(3)	(4)	(5)	(6)
Inst_Financial					0.0031	
					(3.03)***	
Inst_Financial * *Ln*(*MediaCov2*)					−0.0355	
					(2.48)***	
Inst_Law						0.0016
						(2.24)***
Inst_Law * *Ln*(*MediaCov2*)						−0.0209
						(2.21)***
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Year and industry dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8514	8514	8514	8514	8514	8514
Adjusted R^2	0.253	0.252	0.253	0.251	0.252	0.251

Panel A presents regression results of the effect of *MediaCov1* (lagged one period) on *PIN* across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

Panel B presents regression results of the effect of *MediaCov2* (lagged one period) on *PIN* across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

** Statistical significance level at 5%.

* Statistical significance level at 10%.

*** Statistical significance level at 1%.

As shown in Panel B of Table 8, when *PIN* is used as the dependent variable, we find that the coefficients on both media coverage variables are significantly positive at less than the 1% level. This suggests that our results reported in Panel B of Table 3 are robust to the inclusion of the additional control variables. The *PIN*-increasing effect of media coverage holds even after controlling for additional determinants of the probability of informed trading. With respect to the estimated coefficients on the newly included control variables, the following are apparent: First, the amount and availability of private information is greater for Chinese A-share firms that simultaneously issue B-shares for foreign investors than for A-share only firms. Second, consistent with the finding of Chan and Hameed (2006), the coefficient on *ANALYST* is significantly negative, suggesting that the flow of private information about a firm is lower for firms with higher analyst as follows. This is in line with the view that analysts engage more in the acquisition and dissemination of common information than firm-specific information, thereby facilitating the flow of common information in emerging markets such as China. Finally, the coefficients on other control variables are, overall, qualitatively identical with those reported in Panel B of Table 3.

As shown in Panel C (D), the results of the *ALPHA* ($V(s)/V(p)$) regression against the same set of test and control variables are, overall, similar to those reported in Panel B (A) where *PIN* (*SYNCH*) is used as the dependent variable. We therefore do not repeat a full explanation of the regression results again, but we note the following: these results lend further support to H1a, H1b, and H1c, that is, the notion that firm-level media coverage can improve stock price efficiency in Chinese firms.

4.6. Media coverage and non-media institutional development

Consistent with H1a, H1b, and H1c our results provide strong and robust evidence that media coverage generally increases stock price efficiency. However, it is not clear whether the media substitutes/compensates for or complements other non-media institutions, which are underdeveloped in China (Allen et al., 2005). Building on the law and finance literature (La Porta et al., 1998), we examine whether the media in China can play the role of compensating for the relatively underdeveloped non-media investor protection institutions. Specifically, we expect the effect of media coverage on stock price efficiency to be stronger in countries or regions of relatively weaker institutional infrastructures (H2).

Table 11
Media coverage, ALPHA, and institutional development.

Dependent variable = *ALPHA*	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: *ALPHA* and *MediaCov1*						
Constant	−0.7125***	−0.7818***	−0.7075***	−0.8542***	−0.0718	−0.2003**
	(−6.87)	(−7.05)	(−6.99)	(−8.03)	(−0.95)	(−2.05)
Ln (MediaCov1)	0.0244***	0.0349***	0.0214***	0.0483***	0.0282***	0.0231***
	(3.80)	(3.61)	(4.05)	(5.93)	(6.36)	(5.49)
Inst_General * *Ln(MediaCov1)*	−0.0016*					
	(−1.91)					
Inst_General	0.0083**					
	(1.96)					
Inst_Gov * *Ln(MediaCov1)*		−0.0027**				
		(−2.33)				
Inst_Gov		0.0150**				
		(2.55)				
Inst_Private * *Ln(MediaCov1)*			−0.0011*			
			(−1.80)			
Inst_Private			0.0065**			
			(2.10)			
Inst_Protectionism * *Ln(MediaCov1)*				−0.0040***		
				(−4.40)		
Inst_Protectionism				0.0209***		
				(4.68)		
Inst_Financial * *Ln(MediaCov1)*					−0.0011*	
					(−1.79)	
Inst_Financial					0.0047	
					(1.39)	
Inst_Law * *Ln(MediaCov1)*						−0.0009*
						(−1.67)
Inst_Law						0.0036
						(1.25)
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry and year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8514	8514	8514	8514	8514	8514
Adjusted R^2	0.135	0.135	0.135	0.137	0.092	0.092
Panel B: ALPHA and MediaCov2						
Constant	−0.5295***	−0.5810***	−0.5276***	−0.5921***	−0.0544	−0.5020***
	(−5.62)	(−6.00)	(−5.65)	(−6.15)	(−0.82)	(−5.38)
Ln (MediaCov2)	0.0218***	0.0335***	0.0190***	0.0337***	0.0227***	0.0145***
	(3.86)	(3.96)	(4.05)	(4.84)	(6.08)	(4.09)
Inst_General * *Ln(MediaCov2)*	−0.0019***					
	(−2.65)					
Inst_General	0.0059**					
	(2.38)					
Inst_Gov * *Ln(MediaCov2)*		−0.0031***				
		(−3.16)				
Inst_Gov		0.0108***				
		(3.11)				
Inst_Private * *Ln(MediaCov2)*			−0.0014***			
			(−2.71)			
Inst_Private			0.0052***			
			(2.89)			
Inst_Protectionism * *Ln(MediaCov2)*				−0.0030***		
				(−3.94)		
Inst_Protectionism				0.0106***		
				(4.12)		

Table 11 (*continued*)

Dependent variable = *ALPHA*	(1)	(2)	(3)	(4)	(5)	(6)
Inst_Financial * *Ln*(*MediaCov2*)					−0.0009*	
					(−1.78)	
Inst_Financial					0.0022	
					(1.13)	
Inst_Law * *Ln*(*MediaCov2*)						−0.0010**
						(−2.21)
Inst_Law						0.0028*
						(1.69)
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry and year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8464	8464	8464	8464	7899	8464
Adjusted R^2	0.132	0.132	0.132	0.133	0.087	0.132

Panel A presents regression results of the effect of *MediaCov1* (lagged one period) on *ALPHA* across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

Panel B presents regression results of the effect of *MediaCov2* (lagged one period) on *ALPHA* across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of *z* are reported in parentheses. Standard error is clustered by firm.

* Statistical significance level at 10%.

** Statistical significance level at 5%.

*** Statistical significance level at 1%.

We test H2 by evaluating the interaction effect of media coverage and the level of institutional development across different provinces in China. For this purpose, we estimate the following regressions:

$$StockPriceEfficiency = b_0 + b_1 Media\,Coverage + b_2 Institutional\,Development + b_3 Media\,Coverage * Institutional\,Development + \sum_k b_k Control_k + (error) \quad (2)$$

where the dependent variable, *StockPriceEfficiency*, refers to the pricing efficiency of stocks as in Eq. (1). When testing H2a, *StockPriceEfficiency* is proxied by *SYNCH*. When testing H2b, *StockPriceEfficiency* is proxied by *PIN*. When testing H2c, *StockPriceEfficiency* is proxied by $V(s)/V(p)$. We measure the extent of media coverage about a firm using two alternative proxies: *Ln*(*MediaCov1*) and *Ln*(*MediaCov2*). In Eqs. (2), H2a and H2c are supported if $b_1 < 0$ and $b_3 > 0$ when the dependent variable is *SYNCH* or $V(s)/V(p)$. H2b is supported if $b_1 > 0$ and $b_3 < 0$ when the dependent variable is *PIN* or *ALPHA*.

Table 9 presents the regression results for Eq. (2), where *SYNCH* is used as the proxy for stock price efficiency. In Panel A, media coverage is measured by *Ln*(*MediaCov1*), while in Panel B, it is measured by *Ln* (*MediaCov2*). We use six different measures of institutional development constructed by Fan et al. (2009) and described in Section IV C to proxy for the relative non-media institutional development in a region. Results from both Panels A and B of Table 9 show that the effect of media coverage on mitigating synchronicity remains negative and statistically significant (with one exception, in model 6 of Panel B) even after controlling for the strength of regional institutional infrastructures and their interaction with media coverage. This result lends further support for our earlier finding that media coverage facilitates incorporation of firm-specific information into stock prices, thus mitigating synchronicity. In addition, the coefficients on each of six non-media institutional development proxies are all negative and significant for all cases except one (column 4 in Panel B) in both panels. This finding suggests that stocks of firms in regions with more developed non-institutions are less synchronous. This cross-regional evidence within China is consistent with the cross-country evidence reported in Morck et al. (2000) and Kim and Shi (2012).

Table 12
Media coverage, $V(s)/V(p)$, and institutional development.

Dependent variable = $V(s)/V(p)$	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: V(s)/V(p) and MediaCov1						
Constant	0.5318***	0.5392***	0.5544***	0.5764***	0.5465***	0.5472***
	(19.82)	(18.70)	(21.07)	(20.37)	(20.66)	(20.79)
Ln(MediaCov1)	−0.0073***	−0.0102***	−0.0109***	−0.0153***	−0.0084***	−0.0091***
	(−3.26)	(−3.18)	(−6.86)	(−4.51)	(−6.36)	(−7.44)
Inst_General * *Ln(MediaCov1)*	0.0005**					
	(2.04)					
Inst_General	−0.0027*					
	(−1.95)					
Inst_Gov * *Ln(MediaCov1)*		0.0008**				
		(2.34)				
Inst_Gov		−0.0034*				
		(−1.79)				
Inst_Private * *Ln(MediaCov1)*			0.0007***			
			(4.12)			
Inst_Private			−0.0038***			
			(−4.16)			
Inst_Protectionism * *Ln(MediaCov1)*				0.0013***		
				(3.94)		
Inst_Protectionism				−0.0072***		
				(−4.53)		
Inst_Financial * *Ln(MediaCov1)*					0.0005***	
					(2.90)	
Inst_Financial					−0.0031***	
					(−3.04)	
Inst_Law * *Ln(MediaCov1)*						0.0006***
						(3.93)
Inst_Law						−0.0034***
						(−4.10)
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry and Year Effect	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8720	8720	8720	8720	8720	8720
Adjusted R^2	0.675	0.675	0.672	0.676	0.672	0.672
Panel B: V(s)/V(p) and MediaCov2						
Constant	0.5157***	0.5199***	0.5104***	0.5369***	0.5085***	0.5079***
	(19.49)	(19.20)	(19.48)	(20.11)	(19.35)	(19.36)
Ln(MediaCov2)	−0.0085***	−0.0120***	−0.0073***	−0.0128***	−0.0057***	−0.0059***
	(−4.61)	(−4.34)	(−4.45)	(−4.27)	(−4.72)	(−5.00)
Inst_General * *Ln(MediaCov2)*	0.0007***					
	(3.14)					
Inst_General	−0.0023***					
	(−2.79)					
Inst_Gov * *Ln(MediaCov2)*		0.0010***				
		(3.40)				
Inst_Gov		−0.0027**				
		(−2.33)				
Inst_Private * *Ln(MediaCov2)*			0.0005***			
			(2.76)			
Inst_Private			−0.0014**			
			(−2.23)			
Inst_Protectionism * *Ln(MediaCov2)*				0.0011***		
				(3.54)		
Inst_Protectionism				−0.0041***		
				(−4.07)		

Table 12 (*continued*)

Dependent variable = $V(s)/V(p)$	(1)	(2)	(3)	(4)	(5)	(6)
Inst_Financial * *Ln*(*MediaCov2*)					0.0004**	
					(2.34)	
Inst_Financial					−0.0014**	
					(−2.23)	
Inst_Law * *Ln*(*MediaCov2*)						0.0003***
						(2.61)
Inst_Law						−0.0013**
						(−2.47)
Other explanatory variables	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Industry and year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Observations	8720	8720	8720	8720	8720	8720
Adjusted R^2	0.675	0.676	0.675	0.676	0.675	0.675

Panel A presents regression results of the effect of *MediaCov1* (lagged one period) on $V(s)/V(p)$ across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of z are reported in parentheses. Standard error is clustered by firm.

Panel B presents regression results of the effect of *MediaCov2* (lagged one period) on $V(s)/V(p)$ across regions of varying institutional development. The variables are as defined in Appendix A. Other explanatory variables and controls refer to those used in Table 8. Absolute values of z are reported in parentheses. Standard error is clustered by firm.

* Statistical significance level at 10%.

** Statistical significance level at 5%.

*** Statistical significance level at 1%.

More importantly, the coefficients on the interaction of media coverage and each of six non-media institutional development proxies are all positive and significant across six columns in both panels. This finding is consistent with H2, suggesting that the effect of media coverage on mitigating stock price synchronicity is significantly stronger (weaker) in regions of relatively weaker (stronger) institutional development.

Table 10 presents the regression results for Eq. (3), where *PIN* is used as the proxy for stock price efficiency. In Panel A, media coverage is measured by *Ln*(*MediaCov1*), while in Panel B, it is measured by *Ln* (*MediaCov2*). As shown in Panels A and B of Table 10, the coefficient on the media coverage variable is significant with an expected positive sign across all six columns in both panels. This finding lends further support for our earlier finding: the main effect of media coverage on increasing the probability of informed trading remains positive and statistically significant even after controlling for the strength of regional institutional infrastructures and their interaction with media coverage. Moreover, the coefficients on each of six regional non-media institutional development proxies are all significantly positive across all six columns in both panels. This finding suggests that stocks of firms located in regions with more developed non-media institutions have significantly higher PIN, and this result is robust to alternative measures of non-media institutions in each region. Stated another way, we provide evidence suggesting that more firm-specific private information is available for stocks of firms located in regions where non-media institutions are better developed.

More importantly, coefficients on the interaction between media coverage and each of six non-media institutional development proxies are significantly negative across six columns in both Panels A and B. This finding is consistent with H2, suggesting that the effect of media coverage on increasing *PIN* is significantly stronger (weaker) pronounced in regions of stronger (weaker) institutional development. In summary, the results in Tables 8 and 9, taken together, suggest that the media contributes incrementally more to making stock prices more informative in China, where non-media institutions are still less developed. One implication is that the media in China is able to compensate for other less developed non-media investor protection institutions.

Tables 11 and 12 show regression results for Eq. (3), where $ALPHA$ and $V(s)/V(p)$ are used, respectively, as proxies for stock price efficiency. Since results related to $ALPHA$ ($V(s)/V(p)$) are consistent with those related to PIN in Table 9 (with those related to $SYNCH$ in Table 8), we will not repeat the same exposition here. Suffice to say that the results from Tables 9–12, taken together, support our main hypothesis (H2) that the impact of media coverage on stock price efficiency is greater in regions of weaker institutional development within the 31 jurisdictions inside China.

5. Concluding remarks

Based on a sample of over two million newspaper articles and using a large sample of firms listed in China's emerging stock markets, where "investor protection is weak and agency problems are severe" (Allen et al., 2005, p. 90), this study investigates whether the commercialized state-owned media in China has an incremental impact on stock price efficiency. Our main findings can be summarized as follows. First, we find that as media coverage of a firm increases, its stock price synchronicity decreases. Second, we find that as media coverage of a firm increases, both the probability of informed trading and the probability of a private information event occurring in its stock increase. Third, we find that as media coverage of a firm increases, the extent to which its stock price deviates from random walk decreases. These findings, taken together, suggest media coverage improves the efficiency of stock prices. The above results are robust to potential endogeneity and reverse causality as well as to variations in liquidity. Finally, we examine the relative impact of the media across 31 provinces/regions within China. We find that the impact of the media on decreasing stock price synchronicity, increasing the probability of informed trading, and reducing stock pricing error is stronger (weaker) in regions of weaker (stronger) institutional development within China. This finding suggests that media coverage plays a significant role of compensating for the relatively underdeveloped governance institutions in transitional economies such as China.

Overall, our results suggest that a commercialized and increasingly market-driven Chinese media can increase stock price efficiency in China where standard governance institutions to protect investors and maintain corporate transparency are weak. As the first attempt to study the impact of the Chinese media on stock prices, our results are preliminary. As such, they contain one caveat and raise two questions. The caveat is that despite the market-driven commercialization, the watchdog effect of the media would soon cease should the Chinese government (or the Chinese Communist Party) consider curbing corruption and corporate fraud no longer a well-grounded state policy objective (a political survival strategy), although such scenario is luckily unlikely. The first question raised by our results is whether the impact of this type of media (or a more independent media) on stock price efficiency will remain significant in a context when standard governance institutions are developed and strong.[20] The second question is whether the compensating role of the media relative to non-media institutions (where they are less developed) is general or specific only to the pricing efficiency of stocks. We leave these questions to future research.

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[20] Note the inter-regional analysis controls for the difference in institutional development between regions within China but not for the difference in institutional development between China and, say, the US.

Appendix A. Variable definitions

Dependent variables

SYNCH: natural logarithmic transformation of the world market model, computed as Log $[R^2/(1-R^2)]$, where R^2 is the R^2 of the model defined in Appendix B.
PIN: probability of informed trading, computed based on the model defined in Appendix B.
ALPHA: probability of a private information event occurring based on the model defined in Appendix B.
V(s)/V(p): pricing error based on the model defined Appendix B scaled by the standard deviation of intraday (log) transaction prices.

Explanatory variables of interest

MediaCov1: the number of newspaper articles that mention the name of the firm.
MediaCov2: the number of newspaper articles in which the name of the firm in question is mentioned most often (i.e., the number of times the name of the firm is mentioned is highest in an article relative to the names of all other listed firms mentioned in the article).
Ln(MediaCov1): natural logarithmic transformation of *MediaCov1* plus one scaled by two hundred.
Ln(MediaCov2): natural logarithmic transformation of *MediaCov2* plus one scaled by two hundred.
Inst_General: score for overall development of market institutions in the region where the firm is headquartered; a higher score indicates better development.
Inst_Gov: score for government intervention in markets in the region where the firm is headquartered; a higher score indicates lower government intervention.
Inst_Private: score of private enterprise development in the region where the firm is headquartered; a higher score indicates better private enterprise development.
Inst_Protectionism: score of regional protectionism in the region where the firm is headquartered; a higher score indicates lower regional protectionism.
Inst_Financial: score for financial market development in the region where the firm is headquartered; a higher score indicates better financial market development.
Inst_Law: score for legal environment in the region where the firm is headquartered; a higher score indicates better legal environment.

Other explanatory variables

Private Firm: an indicator variable equal to one if the firm is a private sector (rather than government-owned or controlled) firm; zero otherwise.
B_SHARE: indicator variable equal to one if the firm also issues B-shares (denominated in US dollars and set up for foreign investors), zero otherwise.
H_SHARE: indicator variable equal to one if the firm also lists on the Hong Kong stock exchange, zero otherwise.
BIG_4: an indicator variable equal to one if the firm uses a Big 4 international auditor, zero otherwise.
Ln(ANALYST): natural logarithmic transformation of one plus the number of analysts following the firm, zero otherwise.
AD: ratio of selling expense over sales, a proxy for advertising expense.

Control variables

SIZE: natural logarithmic transformation of the book value of a firm's total assets.
LEV: ratio of total liability over total assets.
M/B: market-to-book equity ratio.
INDNUM: natural logarithmic transformation of the number of firms in the industry to which a firm belongs.
INDSIZE: natural logarithmic transformation of year-end total assets of all sample firms in the industry to which a firm belongs.
STDROA: standard deviation of Return-on-assets over the preceding five-year period, including the current year.

VOL: trading volume computed as the total value of daily shares traded scaled by total value of shares outstanding in a day.

Appendix B. Estimation of *SYNCH, PIN*, and *V*(*s*) variables

$SYNCH = \mathrm{Ln}\,(R^2/(1 - R^2))$, where R^2 is the R^2 of a firm-year obtained from the following regression models during that firm-year:

$$RET_{it} = a + \beta_1 MKRET_t + \beta_2 WRDRET_t + \varepsilon_{it}$$

(Firms with A-shares only)

$$RET_{it} = a + \beta_1 MKRET_t + \beta_2 MKRET^B + \beta_3 WRDRET_t + \varepsilon_{it}$$

(Firms with A- and B-shares)

$$RET_{it} = a + \beta_1 MKRET_t + \beta_2 MKRET^H + \beta_3 WRDRET_t + \varepsilon_{it}$$

(Firms with A and H-shares)

where, for firm i at day t, RET_{it} is daily return of firm i trading in either Shanghai or Shenzhen exchange on day t; $MKRET$ is the value weighted market return, based on the domestic composite value weighted index of shares (A-shares). WRDRET is the world market return computed using MSCI World Index. $MKRET^B$ is the composite value weighted market index of B-shares (denominated in US dollars and set up for foreign investors). $MKRET^H$ is the value weighted Hong Kong market return using the value weighted Hang Seng index.

$$PIN = \frac{\alpha\mu}{\alpha\mu + \varepsilon_b + \varepsilon_s}$$

where α ($ALPHA$) is the probability of a private information event occurring; μ is the arrival rate of informed traders when an information event occurs; ε_b is the arrival rate of buy orders, and ε_s is the arrival rate of sell orders. Following Easley et al. (1996), these parameters are estimated using a maximum likelihood estimation procedure, assuming a Poisson arrival process for informed and uninformed traders, and based on actual information about daily buy and sell orders provided by CSMAR for the period 2000–2009. Note that using actual buy and sell orders rather than algorithms to determine these orders increases the precision of our estimates.

$V(s)$ = the lower bound for the standard deviation of dispersion of pricing error, s, based on the model proposed by Hasbrouck (1993) and used by Boehmer and Kelley (2009), where the (log) transaction stock price at transaction time t, p_t, is defined to be the sum of a random component, m_t, and a pricing error, s_t:

$$p_t = m_t + s_t$$

whereas m, the random walk component represents the efficient price, the pricing error, s, captures the extent to which a transaction stock price deviates from its efficient price. $V(s)$ is estimated using a VAR system with five lags over the first difference in p and three trade variables: (1) a trade sign indicator, (2) signed trading volume, and (3) the signed square root of trading volume. Following Boehmer and Kelley (2009), we sign a trade by assuming that a trade is buyer initiated if the price is above the prevailing quote midpoint (and seller initiated for the converse). $V(s)$ is estimated for each stock in the sample for each year based on average weekly trade orders of the stock in the year provided by CSMAR for the period 2000–2009.

Appendix C. From propaganda to commercialization: a discussion of the Chinese media

The media in China is an integral part of the government or of the Chinese Communist Party (CCP). The Propaganda Department of the CCP Central Committee (Central Propaganda Department, or CPD) has been responsible for the overall management of culture, including the mass media, in China since the early 1950s. Most newspapers are administratively linked to the CCP. The CCP Central Committee, for example, publishes *People's Daily*. The model is replicated at local levels (e.g., *Southern Daily* is linked to Guangdong Provincial

Communist Party Committee). For many years prior to the economic reform in 1978, central and local newspaper articles in China were no more than official propaganda documents from the Party-state.

The commercialization of the Chinese media started in the late 1970s in tandem with economic reform for two reasons. First, as the basis of legitimacy for CCP rule changed from one of communist ideology to one of economic performance, the propaganda model of the media no longer suited the Party needs. Second, the financial burden of a nationwide propagandist media also proved too onerous for the government. Thus, media commercialization had the dual aim of reducing state subsidies for the media and yet expanding the media to further Party-state interests in economic development and social stability (Huang, 2001, 2007; Liebman, 2005; Lin, 2006). As the state reduced subsidies, many newspapers had to find other means toward financial self-sufficiency. Continued state ownership, combined with the emphasis on profitability, led to the so-called "marketization" of the media. For example, the state-controlled Southern Daily Group, whose flagship *Southern Daily* remains the official mouthpiece of the Guangdong Provincial Communist Party Committee, also publishes *Southern Metropolitan Daily*, a successful mass market tabloid with a high circulation, and *Southern Weekend*, a weekly paper regarded as one of the most outspoken papers in China. With the exception of the official *Southern Daily*, which receives state support, other papers in the group are *commercial* newspapers in that they must rely entirely on income generated from advertising and private subscription. A portion of profits from the commercially successful *Southern Weekend* and other commercial papers in the Group is used to partly underwrite the cost of publishing the unprofitable mouthpiece *Southern Daily*.

Media commercialization represents a change in the media information paradigm. The number of newspapers in China grew from 186 in 1978 to near 2000 in 2005, and most of them now operate without state subsidies (Esarey, 2005; Lin, 2006). Competition is often fierce between commercial newspapers within and across regions, forcing newspapers to provide more and increasingly diversified information content to readers. The near-complete financial autonomy, the demand for credible and reliable information from the public, and competition among news media force most newspapers in China to become responsive to the information demands of the subscribing public. A newspaper with exclusively old-style *propaganda* reporting typically does not have strong demand from the public. Because readership has become a crucial factor determining profitability (or even survival), the market logic dictates that all commercial newspapers must now carry at least some critical and investigative reporting to maintain and expand market share.

Notwithstanding the extensive commercialization, the media in China remains state-controlled and is subject to a range of extensive formal and informal regulations (Liebman, 2005). As a rule, any direct criticism of the national CCP leadership, the central government, or the military is strictly off-limits. Any reporting that frames these parties as the center of grievance is simply not allowed.[21] Journalists could, however, get away with criticizing regional government officials and corporations if the central government and the CCP leadership are framed to be on the right side. Subject to formal and informal norms in China, a commercialized newspaper caters to the demands of its reading public, on whom it relies for revenue. Although the media avoids ideologically sensitive topics (e.g., democratic reform), it does enjoy a freer hand in exposing wrongdoing at the regional or corporate level (Keatley, 2003). The commercialized financial press (e.g., *Caijing Magazine*) has been instrumental in exposing much corruption/fraud and numerous corporate scandals in firms and financial markets (Lin, 2006).[22]

References

Allen, F., Qian, J., Qian, M., 2005. Law, finance, and economic growth in China. J. Financ. Econ. 77, 57–116.
Acemoglu, D., Johnson, S., 2005. Unbundling institutions. J. Polit. Econ. 113, 949–995.
Boehmer, E., Kelley, E.K., 2009. Institutional investors and the information efficiency of prices. Rev. Financ. Stud. 22, 3563–3594.

[21] Lin (2008) divides the current news reporting in China into three zones: black, gray, and white. *Black zone* reporting covers news reporting that poses direct challenges to the CCP rule and is, without exception, censored by coercive political power. *White zone* reporting covers apolitical soft news and has posed no problems since the early 1990s when the Party retreated from its ideological stand against the *bourgeois life style*. The battleground now is in the *gray zone*, where news reporting covers wrong-doing, fraud, corruption, and other unfair actions at the regional and corporate levels.

[22] See, for example, "Zhengjian Hui Shouxi Guwen Cheng Yulun Jiandu Hen Zhongyao" [Chief Advisor to the Securities Commission States That Popular Opinion Supervision is Very Important], Renmin Wang [People's Daily Online], July 14, 2001.

Chan, K., Hameed, A., 2006. Stock price synchronicity and analyst coverage in emerging markets. J. Financ. Econ. 80, 115–147.
Chen, Q., Goldstein, I., Jiang, W., 2007. Price informativeness and investment sensitivity to stock price. Rev. Financ. Stud. 20 (3), 619–650.
Dyck, A., Volchkova, N., Zingles, L., 2008. The corporate governance role of the media: evidence from Russia. J. Finance 63, 1093–1135.
Dyck, A., Morse, A., Zingles, L., 2010. Who blows the whistle on corporate fraud. J. Finance 65, 2213–2254.
Easley, D., Kiefer, M.N., O'Hara, M., Paperman, J.B., 1996. Cream-skimming or profit sharing? The curious role of purchased order flow. J. Finance 51, 811–833.
Engelberg, J.E., Parsons, C.A., 2011. The causal impact of media in financial markets. J. Finance 66, 67–97.
Esarey, A., 2005. Cornering the market: state strategies for controlling China's commercial media. Asian Perspect. 29, 37–83.
Fan, G., Wang, X.L., Zhu, H.P., 2009. NERI Index of Marketization of China's Provinces. Economics Science Press, Beijing.
Fang, L., Peress, J., 2009. Media coverage and the cross-section of stock returns. J. Finance 64, 2023–2052.
Fernandes, N., Ferreira, M., 2009. Insider trading laws and stock price informativeness. Rev. Financ. Stud. 22 (5), 1845–1887.
Ferreira, M., Ferreira, M.A., Raposo, C.C., 2011. Board structure and price informativeness. J. Financ. Econ. 99, 523–545.
Ferreira, M., Laux, P., 2007. Corporate governance, idiosyncratic risk and information flow. J. Finance 62, 951–989.
Gul, F.A., Kim, J.-B., Qiu, A., 2010. Ownership concentration, foreign shareholding, audit quality, and stock price synchronicity: evidence from China. J. Financ. Econ. 95, 425–442.
Haw, I.-M., Hu, B., Lee, J.J., Wu, W., 2012. Investor protection and price informativeness about future earnings: international evidence. Rev. Account. Stud. 17, 389–419.
Hasbrouck, J., 1993. Assessing the quality of a security market: a new approach to transaction-cost measurement. Rev. Financ. Stud. 6, 191–212.
Hillert, A., Jacobs, H., Muller, S., 2012. Media Makes Momentum. University of Mannheim Working Paper.
Huang, C., 2001. China's state-run tabloids – the rise of city newspapers. Gazette 63, 435–450.
Huang, C., 2007. Trace the stones in crossing the river: media structural changes in post-WTO China. Gazette 69, 413–430.
Huang, Y., 2003. Selling China. Cambridge University Press, Cambridge.
Hutton, A.P., Marcus, A.J., Tehranian, H., 2009. Opaque financial reports, R^2, and crash risk. J. Financ. Econ. 94, 67–86.
Jiang, G., Lee, C.M.C., Yue, H., 2010. Tunneling through intercorporate loans: the China experience. J. Financ. Econ. 98, 1–20.
Jin, L., Myers, S.C., 2006. R^2 round the world: new theory and new tests. J. Financ. Econ. 79, 257–292.
Keatley, R.L., 2003. The Role of the Media in a Market Economy. National Committee on U.S.–China Relations China Policy Series 1, pp. 1–20.
Kim, J.-B., Shi, H., 2012. IFRS reporting, firm-specific information flow, and institutional environments: international evidence. Rev. Account. Stud. 17 (3), 475–517.
Klibanoff, P., Lamont, O., Wizman, T.A., 1998. Investor reaction to salient news in closed-end country funds. J. Finance 53, 673–699.
La Porta, R., Lopez-de-Silanes, F., Shleifer, A., Vishny, R., 1998. Law and finance. J. Polit. Econ. 106, 1113–1155.
La Porta, R., Lopez-de-Silanes, F., Shleifer, A., Vishny, R., 2000. Investor protection and corporate governance. J. Financ. Econ. 58, 3–27.
La Porta, R., Lopez-de-Silanes, F., Pop-Eleches, C., Shleifer, A., 2004. Judicial checks and balances. J. Polit. Econ. 112, 445–470.
Leuz, C., Nanda, D., Wysocki, P.D., 2003. Earnings management and investor protection: an international comparison. J. Financ. Econ. 69, 505–527.
Liebman, B.L., 2005. Watchdog or demagogue? The media in the Chinese legal system. Columbia Law Rev. 105, 1–157.
Lin, F., 2006. Dancing beautifully but with hands cuffed? A historical review of journalism formation during media commercialization in China. Perspectives 7, 79–98.
Lin, F., 2008. Turning Gray: Transition of Political Communication in China, 1978–2008. PhD Dissertation, Department of Sociology, University of Chicago, Chicago.
Merton, R.C., 1987. A simple model of capital market equilibrium with incomplete information. J. Finance 42, 483–510.
Miller, G.S., 2006. The press as a watchdog for accounting fraud. J. Account. Res. 44, 1001–1033.
Morck, R., Yeung, B., Yu, W., 2000. The information content of stock markets: why do emerging markets have synchronous stock price movements? J. Financ. Econ. 58, 215–260.
Pistor, K., Xu, C., 2005. Governing stock market in transition economics: lessons from China. Am. Law Econ. Rev. 7, 184–210.
Piotroski, J.D., Roulstone, D.T., 2004. The influence of analysts, institutional investors, and insiders on the incorporation of market, industry and firm-specific information into share prices. Account. Rev. 79, 1119–1151.
Rosenbaum, P.R., Rubin, Donald B., 1984. Reducing bias in observational studies using subclassification on the propensity score. J. Am. Stat. Assoc. 79, 516–524.
Srinidhi, B., Zhang, H., Zhang, T., 2012. Auditor Choice and Earnings Quality in Politically Connected Firms: Evidence Non-state-controlled Firms in China, Working Paper, City University of Hong Kong.
Tobin, J., 1982. On the efficiency of the financial system. Lloyd's Bank. Rev. 153, 1–15.
Tetlock, P.C., Saar-Tsechansky, M., Macskassy, S., 2008. More than words: quantifying language to measure firms' fundamentals. J. Finance 63, 1437–1467.
Vega, C., 2006. Stock price reaction to public and private information. J. Financ. Econ. 82, 103–133.

