Effect of Privatization on Export through Changes in Productivity and Financial Factors*

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Abstract

This paper examines whether or not privatization of Chinese state-owned enterprises (SOEs) promotes export, and if so, what channels generate this effect. Using firm-level data for the Chinese manufacturing sector for the period 2000-2007, we find that privatized SOEs are more likely to engage in exporting than remaining SOEs. We also find that privatized SOEs improve productivity more, while privatization does not seem to tighten credit constraints. Therefore, we conclude that privatization of SOEs leads to more active exporting through productivity improvement.

Keywords: privatization, export, productivity, credit constraints, China

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1. Introduction

This paper examines whether or not privatization of Chinese state-owned enterprises (SOEs) promotes export, and if so, what channels generate this effect. This is motivated by the following three strands of literature. First, many studies using firm-level data find that firms' exporting decisions are affected by, for example, the firm size, the productivity level, and spillovers from other exporters (Clerides, Lach, and Tybout, 1998;, Bernard and Jensen, 1999 and 2004; and Greenaway and Kneller, 2004, among many others). More recently, some studies including Greenaway, Guariglia, and Kneller (2007), Du and Girma (2007), Muûls (2008), and Feenstra, Li, and Yu (2011), test whether financial conditions of firms affect exporting decisions. A possible reason why financial conditions matter is that since exporting requires initial investment in, for example, marketing abroad and modification of products to foreign preferences, it is more difficult for credit-constrained firms to export. Second, as many SOEs in China were privatized in recent years, many studies look at outcomes of the privatization. For example, Jefferson and Su (2006) and Bai, Lu, and Tao (2009) find that privatization of SOEs improves firm performance such as the productivity level and the firm size. Third, privatization, on the other hand, may lead to tighter credit constraints, since it is argued that state ownership is often associated with soft budget constraints (Qian and Roland, 1996).

Combining the three strands of literature, it is unclear whether or not privatization of SOEs stimulates export, since privatization encourages export through productivity improvement but discourages it through tighter credit constraints. Therefore, we investigate effects of SOEs' privatization on exporting decisions through these two channels, using a rich firm-level data set for the Chinese manufacturing sector for the period 2000-2007. Our estimation procedure takes the following two steps. First, we check standard factors, such as productivity and financial conditions, in fact affect firms' exporting decisions. Second, we evaluate effects of privatization on exporting decisions and possible two channels toward exporting, productivity and financial factors.

In the first step, we find that productivity has a positive effect on exporting decisions while credit constraints have a negative effect, confirming standard results. Then, in the second step, we find that privatization leads to a larger probability of exporting and higher productivity, while not affecting financial factors. Therefore, we conclude that privatized SOEs are more likely to engage in exporting than remaining SOEs, since privatized SOEs improve productivity but does not tighten credit constraints.

2. Empirical Methodology

2.1 Determinants of Export Behavior

There is already a thick literature on determinants of export. Many of the existing studies find a statistically significant effect of previous experiences in exporting, the productivity level, the firm size, financial conditions, and the size of exporters in the same region. An econometric problem of this type of analysis is endogeneity of regressors. For example, productive firms may tend to engage in exporting, but at the same time, exporters may improve productivity. Therefore, correlation between exporting and the productivity level does not necessarily indicate a causal relation from productivity to export.

To correct for possible biases in estimation of determinants of export due to endogeneity, we employ three-stage least squares (3SLS) estimation in which instrumental variables are lagged regressors and the private ownership ratio. We include the private ownership ratio in the set of instruments, since we assume that private ownership affects some of the regressors, such as the productivity level, the firm size, and financial conditions. However, we assume that private ownership does not directly affect firms' exporting behavior, although private ownership does affect it through the regressors.

