Are Chinese Firms Borrowing Externally Too Much? Evidence from Nonparametric Threshold Regression

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Abstract

We investigate the impact of leverage ratio on firms' total factor productivity (TFP) growth to probe the excessive external financing problem in Chinese industrial firms. We propose and empirically test hypotheses on the potential threshold effect of leverage ratio on firms' productivity growth as well as its variation with firm's ownership. To alleviate the risk of inconsistently estimating the threshold effect, we employ a newly developed nonparametric threshold regression model, allowing the unknown threshold level to be estimated without restricting the mean regression structure. We find that Chinese firms on average have leverage ratio exceeding its threshold level, which significantly shrinks its TFP growth. Among firms' ownerships, state-owned enterprises (SOEs) facing the soft budget constraint exhibit significant decline in TFP growth with a higher-than-threshold leverage ratio. As a result, central and western region with higher share of SOEs experience low productivity growth. Our results are supported by a series of nonparametric tests, and are fairly robust to the potential endogeneity of leverage ratio and an alternative measure of external financing.

Keywords: Nonparametric threshold regression, Total factor productivity (TFP), External finance, Chinese industrial firms

JEL Classification: C14, G10, O40.

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1 Introduction

The development of financial markets in China over the past decades are malfunctioning in a sense that firms with different ownerships face unequal degrees of external financing constraints (Guariglia et al., 2011; Chen and Guariglia, 2013). The state-owned enterprises (SOEs) typically have managers with a strong affiliation with the government, allowing a firm's external financing to be made easily at a lower interest rate compared with the one in a competitive market (Bai et al., 2016; Ding et al., 2016). Non-SOEs, such as private and foreign firms, have their external financing with limited access, higher interest rates, and careful monitoring by the financial intermediaries (Bai et al., 2016). In the presence of the malfunctioning financial markets, however, non-SOEs have achieved remarkable productivity growth, thus making the finance-growth nexus in China puzzling. Many studies explain the puzzle by showing that the non-SOEs improve their productivity growth mainly through internal financing, such as cash flow (Song et al., 2011; Guariglia et al., 2011; Chen and Guariglia, 2013).

The limitation of financial markets has induced a set of major reforms in the financial sector, targeting on the improvement of firms' access to external financing through corporate debt (Lin et al., 2003). In recent years, however, firms' corporate debt has increased significantly that makes China as a debt outlier relative to other economies (Ma and Laurenceson, 2017). By 2016, for instance, corporate debt accounts for 145% of GDP compared to government debt (40% of GDP) and household debt (40% of GDP), which is a considerably high level by any international measure (Lipton, 2016). In addition, the corporate debt relative to its total assets, or the leverage ratio, has been growing excessively. As a major indicator of external financing, a higher leverage ratio is strongly indicative of potential growth risks, such as production overcapacity, unsustainable productivity growth, and financial instability (Lipton, 2016; Ma and Laurenceson, 2017). Maliszewski et al. (2016) document that Chinese industries are over-leveraged compared to other developing countries, and thus call for the authority to "deleverage" the economy to mitigate the associated risks.

The surging trend of corporate debt in Chinese firms raises concerns over its impact on Chinese firm's productivity growth. Although a firm's investment behavior under a perfect financial market is independent of its capital structure and financing decisions under the M-M theorem (Modigliani and Miller, 1958), the financial market in China is far from perfect because of its inability to catch up with fast-paced industry reforms and the existence of asymmetric information (Allen et al., 2005). As a result, the external financing provision in China's financial markets becomes relevant to a firm's productivity growth through influencing its decisions on investment, labor, and innovative activities (Allen et al., 2005; Jin et al., 2019). While the positive effect of internal financing on firms' productivity growth is well documented in the literature, the role of external financing on Chinese firms' productivity growth receives relatively less attention. Since the growth of a firm's productivity has been a primary engine of Chinese economy growth (Guariglia et al., 2011), understanding the extent to which the external financing impacts firms' productivity growth can be profound for achieving China's sustainable economy growth.

Indeed, recent literature shows that the external financing through corporate debt may impact firms' productivity growth in a non-monotonic fashion. On one hand, a reasonable level of leverage ratio can lower a firm's waste of cash and external financial constraints, promoting its productivity growth through facilitating the use of uptake technology in the production function (Jin et al., 2019). Firms with a relatively higher leverage ratio can smooth the innovation process more easily to promote productivity growth (Giannetti, 2012). On the other hand, an excessive leverage ratio may create the *soft budget constraint* problem, which induces risk-taking rather than profit-maximizing behaviors of firm managers. As a consequence, funds may be reallocated away from making promising investments, and the decrease of which distorts firms' productivity (Lang et al., 1996; Garriga, 2006). Therefore, an optimal (i.e., a threshold) level of leverage ratio may exist such that a firm's productivity growth can be maximized. Investigating the optimal leverage ratio through a threshold regression analysis has been performed in different countries with micro level data (see, for instance, Coricelli et al. (2012) and Ma et al. (2014)). In case of China, however, it remains questionable regarding the existence of the optimal level of leverage ratio, and if so, whether the leverage ratio in Chinese firms exceeds the optimal level and consequently impedes the growth a firm's productivity.

In addition, the presence of malfunctioning financial markets in China implies that the optimal level of leverage ratio may vary significantly with firm's ownership. Historically, the SOEs do not have full incentives to pursue profit maximization but to provide resources for social stability and fulfill political objectives (Guariglia et al., 2011). As a result, the SOEs, regardless of their firm performance, can easily accumulate corporate debt through borrowing loans from state owned banks. Those borrowed funds are also not subject to close monitoring, since both central and local governments have intensively intervened in the lending process (Riedel et al., 2007). The resulting *soft budget constraint* may consequently bring about moral hazard problem, which negatively influences the productivity growth due to inefficient resource allocation and management discretion (Huang et al., 2006). Kornai et al. (2003) also show that firms with soft budget constraints in transition economies may be motivated to undertake inefficient activities rather than productivity-enhancing investment. In contrast, non-SOEs facing *hard budget constraint* may be effectively disciplined by corporate debt due to its associated financial pressure to non-SOEs managers (Agostino et al., 2018). Hence, the threshold level of leverage ratio for which the productivity growth turns to decline may be higher in non-SOEs than SOEs.

The theoretical arguments regarding the threshold impact of leverage ratio on firm productivity growth in China is ultimately an empirical question. In this paper, we provide, to our best knowledge, the first study in the literature to empirically investigate the impact of *external financing* on Chinese firms' productivity growth. We follow Gomis and Khatiwada (2017) to use firms' leverage ratio as a reasonable proxy of external financing, since the corporate debt in China is mainly consist of banks loans (Xiao, 2005). We follow Coricelli et al. (2012) to construct the measure total factor productivity (TFP) growth by Levinsohn and Petrin (2003), which allows for potential endogeneity of production inputs. Following the discussion above, we mainly propose and empirically test the two following hypotheses:

Hypothesis 1 the leverage ratio may affect a firm's productivity growth non-monotonically in a threshold fashion, thus an optimal level of the leverage ratio may exist for the purpose of productivity growth maximization.

Hypothesis 2 the threshold level may be lower in state-owned enterprises with *soft budget constraint* than other types of firms in the presence of malfunctioning financial market.

To further taking into account the heterogeneity of firms in regions, we perform the threshold analysis in four regions with differences in geography and economic development. We employ a comprehensive Chinese industrial firm level data during 1998-2007 to test our hypotheses through a micro-level perspective.

The identification of the threshold effect of leverage ratio hinges upon a correct functional form of firm's productivity. Such information on the underlying functional form, however, can be rarely given or guided by economic theories. A misspecified empirical model in practice can result in biasing the threshold effect estimates and misleading the results implication, therefore limiting our ability to render solid policy implication. To alleviate the potential risk of model misspecification, we implement a newly developed nonparametric threshold regression by Henderson et al. (2019). Our empirical model does not impose *a priori* model structure assumptions, such as linearity, on the unknown functional form. Instead, our model allows for the threshold effect to be consistently estimated with an unknown structure of mean regression, which permits nonlinear and interactive effects of determinants to be taken into account. Given an identified threshold level of leverage ratio, we further control for its potential endogeneity through a nonparametric analogue of two-stage least square approach by Su and Ullah (2008).

We find that the leverage ratio significantly increases a firm's growth in a threshold fashion: the leverage ratio exhibits a diminishing return on TFP growth after exceeding its threshold level. Comparing with private and foreign firms as non-SOEs, SOEs have the mean of leverage ratio much higher than its threshold level despite of their lower productivity growth. Notably, SOEs, and western region where SOEs dominates, show a clear excessive external borrowing in a sense that the TFP growth becomes negative once the threshold level is exceeded. We support our empirical findings by performing a series of nonparametric tests on functional form specification, the existence of a threshold, and a specific value of threshold. We also demonstrate that our empirical results remain fairly robust after controlling for the endogeneity of leverage ratio and implementing an alternative measure of external financing.

From a policy perspective, our results highlight the importance of effectively regulating the leverage ratio for the policy purpose of maximizing firms' productivity growth. Our results are also related to the literature on the importance of leverage ratio regulation to reduce the future banking credit risk (Drehmann and Juselius, 2014), to prevent financial crisis (Walter and Howie, 2012) and restrain macroeconomic instability (Pettis, 2013). Thus, improving firms' productivity by effectively controlling the leverage ratio would contribute to properly managing the risks associated with China's economic development. As emphasizes by He Liu, the Vice-Premier of China, one major task for the Chinese government in the near future is to effectively control the overall leverage ratio (World Economic Forum, 2018).

The remaining of this paper is organized as follows. Section 2 illustrates our empirical methodology. Section 3 presents the data description, and Section 4 in turn presents our estimation results and related discussions. Section 5 performs a robustness check on our empirical results, and Section 7 concludes.