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Troubled by unequal pay rather than low pay: The incentive effects of a top management team pay gap[☆,☆☆]

Yue Xu [a,1], Yunguo Liu [a,*], Gerald J. Lobo [b]

[a] School of Business, Sun Yat-sen University, China
[b] C.T. Bauer College of Business, University of Houston, United States

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ABSTRACT

We examine the relationships with firm performance of the internal pay gap among individual members of the top management team (TMT) and the compensation level of TMT members relative to their industry peers. We find that pay gap is positively related to firm performance and that this positive relation is stronger when the TMT pay level is higher than the industry median. However, we do not observe such effects in Chinese state-owned enterprises (SOEs), in which both the executive managerial market and compensation are government-regulated. We also document that cutting central SOE managers' pay level can increase firm value, whereas doing so for local SOE managers has the opposite effect. Our findings have important implications for research on TMT compensation as well as for policy makers considering SOE compensation reform.

1. Introduction

Top management team (TMT) incentive-based contract design is a topic of considerable interest to academics and practitioners. It is an especially important issue in China because both the level of TMT

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* Corresponding author. Mobile: +86 186 6608 0107.
E-mail addresses: xinyan1991@126.com (Y. Xu), sysulyg@gmail.com (Y. Liu), gjlobo@uh.edu (G.J. Lobo).
[1] Mobile: +86 135 7030 8719.

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compensation has increased considerably and the compensation differentials among members of the TMT have widened significantly following the introduction of market-oriented reforms in China in 1978. Our focus in this study is on implications of the variation and the level of TMT pay for firm performance.

We define the variation in pay among the TMT as the difference between the CEO's pay and that of the other executives in the TMT, and refer to this difference as "TMT pay gap." TMT pay gap thus concerns the reward that non-CEO executives can expect if they are promoted to CEO. Prior research makes conflicting predictions about whether a large TMT pay gap promotes competition among TMT members and whether it enhances firm performance. Tournament theory, on the one hand, posits that a pay gap among different organizational levels provides a competition incentive such that the larger the pay gap, the better the firm's performance. Social comparison theory, on the other hand, stresses teamwork and cooperation within the TMT, and thus posits that a smaller pay gap can improve satisfaction and willingness to cooperate, thereby boosting firm performance. However, whether a team engages in competition or cooperation is an empirical issue (Harbring and Irlenbusch, 2003), and this relationship is moderated by many factors, including task dependence and the individual incentive system (Shaw et al., 2002; Kepes et al., 2009).

In addition to TMT pay gap, the level of TMT pay is also an important factor that affects firm performance. We define "TMT pay level" as the difference between the average pay of the TMT and the average pay of industry peer TMTs (Gerhart and Milkovich, 1992), and refer to this difference as "TMT pay level." TMT pay level thus reflects the external competitiveness of the firm's compensation policy (He and Hao, 2014). TMT pay level may have both a direct and an indirect effect on firm performance. First, as an important factor in the TMT's incentive system, TMT pay level has direct implications for firm performance. Second, TMT pay level may also have an indirect effect on firm performance because it moderates the relationship between TMT pay gap and firm performance. Excluding the effects of pay level from the model would thus result in a biased estimate of the relationship between TMT pay gap and firm performance.[2] However, most of the related research in China does not consider this interactive effect of TMT pay level (Lin et al., 2003; Chen and Zhang, 2006). Additionally, firms often use industry peers as a benchmark when negotiating contracts with top executives (Jiang, 2011). Different TMT pay levels lead to an external comparison between firms, and top executives and then form corresponding levels of satisfaction with their compensation, which in turn influences the competition–cooperation relationship within the TMT.

As indicated earlier, we conduct our empirical analysis using compensation data from Chinese firms. The market-oriented reforms introduced in China in 1978 have led to substantial increases in the compensation levels of some executives, especially those at monopoly and public welfare firms. Further, the pay gap among Chinese firms' TMTs has considerably widened during this period (Zhang, 2008; Li and Hu, 2012). Concerned about the widening pay gap and increasing compensation level, China's Central Political Bureau passed a resolution on 29 August 2014 to reform the pay system for the responsible persons of centrally managed companies. The program focuses on five main areas: (1) improvement of the reward system, (2) adjustment of the pay structure, (3) strengthening of supervision, (4) regulation of the pay level, and (5) treatment standardization. The last two areas, in particular, are intended to address unreasonably high incomes and pay gaps to promote social justice.

"Pay gap" can refer either to the income gap between executives and general staff or the gap between TMT members' pay. Chinese are generally more sensitive to the former gap, particularly since the round of pay cuts and layoffs in 2008 that saw executives retain high pay levels (Liu and Sun, 2010). However, because the TMT is at the highest managerial level of the firm, the within-team pay gap is related to the distribution of limited compensation among executives, and thus plays an important role in the TMT incentive system. Moreover, if the overall TMT pay level is adjusted, the question is whether and how income should be distributed among team members to ensure the effectiveness of the compensation incentive mechanism. To answer this question, we explore the relationship between a TMT's overall pay level and the pay gap among its members.

Since 2005, it has been mandatory for China's listed firms to disclose their executives' compensation. In this study, we examine whether TMT pay level affects the relationship between a TMT pay gap and firm

[2] For example, Knoeber and Thurman (1994) point out that Ehrenberg and Bognanno (1990a,b) ignore the incentive effect of prize level when using the behavior of professional golfers to examine tournament theory. Because the prize structure was identical across tournaments, larger prize gaps were always the result of higher prize levels.

performance based on executive compensation data from 2005 to 2012. Moreover, in the context of the current reform of China's state-owned enterprises (SOEs), we further analyze the relationship between TMT pay level and pay gap incentive efficiency under different ownership conditions. We find that TMT pay level can moderate pay gap incentive efficiency and that a pay gap within a TMT exerts stronger tournament effects when TMT members' pay level is higher than that of their industry peers. In Chinese SOEs, however, because both the market for executives and their compensation are regulated by the government, even a high pay level fails to stimulate tournament effects in the TMT. In central SOEs, cutting managers' pay level can increase the value of the firm. However, doing so in local SOEs is likely to dampen tournament incentive efficiency. For local SOEs that provide high TMT pay level, a larger TMT pay gap is associated with better firm performance.

This study makes several contributions. First, it enriches the literature on TMT pay incentives. These incentives comprise the entire team's pay level (the first distribution) and its internal pay gap (the second distribution) (Zhang et al., 2012). Many studies discuss these two aspects of incentives separately, but few combine them. Our findings show that the TMT pay level influences the team's internal distribution efficiency. When the pay level is higher than that of industry peers, a larger pay gap induces better tournament incentives and performance. However, when the pay level is lower than that of industry peers, a larger pay gap may be harmful to firm performance.

Second, our study adds to the literature on pay gaps. Tournament theory and social comparison theory do not coincide in terms of their predictions of the pay gap incentive effect. We argue that the applicability of the two theories depends on external equity. When top executives perceive external fairness, tournament theory is more applicable; when they perceive less external fairness, social comparison theory is more applicable. This finding also supplements Brown et al. (2003).

Third, our study contributes to the literature on Chinese SOE reform. Wu et al. (2010) and Li et al. (2014) report that excess compensation does not have incentive effects. Similarly, our findings show that excess pay does not promote internal competition in the TMT. For central SOEs, reducing top managers' pay level can stimulate managerial competition and enhance firm value, which lends support to the recently passed and implemented decision to reduce the compensation of central SOEs' top executives. At the same time, our results suggest that the reform policy for local SOEs should not simply follow than that of central SOEs. Reducing pay levels may not be the optimal decision for local SOEs, although adjusting the internal pay structure may enhance firm performance.

The remainder of the paper is organized as follows. Section 2 presents the literature review and hypothesis development, and Section 3 discusses the research design. Section 4 reports the descriptive statistics and the results of the empirical analyses, while Section 5 discusses the results of robustness tests. Section 6 presents the results of additional analyses and Section 7 concludes the paper.

2. Literature review and hypothesis development

2.1. Literature on pay gap and pay level

There is a sizeable body of research on the relationship between TMT pay gap and firm performance. One strand of research is tournament theory, developed by Lazear and Rosen (1981), which posits that although a non-CEO's salary may double within a day following the promotion to CEO, it would be difficult to argue that his/her ability has also doubled within a day. Thus, it is difficult to explain the pay gap in TMTs through recourse to traditional economic theory, which argues that pay levels are determined by marginal output (e.g., Lin et al., 2003). Tournament theory argues that under the conditions of cooperative effort and task interdependence, it is not feasible to set executives' pay based on their marginal output when monitoring is difficult. Although marginal output-based pay for executives might seem more equitable, a large pay gap can encourage rank-order competition, thereby improving firm performance. Numerous studies provide empirical evidence consistent with the predictions of tournament theory. For example, Ehrenberg and Bognanno (1990a, 1990b) study the behavior of professional golf players, and find that increases in the differential between prizes provide them with an incentive to exert more effort. Main et al. (1993) report a positive relationship between pay gap and return on assets for a sample of more than 200 firms and 2000 executives per year over a five-year period. Rosen (1986), Eriksson (1999), Lambert et al. (1993) and Kale et al. (2009) also provide theoretical and empirical evidence of the theory's efficacy.

Another strand of research is social comparison theory, which, in contrast to tournament theory, argues that people pay close attention to equity in most situations. Particularly when a team is working jointly toward the same objective, team members are likely to not only being concerned about their own income, but also to compare it with those of their fellow team members to judge whether the income distribution is fair. If a TMT's internal pay gap is sufficiently large to seem inequitable to non-CEO executives, dissatisfaction with the income distribution may reduce their willingness to cooperate, thus damaging firm performance. Therefore, social comparison theory places greater stress on the importance of a compressed pay distribution in TMTs (Lazear, 1989; Pfeffer and Langton, 1993; Cowherd and Levine, 1992). O'Reilly et al. (1988) suggest that a compressed pay distribution and smaller pay gap encourage collaboration among employees, and are thus beneficial to firm performance. Drago and Garvey (1998), Bloom (1999), Hibbs and Locking (2000) and Fredrickson et al. (2010) also provide empirical evidence in support of social comparison theory.

The two theories make contradictory predictions about the relationship between a TMT pay gap and firm performance. Tournament theory argues that a larger TMT pay gap provides an incentive for executive competition, whereas social comparison theory posits that a larger TMT pay gap causes non-CEO executives to feel deprived and is not conducive to cooperation. In reality, most members of organizations, particularly top executives, work toward a common objective and require interdependence or collaboration. Cooperation and competition coexist among executives in a TMT, and we thus need empirical evidence to determine which theory is dominant in a specific setting. For example, Lin et al. (2003) propose that in Chinese listed firms, tournament incentives' positive effect exceeds the negative effect brought about by feelings of unfairness and non-cooperative behavior. They thus conclude that tournament theory is more suitable for explaining the relationship between TMT pay gaps and firm performance in Chinese firms.

However, the literature in this area is not limited to studies seeking supportive empirical evidence for these two theories. Scholars have discussed the aforementioned relationship in a variety of contexts (Trevor and Wazeter, 2006). Siegel and Hambrick (2005) examine the interactive effect of technological intensiveness and TMT pay gaps on firm performance, finding that in high-technology firms that require greater technological intensiveness, pay gaps are detrimental to firm performance. Kepes et al. (2009) report a positive relationship between pay gaps and firm performance when those gaps are attributable to the use of performance-based pay because employees feel that the distribution process is fair. Brown et al. (2003) use a large database of hospitals to examine the modulating effect of an organization's pay level on internal pay gap incentive efficiency.

As noted, "pay level" in this paper refers to top executives' average pay level relative to that of their industry peers (Gerhart and Milkovich, 1992). Milkovich and Newman (2002) describe pay levels as leading, matching or lagging the market. The traditional economic literature primarily uses efficiency wage theory, proposed by Akerlof and Yellen (1986), to explain why a firm's pay level affects its performance. If the firm's pay level exceeds that of its peers, it will find it easier to attract, retain and motivate outstanding talent. The Chinese literature explains the influence of pay level on firm performance primarily from the perspective of external equity. Wu et al. (2010) estimate the excess compensation of top managers as a measure of external equity, and find that high pay levels motivate non-SOE managers, but have no effect on SOE top managers. Qi and Zou (2014) use relative quantiles of top executives' average pay as a measure of external pay equity, reporting that a higher pay level always increases managers' perceptions of fairness, and thus motivates them to exert greater effort to boost firm performance. From the perspective of the managerial market, Li et al. (2014) demonstrate that only when a TMT's pay level leads the market does excess compensation for top managers provide positive incentives. Therefore, the TMT's pay level offers the same incentives to top managers as an internal TMT pay gap. Accordingly, in order to focus on the incentives of the latter, we must exclude the direct and indirect effects of pay level. In other words, a study on the effects of pay gaps on firm performance must control for pay level.

The Chinese literature on pay gaps considers both the gap between management and workers and that among managers within TMTs. The Chinese tend to be more sensitive to the former because of the traditional pull of the harmonious society principle, and thus the pay distribution within TMTs has received little attention. Interestingly, however, almost all existing papers on the consequences of an internal TMT pay gap are consistent in supporting tournament theory (Lin et al., 2003; Chen and Zhang, 2006), although they are somewhat limited by their failure to control for pay level. Zhang (2007) examines the relationship between within-TMT pay gaps and firm performance while holding pay level constant, and finds a negative relationship, which is consistent with the prediction of social comparison theory.[3] This result suggests that, for Chinese listed firms, the TMT pay level may be an influential factor in TMT pay gap efficiency.

At present, the Chinese literature on TMT pay level and pay gap is fragmented. Many studies discuss one or the other in isolation, but few consider them together. In a field survey of 376 general workers from 59 departments of Chinese firms, He and Hao (2014) find that within-department pay differences exert a negative effect on workers' emotional commitment only when the overall departmental pay level is lower than that of other departments. In this study, we combine TMT pay levels with a study of TMT pay gap efficiency, in an attempt to contribute to research on top management compensation mechanisms in Chinese listed firms.

2.2. Hypothesis development

As previously discussed, tournament theory emphasizes competition within the organization, whereas social comparison theory stresses cooperation and fairness within the organization. Colquitt et al. (2001) suggest that perceived equity is the result of a subjective judgment process and that subjectivity is reflected primarily in the choice of reference object. Many studies propose that employees compare their salaries with several reference objects (Brown, 2001; Hills, 1980; Law and Wong, 1998). Oldman et al. (1986) divide the comparison with reference objects into internal organizational justice and external organizational justice. A within-TMT pay gap influences internal organizational justice, whereas TMT pay level influences external organizational justice. These two kinds of justice are not independent, but rather interact with each other (Trevor and Wazeter, 2006). It has been mandatory for Chinese listed firms to disclose their executives' compensation since 2005, which makes it easier for top executives to obtain payment information on their peers and to engage in external comparison.

Lazear and Rosen (1981) propose a two-player tournament model in which the players' effort and performance levels are positively related to the size of the reward, but have nothing to do with the prize level. Their model assumes the prize level to be exogenously given, which limits the possibility of players making external comparisons. However, this assumption does not hold for workers' wages in most firms. Efficiency wage theory (Akerlof and Yellen, 1986) states that providing employees with a market-leading wage is beneficial to the firm because employees are more willing to remain in a high-paying organization. Messersmith et al. (2011) find that top executive turnover is less likely when executives receive a higher proportion of overall TMT compensation. Not only is original talent retained, but other outstanding managers are also attracted to the team. Therefore, when there is a within-TMT pay gap, a high TMT pay level exerts a natural incentive effect on the CEO and, more importantly, the negative effect on non-CEOs arising from the internal pay gap is partially offset. Bloom and Michel (2002) suggest that if a firm's payment level is higher than the market's, many of the negative consequences of an internal pay gap are alleviated. Frank (1985) proposes that it is easier for employees to accept an unfair wage distribution when their wages are higher than their marginal output. Trevor and Wazeter (2006) find that when employees are in a lower pay position internally, a higher pay position externally enhances their perception of pay equity. Hence, an internal pay gap is less likely to exert negative effects and more likely to create tournament incentives.

However, if the TMT's pay level lags the market, all of the team's top executives are likely to perceive a low degree of external equity, that is, to perceive themselves at a disadvantage relative to their industry peers. Although a large internal pay gap may somewhat alleviate the CEO's dissatisfaction, it will worsen that of non-CEOs, who are in a poorer pay position relative to both internal and external referents. Such double discontent with their compensation may well reduce their willingness to cooperate (Deutsch, 1985; Pfeffer and Langton, 1993) or even encourage them to desert the firm (Bloom and Michel, 2002). Both outcomes

[3] Chen et al. (2011) also control for the average pay level of the three top executives other than the CEO in their model. Their results are consistent with tournament theory. However, even when these top three executives' pay is controlled for, a larger pay gap always accompanies higher CEO pay, and thus it is still difficult to distinguish whether the positive relationship between a pay gap and firm performance stems from high CEO pay or a large pay gap.

are potentially damaging to firm performance. Moreover, if the managerial market is effective, those who feel underpaid will move to other companies, where they can obtain ability-matched compensation, and those who choose to remain on a poorly paid team may be incompetent and thus struggle to find a better paid job. Messersmith et al. (2011) provide evidence to show that reducing overall TMT compensation increases executive turnover. Therefore, a low TMT pay level and large within-team pay gap induce a number of negative effects, such as a lack of cooperation and higher turnover among non-CEOs, thereby offsetting the positive tournament incentives of the pay gap to a large degree. It is possible that the negative effect of an internal pay gap is dominant when overall pay is below a certain level.

The preceding discussion suggests that a TMT pay gap can produce positive tournament incentives, especially when the TMT pay level is above than that of peer firms. It also suggests that a TMT pay gap can produce negative effects, especially when external equity perceptions brought about by comparison with other TMT pay levels influence internal equity perceptions induced by the pay gap. Although it is unclear whether the positive tournament effects or the negative effects of internal inequities dominate, it is clear from the above discussion that the positive tournament effects of the TMT pay gap will be more positive when the TMT pay level is above than that of peer firms, and the negative effects are more negative when the TMT pay level is below than that of peer firms. This implies that the difference in the relationship between the TMT pay gap and performance will be positive for higher versus lower TMT pay levels. Therefore, we formulate our first hypothesis as follows:

H1. The relationship between TMT pay gap and firm performance is likely to be more positive when TMT pay level is higher than that of peer firms.

Although there is considerable evidence supporting the predictions of tournament theory, its efficacy in China is unclear because of the unique features of China's institutional environment that considerably differ from those of mature market economies. For example, China has many SOEs whose top managers earn higher average pay levels than their non-SOE counterparts, and the TMT pay gap is significantly lower in SOEs than in non-SOEs. The primary reason for this discrepancy is that the compensation of SOE top managers is regulated by the Chinese government.

Given these differences, a question of interest is whether tournament theory is capable of explaining the incentives of top managers in Chinese SOEs. Chinese scholars (e.g., Zhou and Zhu, 2010) suggest that the tournament incentive mechanism is actually encouraged by China's institutional environment. Because the government is the ownership representative of the people, it has a natural information disadvantage. Also, SOE managers are multitaskers. It is thus difficult for the government to find appropriate measures to evaluate SOE managers, and tournament promotion based on relative performance is a widely used measure. Moreover, competence for the CEO position is also judged by political promotion, as most SOE managers are appointed by the government and also have an administrative ranking within the government. Although China has attempted in recent years to implement non-administrative SOE reform, many top managers still treat promotion within the SOE as a way to further their political careers. Yang et al. (2013) study the "quasi-official" promotion mechanism in central SOEs. They assess whether internal promotion within an SOE's TMT is more political than that within a non-SOE's TMT and whether the CEOs of SOEs are able to obtain more political profits,[4] such as political rent-seeking opportunities and transfers to government departments. Therefore, even a very small within-TMT pay gap in an SOE can create tournament incentives and motivate managers.

Of course, under China's tradition of egalitarian thought, the requirement for a fair income distribution may offset the positive effects of such a pay gap to some extent. However, we are not concerned with predicting the direction of a pay gap's net effects, but rather with determining how the TMT pay level, relative to the market or to industry peers, affects TMT pay gap incentives in firms with different ownership types.

First, the compensation that SOE top managers receive is government-regulated and separated from market conditions. Thus, their disclosed pay level is not necessarily indicative of the competitiveness of their

[4] It should be noted that political profits here do not equate to non-monetary income. Non-monetary income refers to such benefits as reputation and status after promotion to CEO, whereas political profits refer specifically to the benefits to the executive's post-promotion political career.

payment capacity. Second, because monetary compensation is regulated, managerial perks have become an alternative means of compensating SOE managers (Chen et al., 2005). Thus, these managers' disclosed compensation does not fully reflect their incentives. They have more implicit incentives that are not sensitive to pay comparisons with peers. For example, Li et al. (2014) suggest that even if SOE managers obtain a higher pay level than their industry peers, excess payment does not bring any added incentives. However, if SOE managers, most of whom are appointed by the government, obtain a lower pay level than their peers, even if they also have an external perception of unfairness, they are unlikely to slack off at work or desert their positions because they are concerned about their careers in the state-owned system. In summary, the TMT pay level in SOEs has little effect on the input and output of top managers.

As previously noted, the premise that the TMT pay level moderates the effects of a TMT pay gap on firm performance is based on the assumption that top managers compare themselves with external referents, thereby exacerbating the perception of internal injustice induced by the pay gap. In non-SOEs, compensation information on top managers is relatively more transparent and focused on monetary incentives (Lu et al., 2012), and the median values of pay in the same industry are more often used as a benchmark (Jiang, 2011). Hence, non-SOE managers are more concerned with peer comparisons than their SOE counterparts. Moreover, because of the relatively high marketization level of non-SOE managers, professional managers can flow freely in the market, and can thus choose job-hopping, slacking off at work or deserting their position in response to a pay gap-induced perception of injustice, which in turn damages firm performance. SOE managers, in contrast, are not sensitive to the competitiveness of their pay level, and thus external compensation comparisons have little effect on internal TMT competition and cooperation. This discussion leads to our second hypothesis:

H2. The effect of TMT pay level on the relationship between TMT pay gap and firm performance will be more pronounced for non-SOEs than for SOEs.

3. Research design

3.1. Data and sample selection

Executive compensation data and other firm characteristic data are sourced from the China Stock Market & Accounting Research (CSMAR) database. It has been mandatory for Chinese listed firms to disclose their executives' compensation details since 2005. Accordingly, our sample period spans 2005–2012, and the sample contains all Chinese listed firms with available data in that period. We define TMT members as all those who occupy a management position. Members of boards of directors and supervisors, who receive only a fixed bonus, are excluded. The top managers in our sample include general managers, deputy general managers, chief engineers, chief accountants, and chief financial officers. The general manager, president or CEO of the firm are defined as the CEO[5] (Liao et al., 2009), and all other managers as non-CEOs. After collecting executive compensation data from the CSMAR database, we screened the data manually according to position titles. The screening steps were as follows: (1) obtain the *comprehensive management file* and the *executive individual compensation file* from the CSMAR database and check the CEO compensation data; (2) drop observations if the CEO changed in the current year[6]; and (3) drop observations if CEO compensation was not disclosed or the CEO received only a fixed bonus. After screening, we had 11,589 valid CEO observations. We then screened non-CEO compensation data from the *executive individual compensation file* by the aforementioned position titles.

After removing (1) 163 financial companies, (2) 566 companies for which we could not calculate the TMT pay gap because CEO compensation was lower than the median compensation for non-CEOs,[7] (3) 193

[5] The titles of the chief managers in Chinese listed firms are not consistent, with some called general manager and some called president. In this paper, we uniformly refer to them as CEOs.

[6] If the CEO changed in the current year, then the reported CEO compensation in that year did not constitute data for a whole year, possibly leading to large deviations in calculating pay gap.

[7] As in Liao et al. (2009), this condition might arise if the CEO receives only part of his or her compensation from this firm or if specialists who are entitled to higher compensation are hired as non-CEOs.

observations whose debt to asset ratio was greater than 1, (4) 1197 observations with missing values,[8] and (5) 284 observations that received special treatment, we were left with 9186 valid firm-level observations for our empirical analysis.

3.2. Variable definitions and model specification

In accordance with the foregoing theoretical analysis and relevant research (Kale et al., 2009; Chen et al., 2011; Zhang, 2007), we use the following model (1) to examine how a firm's TMT pay level influences the relation between an internal TMT pay gap on firm performance:

$$\begin{aligned} PERF_t = {} & \beta_0 + \beta_1 GAP_t + \beta_2 PL_t + \beta_3 GAP_t \times PL_t + \beta_4 SOE_t + \beta_5 LEV_t + \beta_6 SGROW_t + \beta_7 SIZE_t \\ & + \beta_8 TOP1_t + \beta_9 CEOAge_t + \beta_{10} BOARD_t + \beta_{11} INDEPB_t + \beta_{12} DUAL_t + \beta_{13} MKR_t \\ & + \beta_{14} Option_t + \sum Industry_t + \sum Year_t + \varepsilon_t \end{aligned} \tag{1}$$

where *PERF* is either a measure of accounting performance (*ROA*) or market performance (*TOBINQ*).

GAP refers to the TMT pay gap, our main variable of interest. Following Bognanno (2001), Kale et al. (2009), Chen et al. (2011) and Kini and Williams (2012), we define *GAP* as the reward received after a non-CEO executive is promoted to CEO. We use two measures of *GAP*. The first measure, $GAP\_1 = ln$ (*CEO pay–median pay of non-CEOs*),[9] where "ln" refers to the natural logarithm, and the second measure, $GAP\_2 =$ *CEO pay/median pay of non-CEOs*, in accordance with Lin et al. (2003), Zhang (2008) and Hambrick and Siegel (1997).

PL measures the TMT average pay level relative to the median of the average pay levels of all other firms in the same industry. It represents the external competitiveness of the TMT's pay. *PL* equals 1 if the firm's average TMT pay is higher than the yearly median level of firms in the same industry, and 0 otherwise.

The other control variables, including firm characteristics and firm governance, are defined as in the literature (Chen et al., 2011; Lin and Lu, 2009). For example, an *SOE* (ownership status) value of 1 indicates that the firm is state-owned, whereas an *SOE* value of 0 indicates it is not. Xu et al. (2006) suggest that whether the controlling shareholder is state-owned or non-state-owned has a significant influence on firm performance. The liability-asset ratio (*LEV*) is equal to total liabilities divided by total assets. Myers (1977) and Jensen (1986) both find *LEV* to affect firm performance, although the former reports a negative effect and the latter a positive effect. Sales growth (*SGROW*) represents the firm's rate of growth, and is equal to the difference between current-year sales and previous-year sales divided by previous-year sales. The other control variables are defined in Table 1. We also control for industry and year fixed effects.

Endogeneity is a serious concern when studying the relationship between pay gap and firm performance. Li and Hu (2012) suggest that there is a natural positive relationship between these two variables in Chinese SOEs. Kale et al. (2009), Chen et al. (2011), Lin and Lu (2009) and Kini and Williams (2012) all adopt instrumental variable (IV) estimation or two-stage least-squares (2SLS) estimation to control for the endogenous relationship between pay gap and firm performance. We choose lagged TMT pay gap (*LGAP*) and median TMT pay gap in the same industry (*MedianGAP*) as the IVs for *GAP*. Kale et al. (2009) suggest that TMT pay gap is positively associated with the median *GAP* of industry peers.

There is also concern about endogeneity in the relationship between TMT pay level and firm performance, as firms with a higher pay level are likely to have better performance. To address this concern, we follow Fang (2012) and Wu et al. (2010) and estimate a TMT's excess pay level after removing the effects of firm performance, firm characteristics and other factors. We use the following model (2) to estimate the excess pay level (Core et al., 1999; Fang, 2012).

[8] Missing-value observations stem primarily from our need for data from two consecutive years. If a firm was listed in the sample period, the listing year is that firm's first year, and no data are available for the previous year.

[9] The results are similar when we use the mean of non-CEO pay, instead of the median.

Table 1
Variable definitions.

Name	Variable	Definition
Firm Performance	*ROA*	Net income/ending total assets
	TOBINQ	(Market value of tradable shares + net assets + market value of net debt)/ending total assets
TMT Pay Gap	*GAP_1*	ln (CEO pay–median pay of non-CEOs)
	GAP_2	CEO pay/median pay of non-CEOs
	LGAP	Lagged GAP (GAP_1 or GAP_2)
	MedianGAP	Median GAP (GAP_1 or GAP_2) value of each industry in each year
TMT Pay Level	*PL*	Dummy variable that equals 1 if firm *i*'s average TMT pay is higher than the median TMT pay of firms in the same industry each year, and 0 otherwise
	UF	Excess pay of the TMT, estimated as the residuals of model (2)
Firm/Executive Characteristics	*SOE*	Ownership status, a dummy variable that equals 1 if the firm is state-owned, and 0 otherwise
	LEV	Total debt/total assets
	SGROW	Sales growth rate = (sales revenue of current year − sales revenue of previous year)/ sales revenue of previous year
	SIZE	ln (ending total assets)
	CEOAge	ln (CEO age)
	Option	Dummy variable that equals 1 if the top executives have been granted stock options, and 0 otherwise
Firm Governance	*TOP1*	Percentage of outstanding shares held by the firm's largest shareholder
	BOARD	Number of directors on the firm's board
	OUT	Number of independent directors/number of directors of the board
	DUAL	Dummy variable that equals 1 if the CEO is also the chairperson, and 0 otherwise
	MKR	Based on the Chinese regional market index estimated by Fan and Wang (2007), *MKR* is a dummy variable that equals 1 if the firm is in a region with a market index higher than the median value, and 0 otherwise

$$\begin{aligned} LnCOMP_t = {} & \beta_0 + \beta_1 SIZE + \beta_2 LEV_t + \beta_3 ROA_1 + \beta_4 ROA_{t-1} + \beta_5 DUAL_t + \beta_6 BOARD_t + \beta_7 SOE_t \\ & + \beta_8 MHOLD_t + \beta_9 BM_t + \varepsilon_t \end{aligned} \tag{2}$$

where $LnCOMP_t$ is the natural logarithm of the TMT's average pay, ROA_{t-1} is lagged firm performance, *MHOLD* is the TMT's average shareholding ratio, *BM* is the ratio of book value to market value of equity, and the other variables are as defined in Table 1.