2.2 Effect of Privatization of SOEs

When we estimate effects of privatization of SOEs, we also encounter endogeneity problems since privatized SOEs are not chosen in a random manner. To correct for biases due to endogeneity, we employ a propensity score matching (PSM) technique developed by Rosenbaum and Rubin (1983).[1]

In the PSM estimation, we identify the average effect of treatment on the treated (ATT), i.e., the average effect of privatization on the export behavior, productivity, and financial conditions. Let D_{it} be a dummy variable indicating SOE i's privatization in year t. The outcome variable (an indicator variable for exporting, the productivity level, or a financial variable) of firm i in year $t + s$ ($s \geq 0$) is denoted by $Y_{i,t+s}(D_{it})$, which depends on D_{it}. Then, ATT can be defined as

$$ATT = E(Y_{i,t+s}(1) - Y_{i,t+s}(0) | D_{it} = 1, X_{i,t-1}), \quad (1)$$

where $X_{i,t-1}$ denotes characteristics of firm i in year t-1. In words, ATT is the average difference between the outcome of privatized SOEs and their counter-factual outcome if they had not been privatized.

[1] Other methods to estimate ATT include Mahalanobis-metric matching (Rubin, 1980) and weighting by the inversed propensity score (Hirano, Imbens and Ridder, 2003). This study employs PSM, because this is more widely used in the literature.

To identify ATT, we need to assume "strong ignorability," i.e., unconfoundedness and overlap (Rosenbaum and Rubin, 1983). The unconfoundedness assumption is given by

$$Y(1), Y(0) \amalg D \mid X, \quad (2)$$

implying given a set of observable characteristics *X*, potential outcomes are independent of treatment (privatization) assignment. The overlap assumption is given by

$$0 < \Pr(D = 1 \mid X) \equiv P(X) < 1, \quad (3)$$

ensuring a positive probability of privatization and non-privatization. Under the strong ignorability, Rosenbaum and Rubin (1983) show that potential outcomes are also independent of treatment conditional on the probability that the firm is privatized, or the propensity score $P(X)$, and hence that ATT in equation (1) becomes

$$ATT = E(Y_{i,t+s}(1) | D_{it} = 1, P(X_{i,t-1})) - E(Y_{i,t+s}(0) | D_{it} = 0, P(X_{i,t-1})). \quad (4)$$

The first term on the right-hand side of equation (4) is estimated by the average of actual outcomes of participants. Each privatized SOE is matched with a remaining SOE that has a similar propensity score or a set of remaining SOEs weighted by their propensity scores. Then, the second term, the expected outcome of privatized SOEs if they had not been privatized, can be estimated by the average outcome of the matched remaining SOEs.

When panel data are available, as in the case of this paper, one can employ a difference-in-differences (DID) PSM estimator of the ATT proposed by Heckman, Ichimura and Todd (1997, 1998), in which we examine the treatment effect on the *change* in the outcome measure. An advantage of the use of the DID-PSM estimation is that it can eliminate time-invariant effects on the outcome variable. Heckman, Ichimura and Todd (1997, 1998) and Smith and Todd (2005) find that DID estimators perform better than matching estimators without using DID. Formally, the DID-PSM estimator is defined as

$$DID - PSM = \frac{1}{N} \sum_{i \in I_1} \left(\Delta Y_{i,t+s}(1) - \sum_{j \in I_0} W\left(P(X_{i,t-1}), P(X_{j,t-1})\right) \Delta Y_{j,t+s}(0) \right), \quad (5)$$

where $\Delta Y_{i,t+s} \equiv Y_{i,t+s} - Y_{i,t-1}$. I_1 and I_0 are respectively the treatment and the matched control group, and N is the number of observations in the treatment group. W is a weight determined by the distance between propensity scores of the treated and the matched control observations.

We focus on firms which are fully owned by the state and do not engage in exporting in year $t-1$, since by so doing, we can see the change in state ownership and exporting behavior more clearly. To obtain the DID-PSM estimator of the impact of privatization of non-exporting SOEs given the dataset in hand, we first examine how privatization is determined, using a probit model. The covariates used in the probit estimation are similar to those used in Bai, Lu, and Tao (2009): the log of total factor productivity (TFP); the log of the number of workers; the liquidity ratio, defined as firms' current assets less current liabilities over total assets; the long-term liability ratio, defined as the ratio of long-term liabilities to total assets; firms' age; the log of total exports in the region; and dummy variables for industry, region, and year. We also use the square term of the log of TFP, the log of labor, and the age to control for possible non-linear relations.