2 Empirical model

2.1 Threshold regression

2.1.1 Parametric threshold regression

We examine the potential threshold effect of a firm's leverage ratio on TFP growth under a threshold regression framework. To be consistent with the literature, we first employ a popular parametric threshold regression model by Hansen (2000):

$$TFPG_{it} = \alpha_1 L_{it} I(L_{it} \leq \gamma) + \alpha_2 L_{it} I(L_{it} > \gamma) + \mathbf{X}'_{it}\boldsymbol{\beta} + e_{it} \tag{1}$$

where $i = 1, ..., N$ and $t = 1, ..., T$ index a total of N firms over T time periods, and e_{it} is the additive zero-mean error term. $TFPG$ is our dependent variable of firm's TFP growth, L is our key explanatory variable of firm's leverage ratio, $I(\cdot)$ is an indicator function, and $\mathbf{X}$ is a random vector of control variables. Model (1) identifies the existence of threshold level γ by splitting the sample into two "regimes", each of which is used to estimate the coefficient of L that is potentially different across the two regimes ($\alpha_1 \neq \alpha_2$).

For a given value of γ, we obtain the OLS estimates of $\boldsymbol{\beta}(\gamma)'$ and $\alpha(\gamma)'$, and calculate the sum of squared errors $S_n(\boldsymbol{\beta}(\gamma), \alpha(\gamma))$, where $\alpha(\gamma) = [\alpha_1, \alpha_2]'$ defined in (1). Estimation of $(\alpha(\gamma)', \boldsymbol{\beta}(\gamma)')'$ requires γ to belong to a bounded set $\Gamma = [\underline{\gamma}, \bar{\gamma}]$, and we set the upper ($\bar{\gamma}$) and lower bound ($\underline{\gamma}$) of γ, respectively, to be 0.05 and 0.95 percentile of our threshold variable L. Thus, we search over a grid point of $\gamma \in \Gamma$ to obtain the estimator $\hat{\gamma}$ that satisfies:

$$\hat{\gamma} = \underset{\{\gamma \in \Gamma\}}{\operatorname{argmin}} \; S_n(\boldsymbol{\beta}(\gamma), \alpha(\gamma)). \tag{2}$$

We evaluate the significance of the threshold estimates by testing the null hypothesis $H_0 : \alpha_1 = \alpha_2$ for all $\gamma \in \Gamma$ with the feasible test statistic:

$$\hat{I} = \frac{S_0(\hat{\gamma}) - S_1(\hat{\gamma})}{S_1(\hat{\gamma})} \tag{3}$$

where S_0 (S_1) denotes the residual sum of squares in (1) without (with) threshold. Clearly, $\hat{I}$ converges to zero if and only if the null is true. We follow Hansen (2000) to implement a bootstrap version of the test to improve its empirical power.

2.1.2 Nonparametric threshold regression

Model (1) is parametrically specified given a value of γ, where the linear structure is imposed on the unknown functional form *a priori*. While parametric methods possess the advantage of parsimony, the potential disadvantage of model misspecification leads to inconsistent estimates of γ, thus making its nonparametric counterparts alluring. We thereby follow Henderson et al. (2019) to estimate γ nonparametrically by allowing for all covariates to enter the regression via unknown and smooth functions $m_1(\cdot)$ or $m_2(\cdot)$ from

$$TFPG_{it} = m_1(\boldsymbol{Z}_{it}) + e_{it}, \text{ if } L_{it} \leq \gamma, \tag{4}$$

$$TFPG_{it} = m_2(\boldsymbol{Z}_{it}) + e_{it}, \text{ if } L_{it} \geq \gamma, \tag{5}$$

where $\boldsymbol{Z} = [L, \mathbf{X}']'$, and as in (1) we split the sample into regime 1 in (4) or 2 in (5) based on the threshold value γ. The key difference between model (1) and (4)-(5) is that the latter does not impose any functional structure on the mean regression through which $\mathbf{Z}$ impact $TFPG$, therefore robust to potential model misspecification. We rewrite (4)-(5) in a compact manner:

$$TFPG_{it} = \alpha(\boldsymbol{Z}_{it}) + \beta \cdot I(L_{it} > \gamma) + e_{it} \tag{6}$$

where

$$\alpha(\boldsymbol{Z}_{it})= \begin{cases} m_1(\boldsymbol{Z}_{it}), \text{ if } L_{it} \leq \gamma, \\ m_2(\boldsymbol{Z}_{it}) - \beta, \text{ if } L_{it} \geq \gamma. \end{cases}$$

Here, $I(\cdot)$ is the indicator function, and β is the jump size of the regression function at $L_{it} = \gamma$, which may or may not exist. We follow Henderson et al. (2019) to estimate the threshold value of γ by concentrating out both $\alpha(\cdot)$ and β in the following multi-step estimation.

For a given value of γ, we first obtain the estimate of $\alpha(\mathbf{Z})$ in (6), denoted as $\hat{\alpha}_h(\mathbf{Z}; \gamma, \hat{\beta}_h(\gamma)) \equiv \hat{\alpha}_0$ through a local linear estimator evaluated at $\mathbf{Z} \in \{\mathbf{Z}_{it}\}_{i=1,t=1}^{N,T}$ from

$$(\hat{\alpha}_0, \hat{\boldsymbol{\alpha}}_1') = \underset{\{\alpha_0, \boldsymbol{\alpha}_1'\}}{\operatorname{argmin}} \sum_{i=1}^{N} \sum_{t=1}^{T} \left[\tilde{y}_{it} - \alpha_0 - (\mathbf{Z}_{it} - \mathbf{Z})' \boldsymbol{\alpha}_1\right]^2 K_H(\mathbf{Z}_{it}, \mathbf{Z}), \tag{7}$$

where $\tilde{y}_{it} = \left[y_{it} - \hat{\beta}_h(\gamma) \cdot I(\mathbf{Z}_{it} > \gamma)\right]$, $\boldsymbol{\alpha}_1 = [\alpha_{11}, ..., \alpha_{1d}]'$ is a $d \times 1$ vector of the first-order gradient of $\mathbf{Z}$, and

$$K_H(\mathbf{Z}_{it}, \mathbf{Z}) = \frac{1}{|H|} \prod_{s=1}^{d} k\left(\frac{Z_{it}^s - Z^s}{h_s}\right), |H| = h_1 ... h_d,$$

is a product kernel of univariate kernel function $k\left(\frac{Z_{it}^s - Z^s}{h_s}\right)$. The estimate $\hat{\beta}_h(\gamma)$ in $\tilde{y}_{it}$ is obtained from

$$\hat{\beta}_h(\gamma) = \underset{\{\beta_h(\gamma)\}}{\operatorname{argmin}} \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} \left[y_{it} - \hat{\alpha}_h(\mathbf{Z}_{it}; \gamma, \beta_h(\gamma)) - \beta_h(\gamma) I_{it}(\gamma)\right]^2 \hat{f}(\mathbf{Z}_{it})^2,$$

where $\hat{f}(\mathbf{Z}_{it}) = \frac{1}{NT} \sum_{l=1}^{N} \sum_{m=1}^{T} K_H(\mathbf{Z}_{lm}, \mathbf{Z}_{it})$ is the kernel joint density estimates at $\mathbf{Z}_{it}$. The solution of $\hat{\beta}_h(\gamma)$ is given by

$$\hat{\beta}_h(\gamma) = \left(\frac{1}{NT} \sum_{l=1}^{N} \sum_{m=1}^{T} \tilde{I}_{lm}(\gamma)^2\right)^{-1} \frac{1}{NT} \sum_{l=1}^{N} \sum_{m=1}^{T} \tilde{I}_{lm}(\gamma) y_{lm}^* \tag{8}$$

where for $l = 1, ..., N$ and $m = 1, ..., T$, $y_{lm}^* = (1/NT) \sum_{i=1}^{N} \sum_{t=1}^{T} K_H(\mathbf{Z}_{it}, \mathbf{Z}_{lm})(y_{it} - y_{lm})$

with $\mathbf{Z}_{lm}$ and y_{lm} the evaluation points of choice, and

$$\tilde{I}_{lm}(\gamma) = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T} K_H(\mathbf{Z}_{it}, \mathbf{Z}_{lm})\left[I_{it}(\gamma) - I_{lm}(\gamma)\right],$$

with $I_{it}(\gamma) = I(L_{it} > \gamma)$ defined in (4).

The intuition of local linear estimator in (7) is that for each evaluation points of $\mathbf{Z}$ we "locally" fit a linear line through $\mathbf{Z}$ based on the its neighbor observation near to $\mathbf{Z}$. The kernel function $K_H(\mathbf{Z}_{it}, \mathbf{Z})$ assigns higher (lower) weights for points closer to (further away from) $\mathbf{Z}$. We use bandwidth h_s to control the range for which data points are considered as neighbor points around $Z^s \in \mathbf{Z}$. Therefore, larger bandwidths imply more equal weight assigned to the neighbor points of $\mathbf{Z}$, and vice versa. We then connect each linear lines to obtain the unknown function estimate. The rational is similar to obtain the gradients estimates. If the unknown mean function are truly linear, the local linear estimator collapses to OLS estimator (i.e., the local linear regression fits a linear line treating all sample points as neighbor points of $\mathbf{Z}$, implying that $\{h_s\}_{s=1}^{d}$ diverge to infinity).