We first estimate model (2) for each industry-year and use the residuals from this model as our estimate of excess pay of the TMT. We also use this excess pay measure as the IV estimator for TMT pay level.

Returning to model (1), the interaction between TMT pay level and TMT pay gap may also suffer from an endogeneity problem. If we use the IVs for *GAP* and *PL*, the interaction term $GAP \times PL$ should also use the interaction of these IVs. Therefore, we use four IVs, namely *LGAP*, *MedianGAP*, *UF* and $LGAP \times UF$, for the three endogenous variables, *GAP*, *PL* and $GAP \times PL$, in model (1). All of our IVs satisfy relevance and validity criteria.[10]

According to H1, if a higher TMT pay level is accompanied by better TMT pay gap-induced tournament incentives, the regression coefficient on $GAP \times PL$ should be positive and significant,[11] and, according to H2, that positive relationship should be stronger in the non-SOE subsample.

[10] These tests include unidentifiable inspection, the weak identification test and the Sargan test.

[11] Because a TMT pay gap may have positive or negative effects, we do not predict *GAP*'s direction in model (1) as Siegel and Hambrick (2005) do.

Table 2
Descriptive statistics for main variables.

Variable	Number of observations	Mean	Minimum	Median	Maximum	Standard deviation
ROA	9186	0.040	−0.389	0.037	0.208	0.057
TOBINQ	9186	1.727	0.714	1.370	7.222	1.083
PL_DIF	9186	175.098	36.825	149.563	564.233	112.872
PAY_GAP	9186	18.646	0.500	10.847	150.466	24.216
GAP_1	9186	11.552	8.476	11.594	14.366	1.123
GAP_2	9186	1.642	1.000	1.429	4.814	0.658
PL	9186	0.527	0	1	1	0.499
SOE	9186	0.561	0	1	1	0.496
LEV	9186	0.474	0.050	0.489	1.000	0.204
SGROW	9186	0.211	−0.778	0.144	4.317	0.480
SIZE	9186	21.667	18.698	21.522	25.620	1.180
CEOAge	9186	3.853	3.219	3.850	4.317	0.133
TOP1	9186	36.647	9.000	34.795	75.889	15.279
OUT	9186	0.362	0.091	0.333	0.800	0.052
BOARD	9186	9.192	3	9	18	1.856
DUAL	9186	0.185	0	0	1	0.388
MKR	9186	0.784	0	1	1	0.412
Option	9186	0.054	0	0	1	0.227

Note: *PL_DIF* is the range of TMT average pay in each industry, that is, the difference between the highest and lowest average pay. *PL_DIF* and *PAY_GAP* are all in ten thousands Yuan.

Table 3
Correlations between main variables.

	ROA	*TOBINQ*	*GAP_1*	*GAP_2*	*PL*	*SOE*
ROA	1					
TOBINQ	0.246***	1				
GAP_1	0.217***	0.022**	1			
GAP_2	0.027***	0.028***	0.597***	1		
PL	0.205***	−0.030***	0.466***	0.004	1	
SOE	−0.102***	−0.105***	−0.122***	−0.212***	0.097***	1

* Significance at the 0.1 level.
** Significance at the 0.05 level.
*** Significance at the 0.01 level.

4. Empirical results

4.1. Descriptive statistics

Table 2 presents descriptive statistics of the main variables in the 2005–2012 sample period. The mean (median) values of performance measures *ROA* and *TOBINQ* are 0.040 (0.037) and 1.727 (1.370), respectively, and those of *GAP_1* and *GAP_2* are 11.552 (11.594) and 1.642 (1.429), respectively. All of the variables are reasonably distributed without extreme observations. The unscaled values of TMT pay gap (*PAY_GAP*) indicate an average difference between CEO pay and median non-CEO pay of about 180,000 yuan. However, the difference varies widely across the sample, ranging from 5000 yuan to 1,500,000 yuan. The average *PL* value is 0.527, which means 52.7% of firms offer a TMT pay level that is higher than the median pay level in the same industry. The mean across industries in the range of average TMT pay in an industry (*PL_DIF*) is about 1,750,000 yuan, showing that even in the same industry, average TMT pay can differ widely. Table 2 also shows that 56.1% of the sample firms are SOEs, and that 18.5% of the firms have a CEO and chairperson who are the same individual (mean $DUAL = 0.185$). Further, 70% of firms operate in regions with a high rate of marketization, and only 5.4% grant stock options to their top executives (average *Option* value = 0.054).

Table 3 presents the correlation coefficients between the main variables. *GAP* is positively related to firm performance. The correlation coefficient between *GAP_1* and *GAP_2* is 0.597, significant at the 1% level,

Table 4
Univariate comparisons of SOE and non-SOE subsamples.

Variable	Non-SOE		SOE		Mean difference
	No. of observations	Mean	No. of observations	Mean	
Panel A					
CEOPAY	4031	518305.300	5155	522105.800	−3800.443
VPMPAY	4031	290198.800	5155	347235.200	−57037.430***
PL	4031	0.472	5155	0.570	−0.098***
GAP_1	4031	11.707	5155	11.431	0.277***
GAP_2	4031	1.800	5155	1.519	0.281***
ROA	4031	0.047	5155	0.035	0.012***
TOBINQ	4031	1.856	5155	1.626	0.230***
Option	4031	0.102	5155	0.017	0.086***

Variable	Non-SOE				Mean difference
	PL = 1		PL = 0		
	No. of observations	Mean	No. of observations	Mean	
Panel B					
GAP_1	2129	11.168	1902	12.312	−1.144***
GAP_2	2129	1.736	1902	1.871	−0.135***

Variable	SOE				Mean difference
	PL = 1		PL = 0		
	No. of observations	Mean	No. of observations	Mean	
Panel C					
GAP_1	2219	10.840	2936	11.878	−1.037***
GAP_2	2219	1.546	2936	1.498	0.048***

Note: *CEOPAY* is the pay of the CEO, and *VPMPAY* is the median pay of non-CEOs.
* Significance at the 0.1 level.
** Significance at the 0.05 level.
*** Significance at the 0.01 level.

which means that the two measures of pay gap are closely related. Further, TMT pay level *PL* is positively related to *GAP_1* and uncorrelated with *GAP_2*. *SOE* is negatively related to *GAP*.

Table 4 reports the results of univariate analysis for the different ownership subsamples. Panel A presents the comparison for the main variables. Average CEO pay does not differ between the SOE and non-SOE subsamples, although the average non-CEO pay of the non-SOE sample is significantly lower than that of the SOE sample. These results show clearly that the TMT pay gap is much greater in non-SOEs than in SOEs, and are also consistent with the view that TMTs in SOEs are more egalitarian. Further, SOE TMTs enjoy a higher pay level, but exhibit weaker firm performance, than their non-SOE counterparts. Panels B and C show the relationship between TMT pay gap and pay level in non-SOEs and SOEs, respectively. In non-SOEs, a higher TMT pay level is accompanied by a larger pay gap, whereas in SOEs, a higher TMT pay level is associated with a larger absolute pay gap (*GAP_1*) but a smaller relative pay gap (*GAP_2*).

4.2. Top management team pay level and pay gap

Table 5 presents the estimation results of model (1). Panels A and B show the regression results for accounting performance (*ROA*) and market performance (*TOBINQ*), respectively, with columns 4 and 8 reporting the IV regression results.[12] The results in column 1 of Panels A and B show that if TMT pay level (*PL*) is not controlled for, *GAP* is significantly positively related to firm performance at the 1% level, which is consistent

[12] We use the IVs stated above. We present only the second-stage IV regression results here.

Table 5
Relationship between firm performance and TMT pay level and TMT pay gap.

	ROA/GAP_1				ROA/GAP_2			
	OLS			IV	OLS			IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A								
Constant	−0.194***	−0.129***	−0.113***	−0.235***	−0.178***	−0.105***	−0.102***	−0.161***
	(−9.613)	(−6.230)	(−5.074)	(−7.518)	(−8.758)	(−5.158)	(−4.960)	(−6.859)
GAP	0.00666***	0.00374***	0.00232***	0.00723***	0.000225	0.0000444	−0.00202*	−0.00556**
	(13.020)	(6.517)	(2.760)	(4.225)	(0.273)	(0.055)	(−1.654)	(−2.271)
PL		0.0146***	−0.0166	−0.105***		0.0182***	0.0118***	−0.0142**
		(12.476)	(−1.338)	(−4.023)		(17.565)	(4.333)	(−2.173)
GAP × PL			0.00272**	0.00750***			0.00388**	0.00976***
			(2.531)	(3.333)			(2.449)	(2.599)
SOE	−0.00608***	−0.00732***	−0.00724***	−0.00321**	−0.00768***	−0.00837***	−0.00824***	−0.00733***
	(−4.833)	(−5.799)	(−5.736)	(−2.205)	(−5.956)	(−6.547)	(−6.451)	(−5.217)
LEV	−0.120***	−0.118***	−0.118***	−0.120***	−0.123***	−0.119***	−0.119***	−0.119***
	(−29.418)	(−28.960)	(−29.067)	(−34.225)	(−30.038)	(−29.138)	(−29.160)	(−35.232)
SGROW	0.0227***	0.0225***	0.0226***	0.0235***	0.0230***	0.0227***	0.0227***	0.0230***
	(14.495)	(14.479)	(14.510)	(17.650)	(14.406)	(14.431)	(14.416)	(18.040)
SIZE	0.00877***	0.00721***	0.00712***	0.00991***	0.0109***	0.00778***	0.00774***	0.0104***
	(12.073)	(9.825)	(9.656)	(13.587)	(15.216)	(10.635)	(10.577)	(14.776)
TOP1	0.000209***	0.000211***	0.000215***	0.000222***	0.000172***	0.000195***	0.000198***	0.000186***
	(5.695)	(5.806)	(5.900)	(5.264)	(4.656)	(5.350)	(5.436)	(4.554)
CEOAge	0.00146	−0.000143	−0.0000292	0.00489	0.00464	0.000919	0.000846	0.00648
	(0.370)	(−0.036)	(−0.007)	(1.030)	(1.158)	(0.234)	(0.215)	(1.407)
INDEPB	−0.0328***	−0.0350***	−0.0341***	−0.0318***	−0.0335***	−0.0359***	−0.0352***	−0.0350***
	(−2.993)	(−3.214)	(−3.135)	(−2.629)	(−3.051)	(−3.297)	(−3.238)	(−3.002)
BOARD	0.000310	0.0000650	0.0000775	0.000457	0.000321	0.00000914	0.0000199	0.000193
	(1.016)	(0.215)	(0.256)	(1.221)	(1.041)	(0.030)	(0.066)	(0.534)
DUAL	−0.00248*	−0.00188	−0.00196	−0.00549***	−0.000844	−0.000986	−0.00100	−0.00216
	(−1.719)	(−1.318)	(−1.373)	(−3.463)	(−0.581)	(−0.690)	(−0.702)	(−1.416)
MKR	0.00430***	0.00270**	0.00284**	0.00642***	0.00630***	0.00323**	0.00324**	0.00585***
	(3.144)	(1.987)	(2.087)	(4.189)	(4.572)	(2.364)	(2.380)	(3.941)
Option	0.0145***	0.0130***	0.0130***	0.0161***	0.0150***	0.0129***	0.0128***	0.0144***
	(8.068)	(7.272)	(7.234)	(6.538)	(8.259)	(7.153)	(7.104)	(6.046)
N	9186	9186	9186	6624	9186	9186	9186	6784
Adj. *R*-sq	0.284	0.295	0.296	0.234	0.270	0.292	0.292	0.268

	TOBINQ/GAP_1				TOBINQ/GAP_2			
	OLS			IV	OLS			IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel B								
Constant	7.093***	7.886***	8.433***	7.240***	7.192***	7.920***	7.997***	7.979***
	(19.845)	(20.909)	(21.065)	(11.982)	(20.232)	(21.608)	(21.733)	(17.436)
GAP	0.0414***	0.00603	−0.0422***	0.0549*	0.00778	0.00597	−0.0368*	−0.0880*
	(4.141)	(0.543)	(−2.684)	(1.658)	(0.518)	(0.402)	(−1.931)	(−1.843)
PL		0.177***	−0.882***	−1.866***		0.183***	0.0513	−0.476***
		(8.073)	(−3.941)	(−3.702)		(9.309)	(1.039)	(−3.736)
GAP × PL			0.0922***	0.134***			0.0804***	0.232***
			(4.700)	(3.086)			(2.824)	(3.163)
SOE	0.0801***	0.0650***	0.0679***	0.148***	0.0718***	0.0649***	0.0675***	0.118***
	(3.529)	(2.854)	(2.988)	(5.247)	(3.142)	(2.846)	(2.964)	(4.293)
LEV	−0.387***	−0.358***	−0.366***	−0.422***	−0.406***	−0.360***	−0.361***	−0.409***
	(−5.995)	(−5.536)	(−5.671)	(−6.217)	(−6.280)	(−5.539)	(−5.567)	(−6.189)
SGROW	0.0936***	0.0915***	0.0939***	0.109***	0.0955***	0.0916***	0.0918***	0.0996***
	(3.372)	(3.342)	(3.429)	(4.236)	(3.447)	(3.349)	(3.352)	(4.008)

Table 5 (*continued*)

	ROA/GAP_1				ROA/GAP_2			
	OLS			IV	OLS			IV
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
SIZE	−0.308***	−0.326***	−0.329***	−0.295***	−0.295***	−0.326***	−0.327***	−0.290***
	(−22.617)	(−23.201)	(−23.284)	(−20.889)	(−22.042)	(−23.049)	(−23.065)	(−21.120)
TOP1	−0.00416***	−0.00414***	−0.00400***	−0.00415***	−0.00439***	−0.00416***	−0.00410***	−0.00434***
	(−6.587)	(−6.543)	(−6.330)	(−5.080)	(−6.962)	(−6.589)	(−6.500)	(−5.461)
CEOAge	0.0910	0.0716	0.0754	0.112	0.110	0.0723	0.0708	0.0896
	(1.273)	(1.005)	(1.059)	(1.219)	(1.538)	(1.018)	(0.997)	(0.998)
INDEPB	0.259	0.232	0.264	0.260	0.254	0.231	0.244	0.241
	(1.157)	(1.041)	(1.188)	(1.111)	(1.137)	(1.035)	(1.093)	(1.059)
BOARD	0.00536	0.00239	0.00281	0.00570	0.00554	0.00240	0.00262	0.00382
	(1.040)	(0.465)	(0.547)	(0.786)	(1.074)	(0.467)	(0.511)	(0.542)
DUAL	−0.0679**	−0.0607**	−0.0633**	−0.116***	−0.0589**	−0.0603**	−0.0607**	−0.0818***
	(−2.480)	(−2.223)	(−2.326)	(−3.785)	(−2.127)	(−2.194)	(−2.209)	(−2.748)
MKR	−0.0582**	−0.0776***	−0.0731***	−0.0252	−0.0458*	−0.0768***	−0.0764***	−0.0416
	(−2.427)	(−3.231)	(−3.048)	(−0.851)	(−1.916)	(−3.198)	(−3.182)	(−1.440)
Option	0.159***	0.141***	0.139***	0.191***	0.162***	0.141***	0.140***	0.180***
	(3.503)	(3.098)	(3.057)	(4.010)	(3.564)	(3.100)	(3.070)	(3.865)
N	9186	9186	9186	6624	9186	9186	9186	6784
Adj. *R*-sq	0.359	0.364	0.365	0.308	0.358	0.364	0.364	0.321

Note: Panel A shows the effects on *ROA*, and Panel B the effects on *TOBINQ*. We present both the ordinary least squares (OLS) estimation and IV estimation results, although we show only the second-stage results of the latter. The figures in parentheses are robust *t*-statistics adjusted for heteroskedasticity. We also control for year and industry fixed effects in all models.

* Significance level of 0.1.

** Significance level of 0.05.

*** Significance level of 0.01.

with the majority of the Chinese literature. However, after *PL* is added to the model, the regression coefficients of *GAP* decrease, and the *R*-square of the models increases. For example, in column 2 of Panel A, the coefficient relating *GAP* to *ROA* decreases to 0.00374 from 0.00666 (in column 1), and the *R*-square increases from 28.4% to 29.5%. Similarly, in column 2 of Panel B, the coefficient's influence on *TOBINQ* declines from 0.0414 (significant at the 1% level) to 0.00603 (not significant at conventional levels), and the *R*-square rises from 35.9% to 36.4%. These results indicate that, in addition to TMT pay gap, TMT pay level is also positively related to firm performance.

The model in columns 2 and 6 do not consider that, in addition to its main effect, TMT pay level may also relate to firm performance through its interaction with TMT pay gap. The models in columns 3 and 7 of Panels A and B, allow such an interaction. The estimation results indicate that the coefficients of $GAP \times PL$ are significantly positive at the 1% level, which suggests that the pay level can positively affect pay gap-induced incentives. These results are consistent with hypothesis H1. They indicate that the positive tournament effects of the pay gap are greater when the pay level is above the industry median.

In terms of economic significance, taking the sample's average total assets as an example, if a TMT's pay level is higher than that of its industry peers, then every 10,000-yuan increase in the pay gap raises net profits by about 42,600,000 yuan. If, in contrast, a TMT's pay level is lower than the industry average, every 10,000-yuan increase in the pay gap boosts net profits by about 19,800,000 yuan only. These results, which are economically significant, demonstrate that the tournament incentives induced by a pay gap are more pronounced when the TMT receives average pay that exceeds the industry median pay.

Columns 4 and 8 of Panels A and B present the results of the IV regressions. The coefficients of $GAP \times PL$ are again significantly positive, which means that after controlling for the endogeneity of *GAP*, *PL* and $GAP \times PL$, TMT pay level still exerts a positive effect on pay gap efficiency, i.e., a higher pay level is more likely to induce positive pay gap incentives. In summary, Table 5 results are consistent with the prediction of H1.

Table 6
Relationship between firm performance and TMT pay level and TMT pay gap for SOE and non-SOE subsamples.

IV regression	ROA				TOBINQ			
Independent variables	SOE (1)	Non-SOE (2)	SOE (3)	Non-SOE (4)	SOE (5)	Non-SOE (6)	SOE (7)	Non-SOE (8)
Constant	−0.244***	−0.201***	−0.176***	−0.170***	7.122***	9.670***	7.987***	9.979***
	(−5.304)	(−4.015)	(−4.764)	(−3.953)	(8.327)	(8.521)	(12.121)	(11.814)
GAP_1	0.00881***	0.00368*			0.0789	0.0343		
	(3.539)	(1.659)			(1.571)	(0.626)		
GAP_1 × *PL*	0.00578	0.0115***			0.0114	0.280***		
	(1.587)	(3.389)			(0.176)	(3.274)		
GAP_2			−0.00249	−0.00840***			0.00304	−0.136*
			(−0.774)	(−2.738)			(0.040)	(−1.906)
GAP_2 × *PL*			0.00415	0.0150***			−0.00140	0.354***
			(0.895)	(2.873)			(−0.013)	(2.884)
PL	−0.0797*	−0.153***	−0.00218	−0.0273***	−0.427	−3.669***	−0.135	−0.713***
	(−1.874)	(−3.743)	(−0.272)	(−2.852)	(−0.570)	(−3.639)	(−0.776)	(−3.207)
LEV	−0.126***	−0.113***	−0.127***	−0.112***	−1.065***	0.0712	−1.034***	0.0929
	(−20.324)	(−14.524)	(−20.680)	(−15.013)	(−9.828)	(0.524)	(−9.814)	(0.706)
SGROW	0.0233***	0.0235***	0.0226***	0.0232***	0.149***	0.0649	0.137***	0.0584
	(8.285)	(8.673)	(7.960)	(8.885)	(2.893)	(1.443)	(2.768)	(1.366)
SIZE	0.00890***	0.00983***	0.00976***	0.0108***	−0.237***	−0.387***	−0.238***	−0.372***
	(8.342)	(5.865)	(9.011)	(6.609)	(−11.124)	(−11.509)	(−11.416)	(−11.513)
TOP1	0.000168***	0.000271***	0.000104*	0.000287***	−0.00243**	−0.00799***	−0.00276***	−0.00746***
	(2.902)	(4.358)	(1.852)	(4.699)	(−2.288)	(−6.014)	(−2.652)	(−5.803)
CEOAge	0.0117	0.000480	0.0148**	0.00140	−0.0703	−0.0245	−0.0710	−0.0435
	(1.595)	(0.075)	(2.108)	(0.223)	(−0.566)	(−0.184)	(−0.602)	(−0.331)
INDEPB	−0.0476***	−0.00572	−0.0534***	−0.00433	−0.479	0.831*	−0.516*	0.896**
	(−3.228)	(−0.277)	(−3.757)	(−0.212)	(−1.600)	(1.875)	(−1.772)	(2.043)
BOARD	0.000310	0.000944	−0.00000236	0.000902	−0.000224	0.0198	−0.00131	0.0206
	(0.758)	(1.258)	(−0.006)	(1.223)	(−0.032)	(1.437)	(−0.192)	(1.522)
DUAL	−0.00207	−0.00556***	0.00124	−0.00286	−0.0159	−0.0923**	0.0112	−0.0542
	(−0.745)	(−2.649)	(0.462)	(−1.420)	(−0.335)	(−2.190)	(0.235)	(−1.319)
MKR	0.00270	0.0110***	0.00250	0.00987***	0.0500	−0.0287	0.0326	−0.0543
	(1.330)	(3.850)	(1.258)	(3.513)	(1.428)	(−0.481)	(0.965)	(−0.912)
Option	0.0287***	0.0131***	0.0303***	0.0111***	0.470***	0.191***	0.495***	0.171***
	(5.458)	(6.013)	(5.886)	(5.342)	(3.706)	(3.436)	(3.930)	(3.220)
N	3505	3119	3595	3189	3505	3119	3595	3189
Adj. R-sq	0.284	0.193	0.303	0.220	0.362	0.301	0.375	0.316

Note: The figures in parentheses are robust t-statistics adjusted for heteroskedasticity. We control for year and industry fixed effects in all models.
* Significance level 0.1.
** Significance level of 0.05.
*** Significance level of 0.01.

4.3. Influence of ownership status

The results on the relationship between TMT pay gap and firm performance under different ownership conditions are presented in Table 6. To control for endogeneity, the table presents the IV regression results using only the IVs stated above. Columns 1–4 show the results for *ROA*, and columns 5–8 the results for *TOBINQ*.

Table 6 shows that the coefficients of $GAP \times PL$ are significantly positive for the non-SOE subsample but are not significant for the SOE subsample. This result is consistent with H2, and suggests that the TMT pay level in non-SOEs is more likely to moderate the influence of a within-TMT pay gap on firm performance. However, the SOE TMT pay level has no effect on pay gap efficiency, possibly because the top executives of government-regulated SOEs are not sensitive to their external pay standing among their industry peers. Hence, an external compensation comparison does little to change competition and cooperation within SOE TMTs.

Table 7
Relationship between firm performance and TMT pay gap for the lowest and highest deciles of TMT pay level.

Fixed-effects	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Independent variables	ROA		TOBINQ		ROA		TOBINQ	
	SPL = 1	SPL = 10	SPL = 1	SPL = 10	PL = 0	PL = 1	PL = 0	PL = 1
Constant	0.760**	0.234*	36.80***	6.308**	−0.00398	0.0408	18.90***	12.39***
	(2.087)	(1.717)	(7.834)	(2.127)	(−0.052)	(0.770)	(16.564)	(12.662)
GAP_1	−0.0109*	0.00456**	−0.192***	0.0414				
	(−1.920)	(2.392)	(−2.619)	(0.997)				
GAP_2					−0.000155	0.00293**	−0.0763***	0.0339
					(−0.080)	(2.113)	(−2.640)	(1.325)
SOE	−0.0773***	−0.00427	−0.257	0.00177	−0.0291***	0.00330	−0.0573	−0.188**
	(−3.800)	(−0.437)	(−0.978)	(0.008)	(−4.885)	(0.661)	(−0.650)	(−2.035)
LEV	−0.199***	−0.123***	−0.0728	−0.0224	−0.161***	−0.144***	−0.242*	−0.0959
	(−5.710)	(−7.586)	(−0.162)	(−0.063)	(−17.107)	(−19.758)	(−1.735)	(−0.710)
SGROW	0.0200***	0.0285***	0.145*	0.0178	0.0197***	0.0179***	0.0196	0.0666***
	(3.309)	(7.298)	(1.862)	(0.209)	(11.818)	(13.322)	(0.792)	(2.685)
SIZE	−0.0276**	−0.00998*	−1.331***	−0.338***	0.000960	0.000608	−0.837***	−0.570***
	(−2.227)	(−1.951)	(−8.323)	(−3.026)	(0.352)	(0.305)	(−20.684)	(−15.446)
TOP1	0.00157**	−0.000426*	−0.0211**	0.00300	0.000374**	0.000324***	−0.00825***	−0.000829
	(2.045)	(−1.745)	(−2.142)	(0.563)	(2.185)	(2.786)	(−3.245)	(−0.386)
CEOAge	0.0215	0.0153	−1.137	0.451	0.00856	0.00690	0.0177	0.440***
	(0.349)	(0.746)	(−1.429)	(1.010)	(0.674)	(0.849)	(0.094)	(2.932)
INDEPB	0.0118	0.0521	1.207	−0.698	0.0362	−0.0103	0.235	0.0876
	(0.099)	(1.339)	(0.782)	(−0.823)	(1.364)	(−0.563)	(0.597)	(0.259)
BOARD	−0.00689	0.00156	0.00164	0.00319	−0.000178	−0.000336	0.0180	0.000991
	(−1.462)	(1.030)	(0.027)	(0.097)	(−0.169)	(−0.496)	(1.152)	(0.079)
DUAL	0.00624	−0.00539	−0.552**	−0.116	0.00213	−0.00630**	−0.0653	−0.000542
	(0.346)	(−0.897)	(−2.372)	(−0.888)	(0.509)	(−2.090)	(−1.053)	(−0.010)
MKR	0.0105	−0.0312**	−0.260	−0.303	−0.00519	−0.0138*	−0.0804	−0.440***
	(0.277)	(−2.139)	(−0.533)	(−0.950)	(−0.570)	(−1.842)	(−0.595)	(−3.163)
Option	0.0101	−0.00533	0.347	−0.0308	−0.000706	0.00225	0.169*	0.100*
	(0.324)	(−1.128)	(0.860)	(−0.299)	(−0.108)	(0.802)	(1.748)	(1.935)
N	742	1035	742	1035	4348	4838	4348	4838
Adj. *R*-sq	−0.405	−0.241	0.140	−0.094	−0.215	−0.142	0.196	0.153

Note: The figures in parentheses are robust *t*-statistics, and we also control for year and industry fixed effects in all models.
* Significance level 0.1.
** Significance level of 0.05.
*** Significance level of 0.01.

5. Robustness tests

5.1. Subdivision of TMT pay level

Our main focus is on the effects of a TMT pay gap on firm performance under different TMT pay level conditions. As previously noted, firms often use their industry peers as a benchmark when agreeing contracts with top executives (Jiang, 2011). Accordingly, we divide the TMT pay level into two groups, one with a pay level higher than the industry median and the other with a pay level lower than the industry median, and find a stronger positive pay gap effect on firm performance when the TMT pay level is higher than the industry median.

In Table 5, the coefficients of the relative pay gap measure *GAP_2* are all significantly negative, meaning that when TMT members' pay level is lower (higher) than that of their industry peers, *GAP_2* is negatively (positively) related to firm performance. Hence, we can say that the median pay level in a given industry may be the flex point for the effect of a relative pay gap (i.e., *GAP_2*) on firm performance. However, the same cannot be said for the absolute pay gap measure (i.e., *GAP_1*). *GAP_1* is positively related to firm performance in both the higher and lower pay level groups, although more strongly so in the former. To ensure that

Table 8
Relationship between firm performance and TMT pay gap for the lowest and highest deciles of TMT pay level for SOEs and non-SOEs.