Based on the propensity score from the probit estimation, we employ two alternative matching methods to create the matched control observations: caliper and kernel matching. In both methods, we impose a common support condition to satisfy the overlap assumption (equation [3]), dropping observations in the treatment group whose propensity score is higher than the maximum or lower than the minimum score among observations in the control group. In the case of caliper matching, each observation in the treatment group is matched with a control observation that has the closest propensity score to the treated observation's score within the maximum score distance, or the caliper. In this study, the caliper is set at 0.05. In the case of kernel matching, each treated observation is matched with the weighted average of all control observations in the common support region. In the weight function W in equation (5), we use the Epanechnikov kernel function and set the bandwidth at 0.06.[2]

We match treatment observations with control observations in the same year, following Arnold and Javorcik (2005). In the case of evaluation of impacts of a job training program, Heckman, Ichimura and Todd (1997) find that matching estimates perform well when participants and non-participants reside in the same local labor market. Therefore, they argue that geographic mismatches should be avoided in matching estimation. In the case of this paper, time, rather than geographic, mismatches may be more substantial, since the data of this paper contain an eight-year period as explained in detail later and SOEs were privatized throughout the period. Therefore, the time restriction is imposed in this study.

After the matching, the treatment and the control group should have similar characteristics before the privatization. To check whether this is the case, we employ two types of balancing test. First, a simple t test is used to examine whether the mean of each covariate differs between the treatment and the control group after matching. In addition, following Girma and Gorg

[2] Another widely-used kernel is the Gaussian kernel. In addition, a generalized version of kernel matching, called local linear matching, is proposed by Heckman, Ichimura and Todd (1997, 1998). According to Fan (1992), an advantage of local linear estimators over kernel estimators is better adaptation to different data densities. I experimented with Gaussian kernel matching and local linear matching and found qualitatively the same and quantitatively similar ATT estimates as in the case of Epanechnikov kernel matching. However, I also found that these types of matching led to a failure in balancing tests, explained below. Therefore, the benchmark estimation employs Epanechnikov kernel matching.

(2007), the Hotelling's T-squared test is performed to jointly test the equality of the mean between the two groups for all covariates. Second, we run probit using the sample after matching and compare the pseudo-R^2 with that obtained from the probit estimation using the sample before matching. In addition, a likelihood-ratio test is performed to test whether all the estimated coefficients from the after-matching probit estimation are zero. These tests are proposed by Sianesi (2004). If matching is successful, the after-matching probit should have no explanatory power so that the pseudo- R^2 should be low and the estimated coefficients should be close to zero.

Given that the treatment and the control group pass the balancing tests, we compute the DID-PSM estimator using equation (5). To take the advantage of the panel data for this paper which cover an eight-year period from 2000-2007, the length of years between treatment and impact evaluation (s in equation [5]) is set at either zero, one, or two. The standard error of the DID-PSM estimator is obtained by bootstrapping based on 100 replications, following Smith and Todd (2005). Most existing studies use bootstrapping standard errors for PSM estimators, since multiple steps in PSM estimation, including estimation of propensity scores and matching procedures, lead to larger variation in PSM estimators than standard estimators with only one step.

3. Data

3.1 Description of the Data Set

The data utilized in this paper are based on the annual survey of manufacturing firms at the firm level conducted by China's National Bureau of Statistics. The survey targets all SOEs and non-SOEs with annual sales of 5 million Renminbi or more, and the response to the survey is compulsory.

Our sample covers the period 2000-2007. We drop from the sample firms for which reported sales, exports, or the book value of fixed assets are negative, or sales are less than exports in any year. We construct real values of outputs, inputs, and capital stocks, using industry-level deflators built by Brandt, van Biesebroeck, and Zhang (2009).[3]

We use total factor productivity (TFP) for our productivity measure.[4] TFP is obtained from the method developed by Olley and Pakes (1996), in which the labor and capital elasticity is estimated for each 2-digit industry.

The private ownership ratio is defined as the share of any private capital, including collective capital, legal person capital, individual capital, and foreign capital, in the total equity.

[3] These deflators are available at http://www.econ.kuleuven.be/public/n07057/China/.
[4] When labor productivity is used, the main results do not change.

In the propensity score matching estimation, we focus on firms that are fully owned by the state and do not export in year t – 1. These firms are considered to be privatized in year t if the private ownership ratio is more than 50% in years t, t + 1, and t + 2.