Finally, with both estimates in (7) and (8) we obtain an approximate estimate of

$$\hat{M}_n(\gamma) = \frac{1}{NT}\sum_{i=1}^{N}\sum_{t=1}^{T}\left[y_{it} - \hat{\alpha}_h(\mathbf{Z}_{it}; \gamma, \hat{\beta}_h(\gamma)) - \hat{\beta}_h(\gamma) I_{it}(\gamma)\right]^2 W(L_{it}), \tag{9}$$

where $W(\cdot)$ is a trimming function of L to eliminate its extreme value. We repeat steps (7)-(8) until we obtain an optimal threshold estimate $\hat{\gamma}$ by searching over a grid of points $\gamma \in \Gamma$ that satisfies:

$$\hat{\gamma} = \underset{\{\gamma\in\Gamma\}}{\operatorname{argmin}}\ |\hat{M}_n(\gamma)| \tag{10}$$

With $\hat{\gamma}$ obtained, we proceed to split our sample accordingly as in (4)-(5), and estimate the partial effects of our interests $\partial m_j(\mathbf{Z}_j)/\partial \mathbf{Z}_j^s$ with $j = 1, 2$ and $s = 1, ..., d$ via local linear estimator of $\hat{\boldsymbol{\theta}}_{1j}$ from:

$$(\hat{\theta}_{0j}, \hat{\boldsymbol{\theta}}'_{1j})' = \underset{\{\theta_{0j}, \boldsymbol{\theta}'_{1j}\}}{\operatorname{argmin}} \sum_{i=1}^{N_j}\sum_{t=1}^{T_j}\left[y_{jit} - \theta_{0j} - (\mathbf{Z}_{jit} - \mathbf{Z}_j)'\boldsymbol{\theta}_{1j}\right]^2 K_H(\mathbf{Z}_{jit}, \mathbf{Z}_j), \tag{11}$$

where y_j and $\mathbf{Z}_j$ are data in regime j with its corresponding observations N_j and T_j.

2.2 Functional form and threshold tests

In addition to the regression estimation, we implement three nonparametric tests below to draw inference on the nonparametric threshold regression. The three tests are modified versions of F-type test by Ullah (1985):

$$\hat{T} = \frac{RSS_0 - RSS_1}{RSS_1}, \tag{12}$$

where RSS_0 (RSS_1) is the residual sum of squares of the regression under the null (alternative). With the test in (12), we are interested in testing the following three null hypotheses with RSS_0 and RSS_1 changed correspondingly. First, we test if the threshold regression in (6) can be sufficiently parametrized. If so, the estimation efficiency can be increased substantially. We test the null functional form of Hansen's parametric threshold model in (1) against our nonparametric threshold regression alternative in (6), with the null denoted as "H_{01}: *Hansen model*".

Second, we evaluate the significance of our estimated threshold $\hat{\gamma}$. We test the null where the threshold structure in (4)-(5) does not exist (i.e., $m_1(\cdot) = m_2(\cdot)$ for almost all $\mathbf{Z}$) against a nonparametric threshold regression alternative, and we denote the null as $H_{02} : \gamma_{\text{not exist}}$. Finally, we are interested in testing if an estimated threshold value of γ is indistinguishable from an arbitrary value of γ_0 such that $\gamma_0 \neq \gamma$. Based on our empirical results, we test our third null hypothesis where the leverage ratio has its threshold equal to its median value, i.e., $H_{03} : \gamma_0 = L_{50}$. We detail the bootstrap procedure of the test $\hat{T}$ under H_{01-03} in Appendix 1.1-1.3.

3 Data

We source our data from the Chinese Annual Surveys of Industrial Production (ASIP) during 1998-2007, officially referred as the "all state-owned and all above-scale non-state owned industrial enterprise database". Maintained by the National Bureau of Statistics of China (NBSC), it is the firm-level data that covers all state-owned and above-scale, non-state-owned enterprises with annual sales at or above five million RMB. This dataset provides relatively reliable statistic measures (Cai and Liu, 2009), and serves as a important dataset for recent seminal works on China's economy (Brandt et al., 2017; Berkowitz et al., 2017; Song et al., 2011; Hsieh and Klenow, 2009; Zou et al., 2018). The ASIP covers enterprises in 31 provinces and municipalities (i.e., Beijing, Shanghai, Tianjin, and Chongqing), accounting for more than 90% of the gross industrial output of China. See Appendix 3 for a list of the 2-digit Chinese Industrial Classification Code (CICC) and industry names.

Table 1: Descriptive Data Summary

		Whole sample	State	Private	Foreign	Collective	Eastern	Western	Northeastern	Central
TFP growth	25%	-0.200	-0.184	-0.183	-0.211	-0.196	-0.203	-0.205	-0.237	-0.167
	50%	0.066	0.054	0.078	0.062	0.073	0.062	0.062	0.059	0.096
	75%	0.432	0.386	0.432	0.452	0.438	0.423	0.452	0.490	0.468
	mean	0.262	0.253	0.259	0.279	0.257	0.251	0.280	0.315	0.299
Leverage ratio	25%	0.348	0.379	0.428	0.278	0.407	0.366	0.348	0.305	0.348
	50%	0.530	0.558	0.601	0.462	0.584	0.546	0.540	0.506	0.547
	75%	0.693	0.725	0.749	0.631	0.740	0.710	0.703	0.701	0.716
	mean	0.516	0.548	0.579	0.458	0.563	0.532	0.523	0.502	0.528
CF ratio	25%	0.031	0.015	0.032	0.042	0.032	0.032	0.022	0.021	0.029
	50%	0.063	0.038	0.063	0.084	0.068	0.065	0.046	0.053	0.067
	75%	0.119	0.070	0.118	0.144	0.137	0.119	0.087	0.104	0.160
	mean	0.098	0.050	0.106	0.104	0.122	0.095	0.067	0.077	0.140
Age	25%	8.0	15.0	7.0	7.0	9.0	7.0	8.0	8.0	8.0
	50%	11.0	30.0	11.0	9.0	13.0	11.0	13.0	12.0	14.0
	75%	18.0	43.0	17.0	12.0	21.0	16.0	29.0	19.0	30.0
	mean	15.6	29.8	14.7	9.9	16.7	14.4	19.4	17.0	19.4
Total asset	25%	46.2	76.1	35.3	70.8	32.6	45.2	64.0	52.9	32.0
	50%	107.8	200.9	77.4	175.2	63.9	101.2	152.6	127.1	85.2
	75%	302.8	587.7	205.0	500.0	134.8	275.0	410.0	394.6	263.2
	mean	694.2	1734.5	504.2	780.6	232.0	535.7	780.7	1171.3	1023.9
Export	25%	0.000	0.000	0.000	0.005	0.000	0.000	0.000	0.000	0.000
	50%	0.000	0.000	0.000	0.574	0.000	0.013	0.000	0.000	0.000
	75%	0.709	0.000	0.319	0.950	0.000	0.794	0.396	0.820	0.000
	mean	0.299	0.046	0.219	0.515	0.129	0.335	0.246	0.334	0.081
Observations		74,871	12,251	30,626	16,924	10,672	53,232	8,640	4,771	8,151

Note: The table reports 25th, 50th, 75th, and sample mean. TFP growth is calculated using the method by Levinsohn and Petrin (2003); *Leverage ratio* is a firm's total liabilities divided by total assets; *CF ratio* is a firm's cash flow divided by total asset; *Age* is the difference between a firm's observation year and establishing year; *Total asset* is a firm's total assets deflated by price deflators from NBS; *Export* is ??total export divided by total output.

As discussed above, we investigate if the potential threshold effect of leverage ratio varies with firm's ownership. The ASIP contains different types of paid-in capital, including state, collective, legal person, private, Hongkong, Macau and Taiwan, and foreign firms. Following Guariglia et al. (2011) and Chen and Guariglia (2013), we classify state-owned enterprises and collective enterprises, respectively, as firms with more than 50% of the firm's paid-in capital owned by the state and by urban or rural communities; private enterprises are firms with more than 50% paid-in capital owned by private persons and legal persons; and foreign firms are ones with with more than 50 percent of paid-in capital from Hongkong, Macau and Taiwan, or foreign countries.

We follow the literature on Chinese economy to further account for regional heterogeneity, since the substantial differences in geography, natural resources, and local development polices across regions in China may ultimately impact firm's investment decision as well as firm's productivity growth (Zhang et al., 2012; Wu and Zhou, 2013). We classify geographic regions into eastern, northeastern, central, and western area. The eastern area has eight provinces located near or along coastlines, which have been the most developed area of the

Table 2: Distribution of Leverage Ratio among Firms

Percentile	10%	25%	50%	75%	90%
Whole sample	0.187	0.348	0.530	0.693	0.817
By Ownership					
State	0.227	0.379	0.558	0.725	0.849
Private	0.262	0.428	0.601	0.749	0.859
Foreign	0.140	0.278	0.462	0.631	0.765
Collective	0.233	0.407	0.584	0.740	0.856
By Regions					
Eastern	0.203	0.366	0.546	0.710	0.834
Western	0.183	0.348	0.540	0.703	0.833
Northeastern	0.146	0.305	0.506	0.701	0.849
Central	0.172	0.348	0.547	0.716	0.841

economy since the outset of the economic reforms. The northeastern regions contain three provinces in which most of the nation's old state-owned enterprises have started up. The central region with eight provinces is known to lead primarily in the development of heavy industries due to its rich endowment of metal and non-metal resources. The western area having 12 provinces exhibits slowest economic growth due to limited transportation and complex terrain. See Appendix 4 for the specific province names in each region.

We construct our dependent variable *TFPG* by calculating Firm's TFP based on the method by Levinsohn and Petrin (2003).[1] Variables used in TFP calculation include *output* as a firm's real gross industrial product, *capital* as real total fixed capital, *labor* as real sum of wages and benefits, and *material* as the real intermediate material firms used in production.