Independent variables	ΔROA				ΔTOBINQ			
	SOE (1)	Non-SOE (2)	SOE (3)	Non-SOE (4)	SOE (5)	Non-SOE (6)	SOE (7)	Non-SOE (8)
Constant	-0.0804^{**}	-0.0323	-0.0607^{**}	-0.0541	4.053^{***}	6.510^{***}	4.173^{***}	6.168^{***}
	(−2.234)	(−0.751)	(−1.977)	(−1.406)	(8.363)	(9.607)	(8.781)	(9.964)
GAP_1_{t-1}	0.00321^{**}	-0.00168			0.00706	-0.0534^{**}		
	(2.146)	(−1.269)			(0.366)	(−2.397)		
$GAP\_1_{t-1} \times PL_t$	-0.00254	0.00341^{**}			-0.0291	0.0682^{**}		
	(−1.487)	(2.043)			(−1.251)	(2.413)		
GAP_2_{t-1}			0.00279	-0.00251^{*}			0.0596	-0.0359
			(1.630)	(−1.752)			(1.519)	(−1.320)
$GAP\_2_{t-1} \times PL_t$			-0.00316	0.00262			-0.0855^{*}	0.0681^{*}
			(−1.408)	(1.403)			(−1.825)	(1.943)
PL	0.0363^{*}	-0.0273	0.0143^{***}	0.00781^{**}	0.379	-0.624^{*}	0.174^{**}	0.0282
	(1.891)	(−1.429)	(3.917)	(2.155)	(1.437)	(−1.894)	(2.414)	(0.421)
LEV	-0.0679^{***}	-0.0595^{***}	-0.0670^{***}	-0.0603^{***}	-0.354^{***}	0.0984	-0.345^{***}	0.104
	(−10.590)	(−7.824)	(−10.509)	(−7.933)	(−4.166)	(0.887)	(−4.216)	(0.945)
SGROW	0.0208^{***}	0.0252^{***}	0.0206^{***}	0.0248^{***}	-0.0210	-0.0571	-0.0255	-0.0614
	(7.778)	(10.013)	(7.956)	(10.181)	(−0.544)	(−1.213)	(−0.690)	(−1.351)
SIZE	0.00318^{***}	0.00318^{**}	0.00338^{***}	0.00360^{**}	-0.145^{***}	-0.255^{***}	-0.144^{***}	-0.257^{***}
	(3.691)	(2.050)	(3.920)	(2.347)	(−8.935)	(−10.411)	(−9.305)	(−10.694)
TOP1	0.0000157	0.000129^{**}	0.0000177	0.000146^{***}	0.0000156	-0.00102	0.000154	-0.000765
	(0.331)	(2.465)	(0.379)	(2.795)	(0.018)	(−0.940)	(0.187)	(−0.710)
CEOAge	0.00693	0.00346	0.00818	0.00322	-0.0351	-0.0485	-0.0645	-0.0738
	(1.115)	(0.591)	(1.335)	(0.563)	(−0.345)	(−0.483)	(−0.656)	(−0.738)
INDEPB	-0.0251^{**}	-0.0100	-0.0258^{**}	-0.00773	-0.295	0.271	-0.334	0.298
	(−2.079)	(−0.539)	(−2.171)	(−0.420)	(−1.267)	(0.866)	(−1.484)	(0.960)
BOARD	-0.0000926	-0.000317	-0.0000881	-0.000128	-0.00136	0.0129	0.000141	0.0122
	(−0.280)	(−0.488)	(−0.259)	(−0.198)	(−0.246)	(1.345)	(0.026)	(1.283)
DUAL	0.000697	-0.00371^{**}	0.000673	-0.00329^{*}	0.0362	-0.00112	0.0276	-0.00871
	(0.283)	(−1.963)	(0.274)	(−1.755)	(0.992)	(−0.034)	(0.757)	(−0.271)
MKR	0.000559	0.00421	0.000598	0.00371	0.00880	-0.0463	0.00655	-0.0539
	(0.326)	(1.615)	(0.354)	(1.431)	(0.332)	(−1.026)	(0.254)	(−1.189)
Option	0.0174^{***}	0.00661^{***}	0.0179^{***}	0.00580^{***}	0.235^{***}	0.0250	0.253^{***}	0.0366
	(4.659)	(3.898)	(4.833)	(3.410)	(2.592)	(0.662)	(2.769)	(0.980)
ROA_{t-1}	-0.535^{***}	-0.608^{***}	-0.529^{***}	-0.610^{***}				
	(−16.571)	(−16.573)	(−16.617)	(−16.874)				
$TOBINQ_{t-1}$					-0.372^{***}	-0.335^{***}	-0.370^{***}	-0.341^{***}
					(−14.116)	(−13.165)	(−14.332)	(−13.441)
N	3505	3119	3595	3189	3477	2990	3566	3058
Adj. *R*-sq	0.317	0.350	0.315	0.356	0.528	0.502	0.530	0.503

Note: The figures in parentheses are robust *t*-statistics adjusted by heteroskedasticity, and we also control for year and industry fixed effects in all models.
* Significance at the 0.1 level.
** Significance at the 0.05 level.
*** Significance at the 0.01 level.

an absolute pay gap's influence on firm performance also experiences a flex point, we further subdivide TMT pay level. We first arrange it in ascending order for each year and each industry, and then divide the sample into 10 pay level groups. Firms in the first group ($SPL = 1$) have the lowest pay level relative to their peers, i.e., lower than 90% of firms in the same industry, whereas those in the tenth group ($SPL = 10$) have the highest such pay level, i.e., higher than 90% of firms in the same industry. We then estimate model (3) separately for each of the 10 groups to examine the absolute TMT pay gap's influence on firm performance.[13]

[13] The control variables in (3) are the same as those in model (1), and (3) also controls for fixed effects.

$$PERF_t = \beta_0 + \beta_1 GAP_t + \beta_2 SOE_t + \beta_3 LEV_t + \beta_4 SGROW_t + \beta_5 SIZE_t + \beta_6 TOP1_t + \beta_7 CEOAge_t + \beta_8 BOARD_t + \beta_9 INDEPB_t + \beta_{10} DUAL_t + \beta_{11} MKR_t + \beta_{12} Option_t + \sum Industry_t + \sum Year_t + \varepsilon_t, \quad (3)$$

We also estimate model (3) for the $PL = 0$ and $PL = 1$ subsamples separately to examine the relative TMT pay gap's (GAP_2) influence on firm performance. The results are presented in Table 7.

Columns 1–4 of Table 7 show GAP_1's effects on firm performance in the highest and lowest pay level groups. In the lowest group ($SPL = 1$), GAP_1 is negatively related to firm performance, whereas in the highest ($SPL = 10$), it is positively related. These results are consistent with H1. Columns 5–8 show GAP_2's influence on firm performance in the higher- and lower-than-industry-median pay level groups. If a firm's pay level is lower (higher) than the median level, GAP_2 is negatively (positively) related to firm performance. These results are consistent with those in Table 5, and further suggest that the direction of a pay gap's influence on firm performance changes from negative to positive with an increase in pay level, which is consistent with H1.

5.2. Endogeneity of TMT pay level and TMT pay gap

In addition to the 2SLS regression approach with IVs, we use a model with change in performance ($\Delta PERF$) as the dependent variable and lagged pay gap as the independent variable to control for endogeneity. A TMT pay gap in the previous year is less likely to be caused by performance growth in the current year, and the TMT pay level is less endogenous with performance growth than with current year performance. Hence, we adopt model (4) to examine the influence of the previous year's pay gap on performance growth under different ownership conditions.

$$\Delta PERF_t = \beta_0 + \beta_1 GAP_{t-1} + \beta_2 PL_t + \beta_3 GAP_{t-1} \times PL_t + \beta_4 PERF_{t-1} + \beta_5 SOE_t + \beta_6 LEV_t + \beta_7 SGROW_t + \beta_8 SIZE_t + \beta_9 TOP1_t + \beta_{10} CEOAge_t + \beta_{11} BOARD_t + \beta_{12} INDEPB_t + \beta_{13} DUAL_t + \beta_{14} MKR_t + \beta_{15} Option_t + \sum Industry_t + \sum Year_t + \varepsilon_t \quad (4)$$

The variables are defined as before, and we also control for the previous year's performance. The results of model (4) for the different ownership conditions are presented in Table 8. They show that the interaction term $GAP_{t-1} \times PL_t$ is significantly positive for the non-SOE subsample and insignificantly negative for the SOE subsample. Hence, when the pay level of a TMT in a non-SOE is higher than the industry median, a larger TMT pay gap brings about more performance growth. For SOEs, in contrast, a higher pay level does not induce better TMT tournament incentives. When these firms' TMTs receive less pay than their peers, a within-TMT pay gap may be conducive to performance growth.[14]

The results are consistent if we remove observations for which the TMT's pay standing relative to industry peers changed. We also use a lagged pay level with pay gap to re-examine model (1) under different ownership conditions, and the results are consistent with those in Tables 6 and 7. All of these results provide further support for H2.

We also use several other methods to re-measure TMT pay level and find consistent results. To exclude the effects of high CEO pay, we re-define TMT pay level without the CEO's pay and find consistent results. Additionally, we also use only the top three executives' pay or excess TMT pay to measure the TMT pay level, and find consistent results.

6. Additional analyses

According to the descriptive data in Table 4, the TMT pay levels in SOEs are much higher than those in non-SOEs in the same industry. Many Chinese scholars have suggested that the excess pay levels of SOEs are not conducive to firm efficiency because top executives' pay is regulated by the government (Wu et al., 2010; Li

[14] It is later suggested that the interaction's negative coefficients are all in central SOEs.

Table 9
Relationship between firm performance and TMT pay gap and TMT pay level for central SOEs and local SOEs: 2SLS-IV estimation.

IV regression	Central SOE				Local SOE			
Independent variables	ROA		TOBINQ		ROA		TOBINQ	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Constant	−0.386***	−0.209***	2.647	4.269***	−0.259***	−0.198***	8.514***	8.874***
	(−3.575)	(−2.859)	(1.366)	(2.801)	(−5.411)	(−5.158)	(10.234)	(13.111)
GAP_1	0.0174**		0.312**		0.0102***		0.0571	
	(2.535)		(2.534)		(3.718)		(1.200)	
GAP_1 × *PL*	−0.00758		−0.339***		0.00902**		0.134**	
	(−1.042)		(−2.598)		(2.521)		(2.154)	
GAP_2		0.0171		1.295**		−0.00247		0.0355
		(0.775)		(2.371)		(−0.656)		(0.537)
GAP_2 × *PL*		−0.0161		−1.476**		0.00249		−0.0316
		(−0.659)		(−2.446)		(0.337)		(−0.244)
PL	0.0677	0.0324	3.082**	1.740*	−0.125***	−0.000466	−1.798**	0.00833
	(0.862)	(0.861)	(2.186)	(1.866)	(−3.045)	(−0.040)	(−2.508)	(0.041)
LEV	−0.107***	−0.108***	−1.365***	−1.417***	−0.135***	−0.135***	−0.913***	−0.916***
	(−11.623)	(−12.204)	(−8.244)	(−8.177)	(−22.553)	(−23.441)	(−8.744)	(−9.069)
SGROW	0.0227***	0.0210***	0.0666	0.0462	0.0229***	0.0221***	0.182***	0.165***
	(6.732)	(6.506)	(1.102)	(0.739)	(10.277)	(10.481)	(4.684)	(4.439)
SIZE	0.00698***	0.00606***	−0.126***	−0.143***	0.0108***	0.0119***	−0.288***	−0.274***
	(3.800)	(3.608)	(−3.822)	(−4.399)	(9.477)	(10.858)	(−14.528)	(−14.230)
TOP1	0.0000327	0.0000241	−0.00939***	−0.0102***	0.000203***	0.000110*	−0.0000616	−0.00106
	(0.307)	(0.228)	(−4.914)	(−4.992)	(3.007)	(1.734)	(−0.052)	(−0.950)
CEOAge	0.0310**	0.0295**	−0.0458	0.00991	0.00234	0.00976	−0.131	−0.141
	(2.252)	(2.235)	(−0.185)	(0.039)	(0.276)	(1.196)	(−0.890)	(−0.979)
INDEPB	−0.0402	−0.0484*	−0.544	−0.562	−0.0503***	−0.0496***	−0.515	−0.580*
	(−1.364)	(−1.736)	(−1.029)	(−1.033)	(−2.598)	(−2.684)	(−1.529)	(−1.781)
BOARD	−0.000634	−0.000748	−0.00251	−0.0104	0.000978*	0.000379	0.00235	−0.000317
	(−0.750)	(−0.913)	(−0.166)	(−0.655)	(1.770)	(0.725)	(0.245)	(−0.035)
DUAL	0.000292	0.000615	−0.00862	−0.0851	−0.00402	0.000841	−0.00846	0.0413
	(0.049)	(0.103)	(−0.080)	(−0.722)	(−1.332)	(0.291)	(−0.161)	(0.814)
Option	0.00885	0.0110	−0.0404	0.00398	0.0388***	0.0392***	0.667***	0.683***
	(0.822)	(1.052)	(−0.209)	(0.020)	(5.893)	(6.149)	(5.816)	(6.100)
MKR	0.00282	0.00169	0.222***	0.137*	0.00199	0.000422	0.0110	−0.00937
	(0.724)	(0.427)	(3.172)	(1.740)	(0.865)	(0.192)	(0.275)	(−0.242)
N	1012	1038	1012	1012	2493	2557	2493	2557
Adj. *R*-sq	0.231	0.270	0.347	0.308	0.283	0.332	0.359	0.378

Note: The figures in parentheses are robust *t*-statistics, and we also control for year and industry in all models.
* Significance level of 0.1.
** Significance level of 0.05.
*** Significance level of 0.01.

et al., 2014). We find that excess pay levels in SOEs also do not influence internal TMT competition or cooperation, and thus should be addressed as the next step in China's ongoing comprehensive SOE reform. On 29 August 2014, the country's Central Political Bureau passed the *Pay System Reform Program of Central Management Companies' Responsible Person*, which was formally implemented on 1 January 2015. The program requires that the first 72 central SOE top executives receive a pay cut, thereby setting an example for local SOE pay reform. However, the program does not provide the detailed rules on how the pay cuts should be carried out or stipulate whether the internal pay distribution should be considered. Moreover, it is also important to discuss what notifications local SOEs should be given when they consider central SOEs' approach to top executive compensation reform.

As previously noted, if a company's pay incentive system is effective, the TMT pay level will moderate the TMT pay gap's efficiency. Thus, firms should consider choosing a suitable pay gap to ensure a certain pay standing in the industry. Because SOEs are regulated by the government, their industry pay standing is not related to the market. Hence, cutting their top managers' pay should not directly hurt the efficiency of an internal pay gap. However, this is not true for all SOEs because, in China, central and local SOEs are subject to different constraints (Xia and Fang, 2005), different levels of government intervention (Pan et al., 2008), and different pay systems. Although they both belong to the state-owned system, their top managers' preferences concerning monetary compensation differ, leading to different degrees of sensitivity to the firm's external pay standing. Thus, we re-estimate model (1) separately for central SOEs and local SOEs. The results are presented in Table 9. For consistency, we present only the results of the 2SLS regression with the IVs stated above.

Columns 1–4 of Table 9 report the results for central SOEs. The effect of a TMT pay gap on accounting performance (*ROA*) does not differ in the different pay level groups, although the pay gap's effect on market performance (*TOBINQ*) is stronger when the pay level is lower, which suggests that reducing the compensation of top managers in central SOEs is actually beneficial to firm value. When these SOEs implement pay cutting measures, their internal TMT pay gap could be raised.

Columns 5–8 of Table 9 present the results for local SOEs, showing that a higher TMT pay level is beneficial to the incentives induced by an absolute pay gap, but has little effect on those induced by a relative pay gap. These results suggest that in a local SOE, the TMT's pay level influences within-team competition and cooperation to some degree. Thus, when local SOEs engage in top manager compensation reform, direct pay cuts may not be the best approach. For those with a high pay level, increasing the internal pay gap may be a good solution to boost efficiency.

In summary, our results suggest that the pay cutting policy currently being implemented in China's SOEs will not damage the internal efficiency of a central SOE TMT, but local SOEs should not blindly follow the pay cutting decisions of their central counterparts. Because the pay level in local SOE TMTs can also affect within-team pay gap efficiency, the redistribution of payments within the TMT may be the optimal solution. In addition, we find the TMT pay level in local SOEs to have modulating effects in regions with a higher marketization level and in more competitive industries. We do not tabulate the results of these regressions because of space limitations.

7. Conclusion

Although the incentive effects of a TMT pay gap have received considerable research attention, the results of prior research are conflicting. These conflicting results are consistent with the two dominant theories in this area, namely tournament theory and social comparison theory, which make conflicting predictions. One explanation for the conflicting findings is that each of these theories holds, but only under specific conditions. TMT pay level concerns an external comparison and TMT pay gap an internal comparison. Most of the literature examine the efficiency of these two components of the pay system separately, with few studies noting that an external comparison also influences the internal comparison. Moreover, firms often use the market pay level in their industry as a benchmark in designing their own TMT pay contracts.

The mandatory requirement that Chinese listed firms provide information on their top executives' compensation, makes it possible to conduct external comparisons and to also study whether these external comparisons have implications for the managerial incentive and firm performance effects of internal comparisons. Using compensation data from 2005 to 2012, we examine the effects of TMT pay level on within-team pay gap efficiency. Our results show that a higher TMT pay level than that of the team's industry peers stimulates within-team competence, and thus indicates that tournament theory may be more suitable than social comparison theory for explaining the incentives of a pay gap. A lower TMT pay level than industry peers, in contrast, renders top executives more sensitive to an internal pay gap because they perceive their pay to be unfair. Accordingly, social comparison theory provides a more suitable explanation under these circumstances. Therefore, when firms are designing contracts for top executives, it is important that they both adjust the pay level to suit the market and ensure an appropriate within-TMT pay distribution, which is vital to the pay system as a whole. For example, the Chinese listed firm *Everfine Photo* (stock ID: 300306) emphasizes in its 2012 annual report that "the firm has implemented [a] new reward system, [and] the top managers' pay … keep[s] up with the market, and should make dynamic maintenance of the pay gap and pay level every year according to the market level…."

However, because of government regulation, providing SOE top managers with excess compensation is not only unfavorable to firm performance (Wu et al., 2010), but also fails to stimulate internal TMT competition. As noted earlier, a reform of the pay system for central SOEs has been in effect since 1 January 2015, and consequently the majority of these firms' top executives will receive a pay cut. Our findings suggest that such pay cuts will not adversely affect the incentives induced by an internal pay gap and may even boost firm value, although the same does not hold true for local SOEs. The top managers of central and local SOEs face different degrees of government intervention, market competition and other factors, with the former group's pay system being more market-oriented. Directly cutting pay is likely to be detrimental to the incentives induced by an internal TMT pay gap in local SOEs. For high-paying local SOEs, widening the internal pay gap may be a solution to boosting efficiency.

References

Akerlof, G.A., Yellen, J.L., 1986. Efficiency Wage Models of the Labor Market. Cambridge University Press, Cambridge.

Bloom, M., 1999. The performance effects of pay dispersion on individuals and organizations. Acad. Manage. J. 42 (1), 25–40.

Bloom, M., Michel, J.G., 2002. The relationships among organizational context, pay dispersion, and managerial turnover. Acad. Manage. J. 45 (1), 33–42.

Bognanno, M.L., 2001. Corporate tournaments. J. Labor Econ. 19 (2), 290–315.

Brown, M., 2001. Unequal pay, unequal responses? Pay referents and their implications for pay level satisfaction. J. Manage. Stud. 38, 879–896.

Brown, M.P., Sturman, M.C., Simmering, M.J., 2003. Compensation policy and organizational performance: the efficiency, operational, and financial implications of pay levels and pay structure. Acad. Manage. J. 46 (6), 752–762.

Chen, D., Chen, X., Wan, H., 2005. Compensation regulation and job consumption in state-owned enterprises. Econ. Res. J. 2, 100–111 (in Chinese).

Chen, J., Ezzamel, M., Cai, Z., 2011. Managerial power theory, tournament theory, and executive pay in china. J. Corp. Finan. 17 (4), 1176–1199.

Chen, Z., Zhang, M., 2006. Research on top executives' pay distributions. China Account. Rev. 1, 15–28 (in Chinese).

Colquitt, J.A., Conlon, D.E., Wesson, M.J., Porter, C.O., Ng, K.Y., 2001. Justice at the millennium: a meta-analytic review of 25 years of organizational justice research. J. Appl. Psychol. 86 (3), 425–445.

Core, J.E., Holthausen, R.W., Larcker, D.F., 1999. Corporate governance, chief executive officer compensation, and firm performance. J. Financ. Econ. 51 (3), 371–406.

Cowherd, D.M., Levine, D.I., 1992. Product quality and pay equity between lower-level employees and top management: an investigation of distributive justice theory. Adm. Sci. Q. 37, 302–320.

Deutsch, M., 1985. Distributive Justice: A Social–Psychological Perspective. Yale University Press, New Haven, CT.

Drago, R., Garvey, G.T., 1998. Incentives for helping on the job: theory and evidence. J. Labor Econ. 16 (1), 1–25.

Ehrenberg, R.G., Bognanno, M.L., 1990a. Do tournaments have incentive effects? J. Polit. Econ. 98 (6), 1307–1324.

Ehrenberg, R.G., Bognanno, M.L., 1990b. The incentive effects of tournaments revisited: evidence from the European PGA tour. Ind. Labor Relat. Rev. 43 (3), 74S–88S.

Eriksson, T., 1999. Executive compensation and tournament theory: empirical tests on Danish data. J. Labor Econ. 17 (2), 262–280.

Fan, G., Wang, X.R., 2007. Marketization Index in China: A Report on the Process of Marketization in Different Regions. Economic Science Press (in Chinese).

Fang, J.X., 2012. Top executives' excess compensation and corporate governance policy. Manage. World 11, 144–155 (in Chinese).

Frank, R.H., 1985. Choosing the Right Pond: Human Behavior and the Quest for Status. Oxford University Press, New York.

Fredrickson, J.W., Davis-Blake, A., Sanders, W.M., 2010. Sharing the wealth: social comparisons and pay dispersion in the CEO's top team. Strateg. Manage. J. 31 (10), 1031–1053.

Gerhart, B., Milkovich, G.T., 1992. Employee compensation: research and practice. In: Dunnette, M.D., Hough, L.M. (Eds.), Handbook of Industrial and Organizational Psychology, vol. 3. Consulting Psychologists Press, Palo Alto, CA, pp. 481–569.

Hambrick, D., Siegel, P., 1997. Compensation patterns within top management groups: on the harmful effects of executive pay disparities in high-technology firms. Organ. Sci. 16 (3), 259–274.

Harbring, C., Irlenbusch, B., 2003. An experimental study on tournament design. Labour Econ. 10 (4), 443–464.

He, W., Hao, P., 2014. Does departmental pay dispersion really deteriorate employee affective commitment? The moderators of pay level, pay-for-performance intensity and workforce diversity. Nankai Bus. Rev. 4, 13–23 (in Chinese).

Hibbs Jr., D.A., Locking, H., 2000. Wage dispersion and productive efficiency: evidence for Sweden. J. Labor Econ. 18 (4), 755–782.

Hills, F.S., 1980. The relevant other in pay comparisons. Ind. Relat. 19, 345–351.

Jensen, M.C., 1986. Agency costs of free cash flow, corporate finance and takeovers. Am. Econ. Rev. 76 (2), 323–329.

Jiang, W., 2011. Marketization level, industry competition and the growth of managerial compensation. Nankai Bus. Rev. 5, 58–67 (in Chinese).

Kale, J.R., Reis, E., Venkateswaran, A., 2009. Rank-order tournaments and incentive alignment: the effect on firm performance. J. Financ. 64 (3), 1479–1512.

Kepes, S., Delery, J., Gupta, N., 2009. Contingencies in the effects of pay range on organizational effectiveness. Pers. Psychol. 62 (3), 497–531.

Kini, O., Williams, R., 2012. Tournament incentives, firm risk, and corporate policies. J. Financ. Econ. 103 (2), 350–376.

Knoeber, C.R., Thurman, W.N., 1994. Testing the theory of tournaments: an empirical analysis of broiler production. J. Labor Econ. 12 (2), 155–179.

Lambert, R.A., Larcker, D.F., Weigelt, K., 1993. The structure of organizational incentives. Adm. Sci. Q. 38, 438–461.

Law, K.S., Wong, C., 1998. Relative importance of referents on pay satisfaction: a review and test of a new policy-capturing approach. J. Occup. Organ. Psychol. 71, 47–60.

Lazear, E.P., 1989. Pay equality and industrial politics. J. Polit. Econ. 97, 561–580.

Lazear, E.P., Rosen, S., 1981. Rank-order tournaments as optimum labor contracts. J. Polit. Econ. 89 (5), 841–864.

Li, W., Cen, Y., Hu, Y., 2014. Does external pay gap encourage top management? An empirical study based on managerial market and ownership type. Nankai Bus. Rev. 4, 24–35 (in Chinese).

Li, W., Hu, Y., 2012. Who is encouraged by pay dispersion in state-owned enterprises? Econ. Res. 12, 125–136 (in Chinese).

Liao, L., Liao, G., Shen, H., 2009. Operating risk, promotion incentive and corporate performance. China Ind. Econ. 8, 119–130 (in Chinese).

Lin, B., Lu, R., 2009. Managerial power, compensation gap and firm performance—evidence from Chinese public listed companies. Global Financ. J. 20 (2), 153–164.

Lin, J., Huang, Z., Sun, Y., 2003. TMT pay gap, firm performance and corporate governance. Econ. Res. 4, 31–41 (in Chinese).

Liu, C., Sun, L., 2010. A study on relation of salary difference and firm performance: evidence from state-owned enterprises. Nankai Bus. Rev. 2, 40–51 (in Chinese).

Lu, Z., Wang, X., Zhang, P., 2012. Do Chinese state-owned enterprises pay higher wages? Econ. Res. 3, 28–39 (in Chinese).

Main, B.G., O'Reilly III, C.A., Wade, J., 1993. Top executive pay: tournament or teamwork? J. Labor Econ. 11 (4), 606–628.

Messersmith, J.G., Guthrie, J.P., Ji, Y., Lee, J., 2011. Executive turnover: the influence of dispersion and other pay system characteristics. J. Appl. Psychol. 96 (3), 457–469.

Milkovich, G.T., Newman, J.M., 2002. Compensation. Irwin, Chicago, IL.

Myers, S.C., 1977. Determinants of corporate borrowing. J. Financ. Econ. 5 (2), 147–175.

Oldman, G.R., Kulik, C.T., Stepina, L.P., Ambrose, M.L., 1986. Relations between situational factors and comparative referents used by employees. Acad. Manage. J. 29, 599–608.

O'Reilly III, C.A., Main, B.G., Crystal, G.S., 1988. CEO compensation as tournament and social comparison: a tale of two theories. Adm. Sci. Q. 33 (2), 257–274.

Pan, H., Xia, X., Yu, M., 2008. Government intervention, political connections and the mergers of local government-controlled enterprises. Econ. Res. 4, 41–52 (in Chinese).

Pfeffer, J., Langton, N., 1993. The effect of wage dispersion on satisfaction, productivity, and working collaboratively: evidence from college and university faculty. Adm. Sci. Q. 38, 382–407.

Qi, H., Zou, Y., 2014. The empirical research on incentive effect of the external fairness of executive compensation on agents' behavior. Account. Res. 3, 26–32 (in Chinese).

Rosen, S., 1986. Prizes and incentives in elimination tournaments. Am. Econ. Rev. 76 (4), 701–715.

Shaw, J.D., Gupta, N., Delery, J.E., 2002. Pay dispersion and workforce performance: moderating effects of incentives and interdependence. Strateg. Manage. J. 23 (6), 491–512.

Siegel, P.A., Hambrick, D.C., 2005. Pay disparities within top management groups: evidence of harmful effects on performance of high-technology firms. Organ. Sci. 16 (3), 259–274.

Trevor, C.O., Wazeter, D.L., 2006. A contingent view of reactions to objective pay conditions: interdependence among pay structure characteristics and pay relative to internal and external referents. J. Appl. Psychol. 91, 1260–1275.

Wu, L., Lin, J., Wang, Y., 2010. The external fairness of CEO's emolument, the nature of stock rights and company performance. Manage. World 3, 117–126 (in Chinese).

Xia, L., Fang, Y., 2005. Government control, institutional environment and firm value: evidence from the Chinese securities market. Econ. Res. 5, 40–51 (in Chinese).

Xu, L., Xin, Y., Chen, G., 2006. The types of controlling shareholders and the performance of China's listed companies. World Econ. 10, 78–89 (in Chinese).

Yang, R., Wang, Y., Nie, H., 2013. The promotion mechanism of "quasi-official": evidence from Chinese central state-owned enterprises. Manage. World 3, 23–33 (in Chinese).

Zhang, Z.T., 2007. Top management team coordination needs, compensation dispersion and firm performance: a perspective of tournament theory. Nankai Bus. Rev. 2, 4–11 (in Chinese).

Zhang, Z.T., 2008. An empirical study on the impact of the internal pay gap on future performance. Account. Res. 9, 81–87 (in Chinese).

Zhang, Z., Wang, Y., Liu, N., 2012. Study on the design factors and modes of team compensation scheme. Econ. Manage. 8, 89–96 (in Chinese).

Zhou, Q., Zhu, W., 2010. On the incentive effects of SOE tournament. China Econ. Q. 2, 5715–5796 (in Chinese).

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Re-examination of the effect of ownership structure on financial reporting: Evidence from share pledges in China

Zhizhong Huang, Qingmei Xue *

Business School, Nanjing University, China

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ABSTRACT

In this paper, we present evidence that firms with concentrated ownership manage earnings when their large shareholders have an incentive to do so. The large shareholders of Chinese public firms often pledge their shares for loans. Before the split share reform in 2006, loan terms were based on the book value of the firm. Since then, the share price has become critical for share pledged loans. We postulate that the reform triggered large shareholders' incentive to influence financial reports. Using a sample of non-state-owned enterprises, we test the effect of share pledges on earnings smoothing and how this effect changes after the reform. Our results suggest that share pledging firms smooth their earnings more than other firms, but these results are only found after the split share reform. Accordingly, our results provide more direct evidence on the effect of ownership concentration on financial reporting.

1. Introduction

Concentrated ownership creates opportunities for controlling shareholders to expropriate wealth from other shareholders. Most research shows that concentrated ownership is associated with lower earnings quality (Fan and Wong, 2002; Francis et al., 2005a,b). However, Wang (2006) argues that controlling shareholders may provide higher quality reporting because they are long-term investors and care about the reputation, wealth and long-term performance of their firms. These two different effects associated with controlling

* Corresponding author at: Business School, Nanjing University, Nanjing 210093, China.
E-mail address: xueqm@nju.edu.cn (Q. Xue).

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shareholders are summarized as the entrenchment effect and the alignment effect (Fan and Wong, 2002; Wang, 2006). Given the competing views and evidence in the literature, we conclude that the effect of ownership structure on financial reporting behavior is complicated and needs further investigation.

A unique Chinese setting provides an opportunity for us to re-examine this issue. Generally, Chinese public firms have one dominant shareholder whose ownership is much higher than the next largest shareholder. The large shareholders also tend to hold important positions in the management team. In recent years, large shareholders have often used their shares as collateral to obtain short-term loans. Prior to the split share reform in 2006, the pledged shares were mostly non-tradable. Although large shareholders retained their status after the split share reform, their incentives to influence financial reporting have changed. This setting allows us to compare the effects of controlling shareholders' incentives on their firms' discretionary financial reporting decisions.

The findings of numerous studies indicate that managers tend to manage earnings around major financing events, such as IPOs, seasoned equity offerings and seasoned bond offerings (see Leuz et al., 2003; Park and Shin, 2004; Guthrie and Sokolowsky, 2010; Caton et al., 2011). Consistent with the literature, we predict that for share pledge purposes, controlling shareholders tend to manage their firms' accounting performance to increase the value of their collateral.

Before the loans are made, controlling shareholders have incentives to manage earnings to increase their borrowing capacity (e.g., higher loan amounts, lower interest rates, and lower contracting costs) (Ahn and Choi, 2009). As required by law, pledge loan contracts should include a maintenance requirement. After a share pledged loan is made, if the share price drops, the value of the collateral will also decrease, and shareholders (borrowers) will have to make up for the decrease. This scenario is similar to a margin call when buying on margin. To avoid such costly consequences, shareholders will do whatever they can to uphold the share price. Chan et al. (2013) find that pledging firms make repurchases when prices drop. Because financial reporting influences the share price, we expect that shareholders will also manage their earnings to avoid dramatic price drops. Dye (1988) suggests that managers may smooth earnings to increase their firms' share price. Given that most shareholders pledge repeatedly, we predict that share pledged loans increase the degree of earnings smoothing.

However, if the share price is not the primary factor in deciding the loan terms, the shareholder will have no incentive to smooth earnings. In China, listed companies had dual class share ownership until the split share reform in 2006. The shares pledged by large shareholders, which were non-tradable, were evaluated by the book value of the firm. At this time, shareholders were indifferent to the share price. However, the split share reform in 2006 eliminated the discrepancies in the share transfer system. Since then, the value of pledged shares is based on their market price. Therefore, we predict that earnings smoothing is more likely in share pledging firms than other firms *after* the split share reform.