In all estimations, we use industry and region dummies. Industries are classified by the Industrial Classification and Codes for National Economic Activities at the 2-digit level. Regions are classified by the zip code of each firm at the 1-digit level, although the zip code is originally at the 6-digit level. When we construct total exports in the region, we use the modified zip code at the 2-digit level.

3.2 Descriptive Statistics

Figure 1 illustrates the change in the number of Chinese by export and ownership status from 2000 to 2007, using the whole sample. 20 percent of all firms are SOEs in 2000, but their share has declined drastically over time to 2.1 percent in 2007. Thus, our sample covers the period of drastic privatization. Around 20 percent of firms are exporters, and the share of SOEs in exporters is small and declining from 14 percent in 2000 to 1.7 percent in 2007. In stark contrast, the number of private exporters increases from 20,000 in 2000 to 54,000 in 2007.

Our sample for the examination of determinants of exporting decisions consists of 600,204 firm-year observations, including SOEs and non-SOEs and exporters and non-exporters. The number of observations is smaller than the total number of observations in the whole sample (data set) since our estimation requires lagged variables. Summary statistics of the key variables are provided in Table 1.

The sample to examine effects of privatization of non-exporting SOEs consists of 13,991 firm-year observations. Since we define privatization in year t as being privatized in year t and continuing to be privatized up to year t + 2 and use variables in year t – 1 as independent variables, we focus on privatization from 2001 to 2005 although the whole data cover the period 2000-2007. The number of firms by privatization, export status, and year is presented in Table 2. This table shows that in each year, 9 to 18 percent of incumbent SOEs are privatized. Among the privatized SOEs, some immediate start exporting after privatization. The number of observations declines over time, since the number of incumbent SOEs declines as SOEs are privatized.

Summary statistics of the key variables in the sample is shown in Table 3. The second to the fourth rows indicate that 2.4 percent of non-exporting SOEs in year t – 1 are exporters in year t, 3.2 percent in t + 1, and 3.6 percent in t + 2. Note that some of the exporters are remaining SOEs while some are privatized.

4. Results

4.1 Determinants of Export

Table 4 presents results from the 3SLS estimations. As shown in existing studies such as Greenaway and Kneller (2004) and Todo (2011), previous experience in exporting has a large effect on current exporting, explaining 90 percent of exporting behavior. The productivity level measured by TFP and the firm size measured by the number of workers also have a positive and significant effect on the export decision, as often found in the literature. The age of firms has a negative and significant effect, suggesting that older firms are less likely to be engaged in exporting, as found in Du and Girma (2007) in the case of China. In addition, information spillovers from neighboring exporters seems to affect the export decision, as the coefficient on the log of total exports in the same region at the 3-digit level is positive and significant. This finding is consistent with those of Aitken, Hanson, and Harrison (1997), Barrios, Görg, and Strobl (2003), Greenaway, Sousa, and Wakelin (2004), Bernard and Jensen (2004), and Todo (2011).

In each of columns (1)-(3) of Table 4, we use either the liquidity ratio (the ratio of current asset less current debt to the total asset), the leverage ratio (the ratio of current debt to current asset), or the ratio of long-term debt to the total asset as a measure of financial conditions. Among these three measures, the ratio of long-term debt to the total asset has a positive and significant effect. This finding suggests that firms which can increase their long-term debt are more likely to engage in export, since exporting may require long-term investment in product modification or marketing abroad. In other words, firms with softer credit constraints can export more easily. The result is consistent with the findings of Du and Girma (2007) for China and Muûls (2008) for Belgium.

The liquidity ratio or the leverage ratio has no significant effect on the export decision. This is consistent with the finding of Greenaway, Guariglia, and Kneller (2007). The difference between the results using the liquidity ratio and the leverage ratio and those using the long-term debt ratio may come from that exporting requires long-term investment.

In summary, we confirm the results from the existing studies that determinants of export include the previous experience, the productivity level, the firm size, spillovers from regional exporters, and some financial factors.