Our key variable *leverage ratio* (L) is defined as a firm's total liability divided by total assets. We follow the literature to choose the following control variables $\mathbf{X}$ in firm's productivity regression (Coricelli et al., 2012; Chen and Guariglia, 2013): *CF ratio* is defined as a firm's cash flow divided by total assets; *Age* is the difference between a firm's observation year and establishment year; *Firm size* is the logarithm of a firm's real total asset; and *Export* is the export intensity, defined as total export divided by total outputs. Real variables are calculated using ex-factory producer price indexes from various issues of China statistical yearbooks (Guariglia et al., 2011). To mitigate the impact of firm's entry and exit on our analysis, we create a balanced panel data with 8,319 incumbent industrial firms during 1998-2007, giving 83,190 observations in total.

Table 1 provides descriptive statistics for the whole sample and samples classified by

[1]We use the STATA command `Levpet` to obtain the TFP estimate. See Petrin et al. (2004) for more details.

ownership types and regions. For the whole sample, the TFP growth has an average of 0.262, consistent with the findings in Li et al. (2009). The mean of leverage ratio is 0.516 in Chinese industrial firms, a fairly high level compared to 0.42 in Central and Eastern European countries (Coricelli et al., 2012). Data across ownerships demonstrates that firms on average have similar productivity growth rates around 26.2%, with foreign firms having the highest mean of growth rate (0.28), while SOEs having the lowest rate of growth (0.253). Consistent with the findings in Li et al. (2009) and Chen and Guariglia (2013), SOEs on average have considerably lower CF ratio and export intensity, but much larger firm size and age compared with other ownerships. Table 2 provides the distribution of leverage among each sample. We observe that foreign firms have consistently lower leverage ratios than other ownerships in all quantiles of the sample. Given their high (mean) CF ratio of 0.104, it may confirm the hypothesis by Guariglia et al. (2011) that foreign firms face relatively harder external financial constraints, thus have incentive to resort to finance through their own retailed earnings.

Data across regions reveals that the eastern region has the lowest mean of productivity growth but the highest mean of leverage ratio, whereas northeastern region exhibits the highest mean of TFP growth. This observation may seem to be counterintuitive, since northeastern area was initially targeted as the nation's "industrial base", containing many "heavy" industries and SOEs with stagnant productivity growth. However, SOEs in this region have undergone a series of reforms since early 1990s, and many policies have been established to attract high-tech industries such as electronic manufacturing (Smyth et al., 2001). In addition, Zhou and Hall (2017) show that the northeastern region is relatively more marketized, which enhances local entrepreneurship that may facilitate productivity improvement. Finally, eastern region has the highest mean of leverage ratio (0.532), followed by central region (0.528), western region (0.523), and northeastern region (0.502).

4 Empirical Results and Discussions

4.1 Whole sample

Prior to the estimation results discussion, we discuss some of the testing results in *Panel A* of Table 3, which shows the empirical p-values of the test in (12) under H_{01-03}. *Panel A* of Table 3 presents the results for the whole sample. We reject H_{01} that the parametric threshold regression in (1) is correctly specified at 1% level. We also reject H_{02} at 1% level, indicating the existence of a threshold structure in the nonparametric regression. Thus, our tests provide statistical support on the use of the nonparametric threshold regression in our

Table 3: Nonparametric Test Results For the Null Hypothesis H_{01-03}

Panel A	Whole Sample (0.5122)			
H_{01}: *Hansen model*	0.0000			
H_{02}: $\gamma_{\text{not exist}}$	0.0025			
H_{03}: $\gamma = L_{50}$	0.0905			
Panel B		*Ownerships*		
	Collective (0.5154)	Foreign (0.2903)	State (0.2931)	Private (0.5085)
H_{01}: *Hansen model*	0.0960	0.9975	0.7789	0.0005
H_{02}: $\gamma_{\text{not exist}}$	0.0910	–	–	0.0410
H_{03}: $\gamma = L_{50}$	0.0954	–	–	0.0704
Panel C		*Regions*		
	Central (0.4837)	Eastern (0.5137)	Northeastern (0.5099)	Western (0.3944)
H_{01}: *Hansen model*	0.0450	0.0000	0.1497	0.0201
H_{02}: $\gamma_{\text{not exist}}$	0.0910	0.0450	–	0.0000
H_{03}: $\gamma = L_{50}$	0.0603	0.1658	–	0.0050

Note: The threshold tests are based on F-type test by Ullah (1985), with three null hypotheses given by: 1) H_{01}*: Hansen Model* tests the null of a parametric threshold model in (1), 2) H_{02}*:* $\gamma_{not\ exist}$ tests the null of a fully nonparametric model without threshold, and 3) H_{03}*:* $\gamma = L_{50}$ tests the null of the threshold level equal to its median. We report the estimated threshold $\hat{\gamma}$next to each sample's name in parenthesis. All tests above are bootstrap tests with 399 bootstrap repetition.

study.

We present the nonparametric threshold regression for the whole sample in Table 4. The estimated threshold value is given by 0.5122, fairly close to its median of 0.530. Thus, more than half of the observations in our sample possesses leverage ratios higher than the estimated threshold value. The p-value of H_{03} in *Panel A* of Table 3 shows that the estimated threshold value is significantly different from the sample median at 10% level.

Following Henderson et al. (2019), we estimate functions $m_1(\cdot)$ in (4) and $m_2(\cdot)$ in (5) nonparametrically by splitting the sample according to the estimated threshold value. We estimate $\hat{m}_j(\cdot)$ with $j = 1, 2$ by holding all other regressors at their median values, and report the 50^{th} quantile of the estimated gradient distributions along with their asymptotic standard errors. We observe that a one percentage point increase in a firm's leverage ratio significantly rises its TFP growth by 0.361 of a percentage point before the threshold, but only 0.191 after the threshold, all else constant. The result in general confirms our first hypothesis on the threshold effect of leverage on firms' TFP growth, where the marginal impact of leverage is initially positive due to its disciplining rule of debt, but vanishes sharply due to the moral hazard issue when the threshold value is passed.

Other variables also exhibit significant effects on Chinese firms' TFP growth. The CF ratio is positively associated with firms' TFP growth in both regimes. This is consistent with Chen and Guariglia (2013) that Chinese firms facing harder financial constraints are more likely to support their growth through internal financing. Interestingly, the marginal effect of CF ratio drops when the leverage ratio exceeds the estimated threshold value. This

Table 4: Nonparametric Results for Whole Sample and Classification by Ownership

	Whole		**Private**		**Collective**	
Threshold estimate	$\hat{\gamma} = 0.5122$		$\hat{\gamma} = 0.5085$		$\hat{\gamma} = 0.5154$	
	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2
Leverage ratio	0.3610***	0.1910***	0.2699**	0.1284**	0.2198**	0.1957**
	(0.0670)	(0.0577)	(0.0693)	(0.0592)	(0.0933)	(0.0810)
ln(*Age*)	-0.0715**	0.045*	-0.0050	-0.0003	0.0665	-0.0323
	(0.0309)	(0.0231)	(0.0498)	(0.0191)	(0.0672)	(0.0252)
Export	0.3046***	0.1708***	0.7233***	-0.0873**	-0.4513***	0.2591***
	(0.0521)	(0.0390)	(0.0900)	(0.0354)	(0.1121)	(0.0683)
CF ratio	0.3231**	1.0461***	0.0207	0.3400***	0.4231***	2.5366***
	(0.1301)	(0.1011)	(0.2030)	(0.0684)	(0.1677)	(0.1020)
ln(*Firm size*)	0.0171	0.0236**	0.0399	0.0316***	0.0135	-0.0060
	(0.0156)	(0.0114)	(0.0255)	(0.0107)	(0.0308)	(0.0183)
Observations	35,406	39,465	10,936	19,690	4,128	6,544

Note: The estimated threshold effect of leverage ratio ($\hat{\gamma}$) is reported beneath each sample's name. Regime 1 and 2 reports the estimated median derivatives of each variable from $m_1(\cdot)$ and $m_2(\cdot)$, respectively, holding all else variables at the median level. The asymptotic standard errors are in parentheses. ***p<0.01, **p<0.05, *p<0.1.

indicates that firms with excessive leverage ratios may suffer from "bankrupt cost" and have to use the internal earnings to pay off liabilities, thus weakening its impact on rising firms' TFP growth. Engaging in exports exerts positive impact on TFP growth, consistent with the findings in Ma et al. (2014) that exporters are more productive in China. We also observe that a firm's age rises (declines) TFP growth with leverage ratio higher (lower) than the threshold, and a firm's size only improves the TFP growth in the second regime.

4.2 By Ownership

Panel B of Table 3 reports the testing results for samples based on four different firm's ownership. While we reject the parametric threshold regression structure under H_{01} for collective and private firms, we fail to reject it for state and foreign firms. For this reason, we do not proceed to perform the tests on H_{02} and H_{03} for state and foreign groups. Instead, we employ the parametric threshold regression on the two groups, with their estimation results given in Table 5.