In our setting, we posit the same analytical relationship as Tucker and Zarowin (2006), who measure earnings smoothing by the negative correlation between the change in the discretionary-accruals proxy and the change in pre-discretionary income. Our hypothesis concerns the comparative smoothing of the share pledging firms and other firms. Using non-state-owned enterprises (SOEs) listed on the Chinese capital market as our sample, we test the moderation effect of share pledges on the above correlation. Consistent with our hypothesis, we find that share pledging firms smooth earnings more than other firms. This phenomenon is only observed after the split share reform.

Since the split share reform, shareholders have had the option of selling their shares or making a pledged loan. This raises the question of why some shareholders do not sell their shares. First, shareholders may not want to take the risk of losing their control rights. Because it is difficult to obtain approval for an IPO in China, listed companies are themselves valuable resources for capital raising (Liu and Lu, 2007). Second, shareholders may consider the share price to be undervalued, either because they have private information or they are irrationally optimistic about company prospects (Chan et al., 2013). Although we cannot determine the exact reason for each pledge announcement, we find that shareholdings are positively correlated with earnings smoothing. The more shares held by large shareholders, the less they fear losing their control rights, and the more likely the shares are undervalued in the case of a pledge. Therefore, the earnings smoothing that we find in pledged firms is also consistent with the argument of Ronen and Sadan (1981) that the firms are communicating a better future. Further, we find evidence that share pledging firms have higher short-run market returns than the matched sample.

We examine how the change of incentives influences financial reporting decisions in firms with concentrated ownership. Our major conclusion is that ownership concentration only affects financial reporting if large shareholders have such a motivation. Our findings contribute to three strands of the literature. First, our research differs from the literature on ownership concentration and earnings management in that we use a special setting in which large shareholders' incentives are distinct, whereas previous research only postulates different incentives. Moreover, our tests focus on the effect of incentives on income smoothing, whereas previous research mostly tests the association between ownership concentration and earnings quality. Second, we also further our understanding of the effects of the split share reform. Most of the related research find that the split share reform has a positive effect (Hou and Lee, 2014; Li and Zhang, 2011). We show that the split share reform may activate the incentives for earnings management. Finally, we contribute to the growing body of research on the economic consequences of share pledges (Chan et al., 2013; Kao and Wei, 2014; Kao and Chen, 2007; Kao et al., 2004; Yeh et al., 2009). For example, Kao and Chen (2007) and Kao and Wei (2014) find that directors' share collateralization reduces the quality of financial information. However, Tan and Wu (2013) argue that pledged firms' earnings quality is better due to the quality control of financial institutions. We provide more insights into this issue by testing for earnings smoothing before and after the split share reform.

The remainder of this paper is organized as follows. In Section 2, we briefly introduce the Chinese capital market, share pledges, and the split share reform. In Section 3, we review the literature and discuss our hypothesis. In Section 4, we explain our research design. Our empirical results are reported in Section 5. The last section concludes the paper.

2. Institutional background

2.1. Chinese capital market and share pledges

In the Chinese capital market, which was founded in the early 1990s, the majority of listed firms are restructured from SOEs. In recent years, the number of non-SOE firms has increased notably as a result of the reform of the IPO regulations and the privatization of state-controlled listed firms (Xia, 2008). For both SOEs and non-SOEs, the listed companies are normally the strongest part of the company group. All of the other subsidiaries of the group have limited financing sources. The formal and informal procedures of the banking system associated with lending to SOEs rely on collateral and personal relationships. Consequently, the large shareholders of many of the firms that have their shares traded on the security market pledge their shares for loans, especially for non-SOEs, whose large shareholders cannot easily obtain bank loans.

According to the listing rules of the Chinese stock exchanges firms should make announcements when their large shareholders (with more than 5% of shareholdings) pledge their shareholdings for loans. Since 2006, pledging has been active in some of the capital-intensive industries. For example, 348 listed companies reported share pledges in 2010, and the number increased to 697 in 2012. The annual growth rate is about 45%. The value of shares pledged by the shareholders of listed companies in 2012 reached RMB620 billion. Accordingly, it appears to have become common practice that large shareholders of listed firms pledge their shares as collateral for loans.

A pledge is a method of transferring collateral to create a security interest in the collateral. The collateral provider, the borrower, retains legal ownership of the assets. The lender has the right to keep the assets in the event of default by the collateral provider. Although pledges enable shareholders to obtain low cost loans without losing control of their firms, there are a number of risks associated with the pledging of shares. In the case of a default, the financial institution can sell the shares on the open market to recover the dues, which can result in a fall in the share price and erosion of market capitalization. Shareholders also run the risk of losing management control if a large proportion of their holding is pledged. According to the Administrative Measures for the Disclosure of Information of Listed Companies, promulgated by The People's Bank of China and the China Securities Regulatory Commission (CSRC) in 2000, a share pledge agreement should include a maintenance requirement. If the share price drops below the maintenance requirement, the shareholder needs to either provide supplementary pledges or to make earlier payments. This margin call pressure may result in a change of corporate behavior.

2.2. Ownership concentration and split share reform

Ownership of listed companies in China is typically concentrated in the hands of large shareholders. For example, Chen et al. (2009) report that on average, across all listed firms, the largest shareholder owns 43.75% of a firm, while the second largest owns 8.16%. When the Chinese capital market was founded, a unique feature of the ownership structure of listed companies was the split share structure. Under the split share structure, the shares of Chinese listed firms were classified into non-tradable and freely tradable shares. Non-tradable shares exhibited the same voting and cash flow rights as tradable shares, but they could not be traded freely on the stock exchange. Non-tradable shares were held by the founders of the company. Consequently, large shareholders normally held large amounts of non-tradable shares. Although the non-tradable shares were designed to help the government control the SOEs, the same structure existed for non-SOEs. This structure distorted the pricing mechanism in the capital market and caused many corporate governance problems (Li and Zhang, 2011). Large shareholders did not care about share prices, and tunneling by large shareholders was a serious problem for Chinese listed firms.

In 2005, the China Securities Regulatory Commission (CSRC) started the split share reform, which required non-tradable shareholders to pay tradable shareholders to gain the right to trade. By the end of 2006, 95% of China's publicly listed companies were involved in the split share structure reform. The reform measures were fully completed in 2007 (Li and Zhang, 2011). However, there were no fundamental changes in ownership structure after the reform.

Since non-tradable shares became tradable, the use of share pledges has continually increased. Using the market share price as a reference, share pledge loan terms can include a margin call requirement. The split share reform is the cornerstone of the development of the Chinese capital market.

3. Literature review and hypothesis development

3.1. Ownership structure and discretionary earnings reporting

The research on ownership structure and discretionary financial reporting has yielded mixed results. The entrenchment view suggests that concentrated ownership *decreases* earnings quality. As discussed by Shleifer and Vishny (1997), controlling shareholders have incentives to maximize their own benefits at the cost of minority shareholders. Claessens et al. (2002) argue that large shareholdings have an entrenchment effect, which decreases firm value. Fan and Wong (2002) find that earnings are more informative for firms with less concentrated ownership in East Asian countries. Their evidence indicates that greater ownership concentration creates greater agency conflicts and information asymmetry. Francis et al. (2005a,b) provide similar evidence in the U.S. environment, and find lower earnings response coefficients for firms with ownership structures that have unequal voting rights. In the Chinese setting, Liu and Sun (2010) document evidence that controlling shareholders expropriate minority shareholders and lower the quality of financial reports.

The alignment view suggests that concentrated ownership *increases* earnings quality. A long line of research suggests that managers engage in earnings management because of capital market pressures or to avoid violation of contracts (see the review of Dechow and Skinner (2000)). Controlling shareholders are able to discipline managers in the case of opportunistic actions (Fan and Wong, 2002; Shleifer and Vishny, 1997), such as manipulating reported performance for compensation contracts. By aligning the interests of managers with shareholders, managers have less incentive and/or ability to manipulate short-term performance as the controlling shareholders care mostly about the long-term performance of the firm. Warfield et al. (1995) find that low managerial ownership creates a demand for contracts that rely on accounting information to constrain managers' opportunistic behavior. They find that greater managerial ownership is associated with higher earnings quality. Wang (2006) documents that founding families are less likely to expropriate wealth from other shareholders through managing earnings. He explains that founding families have a long-term orientation and seek to protect their reputation by not opportunistically managing earnings.

In summary, both theories agree that controlling shareholders have the *ability* to expropriate from the firm and manipulate earnings to cover their behavior. However, whether they have always had the *incentive* to do so remains an open question. As discussed by Ball et al. (2003), the incentives of financial statement preparers play an essential role in reporting high-quality financial information. The disparity between the two theories is mostly related to incentives. Few studies have provided causal evidence that a concentrated ownership structure leads to discretionary earnings reporting because it is difficult to directly depict the incentives.

We extend the existing research by examining a specific setting in which controlling shareholders develop an incentive to influence financial reports. Specifically, as detailed in the following part, we examine whether share pledging firms are more likely to smooth earnings before and after the split share reform.

3.2. Share pledges and earnings smoothing

Collateral is an important part of most of the loans made in China. When the collateral comprises tangible assets, financial institutions can control it, and the value does not change too much. However, when the collateral is shares, the controlling rights of the firm still belong to the borrower (large shareholder). The large shareholder can influence the value of the collateral by making all kinds of decisions for the firm.

As with other financing events, shareholders seek to improve their bargaining power in negotiating loan covenants with financial institutions. This may induce shareholders to increase the value of their pledges, namely the shareholdings. With other financing events, management tends to increase share value by managing pre-financing earnings. For example, numerous studies have investigated earnings management behavior around significant financing events, such as IPOs (e.g. Aharony et al., 2010; Teoh et al., 1998a), seasoned equity offerings (Teoh et al., 1998b), and initial bond offerings (Caton et al., 2011). Unlike the widely spread ownership in the U.S. and U.K., Chinese listed firms are dominated by large shareholders, who are either government related or individuals. The large shareholders usually have effective control over the firm (Chen et al., 2008), and thus have both the incentive and ability to inflate earnings before share pledges.

After a share pledge, large shareholders may also have an incentive to manipulate the value of the pledged shares. If the share price drops below the maintenance requirement, the shareholders need to meet the margin call. This margin call pressure is similar to the pressure from a debt covenant. Numerous studies have provided evidence that debt is positively associated with income-increasing earnings management when firms want to reduce the probability of debt covenant violations (e.g. DeFond and Jiambalvo, 1994; Dichev and Skinner, 2002; Jaggi and Lee, 2002). A share pledged loan is different because the debt covenant may not directly be based on the accounting numbers of the pledged firm. Moreover, if the share price drops, shareholders have other options to uphold the share price. For example, Chan et al. (2013) find that high pledge companies are more likely to repurchase, especially after a significant drop in share prices. However, because earnings announcements influence share prices, shareholders may also manage their earnings to avoid sharp price drops. If shareholders inflate earnings before a share pledge, the price will be more likely to drop afterward, which will increase the pressure of the margin call. Taken together, for share pledge purposes, large shareholders need to increase the share price, but they will not manipulate earnings in a way which would be hard to continue. Instead, they prefer increasing the share price steadily or at least ensuring a stable share price.

Using data from Taiwan, Kao and Chen (2007) find that the more shares collateralized by the board of directors, the greater the extent of earnings management. Kao et al. (2004) find that collateralized shares have a negative relationship with firms' accounting performance, which could be seen as the reverse effect after earnings management in the year of the pledge. However, Chinese large shareholders may not conduct earnings management in a once and for all manner. Financial institutions usually only provide short-term loans with shares as collateral. If large shareholders need long-term finance, they have to pledge repeatedly, which is normally the case. If they inflate earnings once, the subsequent reversal in earnings will harm their future bargaining power. Thus, they are likely to smooth earnings to facilitate share pledged loans.

Research indicates that firms smooth earnings to meet certain goals (Ronen and Sadan, 1981). For example, managers may smooth earnings to meet a bonus target or to protect their job (Arya et al., 1998; Healy, 1985). In our case, the intention of large shareholders is to increase the share price for share pledged loans.

Smoothing earnings could convince investors that earnings have lower volatility, and hence represent reduced risk, thereby increasing the share price (Beidleman, 1973; Trueman and Titman, 1988).

Other research suggests that only firms with good future prospects can afford to smooth earnings (e.g. Ronen and Sadan, 1981). Managers use their discretion to communicate their assessment of future earnings. As abovementioned, a possible reason why shareholders choose to pledge their shares rather than sell them is that they believe their shares are undervalued. The shareholders of such firms may be more confident about their future and believe they have the ability to smooth. Hence, share pledged firms may have both the motivation and ability to smooth earnings.

Although shareholders may intend to smooth earnings, the financial checking by banks possibly limits this opportunistic practice (Tan and Wu, 2013). However, other research indicates that Chinese banks do not provide the same governance role as Western banks (Hu et al., 2011). In addition, banks make their loan decisions based on the current market value of the collateral. If the market value drops, they can ask for supplementary pledges based on the contract and law. Secured by such terms, banks have less incentive to check the fundamental value of the collateral. Their incentive would be even less after a share pledge.

3.3. Effect of the split share reform

The premise of the above arguments is that share pledged loans are based on the share price. However, before the split share reform, large shareholders often pledged their non-tradable shareholdings. The non-tradable shares could not be publicly traded on the open market, and thus there was no market price. However, the shares could still be transferred by auction or transfer agreement, mostly based on book value (Chen et al., 2008). Financial institutions accepted non-tradable shares as collateral for loans. Because non-tradable shares were less liquid, the banks had to bear higher risk. This is probably why there were much fewer share pledge announcements before the split share reform.

Moreover, Sun (2010) investigates the share pledges between 2001 and 2004, and finds evidence that the value of the collateral is based mostly on the book value per share of the listed firm. The profitability or size of the listed firm has no significant influence on the value of pledged non-tradable shares. Therefore, neither the earnings nor the market price of tradable shares are particularly relevant for shareholders or financial institutions in making their share pledged loan agreements.

The preceding discussion suggests that share pledging firms have different discretionary financial reporting *incentives* than other firms, but only after the split share reform. We formulate our hypothesis as follows:

H1. Share pledging firms smooth earnings more than other firms.

H1a. Share pledging firms smooth earnings NO more than the other firms before the split share reform.

H1b. Share pledging firms smooth earnings more than other firms after the split share reform.

4. Empirical design

4.1. Model of earnings smoothing and share pledges

Our main tests focus on the association between share pledges and earnings smoothing. There are several models for estimating earnings smoothing in the literature (Burgstahler et al., 2006; Dou et al., 2013; Tucker and Zarowin, 2006). As some of the share pledging firms in our sample pledge in one year and not in other years, we are unlikely to be able to calculate the variance, which means that we cannot use the models of Burgstahler et al. (2006) or Dou et al. (2013). Therefore, we follow the specification of Tucker and Zarowin (2006) in testing our earnings smoothing hypothesis. We measure earnings smoothing by the negative correlation between the change in discretionary-accruals proxy ("ΔDA") and the change in pre-discretionary income ("ΔPDI"). The pre-discretionary income (PDI) is calculated as net income (scaled by lagged assets) minus discretionary accruals (PDI = E-DA). Thus, $\Delta E = \Delta\mathrm{DA} + \Delta\mathrm{PDI}$. This measure assumes that there is an underlying pre-managed income series and that managers use discretionary accruals to make the reported series smooth. More income smoothing is evidenced by a more negative correlation between ΔDA and ΔPDI. The regression model is shown as Eq. (1):

$$\Delta \mathrm{DA}_{it} = \delta_0 + \delta_1 \Delta \mathrm{PDI}_{it} + \delta_2 \mathrm{Pledge\_YN}_{i,t} \times \Delta \mathrm{PDI}_{it} + \xi_{i,t}. \quad (1)$$

We postulate that if the firms have their shares pledged, shareholders will have more incentive to smooth earnings. Thus, we use a dummy variable (Pledge_YN) to measure share pledges. Pledge_YN equals 1 if the company's shares have been pledged, and 0 otherwise. We expect that both δ_1 and δ_2 will be significantly negative.

To help illustrate our model, consider the following reasoning:

Because $\Delta E_t = \Delta \mathrm{DA}_t + \Delta \mathrm{PDI}_t$ by replacing ΔDAt with Eq. (1) we have model (1-1):

$$\Delta E_t = \Delta \mathrm{DA}_t + \Delta \mathrm{PDI}_t = \delta_0 + (1 + \delta_1)\Delta \mathrm{PDI}_t + \delta_2 \mathrm{Pledge\_YN}_t \times \Delta \mathrm{PDI}_t + \xi_t, \quad (1\text{-}1)$$

when $\mathrm{Pledge\_YN}_t = 0$, model (1-1) can be rewritten as

$$\Delta E_t = \Delta \mathrm{DA}_t + \Delta \mathrm{PDI}_t = \delta_0 + (1 + \delta_1)\Delta \mathrm{PDI}_t + \xi_t.$$

If we calculate the variance on both sides, we have

$$Var(\Delta E_t) = (1 + \delta_1)^2 Var(\Delta \mathrm{PDI}_t) + \sigma^2.$$

If $\sigma^2 \ll Var(\Delta \mathrm{PDI}_t)$, then we have model (1-2):

$$\text{Income-smoothing coefficient } 1(\mathrm{ISC1}) = Var(\Delta E_t)/Var(\Delta \mathrm{PDI}_t) \approx (1 + \delta_1)^2. \quad (1\text{-}2)$$

If $\delta_1 < 0$, then $Var(\Delta E_t)/Var(\Delta \mathrm{PDI}_t) < 1$, which means that the volatility of actual earnings is lower than that of pre-managed earnings. Therefore, the company will probably smooth its earnings.

- when $\mathrm{Pledge\_YN}_t = 1$, model (1-1) can be rewritten as

$$\Delta E_t = \Delta \mathrm{DA}_t + \Delta \mathrm{PDI}_t = \delta_0 + (1 + \delta_1 + \delta_2)\Delta \mathrm{PDI}_t + \xi_t.$$

If we calculate the variance on both sides, we have

$$Var(\Delta E_t) = (1 + \delta_1 + \delta_2)^2 Var(\Delta \mathrm{PDI}_t) + \sigma^2.$$

If $\sigma^2 \ll Var(\Delta \mathrm{PDI}_t)$, then we have model (1-3):

$$\text{Income-smoothing coefficient } 2(\mathrm{ISC2}) = Var(\Delta E_t)/Var(\Delta \mathrm{PDI}_t) \approx (1 + \delta_1 + \delta_2)^2. \quad (1\text{-}3)$$

If $\delta_2 < 0$, this indicates that the earnings smoothing coefficient for pledged firms (ISC2) is smaller than that of other firms (ISC1). Therefore, pledged firms smooth earnings more than other firms. As a conclusion, model 1 can be used to test whether pledged firms smooth earnings.

To test our hypothesis on the effect of the split-share reform, we partition our sample into two periods: 2004–2006 and 2007–2011.

4.2. Estimation of discretionary accruals

To estimate discretionary accruals (DA) in model 1, we use the cross-sectional version of the Jones model, modified by Kothari et al. (2005):

$$\mathrm{TA}_{i,t} = \delta_0 + \delta_1(1/A_{it-1}) + \delta_2 \Delta \mathrm{REV}_{i_t} + \delta_3 \mathrm{PPE}_{i_t} + \delta_4 \mathrm{ROA}_{i_t} + \mu_{it}, \quad (2)$$

where $\mathrm{TA}_{i,t}$ represents the total accruals of firm i at time t, $\Delta \mathrm{REV}_{i_t}$ is the change in sales scaled by lagged total assets, A_{it-1}, and PPE_{i_t} is net tangible long-term assets scaled by A_{it-1}. ROA_{i_t} is the return on total assets in the current year, which is included because research shows that the Jones model is misspecified for well-performing or poorly performing firms.

Because there are many arguments of DA, we also estimate a year-dummy DA following Bergstresser and Philippon (2006):

$$\mathrm{TA}_{i,t} = \alpha_0 + \alpha_1 \times (1/A_{i,t-1}) + \alpha_2 \times (\Delta\mathrm{REV}_{i,t}/A_{i,t-1}) + \alpha_3 \times (\mathrm{PPE}_{i,t}/A_{i,t-1}) + (\sum_{year=y} \gamma_y) + \varepsilon_{i,t}. \tag{3}$$

Next, we remove the components of accruals that are "nondiscretionary" (NDA). We then estimate nondiscretionary accruals as the fitted value from the regression of total accruals in model 2. The DA are the deviations of actual accruals from NDA:

$$\mathrm{DA} = \mathrm{TA} - \mathrm{NDA}. \tag{4}$$

4.3. Sample selection and descriptive statistics

In the Chinese capital market, there are three types of firms: state-owned firms (SOEs), family firms, and firms controlled by towns and villages (Xia, 2008). Comparatively, the shareholders of SOEs find it easier to get loans from banks (Tsai, 2004). Share pledge announcements are made mostly by non-SOE firms. Further, because SOE firms have different agency problems and more government interference, their earnings management behavior could be influenced by the government. Therefore, we decide to test our hypothesis using non-SOE firms. Because listed companies are required to disclose their ultimate controller, it is easy to determine whether a firm is ultimately controlled by the state. We start with an initial sample of 3272 firm year observations (Table 1). Then we remove firms with missing data in calculating DA, leaving 2632 firm year observations (Table 2).

In this paper, we focus on the share pledge announcements made between 2004 and 2011 because there were very few share pledge announcements before 2004. We also want to observe the effect of the split share reform in

Table 1
Distribution of share pledging firms between 2004 and 2011.

	2004	2005	2006	2007	2008	2009	2010	2011	Total
Panel A: Distribution of firms by year									
Firms	286	308	349	349	388	462	508	622	3272
Pledging firms	56	97	126	124	156	198	221	273	1251

Industry description	Total sample firms	Firms pledged shares	Total number of pledged shares (million)
Panel B: Distribution of firms by industry			
Agriculture industry	80	40	3524.51
Mining industry	7	3	107.60
Textile industry	203	88	11006.74
Papermaking and paper product industry	87	36	3526.28
Petrochemical industry	301	97	11731.77
Industry of rubber and plastic products	47	18	2361.34
General equipment manufacturing	504	195	15182.01
Pharmaceutical industry	296	111	6650.73
Other manufacturing industries	704	209	23576.09
Industry of gas and electric power	23	4	142.19
Construction industry	62	17	1411.10
Transport, storage and postal service industry	36	14	804.69
Industry of information transmission, software and information technology service	258	126	7796.69
Retail industry	188	61	3723.91
Real estate industry	189	117	30334.51
Other service industries	73	16	1105.10
Press and publishing industry	9	7	219.50
Diversified industries	205	92	8670.28
Total	3272	1601	131875.04

Table 2
Descriptive statistics (2004–2011).

	N	Mean	Std. dev	Minimum	Median	Maximum
Year-dummy DA (one year lag)	2632	0.019	0.163	−0.934	0.017	0.955
Kothari model DA (one year lag)	2633	0.018	0.167	−0.971	0.018	1.059
Year-dummy DA	2629	0.017	0.163	−0.934	0.016	0.959
Kothari model DA	2631	0.015	0.163	−0.971	0.013	1.059
Pledge-YN	2687	0.160	0.367	0.000	0.000	1
Pledge-AM (million)	431	107.8	271.7	2.500	48.24	3820
Pledge-Rate	431	0.749	0.572	0.026	0.651	4.047
Block	2685	33.26	14.25	4.440	29.81	89.41
$\ln(A_{i,t-1})$	2685	21.07	1.187	11.35	21.04	24.81
$LEV_{i,t-1}$	2685	0.885	4.749	0.000	0.512	138.4
Growth	2685	0.386	6.331	−10.45	0.060	246.7
StdCFO	2687	0.429	1.901	0.005	0.149	29.95

Year-dummy DA is calculated by models (3) and (4), the Kothari model DA is calculated by models (2) and (4), and Pledge_YN is a dummy variable that equals 1 if the company's shares have been pledged, and 0 otherwise. Pledge_AM represents the number of shares that have been pledged (by million Yuan). Pledge-Rate represents the percentage of pledged shares to the total shares owned by the largest shareholder. $Ln(A_{i,t-1})$ is the log of total assets in year $t-1$. $Lev_{i,t-1}$ is the debt ratio in year $t-1$. Growth is the sales increase rate equal to $(\text{Sales}_t - \text{Sales}_{t-1})/\text{Sales}_{t-1}$. StdCFO is the standard deviation of operational net cash flow between 2004 and 2011.

2006. By the end of 2006, 95% of China's publicly listed companies had participated in the split share structure reform (Li and Zhang, 2011). Therefore, we partition our sample into two periods: 2004–2006 and 2007–2011. Before the reform, there were relatively fewer share pledges (Panel A of Table 1) because the shares held by large shareholders were non-tradable and subject to stricter regulations. However, after the split share reform, the non-tradable shares became tradable, and the number of pledged shares increased dramatically.

The pledge data are available in the WIND database. Other accounting and financial data of the listed companies are obtained from the CCER database.

Panel B of Table 1 provides the distribution of our sample by industry. Companies in the manufacturing sector account for the largest number of pledged shares, followed by the real estate industry and machinery industry. Apparently, pledging is actively conducted in some of the capital-intensive industries probably due to their limited sources of funds.

Table 2 reports descriptive statistics for the sample firm-year observations.

Firms that pledged shares constitute 16 percent of our sample (the mean of Pledge-YN). On average, large shareholders pledged 74.9% of their shares in one year. The level of DA calculated using the Kothari model has a lower mean than year-dummy DA. Table 2 displays the mean and median values of the control variables.

5. Empirical results

5.1. Share pledges and earnings smoothing

Table 3 provides the results of the multivariate regression analysis of the effect of share pledges on earnings smoothing. Columns (1) and (2) are the results for the full sample. Two models are presented, each with one proxy for share pledges: the binary variable (Pledge_YN). Both models are significant at the 1% level. The adjusted R^2 values are 0.45 or higher. The coefficient of ΔPDI is -0.422 with a p-value of <0.01, indicating that earnings smoothing is common in non-SOE firms. The coefficient on Pledge_YN $\times$ ΔPDI is -0.176 with a p-value of <0.01, suggesting that share pledging firms smooth earnings more than the others.

Table 3 shows the results for the two periods: before 2006 and after 2007. In both periods, the coefficients of ΔPDI are significantly negative. However, before the split share reform, the coefficient of Pledge_YN $\times$ ΔPDI is positive, which suggests that pledged firms do not conduct more earnings smoothing than other firms before the split share reform. This finding contrasts with that after the split share reform, in which pledged firms have a significantly negative coefficient of Pledge_YN $\times$ ΔPDI. During 2007–2011, the coefficient on Pledge_YN $\times$ ΔPDI is -0.361 with a p-value of <0.01, which indicates that the share pledges motivated firms

Table 3
The test of incoming smoothing.

Using the absolute value of year-dummy model ΔDA and the Kothari model ΔDA as dependent variables with the pledged sample
$\Delta DA_{it} = \delta_0 + \delta_1 \Delta PDI_{it} + \delta_2 Pledge\_YN_{i,t} \times \Delta PDI_{it} + \xi_{i,t}$

	Total sample		2004–2006		2007–2011	
	Year-dummy model ΔDA	Kothari model ΔDA	Year-dummy model ΔDA	Kothari model ΔDA	Year-dummy model ΔDA	Kothari model ΔDA
ΔPDI	-0.422^{***}	-0.449^{***}	-0.277^{***}	-0.322^{***}	-0.524^{***}	-0.565^{***}
	(−43.37)	(−47.55)	(−16.06)	(−20.20)	(−45.84)	(−51.73)
Pledge_YN × ΔPDI	-0.176^{***}	-0.178^{***}	0.093^{*}	0.097^{**}	-0.361^{***}	-0.371^{***}
	(−6.091)	(−6.468)	1.787	2.007	(−11.19)	(−12.05)
Constant	0.05	−0.003	−0.004	0.002	0.063	−0.001
	(−0.744)	(−0.839)	(−0.022)	0.303	1.004	(−0.218)
N	2586	2588	661	661	1924	1926
Adj. *R*-SQ	0.459	0.524	0.317	0.394	0.608	0.659
F-stat.	366.43	1424.3	52.24	215.92	498.31	1865.4

DA is calculated by models (2), (3), and (4). PDI is pre-discretionary income, which is calculated as net income minus discretionary accruals (PDI = E-DA). E is deflated by the beginning-of-year total assets. Pledge_YN is a dummy variable that equals 1 if the company's shares have been pledged, and 0 otherwise.

* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

Table 4
The effect of ownership structure on income smoothing.

Using the absolute value of year-dummy model ΔDA and the Kothari model ΔDA as dependent variables with the pledged sample
$\Delta DA_{it} = \delta_0 + \delta_1 \Delta PDI_{it} + \delta_2 Block_{i,t} \times \Delta PDI_{it} + \xi_{i,t}$

	Pledging sample		2004–2006 Pledging sample		2007–2011 Pledging sample	
	Year-dummy model ΔDA	Kothari model ΔDA	Year-dummy model ΔDA	Kothari model ΔDA	Year-dummy model ΔDA	Kothari model ΔDA
ΔPDI	−0.229***	−0.312***	−0.011	−0.016	−0.792***	−0.886***
	(−3.686)	(−5.306)	(−0.036)	(−0.053)	(−20.76)	(−27.93)
Block × ΔPDI	−0.011***	−0.010***	−0.006	−0.009	−0.003***	−0.001
	(−6.096)	(−5.777)	(−0.531)	(−0.845)	(−2.800)	(−1.626)
Constant	−0.017**	−0.019***	−0.005	−0.013	−0.006	−0.007**
	(−2.387)	(−2.848)	(−0.174)	(−0.466)	(−1.434)	(−2.166)
N	420	420	59	59	360	360
Adj. *R*-SQ	0.607	0.664	0.120	0.231	0.900	0.934
F-stat.	325.08	416.42	5.034	9.850	1615.4	2567.9

Where DA is calculated by models (2), (3), and (4). PDI is pre-discretionary income, which is calculated as net income minus discretionary accruals (PDI = E-DA). E is deflated by the beginning-of-year total assets. Block is the percentage of shares held by the largest shareholder.

* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

Table 5
Ordinary least-squares regressions predicting one-year post-pledging returns with share pledge and controls.

	Predicting one-year post-pledging returns with pledging level (sample: pledged firms and matched firms)					
Independent variables	Raw returns			Market-adj. return		
Pledge_Dummy	0.085*			0.095*		
	(1.654)			(1.849)		
Ln(Pledge-AM)		0.012**			0.012**	
		(2.076)			(2.133)	
Pledge frequency			0.137**			0.137**
			(2.061)			(1.738)
Ln($Size_{t-1}$)	−0.055*	−0.060*	−0.054*	−0.042	−0.047	−0.041
	(−1.739)	(−1.887)	(−1.719)	(−1.342)	(−1.493)	(−1.306)
Lev_{t-1}	−0.137	−0.135	−0.128	−0.185	−0.185	−0.176
	(−0.939)	(−0.931)	(−0.878)	(−1.258)	(−1.262)	(−1.197)
$BTM_{t\text{-}1}$	0.505***	0.514***	0.491***	0.026	0.034	0.009
	(2.879)	(2.931)	(2.809)	(0.148)	(0.193)	(0.049)
beta	0.009	0.01	0.01	0.008	0.008	0.009
	(0.498)	(0.525)	(0.569)	(0.442)	(0.457)	(0.497)
Year Dummy	Control					
Constant	1.605**	1.695***	1.578**	0.858	0.953	0.829
	−2.539	−2.68	−2.501	−1.379	−1.53	−1.332
n	440	440	440	440	440	440
Adj. R^2	0.288	0.291	0.29	0.018	0.02	0.02

The dependent variables are raw returns and market adjusted returns, calculated by the following models (a) and (b). We apply three measures of pledge (Pledge-Dummy, which equals one for pledged firms and zero for the matched sample; Ln (Pledge-AM) is the natural logarithm of the amount of shares being pledged; and pledge frequency is the number of share pledge announcements made within one year). Ln(Size) is the natural logarithm of total assets. LEV_{t-1} is the debt ratio in year $t-1$. BTM is calculated by the book value per share deflated by the closing price one week before pledge. Beta is based on the regression of the individual share weekly risk premium (50–100 weeks before pledge) on the equity weighted market risk premium (CAPM model).