4.2 Effect of Privatization

Now, we examine whether privatized firms are more likely to export and if so, through which channels export is stimulated, using the propensity score matching method described in Section

2.3. For this purpose, we first run a probit model to estimate how SOEs are chosen for privatization. The results shown in column (1) of Table 5 indicate that own TFP and total exports in the same region have a positive effect on privatization. The number of workers has an inverted U-shaped effect, while the firm age has a U-shaped effect. The liquidity ratio has a negative effect, implying that firms with larger net current asset are less likely to be privatized. The pseudo R squared is 0.118.

Using the propensity score obtained from the probit estimation, we match privatized non-exporting SOEs with remaining non-exporting SOEs within the same year and see how the export behavior in the following years differs between the two groups. The results of the balancing tests shown in Table 6 indicate that although privatized SOEs and remaining SOEs are systematically different before matching, the two groups share very similar characteristics after matching.

The results for the effects of privatization are shown in Table 7. The first set of results indicates that privatized non-exporting SOEs are more likely to start exporting after privatization. The effect is quantitatively large. As shown in Table 3, 2.4 percent of non-exporting SOEs start exporting. According to the PSM estimation, privatization increases the probability of exporting in the year of privatization by 1.8 percentage points, exporting in the next year by 2.6, and exporting two years later by 1.8. Therefore, roughly speaking, privatization doubles the probability of exporting for non-exporting SOEs.

Now, why do privatization stimulate export? As we have seen in Section 4.1, firms are more likely to export when they are more productive, larger, and less credit-constrained. The latter rows of Table 7 show whether privatization affect these factors of the export decision. First, the effect of privatization on productivity growth is positive, statistically significant, and quantitatively large. Privatization improves TFP by 4.1 percent in that year and by 5.1 and 5.8 percent in one and two years, respectively. Second, the effect of privatization on the firm size measured by the number of workers is not clear. The effect on the change in the firm size is positive but not significant for the first two years and significant three years later. Third, privatization does not seem to have a significant effect on any of the three financial variables, the liquidity ratio, the leverage ratio, and the ratio of long-term debt to total asset. In other words, data does not support our hypothesis that SOEs are less credit-constrained. However, this ambiguous effect of privatization on credit constraints is consistent with the theoretical prediction of Lin and Li (2008).

From these results, we conclude that privatized SOEs are more likely to engage in export, mostly because privatization improves productivity.

5. Concluding Remarks

This paper examines whether or not privatization of Chinese state-owned enterprises (SOEs) promotes export and if so, what channels generate this effect. Using firm-level data for the Chinese manufacturing sector for the period 2000-2007, we find that privatized SOEs are more likely to engage in exporting than remaining SOEs. We also find that privatized SOEs improve productivity more, while privatization does not seem to tighten credit constraints. Therefore, we conclude that privatization of SOEs leads to more active exporting through productivity improvement.

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Figure 1: Number of Firms by Export Status and by State Ownership from 2000 to 2007