The sample of SOEs gives out an estimated threshold value of 0.2784 at 1% significant level, indicated by the p-value of bootstrap test in (3) in Table 5. Notably, the threshold value is only about 14.4^{th} percentile in the sample, indicating that more than 85% observations in SOEs have leverage ratio higher than the threshold level. More importantly, the marginal impact of leverage ratio decreases from 0.096 to -0.122 as the leverage ratio rises

Table 5: Parametric Threshold Regression Results

	State	**Foreign**	**Northeastern**
Constant	0.221***	0.5781***	0.528***
	(0.0210)	(0.0033)	(0.0041)
ln(*Age*)	-0.093***	-0.074*	-0.076***
	(0.0110)	(0.0390)	(0.0320)
Export	-2.339***	-0.184***	0.491***
	(0.0940)	-(0.0610)	(0.0970)
CF ratio	2.214***	0.477	-0.139*
	(0.0980)	(0.9610)	(0.0770)
ln(*Firm size*)	-0.028	0.2130***	-0.024
	(0.0560)	(0.1090)	(0.0220)
Threshold Variable: Leverage ratio	$\hat{\gamma} = 0.2784^{***}$	$\hat{\gamma} = 0.3284$	$\hat{\gamma} = 0.4841$
Regime 1	0.096***	0.392***	0.172**
	(0.0341)	(0.0830)	(0.0810)
Regime 2	-0.122***	-0.419	-0.102
	(0.0580)	(0.1140)	(0.3298)
p-value of $\hat{I}$	0.0005	0.6375	0.3125
Observations	12,251	16,924	4,771

Note: The group of State (SOEs), Foreign, and Northeastern are analyzed with parametric threshold regression in (1). The lower panel reports the estimated threshold effect of leverage ratio ($\hat{\gamma}$) in each group, and the estimated coefficient of leverage ratio in Regime 1 and 2. The p-value of the test $\hat{I}$ in (3) is given in the second-to-the-last row. The asymptotic standard errors are in parentheses. ***p<0.01, **p<0.05, *p<0.1.

beyond the threshold, thus providing strong evidence on the excessive external borrowing. In other words, SOEs with higher-than-threshold leverage ratio exhibit a clear negative trend of the firms' TFP growth. Consistent with Jin et al. (2019) and Chen and Guariglia (2013), our results provide empirical evidence on the fact that SOEs face "soft budget constraint", with pervasively high leverage ratios in spite of SOEs' low TFP growth. In contrast, one percentage point rises in the leverage ratio of foreign firms significantly improves their TFP growth by 0.392 without a threshold fashion. Given a relatively harder external financial constraint in foreign firms (Guariglia et al., 2011), we expect that foreign firms may be disciplined efficiently by their limited access to debt to finance productivity-enhancement investment. We note that an intensive export significantly lowers TFP growth. One underlying reason is that a large share of China's exports are processing trade, most of which do not engage in product design or R&D investment in TFP-enhancing activities (Chen and Guariglia, 2013).

Table 4 report the nonparametric regression results for private firms and collective firms. Private firms, accounting for more than 40.9% of all observations in our sample, have an estimated threshold value of 0.5085 (about the 35^{th} percentile in the sample). We also observe a significant threshold effect of leverage on TFP growth in private firms: one percentage point

Table 6: Nonparametric Results for Classification by Region

	Region Classification					
	Eastern		**Western**		**Central**	
Threshold estimate	$\hat{\gamma} = 0.5137$		$\hat{\gamma} = 0.3945$		$\hat{\gamma} = 0.4837$	
	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2
Leverage ratio	0.5927***	0.1327**	0.9233***	0.2097**	-0.3851***	0.0800
	(0.0657)	(0.0614)	(0.1271)	(0.0911)	(0.1075)	(0.0903)
Ln(Age)	-0.0478	0.0222	0.1176	-0.0306	-0.1069	0.0466***
	(0.0359)	(0.0307)	(0.1337)	(0.0305)	(0.0347)	(0.0348)
Export	0.3014***	0.3843***	-1.4682***	-0.8572***	0.6319***	-0.2006**
	(0.0564)	(0.0394)	(0.2416)	(0.0580)	(0.1415)	(0.0945)
CF ratio	0.5822***	0.5465***	-0.9089	0.4608	0.7878***	0.3719***
	(0.1429)	(0.1009)	(0.9810)	(0.2497)	(0.1653)	(0.1105)
Ln(*Firm size*)	0.0540***	0.0421***	-0.0074	-0.0117	0.0082	-0.0087
	(0.0196)	(0.0131)	(0.0691)	(0.0196)	(0.0242)	(0.0144)
Observations	24,874	25,358	2,807	5,833	3,520	4,631

Note: The estimated threshold effect of leverage ratio ($\hat{\gamma}$) from (8) is reported for whole sample and sub-sample based on ownership. Wild-bootstrapped standard errors in parentheses. ***p<0.01, **p<0.05, *p<0.1

increase in leverage ratio improves the TFP growth by 0.27% before reaching the threshold level, but improves it by only 0.128% once the threshold is exceeded. Finally, collective firms reveals an estimated threshold of 0.5154 (about the 49^{th} percentile of L in the sample), with their TFP growth promoted marginally by a magnitude of 0.220 before the threshold level and 0.196 afterwards. Finally, a higher cash flow rises firms' TFP growth throughout all types of ownerships, consistent with the findings in Chen and Guariglia (2013).

Overall, our first hypothesis on the threshold effect of leverage ratio on TFP growth is well supported in all but foreign firms. A comparison of the estimated threshold structure across ownerships in Table 4 and 5 reveals that SOEs have the smallest threshold value of leverage ratio for which the TFP growth changes significantly. After exceeding the threshold, notably, the positive impact of leverage ratio only diminishes in collective and private firms, but turns to be negative in SOEs. Thus, the results are fairly consistent with our second hypothesis on the lower value of threshold in SOEs relative to others, potentially due to the dominating role of moral hazard problem induced by *soft budget constraint*.

4.3 By Region

Panel B of Table 3 reports test statistics for regional groups, showing that the implementation of the nonparametric threshold regression is appropriate for all those samples except for the northeastern region. Thus, we follow the test result to employ parametric threshold regression for northeastern region and report its result in Table 5. We do not find evidence on the threshold structure of leverage ratio in northeastern region, indicated by the insignificant threshold value estimate of 0.4841. Instead, we find that a higher leverage ratio significantly improves the TFP growth only in the first regime. Zhou and Hall (2017) document that the degree of marketization in northeastern region is higher than that of national average, which significantly increases local non-SOE entrepreneurship. This may explain the fact that approximately 40% firms are foreign owned (see Appendix 6), and the positive effect of leverage ratio in foreign firms (see Table 5) may dominant its adverse impact in the northeastern area.

Table 6 presents the results of nonparametric threshold regressions for eastern, western and central region. For the eastern region, we obtain a significant threshold value estimate of 0.5137 at 5% level (see *Panel C* of Table 3), quite close to its median of 0.5304. We continue to observe a significant threshold impact of leverage ratio on TFP growth. Namely, a one percentage point increase in leverage ratio before the threshold raises TFP growth by 0.593, which drops substantially to 0.133 after passing the threshold. In addition, a higher volume of export significantly improves productivity growth in this region. The result may not be surprising, given that most exporting Hi-tech sectors, typically with higher productivity, are located in the eastern region (Chen and Guariglia, 2013). The coefficients of cash flow ratio are positive and highly significant in eastern and central area, again consistent with the internal finance-growth hypothesis by Chen and Guariglia (2013).

Regarding the western region, the estimated threshold value is given by 0.3945, which is about the 32.5^{th} percentile in the sample. Thus, observations of excessive leverage ratio in this region accounts for roughly two-thirds in the sample. As the leverage ratio increases from the first to the second regime, the magnitude of the *TFPG* improvement significantly drops from 0.923 to 0.21. Looking at the empirical results across four different regions, we further observe that while firms in western area have the marginal impact of leverage ratio with the highest magnitude, they exhibit the lowest threshold value. When the endogeneity is taken into account (discussed in the next section), the partial effect of in the second regime further decreases to -0.039. Thus, the result supports the existence of the "excessive leverage" in western area, likely due to the fact that SOEs represent the highest share of firms in this region relative to others (about one-third). Shenggen and Zhang (2004) find that the lower productivity in this region is attributed to the immature level of infrastructure, education, and science and technology. Our results suggest that the excessive leverage ratio may serve as an additional key explanation for the stagnant TFP growth in the western area.

Finally, firms in central region have an estimated threshold value of 0.4837 (close to the

43^{th} percentile in the sample). Unlike other regions that discloses a positive and diminishing partial effect of leverage ratio, the central region have leverage ratio that declines the firm's TFP growth marginally by -0.385 percentage points in the first regime, but exhibits a fairly small and insignificant partial impact (i.e.,0.08) in the second regime. Similar to the western region, we expect that the result in central region may be attributed to the dominating role of SOEs, accounting for 30.3% of the sample. To summarize, we observe a similar pattern of the threshold structure of leverage ratio when taking into account the region heterogeneity, with western and central areas with a relatively higher share of SOEs exhibit severer excessive borrowing than eastern and northwestern area.