Annual raw returns $= \prod_{t=-1}^{49}(1+R_{it}) - 1$ (a)

Annual market adjusted returns $= \prod_{t=-1}^{49}(1+R_{it}) - \prod_{t=-1}^{49}(1+R_{mt})$ (b)

* Significant at 10%.

** Significant at 5%.

*** Significant at 1%.

to smooth their earnings. Therefore, we conclude that the effect of share pledges on earnings smoothing is primarily driven by the sample after the split share reform. Our findings imply that when large shareholders have a strong motivation to manage earnings, the ownership structure facilitates them to do so.

5.2. Additional analysis: Ownership concentration and earnings smoothing

Although we have found evidence supporting our hypothesis that share pledging firms are more likely to smooth earnings, we still do not know whether the ownership structure played any role. Therefore, we use the share pledging firms as a sample, and test the effect of ownership concentration on earnings smoothing. The results are presented in Table 4. The results for the total pledging sample show that the estimated coefficient of ΔPDI is negative and statistically significant at the 1% level. This result is consistent with Table 3 that share pledging firms smooth earnings. We find that the coefficient on Block × ΔPDI is −0.01 with a p-value of <0.01, which suggests that the greater the share ownership, the more the shareholders smooth earnings. As shown in Table 3, we only find significant results during the 2007–2011 period with the year-dummy DA model. This suggests that large shareholders have a similar ability to smooth earnings before and after the split share reform. However, their financial reporting behavior is associated with their motivation to pledge.

5.3. Additional analysis: Share pledges and market return

We predict that when firms make share pledge announcements, they manage earnings to influence the share price. However, if the stock market can observe this earnings management, the share prices will not increase,

Table 6
Sensitivity test of incoming smoothing.

Using the absolute value of year-dummy model ΔDA and the Kothari model ΔDA as dependent variables
$\Delta DA_{it} = \delta_0 + \delta_1 \Delta PDI_{it} + \delta_2 Pledge\_YN_{i,t}(Pledge\_AM_{i,t}) \times \Delta PDI_{it} + X'_{i,t}B + \xi_{i,t}$

	Total sample				2004–2006				2007–2011			
	Year-dummy model ΔDA		Kothari model ΔDA		Year-dummy model ΔDA		Kothari model ΔDA		Year-dummy model ΔDA		Kothari model ΔDA	
ΔPDI	−0.405***	−0.645***	−0.447***	−0.461***	−0.277***	−0.541***	−0.318***	−0.318***	−0.524***	−0.542***	−0.569***	−0.586***
	(−41.19)	(−70.83)	(−47.30)	(−49.24)	(−16.06)	(−28.91)	(−19.12)	(−19.25)	(−45.84)	(−47.99)	(−52.49)	(−54.73)
Pledge_YN × ΔPDI	−0.176***		−0.184***		0.093*		0.073		−0.361***		−0.359***	
	(6.091)		(−6.628)		(1.787)		(1.452)		(−11.19)		(−11.74)	
Ln(Pledge-AM) × ΔPDI		−0.007***		−0.046**		0.007		0.048*		−0.212***		−0.213***
		(−2.890)		(−2.517)		(1.488)		(1.677)		(−8.601)		(−9.138)
Size_1	−0.003	0.001	−0.004	−0.005	0.000	−0.004	0.007	0.007	−0.004	−0.004	−0.006**	−0.007***
	(−0.891)	(0.437)	(−1.549)	(−1.588)	(−0.013)	(−0.699)	(0.854)	(0.861)	(−1.279)	(−1.448)	(−2.361)	(−2.604)
Lev_1	0.011*	0.000	0.009	0.011*	0.016	0.000	0.015	0.015	0.032***	0.033***	0.032***	0.033***
	(1.747)	(−0.364)	(1.564)	(1.812)	(1.087)	(−0.123)	(1.056)	(1.070)	(5.170)	(5.296)	(5.432)	(5.536)
Growth	0.020**	0.000	0.029***	0.026***	0.031	0.000	0.039**	0.038**	−0.007	−0.008	0.001	0.001
	(2.337)	(0.144)	(3.505)	(3.131)	(1.634)	(−0.651)	(2.157)	(2.064)	(−0.860)	(−0.921)	(0.118)	(0.070)
StdCFO	−0.001	0.001	−0.001	−0.001	−0.006	−0.005	−0.005	−0.005	0.001	0.001	0.001	0.001
	(−0.340)	(0.149)	(−0.445)	(−0.473)	(−1.592)	(−0.637)	(−1.502)	(−1.514)	(0.721)	(0.756)	(0.485)	(0.519)
Constant	0.050	−0.015	0.084	0.086	−0.004	0.099	−0.151	−0.152	0.063	0.073	0.117**	0.131**
	(0.744)	(−0.329)	(1.359)	(1.385)	(−0.022)	(0.798)	(−0.897)	(−0.904)	(1.004)	(1.147)	(2.019)	(2.235)
N	2586	2586	2588	2588	661	936	661	661	1924	1924	1926	1926
Adj. *R*-SQ	0.459	0.673	0.526	0.519	0.317	0.544	0.398	0.399	0.608	0.598	0.666	0.657
F-stat.	366.43	1111.8	479.89	66.86	52.24	186.62	73.91	74.10	498.31	477.77	641.45	616.11

Where DA is calculated by models (2), (3), and (4). PDI is pre-discretionary income, which is calculated as net income minus discretionary accruals (PDI = E-DA). E is deflated by the beginning-of-year total assets. Pledge_YN is a dummy variable that equals 1 if the company's shares have been pledged, and 0 otherwise. Pledge_AM represents the number of shares being pledged. $\ln(A_{t-1})$ is the log of total assets in year $t-1$. LEV_{t-1} is the debt ratio in year $t-1$. *Growth* is the sales increase equal to $(Sales_t - Sales_{t-1})/Sales_{t-1}$. *MPB* is the market to book ratio calculated by the average annual price divided by EPS in year $t-1$. StdCFO is the standard deviation of operational net cash flow between 2004 and 2011.

* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

and firms will have no reason to keep manipulating earnings. Teoh et al. (1998a) point out that IPO issuers report unusually high earnings before an IPO, and buyers are guided by such earnings and pay a high price for the shares. Therefore, we need to examine the market reaction and long-term return of share pledging firms to confirm our postulation.

We run an ordinary least squares regression of one-year post-pledging returns with the pledge indicators (see Teoh et al., 1998b), using the sample between 2009 and 2011. We compare the share pledging firms with the matched sample. The results in Table 5 reveal that pledging firms have better returns, and that the higher the value of the shares pledged, and the more frequent the share pledge, the better the returns. Following Teoh et al. (1998b), we use both raw returns and market adjusted returns for the test. Both measures of returns show similar results.

5.4. Robustness tests

We conduct two sensitivity tests. First, we use a continuous variable (Pledge-AM) to measure share pledges. Pledge-AM is the amount that shareholders pledge, in billion RMB. In Table 6, the coefficient on Pledge-AM is −0.046, which suggests that the share pledge level is positively related to smoothing. We also include control variables in our main test model. A number of factors that are important determinants of DA have been identified in the earnings management literature. Research generally shows that total assets, leverage, growth, and operational cash flow may affect the magnitude of earnings management. The regression results are similar to our previous tests.

6. Conclusion

In this paper, we re-examine the relationship between ownership structure and discretionary financial reporting. The literature provides two different views on this relationship. Entrenchment theory holds that ownership concentration is associated with lower earnings quality. In contrast, alignment theory posits that controlling shareholders may provide higher quality reports. We investigate firms whose shares are pledged by their shareholders. Specifically, we test whether these share pledging firms smooth their earnings more than other firms and how their behavior changes with the incentives triggered by the split share reform. Before the split share reform, shareholders held large amounts of non-tradable shares. Pledging shares for loans was the only way to transfer their shareholdings. The share pledges were not based on the share price, but the book value of the firm. Thus, the firms had no incentive to smooth earnings.

After the split share reform, large shareholders could choose to sell their shareholdings or pledge their shareholdings for loans. The share price became the primary factor in deciding the amount of a loan. To facilitate loan financing and avoid violation of debt covenants and maintenance calls, shareholders wish to see an increase in the share price. Therefore, we predict that the market performance of share pledging firms is better than the others. If the share price drops after a share pledge, the large shareholder will do whatever it can to uphold the share price. Managing annual earnings is unlikely to serve as a timely measure. For example, companies may repurchase, release other good news, or manipulate quarterly earnings. However, accounting earnings represent the fundamental profitability of a firm, and annual earnings announcements affect share prices. Therefore, firms may need to strategically report their annual earnings. Firms pledge repeatedly, and thus we predict that share pledging firms smooth their earnings more than other firms.

We use Chinese non-SOE firms in our analysis because non-government controlling shareholders have more difficulty obtaining loans in the Chinese banking system. Comparing SOE and non-SOE firms, we find that non-SOE firms have a significantly higher probability of pledging than SOE firms. Another reason we exclude SOE firms from our sample is that SOE firms sometimes conduct special tasks for political reasons or according to macroeconomic policy.

Our results support our hypothesis. Using Tucker and Zarowin's (2006) model, we find that the relationship between the change in the DA proxy and change in pre-discretionary income is significantly negative for share pledging firms. Furthermore, this effect does not exist before the split share reform, but becomes significant after the reform is implemented. With the 2009–2011 sample, we also find a marginally positive correlation between share pledges and market returns, indicating that share pledging firms have better market performance.

The findings presented in this paper contribute to the literature in several ways. First, we provide more direct evidence on the effect of ownership concentration on financial reporting. Second, our results imply that share pledges can trigger earnings management, which should serve as a reminder for investors and regulators. Like most studies, our paper is not without its limitations. It is difficult to measure earnings smoothing, and most of the existing models could not be used in our setting. Although we use a model with and without control variables, we admit that our measures may not fully capture the dynamics of earnings smoothing. Finally, we acknowledge that share pledging is a very special setting that can only be found in a few countries.

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References

Aharony, J., Wang, J., Yuan, H., 2010. Tunneling as an incentive for earnings management during the IPO process in China. J. Account. Public Policy 29 (1), 1–26.
Ahn, S., Choi, W., 2009. The role of bank monitoring in corporate governance. evidence from borrowers' earnings management behavior. J. Bank. Finance 33 (2), 425–434.
Arya, A., Glover, J., Sunder, S., 1998. Earnings management and the revelation principle. Rev. Acc. Stud. 3 (1–2), 7–34.
Ball, R., Robin, A., Wu, J.S., 2003. Incentives versus standards: properties of accounting income in four East Asian countries. J. Account. Econ. 36 (1–3), 235–270.
Beidleman, C.R., 1973. Income smoothing: the role of management. Account. Rev. 4 (48), 653–667.
Bergstresser, D., Philippon, T., 2006. CEO incentives and earnings management. J. Financ. Econ. 80 (3), 511–529.
Burgstahler, D.C., Hail, L., Leuz, C., 2006. The importance of reporting incentives: earnings management in European private and public Firms. Account. Rev. 5 (81), 983–1016.
Caton, G.L., Chiyachantana, C.N., Chua, C., Goh, J., 2011. Earnings management surrounding seasoned bond offerings: do managers mislead ratings agencies and the bond market? J. Financ. Quant. Anal. 46 (03), 687–708.
Chan, K., Chen, H., Hu, S., Liu, Y., 2013. Shares Pledged and Corporate Repurchase. National Chengchi University.
Chen, G., Firth, M., Xin, Y., Xu, L., 2008. Control transfers, privatization, and corporate performance. efficiency gains in China's listed companies. J. Financ. Quant. Anal. 1 (43), 161–190.
Chen, G., Firth, M., Xu, L., 2009. Does the type of ownership control matter? Evidence from China's listed companies. J. Bank. Finance 33 (1), 171–181.
Claessens, S., Djankov, S., Fan, J.P.H., Lang, L.H.P., 2002. Disentangling the incentive and entrenchment effects of large shareholdings. J. Finance 57 (6), 2741–2771.
Dechow, P.M., Skinner, D.J., 2000. Earnings management: reconciling the views of accounting academics, practitioners, and regulators. Account. Horizons 2 (14), 235–250.
DeFond, M.L., Jiambalvo, J., 1994. Debt covenant violation and manipulation of accruals. J. Account. Econ. 17 (1–2), 145–176.
Dichev, I.D., Skinner, D.J., 2002. Large-sample evidence on the debt covenant hypothesis. J. Account. Res. 40 (4), 1091–1123.
Dou, Y., Hope, O., Thomas, W.B., 2013. Relationship-specificity, contract enforceability, and income smoothing. Account. Rev. 88 (5), 1629–1656.
Dye, R.A., 1988. Earnings management in an overlapping generations model. J. Account. Res. 2 (26), 195–235.
Fan, J.P.H., Wong, T.J., 2002. Corporate ownership structure and the informativeness of accounting earnings in East Asia. J. Account. Econ. 33 (3), 401–425.
Francis, J., LaFond, R., Olsson, P., Schipper, K., 2005a. The market pricing of accruals quality. J. Account. Econ. 39 (2), 295–327.
Francis, J., Schipper, K., Vincent, L., 2005b. Earnings and dividend informativeness when cash flow rights are separated from voting rights. J. Account. Econ. 39 (2), 329–360.
Guthrie, K., Sokolowsky, J., 2010. Large shareholders and the pressure to manage earnings. J. Corp. Finance 16 (3), 302–319.
Healy, P.M., 1985. The effect of bonus schemes on accounting decisions. J. Account. Econ. 7 (1), 85–107.
Hou, W., Lee, E., 2014. Split Share Structure Reform, corporate governance, and the foreign share discount puzzle in China. Eur. J. Finance 7–9 (20), 703–727.
Hu, Y., Li, S., Lin, T.W., Xie, S., 2011. Large creditors and corporate governance: the case of Chinese banks. Rev. Account. Finance 10 (4), 332–367.

Jaggi, B., Lee, P., 2002. Earnings management response to debt covenant violations and debt restructuring. J. Account., Audit. Finance 4 (17), 295–324.

Kao, L., Chen, A., 2007. Directors' share collateralization, earnings management and firm performance. Taiwan Account. Rev. 2 (6), 153–172.

Kao, H., Wei, T., 2014. The effect of IFRS, information asymmetry and corporate governance on the quality of accounting information. Asian Econ. Financ. Rev. 2 (4), 226–256.

Kao, L., Chiou, J., Chen, A., 2004. The agency problems, firm performance and monitoring mechanisms: the evidence from collateralised shares in Taiwan. Corp. Gov.: Int. Rev. 12 (3), 389–402.

Kothari, S.P., Leone, A.J., Wasley, C.E., 2005. Performance matched discretionary accrual measures. J. Account. Econ. 39 (1), 163–197.

Leuz, C., Nanda, D., Wysocki, P.D., 2003. Earnings management and investor protection: an international comparison. J. Financ. Econ. 69 (3), 505–527.

Li, X., Zhang, B., 2011. Has split share structure reform improved the efficiency of the Chinese stock market? Appl. Econ. Lett. 18 (11), 1061–1064.

Liu, Q., Lu, Z.J., 2007. Corporate governance and earnings management in the Chinese listed companies: a tunneling perspective. J. Corp. Finance 13 (5), 881–906.

Liu, G., Sun, J., 2010. Ultimate ownership structure and corporate disclosure quality: evidence from China. Manage. Finance 36 (5), 452–467.

Park, Y.W., Shin, H., 2004. Board composition and earnings management in Canada. J. Corp. Finance 10 (3), 431–457.

Ronen, J., Sadan, S., 1981. Smoothing income Numbers: Objectives, Means, and Implications. Addison-Wesley Publishing Company.

Shleifer, A., Vishny, R.W., 1997. A survey of corporate governance. J. Finance 52 (2), 737–783.

Sun, Q., 2010. Research on the share pledges of public firms and risks of banks. Account. Friends (3), 83–84.

Tan, Y., Wu, J., 2013. Does share pledge have governance effect? Evidence from Chinese public firms. Account. Res. (2), 45–53

Teoh, S.H., Welch, I., Wong, T.J., 1998a. Earnings management and the long-run market performance of initial public offerings. J. Finance 53 (6), 1935–1974.

Teoh, S.H., Welch, I., Wong, T.J., 1998b. Earnings management and the underperformance of seasoned equity offerings. J. Financ. Econ. 50 (1), 63–99.

Trueman, B., Titman, S., 1988. An explanation for accounting income smoothing. J. Account. Res. 26, 127–139.

Tsai, K.S., 2004. Back-alley Banking: Private Entrepreneurs in China. Cornell University Press.

Tucker, J.W., Zarowin, P.A., 2006. Does income smoothing improve earnings informativeness? Account. Rev. 81 (1), 251–270.

Wang, D., 2006. Founding family ownership and earnings quality. J. Account. Res. 44 (3), 619–656.

Warfield, T.D., Wild, J.J., Wild, K.L., 1995. Managerial ownership, accounting choices, and informativeness of earnings. J. Account. Econ. 20 (1), 61–91.

Xia, L., 2008. Founder control, ownership structure and firm value: evidence from entrepreneurial listed firms in China. China J. Account. Res. 1, 31–49.

Yeh, Y., Shu, P., Lee, T., Su, Y., 2009. Non-tradable share reform and corporate governance in the Chinese stock market. Corp. Govern: Int. Rev. 17 (4), 457–475.

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Female CFOs and loan contracting: Financial conservatism or gender discrimination? – An empirical test based on collateral clauses

Xixiong Xu *, Yaoqin Li, Mengmeng Chang

School of Economics & Business Administration, Chongqing University, China

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ABSTRACT

Based on signaling and gender discrimination theory, we examine whether chief financial officer (CFO) gender matters to bank–firm relationships and the designing of collateral clauses in bank loan contracting, and explore the potential path of influence. Data taken from Chinese listed companies between 2009 and 2012 indicate that (1) female-CFO-led firms are less likely to obtain credit loans than male-CFO-led firms; (2) female-CFO-led borrowers are more likely to be required to provide collateral for loans than male-CFO-led borrowers; and (3) banks are more inclined to claim mortgaging collateral when lending to female-CFO-led firms and prefer to guarantee collateral when lending to male-CFO-led firms. Female-CFO-led borrowers seem to be granted more unfavorable loan terms than male-CFO-led borrowers, supporting the hypothesis that female CFOs experience credit discrimination. Further analysis reveals that regional financial development helps to alleviate lending discrimination against female CFOs. Furthermore, female CFOs in SOEs are less likely than their non-SOE counterparts to experience gender discrimination in the credit market.

* Corresponding author.
E-mail address: xuxixiong@cqu.edu.cn (X. Xu).

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1. Introduction

Many studies have shown that financial development plays an important role in promoting economic growth (Rajan and Zingales, 1996; Levine, 2005). However, compared with Western countries, the security market in China is less developed and the financial system is dominated by the banking sector (Chen et al., 2013; Gou et al., 2014). Under these conditions, bank loans are major sources of corporate financing, even for large public companies. For instance, the average ratio of the volume of bank loan financing to GDP from 1990 to 2008 in China was 82.4%, and the average ratios of equity volume and bond financing to GDP were only 0.81% and 0.76%, respectively.[1] In fact, even in the United States, the volume of bank loan financing is much larger than that of equity and bond financing combined. Given the economic significance of bank loans in allocating capital to corporations, considerable effort has been made to investigate the determinants of bank loan contracting (e.g., Qian and Strahan, 2007; Graham et al., 2008; Chava et al., 2009; Hasan et al., 2012; Ge et al., 2012). Different from the current literature, this study focuses on the effect of CFO gender on bank–firm relationships and attempts to explore whether and how female CFOs exert a significant influence on the designing of bank loan terms.

Over the past decade, the number of female executives (especially female CFOs) has increased dramatically.[2] As their education levels and social status improve, female executives are playing an increasingly important role in corporate decisions. The significant increase in female CFOs has also attracted considerable attention from academics. An emerging stream of literature is beginning to investigate the systematic differences between male and female CFOs in terms of their accounting, financing and investment decisions (Francis et al., 2013). For example, studies have indicated that firms with female CFOs adopt more conservative accounting policies (Francis et al., 2014), are less likely to manipulate earnings (Chava and Purnanandam, 2010; Liu et al., 2015) and are less likely to make significant acquisitions but more likely to decrease leverage levels than firms with male CFOs (Huang and Kisgen, 2013).

In the bank loan literature, accounting information is an important standard that banks rely on to assess borrowers' credit risk. Earlier studies find that when banks initiate private debt, they are very sensitive to various accounting information attributes, such as operating accruals (Bharath et al., 2008) and conservatism (Sunder et al., 2009; Zhang, 2008). Assuming that female CFOs are more likely to report high-quality and conservative earnings than male CFOs, as documented in earlier studies, banks should realize the benefits of female CFOs in providing more reliable and conservative accounting information to lenders (Francis et al., 2014) and thus reward borrowers under the control of female CFOs with favorable loan terms.

However, finance studies have long suggested the notion of credit discrimination, which indicates that discrimination in the credit market occurs when the decisions lenders make in relation to loan applications are influenced by personal characteristics that are not relevant to the transactions, such as the gender or race of the top executives of the borrowing firm. In the classical model adopted by Becker (1957), discrimination arises due to the taste-based preferences of the lender, who is willing to pay a price to avoid being associated with certain groups of borrowers. Alesina et al. (2013) provide evidence that female-managed firms are less likely than their male-managed counterparts to obtain bank loans and are charged higher interest rates when loan applications are approved. Similar results are presented by Bellucci et al. (2010), who find that female entrepreneurs face tighter credit availability. Thus, the presence of credit discrimination against women reveals that banks may charge higher loan prices and require stricter non-price terms when lending to female-CFO-led companies.

An important implication of the preceding discussion is that female CFOs may not necessarily obtain favorable or unfavorable loan contracts. Therefore, the direction of the relationship between the two remains an open empirical question. The main purpose of this study is to explore whether and how CFO gender influences the designing of collateral clauses in bank loan contracts. Examining this issue in the Chinese setting is interesting and important for several reasons. First, bank loans are a major source of external financing in

[1] *Data source*: http://www.pbc.gov.cn/publish/html/kuangjia.htm?id=2013s01.htm.

[2] For example, among the major U.S. corporations in 2005, 7.5% of chief financial officers (CFOs) and 1.5% of chief executive officer (CEOs) were women, versus 3.0% and 0.5% in 1994, respectively. According to Grant Thornton's "Report on International Business Questionnaire," the ratio of female executives in Chinese companies was about 25%, an amount larger than the global average (21%).

China. Thus, it is economically important to investigate the effect of CFO gender on bank–firm relationships and the designing of bank loan contracts. Second, this question has attracted some attention in the bank loan literature, which broadly focuses on interest rates, maturities, loan size and possible covenants, while leaving the nature of collateral clauses in the background. However, due to the weakness of the credit environment in transitional China, collateral requirements are common terms in bank loan contracts. Therefore, this paper helps us to gain additional insights into the determinants of loan contracts in emerging and transitional economies such as China, which may differ systematically from those predicted in developed economies.

From the perspectives of signaling theory and gender discrimination theory, we investigate the effect of CFO gender on the designing of collateral clauses in bank loan contracts and explore the potential path of influence. Using a sample of Chinese listed companies from 2007 to 2012, we find that CFO gender does affect bank loan contracts and that the collateral clauses granted to female-CFO-led companies are much stricter than those granted to their male-CFO-led counterparts. In our sample, firms with female CFOs are less likely to gain credit loans than firms with male CFOs and are more likely to be required to provide collateral if their loan applications are approved. Furthermore, loans given to female-CFO-led companies are more likely to be required to provide stricter pledging collateral or mortgaging collateral. On the contrary, loans given to male-CFO-led companies are more likely to provide looser guaranteeing collateral. Our empirical results provide support for the hypothesis of credit discrimination against female CFOs. Further analysis reveals that reform through market-oriented financial development helps to alleviate lending discrimination against female CFOs in the Chinese credit market. In addition, compared with female CFOs in non-SOEs, female CFOs in SOEs are less likely to suffer from gender discrimination in the credit market.

Our study makes two main contributions to the literature. First, it extends studies of the relationship between accounting information and bank loan contracting. It is well established that banks are the most important providers of external capital to corporations around the world. Accordingly, finance scholars have devoted considerable effort to understanding the factors that affect bank loan terms. Studies show that bank loan contracts are significantly influenced by borrowers' accounting quality and financial reporting transparency (Bharath et al., 2008; Armstrong et al., 2010), conservatism in financial policies (Sunder et al., 2009) and internal control weaknesses (Kim et al., 2011a), and country-level legal environments such as creditor rights and legal enforcement (Qian and Strahan, 2007; Bae and Goyal, 2009; Ge et al., 2012). This study attempts to relate bank loan decisions to the gender of the top executives of borrowing firms, and focuses on the effects of female CFOs on the designing of collateral clauses from the perspectives of both signaling theory and gender discrimination theory. Consistent with the gender discrimination hypothesis, our empirical results suggest that bank loan contracts granted to female-CFO-led companies are more unfavorable than those granted to male-CFO-led companies. In addition, regional financial development is helpful in mitigating lending discrimination against female CFOs. Our study expands and deepens the theoretical understanding of the determinants of bank loan contracts and provides evidence of the existence of gender discrimination in the Chinese credit market.

Second, this study contributes to the gender literature. A proliferation of studies mainly focuses on the systematic differences between male and female executives in terms of their risk attitudes (Dwyer et al., 2002; Atkinson et al., 2003), accounting and financial policies (Huang and Kisgen, 2013; Francis et al., 2014), leadership styles (Matsa and Miller, 2011) and agency costs and governance efficiency (Chava and Purnanandam, 2010; Adams and Ferreira, 2009). Rather than concentrating on internal corporate decisions, this study relates the gender of top executives to banks' lending decisions from an outside perspective and provides evidence that CFO gender also greatly influences loan contracts, thereby expanding and furthering our understanding of the economic consequences of female executives. In addition, despite increasing concerns over whether gender plays a role in corporate decisions, the research has mainly focused on chief executive officers (CEOs) or chairmen. To the best of our knowledge, this study is one of the few to concentrate on the role of female CFOs.

The remainder of this study is organized as follows. The next section describes the theoretical model and presents hypotheses for empirical testing. Section 3 outlines the sample and empirical research design. Section 4 presents the results. Section 5 discusses the implications of the findings and limitations of the study.

2. Theoretical analysis and hypotheses

2.1. Bank loan contracts

Banks are the dominant suppliers of external finance in most economies around the world (Graham et al., 2008; Kim et al., 2011b) and thus play an important role in allocating capital to corporations (Francis et al., 2012). The security market in transitional China is especially underdeveloped. As a result, the banking system is much larger than the equity market, and firms rely heavily on bank loans for their external financing needs. Due to information asymmetry in the credit market, banks continually face severe default risk from borrowers. In the bank loan literature, accounting information is the primary resource for banks to evaluate and predict the riskiness of borrowers (Anderson et al., 2004; Armstrong et al., 2010). For example, earlier studies show a negative relationship between bank loan pricing and earnings quality (Bharath et al., 2008) and conservatism (Zhang, 2008). Hasan et al. (2012) find that firms with more predictable earnings have more favorable loan contract terms. At the country level, Qian and Strahan (2007) indicate that borrowers in countries with strong creditor-protection environments enjoy loans with lower interest rates and longer maturities.

Although interest is an effective way to price credit risk, it also presents negative moral hazard problems (Stiglitz and Weiss, 1981). Therefore, collateral requirements along with interest rates, maturities, loan size and possible covenants are common terms in bank loan contracts that compensate for higher default risks, facilitate monitoring and limit potential losses (Jiménez et al., 2006; Qian and Strahan, 2007). In practice, bank loans can be divided into two categories, including credit loans and secured loans,[3] based on whether the borrowers are required to provide collateral. Collateral decreases the risk of debt in two important ways. First, it facilitates enforcement against a defaulting debtor. In case of a default, the creditor can seize the secured assets and satisfy the obligation. Second, it offers protection against competing claims by other creditors when an insolvent debtor faces liquidation. Collateral is therefore an important contractual device that influences the behavior of borrowers and lenders and the design of debt contracts (Cerqueiro et al., 2015). In fact, the use of collateral has become increasingly popular (Francis et al., 2012, 2013). According to Chen et al. (2013), 74% of the bank loans granted to Chinese listed companies have introduced collateral terms. A series of empirical results also show that collateral use is beneficial in decreasing the information asymmetry risk faced by commercial banks (Graham et al., 2008; Kim et al., 2011a).

2.2. Female CFOs and bank loan contracts: based on signaling theory

The behavioral differences exhibited between women and men are generally non-controversial. Gender differences in attitudes toward risk and risk-related behavior have long been studied in the sociology, psychology and economics fields. Most studies have supported the notion that women are more risk-averse than men (Knight, 2002). This gender difference is evidenced by the typically safer play behavior of girls relative to boys and by women's more cautious behavior relative to men in terms of sex, recreational drug use, alcohol consumption, gambling and driving (Sapienza et al., 2009). According to "behavioral consistency theory," individuals behave consistently across different situations. Therefore, it is reasonable to suppose that firms behave consistently based on how their top executives behave personally. For example, looking at professional money managers, Dwyer et al. (2002) find that female managers take fewer risks than male managers in their mutual fund investments.

Given the dramatic increase in female corporate executives over the past decade, a stream of studies considers the effect of the gender of top executives on various corporate decisions (Graham et al., 2013). For instance, Ahern and Dittmar (2012) document that the introduction of mandatory board member gender quotas led to increases in acquisitions and performance deterioration in Norwegian publicly traded firms. Faccio et al. (2015) find that firms run by female CEOs have lower leverage, less volatile earnings and a higher chance

[3] In light of the variance in binding strength of collateral agreements, there are three types of collateral forms: (1) guaranteeing collateral, in which case a third party promises to bear the joint obligation when a borrower fails to repay a loan; (2) mortgaging collateral, which requires the borrower or a third party to provide real property as collateral for the loan; and (3) pledging collateral, which requires the borrower or a third party to provide movable property or a warrant as collateral for the loan.

of survival than similar firms run by male CEOs. Huang and Kisgen (2010) concentrate on the effect of CFO gender on corporate financial decisions and indicate that firms under the control of female CFOs are less likely to make significant acquisitions and issue long-term debt. Furthermore, female CFOs are more likely to decrease leverage levels than male CFOs. A more recent study by Francis et al. (2014) documents that female CFOs tend to report more conservative accounting numbers than male CFOs. In addition, female CFOs appear to make less risky financing and investment decisions than their male counterparts. These results provide some supportive evidence that female CFOs are more risk-averse and conservative than male CFOs when making corporate financial decisions.