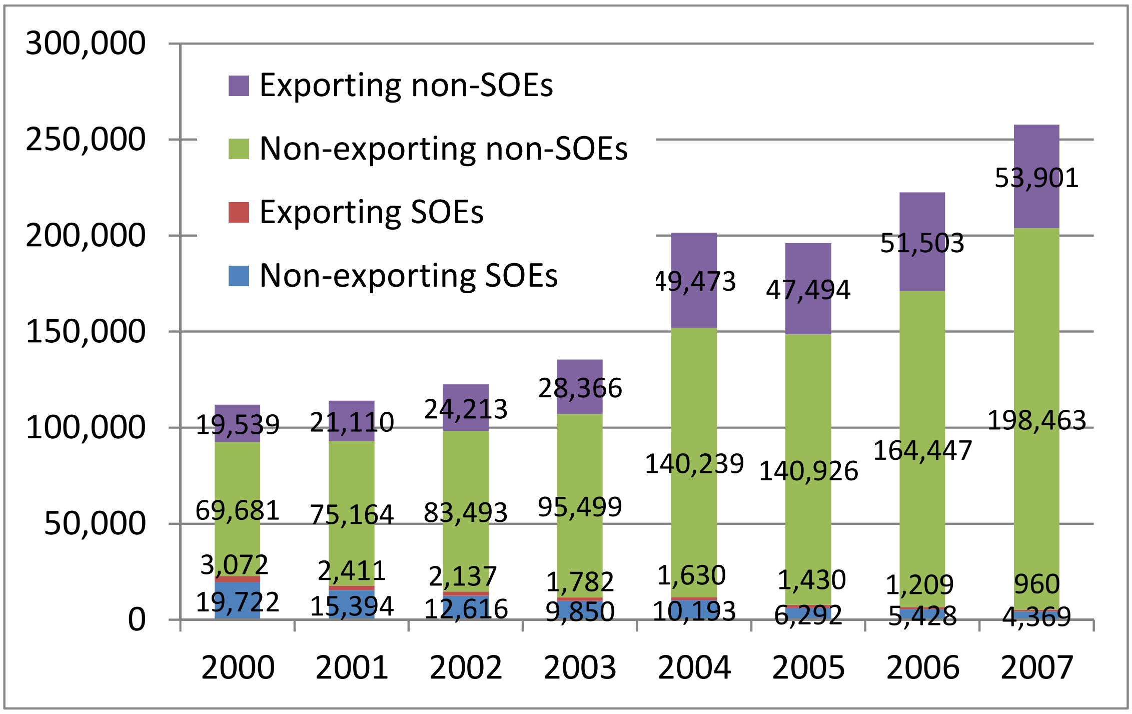

Table 1: Summary Statistics for 3SLS

Variable	Mean	Standard Deviation	Minimum	Maximum
Export dummy	0.266	0.442	0.000	1.000
Log of TFP	1.39	0.320	-0.598	2.87
Log of labor	4.856	1.084	0.000	11.907
Private ownership ratio	0.924	0.250	0.000	1.000
Liquidity ratio	0.066	0.271	-1.107	0.796
Leverage ratio	0.991	0.689	0.000	8.199
Long-term debt to asset ratio	0.046	0.104	0.000	0.805
Age	11.168	10.486	0.000	57.000
Log of total exports in the regions	17.905	1.776	6.218	20.559

Notes: The number of observations is 600,204. All variables except for the dummy for a switcher and the age is lagged one year.

Table 2: Number of Non-exporting SOEs by Subsequent Privatization and Export Status

Year t	SOEs in years t-1 and t				SOEs in years t-1 and privatized in t				Total
	Non-exporters		Exporters		Non-exporters		Exporters		
2001	3,644	(88.4)	56	(1.4)	405	(9.8)	18	(0.44)	4,123
2002	2,661	(90.2)	35	(1.2)	245	(8.3)	9	(0.30)	2,950
2003	2,223	(88.4)	29	(1.2)	260	(10.3)	3	(0.12)	2,515
2004	1,956	(81.1)	51	(2.1)	380	(15.7)	26	(1.08)	2,413
2005	1,555	(78.1)	76	(3.8)	331	(16.6)	28	(1.41)	1,990
Total	12,039	(86.1)	247	(1.8)	1,621	(11.6)	84	(0.60)	13,991

Notes: This table shows the number of non-exporting fully-state-owned enterprises in year t by privatization and export status in year $t + 1$. SOEs are defined as firms fully owned by the state. SOEs are defined to be privatized in year t if the private ownership ratio is more than a half in year $t + 1$, $t + 2$, and $t + 3$.

Table 3: Summary Statistics for non-exporting SOEs

Variable	Mean	Standard Deviation	Minimum	Maximum
Privatization dummy (*t*)	0.122	0.327	0.000	1.000
Export dummy (*t*)	0.0237	0.152	0.000	1.000
Export dummy (*t*+1)	0.0324	0.177	0.000	1.000
Export dummy (*t*+2)	0.0364	0.187	0.000	1.000
Log of TFP (*t*-1)	4.208	1.125	-1.687	9.520
Log of number of workers (*t*-1)	5.096	1.358	0.000	10.428
Long-term liability ratio (*t*-1)	0.098	0.143	0.000	0.811
Liquidity ratio (*t*-1)	-0.013	0.283	-1.112	0.794
Age	26.021	16.089	0.000	54.000
Log of regional exports (*t*-1)	16.041	1.775	6.218	20.015

Notes: The number of observations is 13,991.