5 Robustness Check

5.1 Endogeneity

The leverage ratio (L) may be endogenous in the mean regression of TFPG. For instance, firms with higher TFP growth may be more likely to meet their liability of debt, thus inclining to rises corporate debt more than firms with lower TFP growth. For this reason, we apply a two-stage local polynomial estimator by Su and Ullah (2008) to cope with the potential endogenous problem of the leverage ratio. With our estimated threshold value, we estimate the mean regressions $m_j(\cdot)$ in (4)-(5) individually under a triangular system of equations, with $j = 1, 2$. For notation brevity, we drop the subscript j in the following discussion. To be more specific, in each group we have:

$$y_{it} = m(L_{it}, \mathbf{Z}^1_{it}) + u_{it} \tag{13}$$

$$L_{it} = g(\mathbf{Z}^1_{it}, \mathbf{Z}^2_{it}) + \epsilon_{it} \tag{14}$$

where L is the endogenous variable, $\mathbf{Z}^1$ is a vector of exogenous variables, and $\mathbf{Z}^2$ contains a relevant and valid instrumental variable. We choose the one-year lag of leverage ratio (i.e., $L_{i,t-1}$) as a common choice of instrument variable in the literature (See, for example, Coricelli et al. (2012)). Assuming $E(\epsilon|\mathbf{Z}^1, \mathbf{Z}^2) = 0$ and $E(u|\mathbf{Z}^1, \mathbf{Z}^2, \epsilon) = E(u|\epsilon)$, we have

$$\begin{aligned}
E(y|L, \mathbf{Z}^1, \mathbf{Z}^2, \epsilon) &= m(L, \mathbf{Z}^1) + E(u|L, \mathbf{Z}^1, \mathbf{Z}^2, \epsilon) \\
&= m(L, \mathbf{Z}^1) + E(u|L - g(\mathbf{Z}^1, \mathbf{Z}^2), \mathbf{Z}^1, \mathbf{Z}^2, \epsilon) \\
&= m(L, \mathbf{Z}^1) + E(u|\mathbf{Z}^1, \mathbf{Z}^2, \epsilon) \\
&= m(L, \mathbf{Z}^1) + E(u|\epsilon),
\end{aligned}$$

which is in fact an additive model such that $E(y|L, \mathbf{Z}^1, \mathbf{Z}^2, \epsilon) = M(L, \mathbf{Z}^1, \epsilon)$, where $M(L, \mathbf{Z}^1, \epsilon) = m(L, \mathbf{Z}^1) + w(\epsilon)$, with $w(\epsilon) \equiv E(u|\epsilon)$. The identification of our interest function $m(L, \mathbf{Z}^1)$ requires that $E(u) = 0$, so

$$m(L, \mathbf{Z}^1) = \int_{\epsilon} M(L, \mathbf{Z}^1, \epsilon) dF(\epsilon). \tag{15}$$

Estimation of (15) involves three steps. We first obtain the function estimate $\hat{g}(\mathbf{Z}^1_{it}, \mathbf{Z}^2_{it})$ in (14) through local linear regression and get the residuals $\hat{\epsilon}_{it} = L_{it} - \hat{g}(\mathbf{Z}^1_{it}, \mathbf{Z}^2_{it})$. Here, we denote H_1 as the bandwidth sequence in the first stage. Next, we estimate (15) with marginal integration by Linton and Nielsen (1995) by replacing it with its sample analogue:

$$\hat{m}(L, \mathbf{Z}^1) = \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{M}(L, \mathbf{Z}^1, \hat{\epsilon}_{it}), \tag{16}$$

where $\hat{M}(L, \mathbf{Z}^1, \hat{\epsilon})$ is the local constant estimator:

$$\hat{M}(L, \mathbf{Z}^1, \hat{\epsilon}) = \frac{\sum_{i=1}^{N} \sum_{t=1}^{T} K_H(W_{it}, W) y_{it}}{\sum_{i=1}^{N} \sum_{t=1}^{T} K_H(W_{it}, W)} \tag{17}$$

with $W_{it} = [L_{it}, \mathbf{Z}^1_{it}, \hat{\epsilon}_{it}]'$, and W is the evaluation set of points. In a similar fashion, we obtain the interested gradients estimate by replacing function estimate in (17) with its consistent derivate estimator:

$$\hat{m}'(L, \mathbf{Z}^1) = \frac{1}{NT} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{M}'(L, \mathbf{Z}^1, \hat{\epsilon}_{it}) \tag{18}$$

and $M'(L, \mathbf{Z}^1, \hat{\epsilon})$ is the local constant derivative estimate. We purposely choose local linear in step 1 and local constant estimator in step 2 to fulfill the theoretical requirement in Su and Ullah (2008) for consistent estimators of both unknown function and derivate estimates. The theoretical justification on our choice of local polynomial regression order are provided in Appendix 2.

Table 7 reports the median of the distribution of estimated partial effects of L under each sample.[2] For brevity, we only report results for groups whose threshold effects need to be estimated nonparametrically based on our test results in Table 3. In the whole sample as well as in private, collective, and eastern group, we observe that the significant threshold effect of leverage ratio on TFP growth continuously exists, indicated by a clear reduction in the magnitude of the L's marginal impact once the threshold value is exceeded. The leverage ratio in western area clearly exhibits an inverted U-relationship with TFP growth,

[2] We omit the effects of control variables to save space, and make the results available upon request.

Table 7: Results of Nonparametric Partial Effects of the Endogenous Leverage Ratio

	Whole		**Private**		**Collective**	
Threshold estimate	$\hat{\gamma}$=0.5122		$\hat{\gamma}$=0.5085		$\hat{\gamma}$=0.5154	
	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2
Leverage ratio	0.309***	0.132***	0.259***	0.104*	0.189**	0.176**
	(0.069)	(0.050)	(0.062)	(0.056)	(0.096)	(0.079)
Observations	31,704	34,848	9,090	16,382	3,295	4,742
	Eastern		**Western**		**Central**	
Threshold estimate	$\hat{\gamma}$=0.5137		$\hat{\gamma}$=0.3945		$\hat{\gamma}$=0.4837	
	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2
Leverage ratio	0.325***	0.119**	0.663***	-0.039*	-0.446***	-0.055
	(0.051)	(0.045)	(0.109)	(0.022)	(0.091)	(0.078)
Observations	22,319	24,978	2,492	5,178	3,120	4,108

while only the negative impact of leverage ratio shows up in the first regime of central area. A comparison of the estimation results difference between Table 4 and 7 reveals that the threshold structure of leverage ratio on TFP growth in general does not exhibit significant qualitative changes after controlling for the endogeneity of L.

5.2 An alternative indicator of External financing

Although a firm's leverage ratio has been served as a major indicator for external financing in the literature, it may not be a perfect indicator for policy implications. Namely, firms can purchase more assets and borrow more external fundings simultaneously to purposely maintain the leverage ratio at a certain level. For this reason, we use the annul growth rate of leverage ratio (LG) as an alternative measure of external financing to perform our empirical analysis. See Appendix 5 for the summary statistics of leverage growth across different samples.

We present the estimation results in Table 8, and we narrow down our focus only on the impact of LG for brevity. The whole sample gives an estimated threshold values of 0.1037 in LG, accounting for the 59.3^{th} of the observations in our sample. Similar to the leverage ratio, we clearly see that the LG possesses a significant threshold effect, where a one percentage increase in LG rises significantly the TFP growth by 0.105 and 0.038 of a percentage point before and after the threshold level, respectively. The similar pattern of the threshold effect also shows up in firms with different ownerships. The corresponding threshold values of LG are given by 0.0343 (49^{th} percentile) for SOEs, 0.1035 (57^{th} percentile) for private firms, 0.1137 (60^{th} percentile) for foreign firms, and 0.0673 (58^{th} percentile) for collective firms.

Table 8: Nonparametric Threshold Regression Results: Leverage Growth

	Whole sample							
Threshold estimate	$\hat{\gamma}$=0.1037							
	Regime 1	Regime 2						
Leverage growth	0.1051***	0.0376*						
	(0.0201)	(0.0208)						
Observations	39,070	27,336						
	By ownership							
	State		Private		Foreign		Collective	
Threshold estimate	$\hat{\gamma}$=0.0343		$\hat{\gamma}$=0.1035		$\hat{\gamma}$=0.1137		$\hat{\gamma}$=0.0673	
	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2
Leverage growth	0.0628**	-0.0720***	0.0381***	0.0250***	0.1694***	0.0419*	0.025***	0.0082***
	(0.0284)	(0.0190)	(0.0014)	(0.0018)	(0.0271)	(0.0216)	(0.0063)	(0.0024)
Observations	4,818	5,022	14,701	11,151	6,068	3,992	4,807	3,626
	By region							
	Eastern		Western		Central		Northeastern	
Threshold estimate	$\hat{\gamma}$=0.1037		$\hat{\gamma}$=0.1004		$\hat{\gamma}$=0.1093		$\hat{\gamma}$=0.1256	
	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2	Regime 1	Regime 2
Leverage growth	0.0539***	0.0209**	0.0621*	0.0327**	0.0157**	-0.0312*	0.0443*	0.0270*
	(0.0036)	(0.0091)	(0.0320)	(0.0146)	(0.0069)	(0.0180)	(0.0262)	(0.0140)
Observations	27,541	19,713	4,483	3,137	4,497	2,714	2,639	1,594

We also observe that the leverage growth exceeding its threshold value declines significantly the TFP growth of SOEs, again posing an inverted-U relationship that indicate the excessive borrowing issue. The estimated threshold value of SOEs is the smallest among all ownerships, fairly consistent with our previous finding.

The respective threshold values among regions are given by 0.1037 (58.3^{th} percentile) for the eastern region, 0.1004 (58.6^{th} percentile) for the western region, 0.1093 (62^{th} percentile) for the central region, and 0.1256 (62^{th} percentile) for the northeastern region. We observe that the partial effect of LG diminishes by roughly one-half in magnitude for most regions in the second regime, except that central region now shows an inverted-U relationship between LR and $TFPG$. In all, our empirical evidence on our hypotheses remain fairly robust with the alternative indicator of external financing.

6 Conclusion

Inspired from the recent literature on external finance-growth nexus and the excessive credit growth problem in Chinese industrial firms, in this paper we hypothesize and empirically investigate the threshold effect of leverage ratio on firms' TFP growth. Given the presence of malfunctioning financial market in China, we also hypothesize the variation of the threshold effect conditioning on firms' ownership. The understanding of the existence of the optimal leverage ratio for firms' productivity is non-trivial for policy implication on maintaining China's economy growth.