As mentioned previously, accounting information is a persistent standard that banks rely on to assess borrowers' credit risk. Studies have shown that when banks make lending decisions, they are very sensitive to various attributes of accounting information, such as operating accruals (Bharath et al., 2008) and conservatism (Sunder et al., 2009; Zhang, 2008). Based on signaling theory, assuming that female CFOs are more likely to report high-quality earnings and adopt conservative financial policies than male CFOs, as documented in earlier studies, banks should realize the benefits of female CFOs in providing more reliable and conservative accounting information to lenders. Thus, we conjecture that banks should acknowledge the positive gender effect and reward borrowing firms under the control of female CFOs with more favorable lending terms.

In addition to the risk associated with the availability of credible information, bank loan decisions are influenced by agency risk, which is mainly derived from the opportunistic activities of borrowers (Chava et al., 2009; Ge et al., 2012). Corporate executives are supposed to act in the interests of shareholders, but not necessarily in the interests of debt holders. Agency theory of debt implies that shareholders have incentives to transfer wealth from debt holders to shareholders (Jensen and Meckling, 1976). The business ethics literature normally suggests that women are more ethical than men in their attitudes and behavior (Ruegger and King, 1992; Nguyen et al., 2008). As a result of these personality characteristics, female executives are generally less likely to exhibit opportunistic behavior than their male counterparts (Srinidhi et al., 2011). Based on this logic, borrowers with female CFOs enjoy more favorable loan terms than those without female CFOs, as female CFOs help to mitigate conflict between shareholders and debt holders and thus decrease the default risk *ex post*.

Banks should collectively recognize the benefits of female CFOs in decreasing information risk *ex ante* and default risk *ex post* and reward borrowers under the control of female CFOs with more favorable loan contract terms. The empirical research also provides proof of this judgment. For example, using a sample of S&P 1500 companies from 1994 to 2006, Francis et al. (2013) find that firms with female CFOs enjoy bank loan prices about 14 basis points lower on average than firms with male CFOs. In addition, loans granted to female-CFO-led companies have 9% longer maturities and are about 8% less likely to be required to provide collateral than loans granted to male-CFO-led companies.

Based on the aforementioned analysis, we argue that the transmission of financial conservatism signals in the credit market is a possible mechanism through which CFO gender influences bank lending decisions. Thus, we propose our first research hypothesis as follows.

Hypothesis 1. According to signaling theory, all else being equal, banks grant borrowers with female CFOs more favorable loan contract terms than borrowers with male CFOs.

2.3. *Female CFOs and bank loan contracts: based on gender discrimination theory*

Although default risk is the key determinant of loan contracts, many studies have also shown that females are discriminated against in the credit market.[4] Becker (1957) uses the term "taste-based" discrimination, whereby minorities, low-income families and women, among others, receive disparate and less advantageous treatment than their counterparts. A large number of investigations have focused on the role of race, ethnicity and gender as determinants of credit applications, loan denials, charged interest rates and other types of

[4] Many studies have considered discrimination in the credit market. For a detailed survey of this literature, see a study by Blanch et al. (2003).

restricted access to finance (Cavalluzzo et al., 2002; Blanch et al., 2003; Cavalluzzo and Wolken, 2005). In the classical model proposed by Becker (1957), discrimination arises due to the taste-based preferences of the lender, who is willing to pay a price to avoid being associated with certain groups of borrowers. He also notes that such discrimination tends to vanish with competition between lenders, as they are no longer able to bear the cost of the non-economically motivated choices.

In terms of gender-based discrimination in the credit market, Cavalluzzo et al. (2002) report evidence of a credit access gap between firms owned by white males and white females in the United States. Using the cross-country Business Environment and Enterprise Performance Survey (BEEPS), Muravyev et al. (2009) also note that female-managed firms are less likely to obtain bank loans than their male-managed counterparts. In addition, their analysis suggests that female entrepreneurs are charged higher interest rates when their loan applications are approved. Extending these studies, Ongena and Popov (2013) identify the causal effect of gender bias on access to finance through a cross-country sample and find that female-owned firms in countries with a higher gender bias are more frequently discouraged from applying for bank credit and more reliant on informal finance. They emphasize that these results are not driven by credit risk differences between female- and male-owned firms in countries with high gender biases.

Gender discrimination is also a severe social problem in contemporary China. Despite the declaration that women hold up half the sky, China has followed patrilineal rule and a male power system since ancient times. As a result, women can only occupy a subordinate position in both their families and the social system. Although the women's revolution that accompanied the establishment of the People's Republic of China led to a great improvement in their social status, women continue to experience an obvious gender bias in terms of job availability, promotion opportunities and social identity (Qing and Zheng, 2013). In fact, as a defective social norm in China, gender discrimination has permeated into various social and economic activities. As a result, women must expend more effort than their counterparts to achieve the same success. The gender bias hypothesis simply implies that banks may charge higher loan prices and require tighter non-price terms when lending to female-CFO-led companies.

However, social capital also plays an important role in financial development (Guiso et al., 2004). Due to the prevalence of "guanxi" and the weaker institutional environment in transitional China (Peng and Luo, 2000), the effect of relational networks and social capital on corporate financing behavior is more severe and explicit (Talavera et al., 2012). In contrast to their male counterparts, women are much shier and more reluctant to build relationship networks, which decreases the chances of success for female CFOs competing for scarce financial resources.

Taken together, we argue that gender bias is another possible substitutive mechanism through which CFO gender affects bank lending decisions. Thus, we propose our second research hypothesis as follows.

Hypothesis 2. According to gender discrimination theory, all else being equal, banks grant borrowers with female CFOs more unfavorable loan contract terms than borrowers with male CFOs.

3. Research design

3.1. Data and sample

The sample comprises A-share listed companies in the Shanghai and Shenzhen stock markets from 2007 to 2012. According to our research objectives and consistent with similar studies, we select our sample based on the following criteria. First, due to their specific capital structures and debt covenants, we exclude companies in the banking, insurance and other financial industries. Second, we exclude companies that also issue B and H stocks to eliminate the effects of financial environment and regulatory policy on bank loan contracts. Third, we exclude firm years in which no loans were provided by banks. Fourth, we exclude firms whose CFOs changed in a given year. Fifth, we exclude firm years if the data required to measure the firm-specific control variables are unavailable. These criteria ultimately yield 5312 firm-year observations. The gender information of top executives is collected manually from annual reports, and the bank loan information is collected by hand from borrowers' annual report notes. Other accounting and financial data are taken from the CSMAR and WIND databases.

Table 1
Variable definitions.

Variable	Definition
Dependent variables	
Loan	The total bank loan amount in billions of yuan at year-end
Collateral	A dummy variable that equals 1 if the company obtains credit loans from banks in a given year and 0 otherwise
Collateral-ratio	The ratio of the volume of secured loans to the total volume of bank loans
Pled-collateral	The ratio of the volume of pledging collateral loans to the total volume of secured loans
Mort-collateral	The ratio of the volume of mortgaging collateral loans to the total volume of secured loans
Guar-collateral	The ratio of the volume of guaranteeing collateral loans to the total volume of secured loans
Independent test variables	
CFO-gender	A dummy variable that equals 1 if the CFO gender is female and 0 otherwise
Chair-gender	A dummy variable that equals 1 if the chairman gender is female and 0 otherwise
CEO-gender	A dummy variable that equals 1 if the CEO gender is female and 0 otherwise
Control variables	
Size	A proxy variable for firm size, measured as the natural logarithm of the firm's market value of equity at year-end
Lev	A proxy variable for the leverage ratio, measured as the ratio of total liabilities to total assets
Tangible	A proxy variable for the characteristics of the firm's assets, measured as fixed assets divided by total assets
Growth	A proxy variable for growth opportunities, measured as the market value of equity plus the book value of debt divided by total assets
ROA	A proxy variable for profitability, measured as the ratio of net income to total assets at year-end
INDDUM	Indicator variables for industry membership, following the CSRC Industry Code (2001)
YDUM	Year indicator variables that proxies for year-specific effects
Conditioning variables	
State	A proxy variable for property rights that equals 1 if the ultimate controlling shareholder of a listed firm is a (central or local) government agency or government-controlled SOE and 0 otherwise
Market	A proxy variable for the degree of province-level financial marketization in China, which is one sub-index of the Marketization index (Fan et al., 2011)

3.2. Variable measurement

Bank loan contracts include not only price terms (i.e., loan interest rates) but also non-price terms, such as loan size, maturity, collateral requirements and restrictive covenants. In this study, we focus on the effect of gender on collateral terms. In particular, we construct three dependent variables to capture the characteristics of the collateral clauses of bank loan contracts, denoted by *Collateral*, *Collateral-ratio* and *Collateral-intensity*, respectively. The first dependent variable is *Collateral*. It is an indicator variable that takes the value of 1 if a company obtains credit loans from banks in a given year and 0 otherwise. The second dependent variable is *Collateral-ratio*, which is defined as the proportion of the volume of secured loans to the total volume of bank loans. Finally, the third dependent variable is *Collateral-intensity*, which represents the restrictive intensity of the clauses of three different types of collateral. In our main tests, we adopt *Guar-collateral*, *Mort-collateral* and *Pled-collateral* to proxy for *Collateral-intensity*, which represents the ratio of the volume of guaranteeing, mortgaging and pledging collateral loans to the total volume of secured loans, respectively. In our study, the main independent variable is *CFO-gender*,[5] which equals 1 if the CFO position is held by a female executive and 0 otherwise. All of the variables are defined in Table 1.

3.3. Model specification

We specify the following regression model to evaluate the effect of CFO gender on the collateral terms of bank loan contracts:

$$LoanContract = \alpha + \beta_1 Gender + \beta_2 Control + \varepsilon \tag{1}$$

[5] In practice, CEOs and chairmen take ultimate responsibility for corporate decisions. Therefore, in our robustness checks, we introduce *Chair-gender* and *CEO-gender* as alternative testing variables to explore whether CEO and chairman gender influences bank loan contracts. However, we do not find that the presence of female CEOs and chairmen affects the collateral terms considered in our study.

Table 2
Descriptive statistics.

Variable	Observations	Mean	Median	Std.	Q1	Q3
Panel A: Loan facility characteristics						
Loan	5312	1.341	0.395	3.111	0.13	1.141
Collateral	5312	0.619	1	0.485	0	1
Collateral-ratio	5312	0.714	0.912	0.353	0.468	1
Mort-collateral	4912	0.402	0.235	0.358	0	0.629
Pled-collateral	4912	0.113	0	0.237	0	0.097
Guar-collateral	4912	0.485	0.505	0.384	0.026	0.868
Panel B: Borrowing firm characteristics						
CFO-gender	5312	0.274	0	0.446	0	1
CEO-gender	5312	0.056	0	0.230	0	0
Chair-gender	5312	0.042	0	0.200	0	0
Size	5312	21.696	21.562	1.099	20.905	22.361
Leverage	5312	0.482	0.485	0.193	0.337	0.625
ROA	5312	0.051	0.047	0.087	0.021	0.081
Tangible	5312	0.255	0.228	0.165	0.124	0.366
State	5312	0.498	0	0.500	0	1
Growth	5312	1.738	1.469	0.889	1.167	2.000

where all of the variables are as defined in Table 1. The dependent variable, *LoanContract*, refers to one of the following three bank loan collateral measures: *Collateral*, *Collateral-ratio* or *Collateral-intensity*. The main test variable, *CFO-gender,* represents the gender of the CFO of the borrowing firms. In a robustness test, we also examine the effects of CEO and chairman gender on the collateral terms.

Following studies of bank loan contracts (e.g., Francis et al., 2013; Chen et al., 2013), we control for several firm characteristics that may affect collateral terms in the regressions. We begin by using the natural logarithm of a firm's total assets to measure firm size. We use *Growth*, which is measured as the market value of equity plus the book value of debt divided by the total assets, as a proxy for the firm's growth opportunities. We also control for *Leverage*, which is measured as long-term debt plus current debt divided by total assets; *ROA*, which is measured as the ratio of net income to total assets; and *Tangible*, which is measured as fixed assets divided by total assets. We also introduce *State* into the regression model to control for the influence of the nature of the ultimate owners on the bank loan contracts. In addition, to address potential endogeneity concerns, all of the control variables used in this model take a lagged value of one year. Finally, we use industries dummies to control for the potential differences in loan contracting across industries.

4. Empirical results

4.1. Descriptive statistics

Table 2 reports the summary statistics for the main variables used in this study. As shown in Panel A of Table 2, in our sample, the average loan amount is 1.341 billion yuan, and 61.9% of Chinese listed companies obtained credit loans from banks. Looking at the bank loan characteristics, 71.4% of the loans in our sample use collateral (security). Furthermore, the means of the *Collateral-intensity* measures such as *Guar-collateral*, *Mort-collateral* and *Pled-collateral* are 0.485, 0.402 and 0.113, respectively. This suggests that guaranteeing collateral is the most frequently used collateral mode and that pledging collateral is the least used.

According to Panel B of Table 2, the mean values of *CFO-gender*, *CEO-gender* and *Chair-gender* are 0.274, 0.056 and 0.042, respectively. These numbers indicate that 27.4% of the CFO positions of the companies in our sample are held by woman, a figure much larger than the 4.9% figure found in the U.S. setting by Francis et al. (2013). In addition to female CFOs, 5.6% of CEO positions and 4.2% of chairman positions are held by female executives, respectively. Female executives account for more than one third of top management teams, implying that female executives are playing an increasingly important role in corporate decisions.

Table 3
Correlation matrix.

Variable	1	2	3	4	5	6	7	8	9	10	11	12	13	14	15
Loan	1.000														
Collateral	0.324***	1.000													
Collateral-ratio	−0.071***	−0.632***	1.000												
Guar-collateral	0.148***	0.100***	−0.035**	1.000											
Mort-collateral	−0.180***	−0.141***	0.089***	−0.806***	1.000										
Pled-collateral	0.047***	0.063***	−0.084***	−0.328***	−0.293***	1.000									
CFO-gender	−0.042***	−0.016*	0.029**	−0.048***	0.047***	0.003	1.000								
CEO-gender	−0.028**	−0.061***	0.062***	−0.037***	0.040***	−0.004	0.052***	1.000							
Chair-gender	−0.028**	−0.039***	0.037***	0.011	0.004	−0.005	0.031**	0.202***	1.000						
Growth	−0.277***	−0.091***	0.001	−0.031**	0.052***	−0.032**	0.001	0.017	−0.001	1.000					
Lev	0.604***	0.105***	0.103***	0.057***	−0.055***	−0.004	−0.065***	0.010	0.003	−0.149***	1.000				
Tangible	0.219***	0.113***	−0.065***	0.065***	−0.040***	−0.042***	−0.001	−0.031**	−0.046***	−0.001	0.093***	1.000			
Size	0.726***	0.334***	−0.215***	0.136***	−0.222***	0.135***	−0.051***	−0.056***	−0.021	−0.328***	0.364***	0.042***	1.000		
State	0.252***	0.261***	−0.242***	0.128***	−0.151***	0.034**	−0.057***	−0.073***	−0.032**	−0.032**	0.252***	0.162***	0.319***	1.000	
ROA	−0.084***	0.045***	−0.113***	−0.028	−0.003	0.051***	−0.001	0.005	0.016	0.119***	−0.243***	−0.134***	0.095***	−0.050***	1.000

Note: This table presents Spearman correlations of the main variables used in our analysis. All of the variables are defined in Table 1.
[*] Represent significance at the 10% level.
[**] Represent significance at the 5% level.
[***] Represent significance at the 1% level.

Table 4
Univariate testing.

Variable	Male CFO	Female CFO	Test of difference	
	$N = 3856$	$N = 1456$	Difference	t-statistic
Loan	1.453	1.043	0.410	4.289***
Collateral	0.624	0.606	0.018	1.224
Collateral-ratio	0.708	0.731	−0.023	−2.135**
Mort-collateral	0.390	0.430	−0.040	−3.322***
Pled-collateral	0.112	0.114	−0.001	−0.212
Guar-collateral	0.496	0.454	0.041	3.415***
Size	21.731	21.603	0.127	3.775***
State	0.516	0.451	0.065	4.222***
Lev	0.490	0.462	0.028	4.757***
Growth	1.737	1.741	−0.004	−0.135
Tangible	0.255	0.255	0.000	0.072
ROA	0.051	0.051	0.000	0.096

Note: This table reports t-statistics for the differences in means of each variable across firms with male and female CFOs.
* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

Table 3 reports a Spearman correlation matrix for the variables included in our regression analyses. Our test variable *CFO-gender* is significantly negatively correlated with our measure of bank loan amount (*Loan*) and the indicator of the likelihood to obtain credit loans (*Collateral*). However, it displays a positive correlation with the use of collateral requirements (*Collateral-ratio*), suggesting that firms with female CFOs are less likely to obtain bank loans and especially credit loans than firms with male CFOs. In terms of the correlations between *CFO-gender* and *Collateral-intensity*, *CFO-gender* is significantly and negatively associated with *Guar-collateral* but significantly and positively associated with *Mort-collateral*. The coefficient between *CFO-gender* and *Pled-collateral* is also positive but not significant. All of these results indicate that the loan contracts granted to female-CFO-led companies are more rigid than those granted to male-CFO-led companies, which provides preliminary support for the hypothesis about discrimination against female CFOs in bank lending decisions.

In addition, Table 3 shows that although *Loan* has significantly positive correlations with both *Size* and *State*, *Collateral* has significantly negative correlations with both *Size* and *State*. This implies that compared with small-scale enterprises and non-SOEs, large-scale enterprises and SOEs obtain more bank loans and are less likely to be required to provide collateral.

4.2. Univariate testing

Table 4 provides univariate comparisons of the bank loan terms between firms with female CFOs and their counterparts without female CFOs. Consistent with Hypothesis 2, the mean of the loan amount for firms with female CFOs is 1.043 billion yuan and that for firms with male CFOs is 1.453 billion yuan. The mean difference is 0.410 and is significant at the 1% level. The *Collateral-ratio* means differ significantly between the female- and male-CFO-led firms. On average, the borrowers with female CFOs are less likely to obtain credit loans and thus have bigger proportions of secured loans than the borrowers with male CFOs. More importantly, the firms with female CFOs are significantly more likely to use mortgaging collateral, and the firms with male CFOs are more inclined to provide guaranteeing collateral. In terms of pledging collateral, we find no significant differences between the two types of sample firms. Table 4 also indicates no significant difference in key firm characteristics such as asset size, growth opportunities and earnings capacity across the subsamples, which are the main predictors of loan default risk.

In summary, the univariate tests suggest that banks provide more favorable loan terms to borrowers with male CFOs. The results also indicate that the difference in loan contracts between borrowers with and without female CFOs is not caused by variance in the loan default risk, but is rather derived from banks' lending discrimination against female CFOs. This provides further support for Hypothesis 2.

4.3. Multivariate test results

We begin our multivariate regression tests by testing how female CFOs affect bank loan contracts. We then examine how regional financial marketization relieves banks' lending discrimination against firms with female CFOs. Furthermore, we attempt to investigate whether the state-owned property right of borrowers is helpful in alleviating credit discrimination against female CFOs. Finally, we conduct a series of robustness checks using different statistical methods, including Heckman's two-stage approach, a propensity score-matching approach and the differences-in-differences approach.

4.3.1. Female CFOs and bank loan contracts

Based on the preceding basic model, we explore the effect of female CFOs on collateral terms using several different dependent variables. Table 5 reports the regression analysis results.

In Column (1) of Table 5, we test how the presence of a female CFO affects a firm's likelihood of obtaining credit loans, using an OLS regression. The estimated coefficient of *CFO-gender* equals −0.016 and is significant at the 5% level, indicating that firms with female CFOs are less likely to receive unsecured loans from banks than firms with male CFOs. Using *Collateral-ratio* as the dependent variable, Column (3) of Table 5 shows that the regression coefficient of *CFO-gender* is 0.026 and significant at the 1% level. Therefore, the secured loan ratio of firms with female CFOs is significantly larger than that of firms with male CFOs. In addition, to further investigate the influence of female CFOs on the *Collateral-intensity* of collateral terms, we introduce the proxies of three collateral forms (i.e., *Guar-collateral*, *Mort-collateral* and *Pled-collateral*) individually as dependent variables. As shown in Columns (6) and (7) of Table 5, the coefficients of *CFO-gender* are −0.031 and 0.026, respectively, and both are significant at the 1% level. Looking at *Pled-collateral*, the coefficient of *CFO-gender* is also positive but insignificant, perhaps because the number of pledging collateral cases in our sample is too small.[6] These results imply that firms with female CFOs are more likely than firms with male CFOs to be required to provide mortgaging collateral, but less likely to use guaranteeing collateral. Therefore, the restriction of collateral clauses granted to female-CFO-led companies is more rigid than those granted to male-CFO-led companies. Thus, Hypothesis 2 is confirmed.

In terms of the control variables, *State* is significantly positively related to *Collateral* but significantly negatively related to *Collateral-ratio*. At the same time, the association between *State* and *Guar-collateral* is significantly negative and that between *State* and *Mort-collateral* is significantly positive. These results mean that the property characteristics of borrowing firms have an important effect on the designing of collateral terms in bank loan contracts. In contrast to non-SOEs, it is much easier for SOEs to apply for credit loans from banks, and SOEs also enjoy the benefits of better collateral clauses. Our findings are consistent with those of Fang (2007), who argues that "[t]here exists credit discrimination against private firms in Chinese bank loan decisions." Lu et al. (2012) also suggest that the big four state-owned banks still dominate the banking sector in China, favor SOEs and private firms with political connections in their lending decisions and discriminate against other non-SOEs such as small-town and village enterprises and other private firms.

In addition, as firm size increases, the probability of borrowers obtaining credit loans increases significantly and their secured loan ratios decrease significantly. Moreover, large-size borrowing firms are inclined to adopt guaranteeing collateral with looser restrictions. In contrast, small-size borrowing firms are more likely to adopt mortgaging collateral with stronger constraints. These results are in line with the findings of Lin and Sun (2005) and provide some evidence of the financing dilemma faced by medium-sized and small enterprises in China.

In the previous OLS regression, although we control for various observable firm characteristics that have been widely used in other studies, unobservable time-invariant factors could nevertheless affect the collateral

[6] Relative to the frequent use of mortgaging and guaranteeing collateral, pledging collateral is used far less often. In our sample, only 11.3% of the secured loans adopt pledging collateral.

Table 5
Effect of female CFOs on the collateral terms of bank loan contracts.

Variables	Collateral		Collateral-ratio		Collateral-intensity		
					Guar-collateral	Mort-collateral	Pled-collateral
	OLS	Fixed effect	OLS	Fixed effect	OLS	OLS	OLS
Constant	−15.918*** (−17.49)	−12.403*** (−14.88)	2.323*** (20.54)	4.080*** (18.88)	−0.617*** (−4.61)	2.155*** (16.75)	−0.537*** (−6.49)
CFO-gender	**−0.016** (−2.36)**	**−0.038** (−2.44)**	**0.036*** (2.97)**	**0.057*** (4.52)**	**−0.031*** (−2.67)**	**0.026*** (3.34)**	**0.004 (0.68)**
Growth	0.024 (0.60)	0.011 (0.60)	−0.023*** (−3.92)	−0.011*** (−3.47)	0.003 (0.53)	−0.013** (−1.99)	0.009** (2.24)
Tangible	0.424* (1.83)	0.206*** (2.80)	−0.135*** (−4.94)	−0.168*** (−5.08)	0.028 (0.76)	0.151*** (4.17)	−0.180*** (−7.72)
Size	0.750*** (17.94)	0.573*** (14.28)	−0.073*** (−14.23)	−0.055*** (−12.18)	0.042*** (7.09)	−0.071*** (−12.37)	0.029*** (7.76)
Lev	−0.058 (−0.30)	−0.064 (−0.48)	0.358*** (−12.93)	0.229*** (−9.56)	0.048 (1.47)	−0.038 (−1.22)	−0.009 (−0.48)
State	0.863*** (12.43)	1.026*** (16.55)	−0.163*** (−16.39)	−0.144*** (−13.82)	0.087*** (7.40)	−0.080*** (−7.10)	−0.006 (−0.93)
ROA	0.866** (2.19)	0.744*** (3.87)	−0.195*** (−3.47)	−0.173*** (−3.20)	−0.113* (−1.75)	0.093 (−1.49)	0.020 (0.51)
Industry	Yes	No	Yes	No	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs.	5312	5312	5312	5312	4912	4912	4912
Adj/Pseudo R^2	0.151	0.122	0.159	0.225	0.077	0.127	0.077

Note: This table presents the OLS and firm- and year-fixed-effect regression results for the effect of female CFOs on the collateral terms. All of the variables are defined in Table 1. The absolute values of the heteroskedasticity robust *t*-statistics or *z*-statistics are in parentheses.

* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

terms of bank loan contracts. For example, firms with female CFOs probably behave worse in terms of profitability and thus have greater default risks, inducing banks to decrease their credit loans and claim stronger collateral constraints. This effect is caused by the borrowers' characteristics rather than gender discrimination in the credit market. To deal with this issue, we apply a fixed-effect regression to Eq. (1). The results are shown in Columns (2) and (4) of Table 5. After controlling for the firm and year fixed effects, the effect of female CFOs on firms' likelihood of obtaining credit loans increases to 0.038 from 0.016, and its effect on the secured loan ratios increases to 0.057 from 0.036. Both effects remain economically and statistically significant.

Taken together, the results in Table 5 indicate that CFO gender *does* matter to bank loan contracts. Banks tend to discriminate against female CFOs, grant firms with female CFOs unfavorable loan contracts and especially impose stronger restrictions on collateral clauses. The empirical results are consistent with Hypothesis 2.

4.3.2. Does financial development alleviate credit discrimination against female CFOs?

Many researchers concentrate on how cross-country differences in laws and their enforcement affect bank loan contracting and find that institutional environments have an important effect on bank loan terms (e.g., Bae and Goyal, 2009; Qian and Strahan, 2007; Ge et al., 2012). Although our previous test provides preliminary support for the hypothesis about credit discrimination against female CFOs, such discrimination should ultimately disappear with competition between lenders, as they are no longer able to bear the cost of non-economically motivated choices (Becker, 1957; Muravyev et al., 2009). Consistent with the view that firms facing less product market competition are more likely to discriminate (Orley and Hannan, 1986; Kuhn and Shen, 2013), we have good reason to believe that the enhancement of regional financial marketization and competition in the banking industry leads to a decrease in lending discrimination against female CFOs. The degree of financial marketization in different regions (provinces) of China varies greatly and derives from their great differentiation in geographical location, resource endowment and even public policy (Lu and Yao, 2004; Fan et al., 2011). This provides a unique scenario for determining whether improved financial marketization is beneficial in decreasing gender discrimination in bank loan decisions.

To examine whether the association between female CFOs and the collateral features of loan contracts is conditional on the development of a region's financial marketization, we add the interaction term of *CFO-gender* and *Market*, which is the indicator variable for provincial-level financial development, to Eq. (1). Table 6 summarizes the results when we include the interaction term *CFO-gender* * *Market* in Eq. (1). As shown in Column (1), *CFO-gender* is negatively associated with *Collateral*, and the association

Table 6
Female CFOs, regional financial development and loan collateral.

Variable	Collateral	Collateral-ratio	Collateral-intensity		
			Guar-collateral	Mort-collateral	Pled-collateral
Constant	-16.063^{***} (-17.38)	2.306^{***} (20.00)	-0.725^{***} (-5.31)	2.236^{***} (17.04)	-0.446^{***} (-5.34)
CFO-gender	**-0.110^{*} (-1.76)**	**0.042^{**} (1.95)**	**-0.017^{**} (-2.34)**	**0.040^{***} (2.80)**	**0.054 (1.04)**
Market	**0.024^{**} (2.02)**	**-0.001^{***} (-3.72)**	**0.014^{***} (4.66)**	**-0.011^{***} (-3.84)**	**-0.007^{***} (-2.67)**
CFO-gender * Market	**0.011^{**} (2.34)**	**-0.002^{***} (-3.96)**	**0.001^{**} (2.18)**	**-0.001^{**} (-2.28)**	**0.005 (1.15)**
Growth	0.028 (0.49)	-0.023^{***} (-3.87)	0.006 (0.88)	-0.015^{**} (-2.29)	0.008^{**} (2.04)
Tangible	0.466^{**} (1.99)	-0.027 (-0.86)	0.051 (1.34)	0.133^{***} (3.64)	-0.184^{***} (-7.87)
Size	0.746^{***} (17.85)	-0.073^{***} (-14.22)	0.041^{***} (6.90)	-0.070^{***} (-12.19)	0.029^{***} (7.78)
Lev	-0.036 (-0.19)	0.359^{***} (-12.91)	0.062^{*} (1.89)	-0.050 (-1.60)	-0.011 (-1.18)
State	0.882^{***} (12.53)	-0.162^{***} (-16.12)	0.096^{***} (8.10)	-0.088^{***} (-7.70)	-0.008 (-0.56)
ROA	0.839^{**} (2.12)	-0.196^{***} (-3.49)	-0.124^{*} (-1.89)	0.102^{*} (1.64)	0.026 (0.65)
Industry	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes
Obs.	5312	5312	4912	4912	4912
Adj/Pseudo R^2	0.152	0.161	0.082	0.130	0.078

Note: The absolute values of the heteroskedasticity robust *t*-statistics or *z*-statistics are in parentheses.
* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

is marginally significant at the 10% level. The estimated coefficient of the interaction term, *CFO-gender* * *Market*, is positive and significant at the 5% level. The results suggest that banks are less likely to offer credit loans to firms with female CFOs, and the effect of female CFOs on firms' likelihood of obtaining unsecured loans decreases with the development of regional financial marketization. In Column (2) of Table 6, the coefficient of *CFO-gender* is 0.042, significant at the 5% level, and the coefficient of the interaction term *CFO-gender* * *Market* is −0.002, significant at the 1% level. The results indicate that banks tend to claim collateral when lending to firms with female CFOs, and the effect of female CFOs on the proportion of secured loans is stronger in provinces with weak financial marketization than in provinces with strong financial marketization.

Looking at *Collateral-intensity*, Columns (3)–(5) of Table 6 report the regression analysis results using *Guar-collateral*, *Mort-collateral* and *Pled-collateral* as the dependent variables, respectively. The coefficient of the interaction term *CFO-gender* * *Market* is 0.001 in the *Guar-collateral* regression model, significant at the 5% level, and −0.001 in the *Mort-collateral* regression model, significant at the 5% level. These results indicate that when the level of regional financial marketization is low, the negative association between female CFOs and the use of guaranteeing collateral and the positive association between female CFOs and the use of guaranteeing collateral are significantly enhanced. According to Column (5), the coefficients of *CFO-gender* and the interaction term are positive but not statistically significant, suggesting that the effect of female CFOs on the use of pledging collateral does not differ significantly between firms in regions with different financial development.