Table 4: Determinants of Export

Dependent variable: export dummy

	(1) 3SLS	(2) 3SLS	(3) 3SLS
Export dummy (*t*-1)	0.895***	0.895***	0.895***
	(0.00115)	(0.00115)	(0.00115)
Log of TFP (*t*-1)	0.0151***	0.0154***	0.0157***
	(0.00262)	(0.00260)	(0.00253)
Log of labor (*t*-1)	0.0134***	0.0133***	0.0132***
	(0.000422)	(0.000418)	(0.000416)
Age	-0.000406***	-0.000406***	-0.000414***
	(3.51e-05)	(3.51e-05)	(3.53e-05)
Log of exports in the region (*t*-1)	0.00473***	0.00474***	0.00479***
	(0.000302)	(0.000302)	(0.000303)
Liquidity ratio (*t*-1)	0.000801		
	(0.00191)		
Leverage ratio (*t*-1)		0.000203	
		(0.000878)	
Ratio of long-term debt to asset (*t*-1)			0.0117**
			(0.00520)
Industry dummies	Yes	Yes	Yes
Region dummies	Yes	Yes	Yes
Year dummies	Yes	Yes	Yes
Number of observations	600,204	600,016	600,204
R-squared	0.652	0.652	0.652

Note: Standard errors are in parentheses. *, **, and *** denote statistical significance at the 10, 5, and 1 percent levels, respectively. Year, industry, and region dummies are included, but results are not presented. All dependent variables except for the age are instrumented by the dependent variables and the private ownership ratio in year *t* - 2.

Table 5: Probit Estimation

Dependent variable: privatization dummy

	(1) Before matching	(2) After matching
Log of TFP	0.474***	0.0173
	(0.161)	(0.273)
Log of TFP squared	-0.0771	-0.0159
	(0.0646)	(0.106)
Log of the number of workers	0.857***	0.0979
	(0.0777)	(0.124)
Log of the number of workers squared	-0.0734***	-0.00789
	(0.00724)	(0.0116)
Ratio of long-term debt to total asset	-0.119	-0.0419
	(0.108)	(0.162)
Liquidity ratio	-0.163***	0.0232
	(0.0551)	(0.0806)
Age	-0.0422***	-0.00625
	(0.00369)	(0.00532)
Age squared	0.000518***	0.000107
	(7.09e-05)	(0.000107)
Log of total exports in the region	0.0415***	-0.0208
	(0.0112)	(0.0170)
Industry dummies	Yes	Yes
Region dummies	Yes	Yes
Year dummies	Yes	Yes
Number of Observations	13,991	3,398
log likelihood	1218.95	22.3
Pseudo R squared	0.1175	0.0047

Note: Standard errors are in parentheses. *, **, and *** denote statistical significance at the 10, 5, and 1 percent levels, respectively. Year, industry, and region dummies are included in the probit estimation, but results are not presented. All covariates except for the year, industry, and region dummies are first lagged.

Table 6: Balancing Tests

	Sample before matching			Sample after caliper matching		
Covariate	Mean for treatment group	Mean for control group	*t* statistics	Mean for treatment group	Mean for control group	*t* statistics
Log of TFP	1.27	1.15	12.4***	1.27	1.27	0.121
Log of TFP squared	1.72	1.46	11.6***	1.72	1.72	0.0702
Log of labor	5.29	5.07	6.37***	5.29	5.27	0.412
Log of labor squared	29.3	27.6	4.73***	29.3	29.2	0.283
Ratio of long-term debt to asset	0.0884	0.0990	2.86***	0.0886	0.0895	0.187
Liquidity ratio	-0.0326	-0.0106	3.00***	-0.0317	-0.0360	0.447
Age	19.3	27.0	18.6***	19.4	19.6	0.445
Age squared	635	978	15.3***	638	643	0.192
Log of total exports in the region	16.3	16.0	6.83***	16.3	16.4	0.139
N	1,705	12,286		1,699	1,699	

Note: This table compares covariates in year $t - 1$ between the treatment groups, i.e., former SOEs privatized in year t, and the control group, i.e., remaining SOEs, using t tests. *, **, and *** denote statistical significance at the 10, 5, and 1 percent levels, respectively.