To alleviate the potential problem of model misspecification when identifying the threshold value, we test our hypotheses through employing a newly developed nonparametric threshold regression by Henderson et al. (2019), which allows us to impose no functional structure restrictions on the unknown mean regression of TFP growth. Our empirical results are fairly consistent with our hypotheses, showing that the leverage ratio lower (higher) than its threshold value exhibits a larger (smaller) partial effect on TFP growth. Relative to non-SOEs, SOEs and western region where SOEs dominate experience a much lower threshold value, with their TFP growth negatively affected by their excessive leverage ratio (i.e., an inverted U-shape). Given that a large portion of observation in SOEs, western and central region have excessive leverage ratio relative to the threshold level, we observe that the excessive borrowing problem is severe in firms or regions with soft budget constraint. Our results remain fairly robust when taking into account endogeneity and alternative indicator of external financing.

Our empirical analysis provides important inference on the optimal leverage ratio in a sense of maximizing firms' productivity growth. We thus render policy implications with an particular emphasis on identifying the excessive leverage, which has been a major concern for Chinese firms in recent years. The future research may focus more on the identification of the underlying mechanisms through which firms' corporate debt impacts TFP in China from a micro level foundation.

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Appendix 1.1 Bootstrap procedure for testing parametric threshold function model specification under $H_{01}$by Ullah (1985)

1. Given the original sample $\{y_{it}, L_{it}, \mathbf{X}'_{it}\}_{i=1,t=1}^{N,T}$, obtain RSS_0 from the parametric threshold regression model in (1) and RSS_1 from nonparametric threshold regression model in (4)-(5). Then calculate $\hat{T}$ in (12).
2. Generate centered bootstrapped residual $\{\hat{u}^*_{it}\}_{i,t=1}^{N,T}$, where for each observation $i = 1, ..., N$ and $t = 1, ..., T$, $\hat{u}^*_{it} = \frac{1-\sqrt{5}}{2}(\hat{u}_{it} - \bar{\hat{u}})$ with probability $\frac{1+\sqrt{5}}{2\sqrt{5}}$ and $\hat{u}^*_{it} = \frac{1+\sqrt{5}}{2}(\hat{u}_{it} - \bar{\hat{u}})$ with probability $\frac{1-\sqrt{5}}{2\sqrt{5}}$, where $\bar{\hat{u}}$ refers to the mean of $\hat{u}$. Then construct the bootstrapped independent variable $y^*_{it} = \hat{\alpha}_1 L_{it} I(L_{it} \leq \hat{\gamma}) + \hat{\alpha}_2 L_{it} I(L_{it} > \hat{\gamma}) + \mathbf{X}'_{it}\hat{\boldsymbol{\beta}} + u^*_{it}$, where the estimated parameters $(\hat{\alpha_1}, \hat{\alpha_2}, \hat{\boldsymbol{\beta}}', \hat{\gamma})$ are obtained from estimation in (1). Call $\{y^*_{it}, L_{it}, \mathbf{X}'_{it}\}_{i,t=1}^{N,T}$ the bootstrapped sample.
3. Compute $\hat{T}^*$ in the same way as in step 1, except we replace the original sample $\{y_{it}, L_{it}, \mathbf{X}'_{it}\}_{i=1,t=1}^{N,T}$ with the bootstrapped sample.
4. Repeat step 2-3 B times to obtain $\hat{T}^*_1$,$\hat{T}^*_2$,...,$\hat{T}^*_B$. Reject H_0 if $\hat{T} > \hat{T}^*_{1-\alpha}$, where $\hat{T}^*_{1-\alpha}$ is the upper $(1-\alpha)$ percentile value of its empirical distribution with α the significance level.

We set $\alpha = 0.05$ and $B = 399$, and implement the adaptive rule-of-thumb bandwidth selection to reduce computation burden.

Appendix 1.2 Bootstrap procedure for testing the existence of threshold under H_{02} by Ullah (1985)

1. Let $\mathbf{Z} = [L, \mathbf{X}']'$, and given the original sample $\{y_{it}, \mathbf{Z}'_{it}\}_{i=1,t=1}^{N,T}$ to obtain RSS_0 from nonparametric regression model *without* threshold, and using the original sample to obtain RSS_1 from nonparametric regression model *with* the threshold structure. Then calculate $\hat{T}$ in (12).
2. Generate centered bootstrapped residual $\{\hat{u}^*_{it}\}_{i,t=1}^{N,T}$, where for each observation $i = 1, ..., N$ and $t = 1, ..., T$, $\hat{u}^*_{it} = \frac{1-\sqrt{5}}{2}(\hat{u}_{it} - \bar{\hat{u}})$ with probability $\frac{1+\sqrt{5}}{2\sqrt{5}}$ and $\hat{u}^*_{it} = \frac{1+\sqrt{5}}{2}(\hat{u}_{it} - \bar{\hat{u}})$ with probability $\frac{1-\sqrt{5}}{2\sqrt{5}}$, where $\bar{\hat{u}}$ refers to the mean of $\hat{u}$. Then construct the bootstrapped independent variable $y^*_{it} = \hat{m}(\mathbf{Z}_{it}) + u^*_{it}$, where the function estimate $\hat{m}(\cdot)$ estimated by local linear estimator in (7). Call $\{y^*_{it}, \mathbf{Z}'_{it}\}_{i,t=1}^{N,T}$ the bootstrapped sample.
3. Compute $\hat{T}^*$ in the same way as in step 1, except we replace the original sample $\{y_{it}, \mathbf{Z}'_{it}\}_{i=1,t=1}^{N,T}$ with the bootstrapped sample.
4. Repeat step 2-3 B times to obtain $\hat{T}^*_1$,$\hat{T}^*_2$,...,$\hat{T}^*_B$. Reject H_0 if $\hat{T} > \hat{T}^*_{1-\alpha}$, where $\hat{T}^*_{1-\alpha}$ is the upper $(1-\alpha)$ percentile value of its empirical distribution with α the significance level.

We set $\alpha = 0.05$ and $B = 399$, and implement the adaptive rule-of-thumb bandwidth selection to reduce computation burden.

Appendix 1.3 Bootstrap procedure for testing specific value of the threshold under $H_{03}$by Ullah (1985)

1. Let $\mathbf{Z} = [L, \mathbf{X}']'$, and given the original sample $\{y_{it}, \mathbf{Z}'_{it}\}_{i=1,t=1}^{N,T}$ to obtain RSS_0 from nonparametric regression model with threshold variable L equal to its median level L_{50} (i.e., $\gamma_0 = L_{50}$), and RSS_1 from nonparametric threshold regression model with threshold equal to the estimated value in (10). Then calculate $\hat{T}$ in (12).

2. Generate centered bootstrapped residual $\{\hat{u}^*_{it}\}_{i,t=1}^{N,T}$, where for each observation $i = 1, ..., N$ and $t = 1, ..., T$, $\hat{u}^*_{it} = \frac{1-\sqrt{5}}{2}(\hat{u}_{it} - \bar{\hat{u}})$ with probability $\frac{1+\sqrt{5}}{2\sqrt{5}}$ and $\hat{u}^*_{it} = \frac{1+\sqrt{5}}{2}(\hat{u}_{it} - \bar{\hat{u}})$ with probability $\frac{1-\sqrt{5}}{2\sqrt{5}}$, where $\bar{\hat{u}}$ refers to the mean of $\hat{u}$. Then construct the bootstrapped dependent variable $y^*_{it} = \hat{\alpha}(\boldsymbol{Z}_{it}) + \hat{\beta} \cdot I(\boldsymbol{Z}_{it} > \hat{\gamma}_0) + u^*_{it}$, where all estimates $[\hat{\alpha}(\cdot), \hat{\beta}]'$ are from nonparametric regression model where $\hat{\gamma} = \hat{\gamma}_0$. Call $\{y^*_{it}, \mathbf{Z}'_{it}\}_{i,t=1}^{N,T}$ the bootstrapped sample.

3. Compute $\hat{T}^*$ in the same way as in step 1, except we replace the original sample $\{y_{it}, \mathbf{Z}'_{it}\}_{i=1,t=1}^{N,T}$ with the bootstrapped sample $\{y^*_{it}, \mathbf{Z}'_{it}\}_{i,t=1}^{N,T}$.

4. Repeat step 2-3 B times to obtain $\hat{T}^*_1$,$\hat{T}^*_2$,...,$\hat{T}^*_B$. Reject H_0 if $\hat{T} > \hat{T}^*_{1-\alpha}$, where $\hat{T}^*_{1-\alpha}$ is the upper $(1-\alpha)$ percentile value of its empirical distribution with α the significance level.

We set $\alpha = 0.05$ and $B = 399$, and implement the adaptive rule-of-thumb bandwidth selection to reduce computation burden.

Appendix 2 Theoretical Justification on the Polynomial Regression Order in Section 5.1.

To justify the validity of our choice for the order of polynomial regression in the first two-step procedure, recall that in the triangular system equation,

$$TFPG = m_j(L, \mathbf{Z}_1) + u \tag{19}$$

$$L = g_j(\mathbf{Z}_1, \mathbf{Z}_2) + \epsilon, \tag{20}$$

where L is endogenous so that $E(u|L) \neq 0$, $Z_1 \in \mathbb{R}^{q_1}$ is a vector of exogenous regressors, $Z_2 \in \mathbb{R}^{q_2}$ is a vector of instrumental variables, and $j = 1, 2$. Denote $H_1 \in \mathbb{R}^{q_1+q_2}$ and $H_2 \in \mathbb{R}^{q_1+2}$ be a set of bandwidth used in the first and second step, respectively, and let p_1 and p_2 be the order of polynomial regression in the first and second step, respectively. Note that local linear and local constant regression are equivalent to the polynomial regression of order one or zero, respectively, thus we have $p_1 = 1$ and $p_0 = 0$.