Taken together, the results in Table 6 indicate that the discrimination effect against female CFOs on collateral terms is more pronounced in regions with weak financial marketization. In this sense, our findings suggest that financial marketization reform is beneficial in alleviating gender discrimination against firms with female CFOs in the Chinese credit market.

4.3.3. Do property rights matter to credit discrimination against female CFOs?

Although the number of collective, private and foreign banks has continued to grow in the past decade, the banking sector in China is characterized by a lack of competition, the dominance of state-owned banks,[7] lending discrimination against non-SOEs and a lending preference exhibited toward SOEs (Kim et al., 2015). Many studies have noted that because of state or government interventions in the financial markets, state-owned banks favor SOEs and private firms with political connections in their lending decisions while discriminating against other non-SOEs such as small-town and village enterprises and other private firms (Fang, 2007; Lu et al., 2012). This means that banks consider the property right status of borrowers as a key factor in their lending decisions and tend to offer SOEs more favorable terms. Thus, we divide the sample into two categories and go further to determine whether there is a significant difference in the gender discrimination effect between SOE and non-SOE borrowers.

Table 7 presents the regression results by group. In the non-SOE group, the regression coefficient of *CFO-gender* to *Collateral* and *Collateral-ratio* is −0.153 and 0.026, significant at the 5% and 1% levels, respectively. However, in the SOE group, neither of the coefficients is significant. This means that non-SOEs with female CFOs have a significantly lower likelihood of obtaining unsecured loans than those with male CFOs. However, there is no significant difference between SOEs with female CFOs and those with male CFOs. When the dependent variables are *Guar-collateral*, *Mort-collateral* and *Pled-collateral*, all of the regression coefficients of *CFO-gender* are significant in the non-SOE subsample but insignificant in the SOE subsample, which implies that the state-owned property of borrowing firms is also helpful in decreasing the gender discrimination in collateral clauses faced by female CFOs.

In summary, our findings indicate that the property right status of borrowing firms has an important influence on the gender discrimination effect in the credit market. Different from the obvious discrimination shown against female CFOs in non-SOE borrowers, the gender discrimination effect in SOE borrowers is insignificant.

[7] In China, the big four state-owned banks still dominate the banking sector, which include the Agricultural Bank of China (ABC), Bank of China (BOC), Industrial and Commercial Bank of China (ICBC) and People's Construction Bank of China (CBC).

Table 7
Female CFOs, property rights and loan collateral.

Variable	State-owned enterprises ($N = 2647$)					Non-state-owned enterprises ($N = 2665$)				
	Collateral	Collateral-ratio	Collateral-intensity			Collateral	Collateral-ratio	Collateral-intensity		
			Guar-collateral	Mort-collateral	Pled-collateral			Guar-collateral	Mort-collateral	Pled-collateral
Constant	−22.318*** (−14.34)	2.702*** (16.33)	−0.213 (−1.10)	1.977*** (10.89)	−0.762*** (−6.36)	−11.48*** (−9.57)	1.636*** (10.09)	−1.020*** (−5.21)	2.306*** (11.91)	−0.286** (−2.39)
CFO-gender	**0.189 (1.63)**	**0.003 (0.24)**	**−0.029 (−1.62)**	**0.279 (1.63)**	**0.001 (0.14)**	**−0.153*** **(−1.71)**	**0.026***** **(3.02)**	**−0.042***** **(−3.75)**	**0.025***** **(2.68)**	**0.016***** **(2.78)**
Growth	0.165** (2.40)	−0.039*** (−4.08)	0.002 (0.25)	−0.024** (−2.33)	0.022*** (3.14)	−0.023 (−0.44)	−0.008 (−1.10)	0.004 (0.43)	−0.006 (−0.68)	0.002 (0.39)
Tangible	1.407*** (3.91)	−0.088* (−1.95)	0.029 (0.56)	0.135*** (2.71)	−0.165*** (−5.01)	0.004 (0.01)	0.017 (0.37)	0.045 (0.82)	0.150***(2.75)	−0.195*** (−5.79)
Size	1.088*** (15.63)	−0.097*** (−13.52)	0.028*** (3.44)	−0.065*** (−8.38)	0.037*** (7.13)	0.534*** (9.60)	−0.041*** (−5.61)	0.060*** (6.78)	−0.079*** (−8.91)	0.018*** (3.33)
Lev	0.026 (0.09)	0.427*** (9.90)	0.029 (0.58)	−0.051 (−1.08)	0.022 (0.70)	0.013 (0.050	0.290*** (7.94)	0.057 (1.29)	−0.046 (−1.06)	−0.011 (−0.40)
ROA	0.345 (0.61)	−0.132*** (−1.69)	−0.193** (−2.19)	0.113 (1.37)	0.078 (1.46)	1.693*** (2.65)	−0.271*** (−3.32)	0.013 (0.14)	0.044 (0.46)	−0.058 (−0.98)
Industry	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adj/Pseudo R^2	0.176	0.142	0.093	0.095	0.121	0.072	0.073	0.077	0.118	0.050

Note: The absolute values of the heteroskedasticity robust t-statistics or z-statistics are in parentheses.
* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

Table 8
Female CFOs and loan collaterals: Subsample regressions based on regional discrimination culture.

Variable	Low-level discrimination regions				High-level discrimination regions			
	Collateral-ratio	Guar-collateral	Mort-collateral	Pled-collateral	Collateral-ratio	Guar-collateral	Mort-collateral	Pled-collateral
Constant	2.630***	−0.605***	2.018***	−0.413***	2.057***	−482**	2.147***	−0.664***
	(16.70)	(−3.30)	(11.42)	(−3.74)	(12.32)	(−2.48)	(11.83)	(−4.84)
CFO-gender	0.011	−0.010	0.015	0.005	0.017**	−0.056***	0.039***	0.018*
	(0.85)	(−1.55)	(0.96)	(0.54)	(2.17)	(−3.31)	(3.45)	(1.65)
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Industry/Year	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Obs.	2,600	2,387	2,387	2,387	2,712	2,448	2,448	2,448
Adj-R^2	0.190	0.092	0.133	0.059	0.138	0.085	0.134	0.085

Note: Due to limited space, the regression results of the control variables are omitted.
* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

4.3.4. *Robustness checks*

We conduct the following robustness checks to improve the reliability of our empirical results.

4.3.4.1. Further analysis of the gender discrimination effect in bank loan contracts. Our earlier analysis indicates that banks tend to offer unfavorable collateral terms to borrowing firms under the control of female CFOs. However, this may not be caused by the gender discrimination against female CFOs, but rather the banks' rational evaluation of the differences in the decision-making abilities and default risk between male and female CFOs. To address this concern and confirm the existence of a gender discrimination effect, we draw regional discrimination culture into our analysis framework. As mentioned by Gao and Lin (2014), corporate decisions are likely to be influenced by local cultural norms. In line with this argument, we conjecture that borrowing firms in provinces with stronger gender discrimination cultures should exhibit more credit discrimination against female CFOs.

Following the approach of Gao and Lin (2014), we use the province-level sex ratio at birth as a proxy of local discrimination culture[8] and divide our sample into two groups based on the mean value (high-level discriminative regions vs. low-level discriminative regions). As shown in Table 8, the various collateral terms granted to borrowers with female CFOs are more unfavorable than those granted to borrowers with male CFOs in provinces with stronger gender discrimination. However, the difference between borrowers with female CFOs and those with male CFOs in provinces with low-level gender discrimination is insignificant. Sub-group regressions further indicate that the effect of CFO gender on bank loan contracts is mainly a result of the gender discrimination effect, which is also consistent with the preceding prediction and provides further support of Hypothesis 2.

4.3.4.2. Test of potential endogeneity problems. Endogeneity is a large concern in the study of gender issues. For instance, female CFOs may not be randomly assigned to firms. Firms with conservative financial policies and good credit records may be more likely to hire female CFOs. In addition, unobservable time-variant or

[8] Gender discrimination is deeply rooted in the traditional Chinese culture norms of "parental preference for sons so as to carry on the family line." At the same time, due to the strict family planning policy, many parents adopt inexpensive technology to screen the sex of a fetus (Ultrasound B in particular) and perform sex-selective abortions, which leads to a severe sex ratio imbalance. More importantly, there is great variation in the gender discrimination cultures across different provinces in China. Relative to north China, discrimination against girls is much more serious in south China, where the patriarchal clan culture prevails. Following the approach of Gao and Lin (2014), we use the newborn male-to-female ratio as our primary measure of gender discrimination and find that the top two provinces with the highest sex ratios are Jiangxi (1.257) and Hainan (1.233) and that the two provinces with the lowest sex ratios are Xinjiang (1.060) and Shaanxi (1.071).

Table 9
Female CFOs and loan collateral: Propensity score-matching results.

Variable	Collateral	Collateral-ratio	Collateral-intensity		
			Guar-collateral	Mort-collateral	Pled-collateral
Constant	-13.044^{***} (-11.75)	1.965^{***} (13.11)	-0.252^{***} (-3.39)	1.833^{***} (10.94)	-0.783^{***} (-6.83)
CFO-gender	**-0.055^{**} (-2.37)**	**0.019^{***} (2.80)**	**-0.025^{***} (-2.93)**	**0.031^{***} (3.29)**	**-0.001 (-0.49)**
Growth	-0.052 (-1.09)	-0.005 (-0.80)	0.009 (1.01)	-0.011 (1.44)	0.002 (0.52)
Tangible	1.445^{***} (5.33)	-0.106^{***} (-2.77)	0.052 (1.11)	-0.014 (-0.33)	-0.068^{**} (-2.29)
Size	0.604^{***} (11.60)	-0.060^{***} (-8.67)	0.031^{***} (3.70)	-0.067^{***} (-8.66)	0.045^{***} (8.47)
Lev	-0.578^{**} (-2.45)	0.422^{***} (11.92)	0.013 (0.30)	0.085^{**} (2.12)	-0.139^{***} (-5.09)
State	0.876^{***} (9.98)	-0.172^{***} (-12.92)	0.061^{***} (3.76)	-0.093^{***} (-6.16)	0.005 (0.54)
ROA	0.735 (1.56)	-0.203^{***} (-2.83)	-0.139 (-1.62)	0.110 (1.39)	-0.025 (-0.48)
Industry	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes
Obs.	2912	2912	2696	2696	2696
Adj/Pseudo R^2	0.112	0.123	0.048	0.058	0.033

Note: This table presents the OLS and logit regressions results for the effect of female CFOs on the various collateral terms of bank loans by applying a propensity score-matching approach. The absolute values of the heteroskedasticity robust t-statistics or z-statistics are in parentheses.

* Indicate significance at the 10% level.

** Indicate significance at the 5% level.

*** Indicate significance at the 1% level.

time-invariant factors may be correlated with the collateral terms of bank loans. Furthermore, the causality problem makes our results hard to interpret. We conduct a series of econometric analyses to address potential endogeneity issues.

First, to compare the two groups of firms (male- and female-CFO-led firms) fairly, we construct matched male CFO firms using a propensity score-matching approach. Following Francis et al. (2013), we begin our matching with a logistic regression of the female CFO dummy variable on industry, year, firm size and leverage. We then use the propensity scores obtained from the logistic estimation and perform a one-to-one nearest-neighbor match with replacements. This procedure ensures that each female CFO firm is paired with a male CFO firm. We then obtain a new pooled sample that includes 1456 observations with female CFOs and 1456 matched observations with male CFOs. As shown in Table 9, the collateral clauses given to firms with female CFOs are more rigorous than those given to firms with male CFOs. The results are consistent with those reported in Table 5, implying that the main conclusions remain qualitatively unchanged even when we use alternative matching methods.

Second, we use Heckman's two-stage self-selection model to control for the self-selection bias induced in a firm's choice to hire a female CFO. According to Gao and Lin (2014), firms in the provinces with stronger gender discrimination have fewer female top managers. Similarity/attraction theory and many studies have also suggested that firms are more likely to hire female CFOs when their boards of directors are dominated by women (Adams and Ferreira, 2009). Hence, we choose the sex ratio at birth (*Sex-ratio*) of the province in which the firm is headquartered and the ratio of women members on the firm's board of directors (*Female-director*) as instrumental variables. We then estimate the 2SLS model based on Eq. (1).

In the first stage, we run a probit regression using *Female CFOs* as the dependent variable. According to the first-stage regression results, the estimated coefficients of *Sex-ratio* and *Female-director* are -0.507 and 0.817, respectively. Both are significant at the 1% level, suggesting that our selection of instrumental variables is reasonable. In the second stage, we run OLS and logit regressions. As shown in Table 10, the estimated coefficients of *CFO-gender* in Columns (1)–(5) are -0.193, 0.406, -0.648, 1.246 and -0.597, respectively. They are significant at the 1% level, suggesting that our results hold after considering the endogeneity of the choice of female CFOs.

Third, similar to Francis et al. (2013, 2014), we trace the firms that changed their CFOs from male to female (treated group) and from male to male (control group) and apply the difference-in-difference approach to mitigate any unobservable time-variant factors that may affect the estimated influence of female CFOs.

Table 10
Female CFOs and loan collateral: Heckman two-stage model results.

Variable	Collateral	Collateral-ratio	Collateral-intensity		
			Guar-collateral	Mort-collateral	Pled-collateral
Constant	1.615*** (7.30)	0.673*** (9.86)	0.722*** (7.67)	−0.130*** (−2.99)	−0.408*** (−5.50)
CFO-gender	**−0.193*** (−5.40)**	**0.406*** (3.24)**	**−0.648*** (−3.75)**	**1.246*** (5.12)**	**−0.597*** (−4.38)**
Growth	−0.058*** (−3.00)	0.006 (1.10)	−0.015* (−1.82)	0.018 (1.60)	−0.003 (−0.54)
Tangible	0.319*** (2.80)	−0.096*** (−2.78)	0.099** (2.03)	0.046 (0.69)	−0.146*** (−3.81)
Size	0.135*** (15.33)	−0.082*** (−16.68)	0.044*** (6.66)	−0.065*** (−7.95)	0.021*** (2.84)
Lev	0.003 (0.53)	0.195*** (5.50)	0.050 (1.05)	0.068 (1.15)	−0.118** (−2.23)
State	0.146*** (3.79)	−0.129*** (−10.84)	0.098 *** (6.00)	−0.081*** (−3.52)	−0.017 (−1.34)
ROA	0.393** (2.17)	−0.366*** (−6.54)	−0.033 (−0.43)	−0.044 (−0.41)	0.077 (1.27)
Industry	Yes	Yes	Yes	Yes	Yes
Year	Yes	Yes	Yes	Yes	Yes
Obs.	4835	4835	4835	4835	4835
Adj/Pseudo R^2	0.018	0.034	0.044	0.051	0.032

Note: This table presents the OLS and logit regressions results for the effect of female CFOs on the various collateral terms of bank loans by applying a 2SLS regression model. The absolute values of the heteroskedasticity robust *t*-statistics or *z*-statistics are in parentheses.
* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

Table 11
Female CFOs and loan collateral: Difference-in-difference regression results.

Variable	Collateral	Collateral-ratio	Collateral-intensity		
			Guar-collateral	Mort-collateral	Pled-collateral
Post	0.012 (0.58)	0.010 (0.67)	0.108 (1.22)	0.088 (0.40)	0.055 (0.72)
Post * CFO-gender	**−0.103** (−2.33)**	**0.092*** (3.84)**	**−0.612*** (−3.55)**	**1.104*** (4.83)**	**0.511*** (4.03)**
Controls					Omitted
Obs.	310	310	310	310	310
Adj/Pseudo R^2	0.056	0.101	0.094	0.088	0.067

Note: This table presents the OLS and logit regressions results for the effect of female CFOs on the various collateral terms of bank loans by applying the difference-in-difference approach. *Post* is a dummy variable that equals 1 if the year is after the CFO transition year and 0 if the year is before the CFO transition year. *CFO-gender* is a dummy variable that equals 1 if the CFO is female and 0 otherwise. Due to limited space, all of the control variable regression results are omitted.
* Indicate significance at the 10% level.
** Indicate significance at the 5% level.
*** Indicate significance at the 1% level.

As reported in Column (1) of Table 11, the estimated coefficient of *Post*, which captures the effect of male-to-male CFO transition on *Collateral*, is insignificant, indicating that there are no significant differences in the likelihood of obtaining unsecured loans between the pre- and post-transition periods for the firms in the control group. The estimated coefficient of the interaction term between *Post* and *Female CFOs*, which captures the incremental effect of male-to-female CFO transition on *Collateral*, is −0.103 and significant at the 5% level. Hence, compared with male CFOs, female CFOs decrease their firms' probability (10.3%) of gaining credit loans significantly after CFO transitions. Consistent results are also found in Columns (2)–(5) when we replace the dependent variable with *Collateral-ratio*, *Guar-collateral*, *Mort-collateral* and *Pled-collateral*, respectively. The results of the difference-in-difference approach demonstrate that our findings related to credit discrimination against female CFOs in collateral terms hold after considering the time-variant omitted variable bias. The results also suggest that female CFOs decrease their firms' likelihood of receiving unsecured loans and result in much tighter collateral terms in their bank loan contracts.

4.3.4.3. Effect of chairman and CEO gender. Considering that board chairmen and CEOs are the ultimate decision makers on important corporate affairs, we also confirm the robustness of our results in a change regression, in which we investigate the effect of chairman and CEO gender on the collateral terms of bank loans. However, we do not find that the presence of female chairmen and CEOs significantly affects the collateral terms considered in our study, a finding similar to that of Huang and Kisgen (2013). In contrast with chairmen or CEOs, CFOs are the real executants of corporate financial policy and participate directly in negotiations with lending officers. As such, banks view CFOs rather than CEOs or chairmen as the primary executives who determine the quality of accounting information and default risk, which in turn affects their final lending decisions. The results are consistent with recent studies that have found a strong relationship between CFOs and the quality of firms' accounting information and leverage levels (e.g., Jiang et al., 2010). Of course, the number of firms led by female chairman or CEOs in our sample is limited, which may also have some adverse effect on the empirical results. Therefore, our research conclusions on this point must be treated cautiously.

4.3.4.4. Additional collateral requirement terms. When signing loan contracts with borrowing firms, banks typically use collateral requirements together with other price and non-price terms, including interest rates, maturities, loan size and possible covenants. At the same time, borrowers probably actively provide more collateral to banks to seek favorable loan prices. To address this issue, we conduct an additional robustness test using the "firm-year-bank loan case" observations as our sample. As firms do not typically disclose the detailed information of each loan, we gather only 2854 observations. Although not tabulated for the sake of brevity, the related results are qualitatively similar to those reported previously and thus provide additional support for Hypothesis 2.

4.3.4.5. Monetary policy. Bank loan decisions are influenced by a country's monetary policy to a great extent. Thus, we introduce the variable of monetary policy into our analytical framework and investigate the effect of CFO gender on bank loan contracts under different monetary policy conditions. Referring to the related literature, we adopt the annual rate of growth in broad money supply as our proxy for monetary policy. During a period of tightening monetary policy, the loan contracts granted to firms with female CFOs are more unfavorable than those granted to firms with male CFOs. On the contrary, CFO gender is not significantly associated with loan terms during a period of easing monetary policy. These results reflect that the loosening of monetary policy can relieve the lending discrimination faced by female CFOs to some extent.

4.3.4.6. Marketization in allocation of credit funds. We also use the "marketization in allocation of credit funds" index as a substitutive measurement for competition intensity in the banking industry. The regression results (not tabulated for the sake of brevity) show that our inferences remain unchanged.

5. Conclusion

In this study, we examine whether CFO gender affects bank–firm relationships and the designing of collateral clauses in bank loan contracting. Based on both signaling theory and gender discrimination theory, we propose two opposing hypotheses to explain the effect of CFO gender on bank loan contracts and the potential paths of influence and then conduct empirical testing using data from Chinese listed companies. Our empirical results support the hypothesis about gender discrimination in the credit market. In our sample, compared with firms with male CFOs, firms with female CFOs face tighter credit availability and are more likely to be required to provide collateral despite the approval of their loan applications. Furthermore, banks prefer to claim mortgaging collateral when lending to companies under the control of female CFOs and are more inclined to require guaranteeing collateral when lending to companies under the control of male CFOs. This suggests that the collateral clauses granted to female CFOs are much more restricted than those granted to male CFOs. Further analysis reveals that regional financial development and the enhancement of competition in the banking industry are beneficial in alleviating lending discrimination against female CFOs. In addition, female CFOs of SOEs are less likely to suffer from gender discrimination in the credit market than female CFOs of non-SOEs.

Overall, our results suggest that banks do not recognize the benefits of female CFOs in decreasing the information risk ex ante and default risk ex post and reward them with more favorable loan contracts. On the contrary, banks discriminate against female CFOs in their loan decisions and thus claim much stricter collateral when lending to firms under the control of female CFOs. Gender discrimination is a serious obstacle to increasing the efficiency with which scarce credit resources are allocated and hurts the improvement of the financial market. In this sense, in addition to enhancing the legal institutions and financial marketization, eliminating the negative influence of gender discrimination and other unhealthy social norms is an urgent task in the process of improving China's financial reforms. This study deepens and expands our understanding of the determinants of bank loan contracts and contributes new knowledge to the gender literature. In addition, our study of the effect of gender discrimination provides a strong case for comprehending the important role informal rules play in the corporate decisions made within China's transitional economy.

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References

Adams, R.B., Ferreira, D., 2009. Women in the boardroom and their impact on governance and performance. J. Financ. Econ. 94 (2), 291–309.

Ahern, K.R., Dittmar, A.K., 2012. The changing of the boards: the impact on firm valuation of mandated female board representation. Quart. J. Econ. 127, 137–197.

Alesina, A.F., Lotti, F., Mistrulli, P.E., 2013. Do women pay more for credit? Evidence from Italy. J. Eur. Econ. Assoc. 11 (s1), 45–66.

Anderson, R.C., Mansi, S.A., Reeb, D.M., 2004. Board characteristics, accounting report integrity, and the cost of debt. J. Account. Econ. 37 (3), 315–342.

Armstrong, C., Guay, W.R., Weber, J., 2010. The role of information and financial reporting in corporate governance and debt contracting. J. Account. Econ. 50 (2–3), 179–234.

Atkinson, S.M., Boyce Baird, S., Frye, M.B., 2003. Do female mutual fund managers manage differently? J. Financ. Res. 26 (1), 1–18.

Bae, K.H., Goyal, V.K., 2009. Creditor rights, enforcement and bank loans. J. Finance 64 (2), 823–860.

Becker, G.S., 1957. The Economics of Discrimination. Chicago University Press.

Bellucci, A., Borisov, A., Zazzaro, A., 2010. Does gender matter in bank–firm relationships? Evidence from small business lending. J. Banking Finance 34, 2968–2984.

Bharath, S.T., Sunder, J., Sunder, S.V., 2008. Accounting quality and debt contracting. Account. Rev. 83 (1), 1–28.

Blanch, D.G., Levine, P.B., Zimmerman, D.J., 2003. Discrimination in the small-business credit market. Rev. Econ. Statistics 85 (4), 930–943.

Cavalluzzo, K., Cavalluzzo, L., Wolken, J., 2002. Competition, small business financing, and discrimination: evidence from a new survey. J. Bus. 75 (4), 641–680.

Cavalluzzo, K., Wolken, J., 2005. Small business loan turndowns, personal wealth and discrimination. J. Bus. 78 (6), 2153–2177.

Cerqueiro, G., Ongena, S., Roszbach, K., 2015. Collateralization, bank loan rates and monitoring. J. Finance. http://dx.doi.org/10.1111/jofi.12214.

Chava, S., Livdan, D., Purnanandam, A.K., 2009. Do shareholder rights affect the cost of bank loans? Rev. Financ. Stud. 22 (8), 2973–3004.

Chava, S., Purnanandam, A., 2010. CEOs versus CFOs: incentives and corporate policies. J. Financ. Econ. 97 (2), 263–278.

Chen, D., Xiao, Z., Dong, Z., 2013. The composition of the clannish control rights and the credit contract of banks: is it rent-seeking or efficiency? Econ. Res. J. 9, 130–142 (in Chinese).

Dwyer, P.D., Gilkeson, J.H., List, J.A., 2002. Gender differences in revealed risk taking: evidence from mutual fund investors. Econ. Lett. 76 (2), 151–158.

Faccio, M., Marchica, M.T., Mura, R., 2015. CEO Gender, Corporate Risk-Taking, and the Efficiency of Capital Allocation. Working Paper. <http://papers.ssrn.com/sol3/papers.cfm?abstract_id=2021136>.

Fan, G., Wang, X., Zhu, H., 2011. NERI Index of Marketization of China's Provinces 2011 Report. Economic Science Press.

Fang, J., 2007. Ownership, institutional environment and capital allocation. Econ. Res. J. 12, 82–92 (in Chinese).

Francis, B., Hasan, I., Koetter, M., Wu, Q., 2012. Corporate boards and bank loan contracting. J. Financ. Res. 35 (4), 521–552.

Francis, B., Hasan, I., Park, J.C., Wu, Q., 2014. Gender Differences in Financial Reporting Decision-Making: Evidence From Accounting Conservatism. Working Paper. Bank of Finland Research Discussion Papers.

Francis, B., Hasan, I., Wu, Q., 2013. The impact of CFO gender on bank loan contracting. J. Account. Audit. Finance 28 (1), 53–78.

Gao, H., Lin, Y., 2014. Sex Discrimination and Female Top Managers: Evidence from China. Working Paper. <http://www.ssrn.com>.

Ge, W.X., Kim, J.B., Song, B.Y., 2012. Internal governance, legal institutions and bank loan contracting around the world. J. Corporate Finance 18 (3), 413–432.
Gou, Q., Huang, Y., Liu, X., 2014. Do banks really discriminate against private firms? Manage. World 1, 16–26 (in Chinese).
Graham, J.R., Harvey, C.R., Puri, M., 2013. Managerial attitudes and corporate actions. J. Financ. Econ. 109 (1), 103–121.
Graham, J.R., Li, S., Qiu, J., 2008. Corporate misreporting and bank loan contracting. J. Financ. Econ. 89 (1), 44–61.
Guiso, L., Sapienza, P., Zingales, L., 2004. The role of social capital in financial development. Am. Econ. Rev. 94 (3), 526–556.
Hasan, I., Park, J.C., Wu, Q., 2012. The impact of earnings predictability on bank loan contracting. J. Business Finance Account. 39 (7–8), 1068–1101.
Huang, J., Kisgen, D.J., 2013. Gender and corporate finance: are male executives overconfident relative to female executives? J. Financ. Econ. 108 (3), 822–839.
Jensen, M., Meckling, W., 1976. Theory of the firm: managerial behavior, agency costs and ownership structure. J. Financ. Econ. 32 (3), 305–360.
Jiang, J., Petroni, K.R., Wang, Y., 2010. CFOs and CEOs: who has the most influence on earnings management? J. Financ. Econ. 96 (3), 513–526.
Jiménez, G., Salas, V., Saurina, J., 2006. Determinants of collateral. J. Financ. Econ. 81 (2), 255–281.
Kim, J.B., Ma, L.Z., Wang, H.P., 2015. Financial development and the cost of equity capital: evidence from China. China J. Account. Res. 8 (4), 243–277.
Kim, J.B., Song, B.Y., Zhang, L., 2011a. Internal control weakness and bank loan contracting: evidence from SOX Section 404 disclosures. Account. Rev. 86 (4), 1157–1188.
Kim, J.B., Tsui, J., Yi, C.H., 2011b. The voluntary adoption of international financial reporting standards and loan contracting around the world. Rev. Acc. Stud. 16 (4), 779–811.
Knight, J., 2002. Sexual stereotypes. Nature 415, 254–256.
Kuhn, P., Shen, K., 2013. Gender discrimination in job ads: evidence from China. Quart. J. Econ. 128, 287–336.
Lin, Y., Sun, X., 2005. Information, informal finance and SME financing. Econ. Res. J. 7, 35–44 (in Chinese).
Liu, Y., Wei, Z., Xie, F., 2015. CFO gender and earnings management: evidence from China. Rev. Quant. Finance Account. 8. http://dx.doi.org/10.1007/s11156-014-0490-0.
Lu, F., Yao, Y., 2004. Legality, financial development and economic growth under financial repression. Social Sci. China 1, 42–55 (in Chinese).
Lu, Z., Zhu, J., Zhang, W., 2012. Bank discrimination, holding bank ownership, and economic consequences: evidence from China. J. Bank. Finance 36, 341–354.
Levine, R., 2005. Finance and growth: theory and evidence. Handbook of Economic Growth.
Matsa, D.A., Miller, A.R., 2011. Chipping away at the glass ceiling: gender spillovers in corporate leadership. Am. Econ. Rev. 101 (3), 635–639.
Muravyev, A., Talavera, O., Schafer, D., 2009. Entrepreneurs' gender and financial constraints: evidence from international data. J. Comp. Econ. 37 (2), 270–286.
Nguyen, N.T., Basuray, M.T., Smith, W.P., 2008. Moral issues and gender differences in ethical judgment using Reidenbach and Robin's (1990) multidimensional ethics scale: implications in teaching of business ethics. J. Bus. Ethics 77 (4), 417–430.
Ongena, S., Popov, A., 2013. Gender Bias and Credit Market Barriers for Female Entrepreneurs. Working Paper. <http://www.ssrn.com>.
Orley, A., Hannan, T., 1986. Sex discrimination and product market competition: the case of the banking industry. Quart. J. Econ. 101, 149–173.
Peng, M., Luo, Y., 2000. Managerial ties and firm performance in a transition economy: the nature of a micro-macro link. Acad. Manage. J. 43 (3), 486–501.
Qian, J., Strahan, P.E., 2007. How law and institutions shape financial contracts: the case of bank loans. J. Finance 62 (6), 2803–2834.
Qing, S., Zheng, J., 2013. Unequal pay or unequal job: effect of hierarchical segregation on gender earning differentials. China Econ. Quart. 12 (2), 735–756 (in Chinese).
Rajan, R.G., Zingales, L., 1996. Financial dependence and growth. Am. Econ. Rev. 83 (3), 559–586.
Ruegger, D., King, E.W., 1992. A study of the effect of age and gender upon student business ethics. J. Bus. Ethics 11 (3), 179–186.
Sapienza, P., Zingales, L., Maestripieri, D., 2009. Gender differences in financial risk aversion and career choices are affected by testosterone. Proc. Natl. Acad. Sci. USA 106 (8), 15268–15273.
Srinidhi, B., Gul, F.A., Tsui, J., 2011. Female directors and earnings quality. Contemp. Account. Res. 28 (5), 1610–1644.
Stiglitz, J.E., Weiss, A., 1981. Credit rationing in markets with imperfect information. Am. Econ. Rev. 71 (3), 393–410.
Sunder, J., Sunder, S.V., Zhang, J., 2009. Balance Sheet Conservatism and Debt Contracting. Working Paper. Northwestern University.
Talavera, O., Xiong, L., Xiong, X., 2012. Social capital and access to bank financing: the case of Chinese entrepreneurs. Emerg. Markets Finance Trade 48 (1), 55–69.
Zhang, J., 2008. The contracting benefits of accounting conservatism to lenders and borrowers. J. Account. Econ. 45 (1), 27–54.