Table 7: Effect of Privatization

Outcome variable (Y)	Time difference	Mean for treatment group	Mean for control group	Difference	t statistics
Before matching					
Export dummy	$Y_t - Y_{t-1}$	0.0493	0.0201	0.0292	7.45***
	$Y_{t+1} - Y_{t-1}$	0.0587	0.0287	0.0300	6.56***
	$Y_{t+2} - Y_{t-1}$	0.0581	0.0333	0.0247	5.12***
Log of TFP	$Y_t - Y_{t-1}$	0.0331	0.0170	0.0164	1.92*
	$Y_{t+1} - Y_{t-1}$	0.0742	0.0548	0.0194	2.02**
	$Y_{t+2} - Y_{t-1}$	0.111	0.103	0.00886	0.864
Log of labor	$Y_t - Y_{t-1}$	-0.0274	-0.0582	0.0307	2.77***
	$Y_{t+1} - Y_{t-1}$	-0.0745	-0.118	0.0430	3.27***
	$Y_{t+2} - Y_{t-1}$	-0.0996	-0.192	0.0922	6.11***
Liquidity ratio	$Y_t - Y_{t-1}$	0.00410	-0.00201	0.00611	1.30
	$Y_{t+1} - Y_{t-1}$	0.00655	-0.00510	0.0117	2.08**
	$Y_{t+2} - Y_{t-1}$	0.0184	-0.00893	0.0273	4.26***
Leverage ratio	$Y_t - Y_{t-1}$	0.00171	0.00149	0.00022	0.0138
	$Y_{t+1} - Y_{t-1}$	-0.0176	0.0103	-0.0280	1.48
	$Y_{t+2} - Y_{t-1}$	-0.0508	0.0255	-0.0764	3.57***
Long-term liability ratio	$Y_t - Y_{t-1}$	-0.00575	-0.00429	-0.00147	0.621
	$Y_{t+1} - Y_{t-1}$	-0.0142	-0.00964	-0.00457	1.65*
	$Y_{t+2} - Y_{t-1}$	-0.0175	-0.0139	-0.00365	1.18
After matching					
Export dummy	$Y_t - Y_{t-1}$	0.0495	0.0318	0.0177	2.01**
	$Y_{t+1} - Y_{t-1}$	0.0589	0.0330	0.0259	3.06***
	$Y_{t+2} - Y_{t-1}$	0.0583	0.0400	0.0183	1.91*
Log of TFP	$Y_t - Y_{t-1}$	0.0353	-0.00580	0.0411	3.22***
	$Y_{t+1} - Y_{t-1}$	0.0764	0.0256	0.0508	3.39***
	$Y_{t+2} - Y_{t-1}$	0.114	0.0559	0.0578	3.68***
Log of labor	$Y_t - Y_{t-1}$	-0.0285	-0.0591	0.0306	1.45
	$Y_{t+1} - Y_{t-1}$	-0.0757	-0.113	0.0374	1.53
	$Y_{t+2} - Y_{t-1}$	-0.101	-0.164	0.0633	2.58***
Liquidity ratio	$Y_t - Y_{t-1}$	0.00325	0.00302	0.000224	0.0239
	$Y_{t+1} - Y_{t-1}$	0.00618	0.00334	0.00284	0.271
	$Y_{t+2} - Y_{t-1}$	0.0176	0.00244	0.0151	1.26
Leverage ratio	$Y_t - Y_{t-1}$	0.00365	-0.0140	0.0177	0.615
	$Y_{t+1} - Y_{t-1}$	-0.154	-0.0218	0.00644	0.190
	$Y_{t+2} - Y_{t-1}$	-0.0488	-0.0495	-0.0438	1.17
Long-term liability ratio	$Y_t - Y_{t-1}$	-0.00590	-0.00694	0.00103	0.267
	$Y_{t+1} - Y_{t-1}$	-0.0145	-0.0131	-0.00137	0.274
	$Y_{t+2} - Y_{t-1}$	-0.0179	-0.0127	-0.00521	0.930

Notes: This table shows the effect of privatization in year t on the change in the outcome variable from t - 1 to either t, t + 1, or t + 2. *, **, and *** denote statistical significance at the 10, 5, and 1 percent levels, respectively.