Assumption A5 and Remark 5 in Su and Ullah (2008) derive the specific rate of convergence of H_1 and H_2 that allow for function and gradient estimates of $m(L, \mathbf{Z}_1)$ to be consistently estimated. Specifically, we have:

$$H_2 \propto n^{-\frac{1}{\gamma_2}} \tag{21}$$

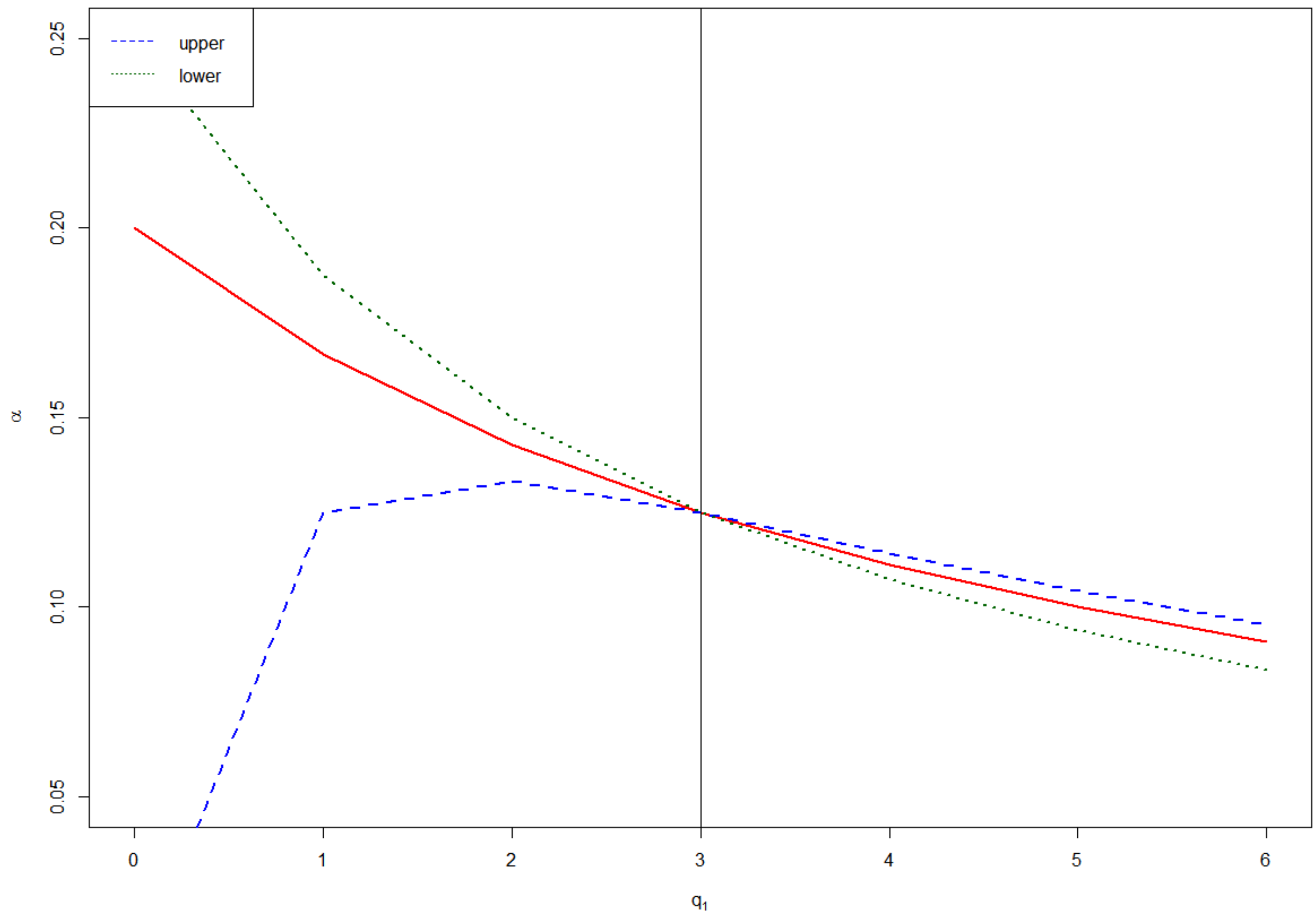
$$H_1 \propto n^{-\frac{1}{\gamma_1}} \tag{22}$$

where $\gamma_2 = 2(p_2 + 1) + q_1 + 1$ and $\gamma_1 = 2(p_1 + 1) + (q_1 + q_2)$. Furthermore, let $\alpha = 1/\gamma_1$, they show that the consistent estimator also requires the choice of p_1 and p_2 to satisfy $H_1 \propto n^{-\alpha}$, where

$$\underline{\alpha} < \alpha < \bar{\alpha} \tag{23}$$

with $\underline{\alpha} = \frac{1}{\gamma_2}\max\{\frac{p_2+1}{p_1+1}, \frac{p_2+3}{2(p_1+1)}\}$ and $\bar{\alpha} = \frac{1}{\gamma_2}\left(\frac{p_2+q_1}{q_1+q_2}\right)$. In other words, we need to choose p_1 and p_2 to ensure H_1 to converges at a rate bounded by the required interval in (23). Since we have $q_1 = 4$ and $q_2 = 1$, our choice of $p_1 = 1$ (i.e., local linear estimator) and $p_2 = 0$ (i.e., local constant estimator) is theoretically justified. This can be seen clearly from Figure 1 below, which plots the dimension of $\mathbf{Z}^1$ (i.e., q_1) and α, with $q_2 = 1$. The green dot line and blue dash line represents $\underline{\alpha}$ and $\bar{\alpha}$, respectively. Clearly, our choice of $p_1 = 1$ and $p_2 = 0$ maintain the inequality in (23) so long as $q_1 > 3$, which is true in our case as $q_1 = 4$.

Figure 1: The rate of convergence H_1 given $p_1 = 1$ and $p_2 = 0$



Note: q_1 is the dimension of exogenous variables $\mathbf{Z}^1$ and $\alpha = 1/\gamma_1$. The value of α given each value of q_1 is calculated by setting $p_1 = 1$, $p_2 = 0$, and $q_2 = 1$.

Appendix 3 Chinese Industrial Two-Digit Classification Codes

CICC	**Industry Name**
6	Exploration and abstention of coal and char industry
7	Exploration of oil and natural gas industry
8	Picking of ferrous metal mine industry
9	Picking of non-ferrous metal mine industry
10	Picking of nonmetal mine industry
11	Other mining industry
13	Agriculture and food processing industry
14	Foodstuff manufacturing industry
15	Soft drink manufacturing industry
16	Tobacco manufacturing industry
17	Textile industry
18	Waving costume, shoes and cap manufacturing industry
19	Leather, fur and feather manufacturing industry
20	Wood working, and wood,bamboo,bush rope,palm, and straw manufacturing industry
21	Furniture manufacturing industry
22	Paper making and paper products industry
23	Print and copy of record vehicle industry
24	Stationary and sporting goods manufacturing industry
25	Oil processing, coking and nuclear manufacturing industry
26	Chemical material and chemical product manufacturing industry
27	Medicine manufacturing industry
28	Chemical fiber manufacturing industry
29	Rubber product industry
30	Plastics product industry
31	Nonmetallic mineral product industry
32	Ferrous metal refining and calendaring processing industry
33	Non-ferrous metal refining and calendaring processing industry
34	Metal product industry
35	Universal equipment manufacturing industry
36	Task equipment manufacturing industry
37	Transport and communication facilities manufacturing industry
39	Electric machine and fittings manufacturing industry
40	Communication apparatus ,computer and other electric installation manufacturing industry
41	Instrument and meter, stationery machine manufacturing industry
42	Handicraft and other manufacturing industry
43	Removal and processing of obsolete resource and material industry
44	Electricity and thermal manufacturing and supplying industry
45	Combustion gas manufacturing and supplying industry
46	Water manufacturing and supplying industry

Note: CICC refers to two-digit industry classification code based on NBSC. Industries are mainly classified as manufacturing (CICC 13-43), mining (CICC 6-11), and public utility industries (CICC 44-46).

Appendix 4 Geographic-Based Region Classification

Eastern Region	Beijing	Tianjin	Hebei	Shanghai	Jiangsu
	Zhejiang	Shandong	Fujian	Guangdong	Hainan
Northeastern Region	Liaoning	Jilin	Heilongjiang		
Central Region	Shanxi	Anhui	Jiangxi	Henan	Hubei
	Hunan				
Western Region	Neimenggu	Guangxi	Chongqing	Sichuan	Guizhou
	Yunnan	Tibet	Shǎnxi	Gansu	Qinghai
	Ningxia	Xinjiang			

Note: Region Classification is on the basis of National Bureau of Statistics of China.

Appendix 5 Leverage Growth Summary

Percentile	10%	25%	50%	75%	90%	mean
whole sample	-0.3176	-0.112	0.0416	0.2825	0.7378	0.5041
By ownership						
state	-0.235	-0.071	0.0392	0.2109	0.5591	0.372
private	-0.2738	-0.0905	0.0543	0.2996	0.7349	0.3603
foreign	-0.3967	-0.1645	0.0344	0.3165	0.8393	0.7239
collective	-0.293	-0.0978	0.0263	0.229	0.6411	0.3393
By region						
eastern	-0.3173	-0.1165	0.0438	0.2866	0.7225	0.4377
western	-0.3256	-0.1033	0.0434	0.2869	0.781	0.9073
northeastern	-0.3501	-0.1316	0.0334	0.2839	0.8471	0.3173
central	-0.2887	-0.0813	0.031	0.248	0.7835	0.6228

Appendix 6 Percentage of Firm ownership Types across Regions

	Eastern region	Western region	Central region	Northeastern region
State firms	9.6%	31.8%	30.3%	20.4%
Private firms	41.5%	27.6%	42.5%	27.8%
Collective firms	14%	11.7%	18.6%	10.3%
Foreign firms	34.9%	28.9%	8.6%	41.5%