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Firm-level evidence from Chinese state-owned enterprises

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Abstract

This paper examines whether or not privatization of Chinese state-owned enterprises (SOEs) increases the probability of exporting, and, if so, what channels generate such effect. Using firm-level data for the Chinese manufacturing sector for the period 2000-2007, we find that privatization has a positive effect on exporting decisions, productivity, and firm size and a negative effect on firms' long-term debt. We also find that Chinese firms are more likely to engage in export when the productivity level, firm size, or long-term debt is larger. These two sets of results suggest that privatization has positive effects on exporting decisions through improving productivity and firm size and a negative effect through lowering debt. However, quantitative analysis reveals that effects of privatization through these three channels are small. Therefore, we conclude that the positive effect of privatization on exporting decisions comes mostly from other unobservable factors, most probably changes in attitude toward profits and risks associated with privatization.

1.   Introduction

This paper examines whether or not privatization of Chinese state-owned enterprises (SOEs) increases the probability of exporting, 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 et al. (2007), Du and Girma (2007), Muuls (2008), and Feenstra et al. (2011) test whether financial conditions of firms affect exporting decisions. A possible reason why financial conditions matter is that since starting to export 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. In particular, using firm-level data for China, Du and Girma (2007) show that Chinese firms are more likely to export when they are more productive and larger and receive more bank loans. They interpret this result as evidence showing that access to credits encourages firms to start exporting.

Second, as many SOEs in China were privatized in recent years, many studies look at outcomes of privatization. For example, Jefferson and Su (2006) and Bai et al. (2009) find that privatization of SOEs improves firm performance such as the productivity level and the firm size.

Third, privatization may influence credit constraints of firms. For example, Qian and Roland (1996) argue that privatization tightens credit constraints, since state ownership is often associated with soft budget constraints. According to Lin and Li (2008), however, when the government imposes policy burdens on firms, privatized SOEs demand more subsidies than SOEs to be profitable, and hence, privatization will only soften budget constraints. Since these two studies are mostly theoretical, whether privatization tightens or softens credit constraints of former SOEs in China remains an empirical question.

Combining the three strands of literature, it is unclear whether or not privatized SOEs are more likely to start exporting that remaining SOEs. Therefore, we investigate effects of SOEs' privatization on exporting decisions through these three channels, using a rich firm-level data set for the Chinese manufacturing sector for the period 2000-2007.

We employ propensity score matching estimation and find that that privatization leads to a larger probability of exporting, higher productivity, and larger size and to a smaller ratio of long-term debts to total assets. We also find that Chinese firms are more likely to engage in export, when the productivity level, firm size, or long-term debt is larger, as Du and Girma

(2007) found. These two sets of results suggest that privatization has positive effects on exporting decisions through improving productivity and firm size and a negative effect through lowering long-term debts. However, quantitative analysis reveals that effects of privatization through these three channels are in fact quantitatively very small. Therefore, we conclude that the positive effect of privatization on exporting decisions comes mostly from other unobservable factors such as changes in attitude toward profits and risks associated with privatization.

2.   Empirical Methodology

2.1   Empirical Framework

When we estimate effects of privatization of SOEs, we 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 exporting decisions, productivity, firm size, 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, the firm size, 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

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

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.

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

$$Y(1), Y(0) \coprod D | X, \tag{2}$$

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

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

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

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

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

The first term on the right-hand side of equation (4) is estimated by the average of actual outcomes of privatized SOEs. Each privatized SOE is matched with a remaining SOE that has a similar propensity score or the weighted average of remaining SOEs using their propensity scores to construct the weights. 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 et al. (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 et al. (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

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

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.

2.2   Practical Procedure

To see effects of privatization on initiating export more clearly, we focus on firms which are fully owned by the state and do not engage in exporting in year t-1. In other words, our treatment group consists of SOEs which are fully state-owned and not exporting in year t-1 and are privatized in year t, while our control group consists of SOEs which are fully state-owned and not exporting in year t -1 and remain state-owned in year t.

To obtain the DID-PSM estimator of the impact of privatization of non-exporting SOEs given the data set 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 et al. (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.

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 et al. (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 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 et al. (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 significance of coefficients and the pseudo-R^2 with those obtained from the probit estimation using the sample before matching. 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   Privatization in China

SOEs in China have been gradually privatized since the early 1990s, as the central government started to be concerned about large debts of SOEs. In 1995, the central government endorsed the "retain the large, release the small" (*Zhua Da Fang Xiao*) policy, and since then small and medium SOEs are more likely to be privatized, whereas larger SOEs remain stated-owned due to their economic and strategic importance. Examples of such large SOEs include China Faw Group Co., Ltd and Dongfeng Automobile Co., Ltd in the automobile industry, China Petrochemical Co., Ltd in the petrochemical industry, and State Grid Corporation of China in the power industry.

Privatization takes a variety of forms, including reorganization, mergers and takeovers, leasing and management contracts, and conversion to shareholding companies. Some SOEs are completely privatized, while others are done so only partially. In the latter case, some of the

partially privatized SOEs are still controlled by the government even after privatization.

Figure 1 illustrates the change in the number of Chinese by export and ownership status from 2000 to 2007, using the whole sample of 1,361,776 firm-year observations. 44 percent of all firms are SOEs, i.e., firms with a state-ownership ratio of 50 percent or more, in 2000, but their share has declined drastically over time to 6.1 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 32 percent in 2000 to 3.7 percent in 2007. In stark contrast, the number of private exporters increases from 15,000 in 2000 to 53,000 in 2007.

3.2 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 et al. (2011).[2]

We use total factor productivity (TFP) for our productivity measure.[3] 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 state ownership ratio of a firm is defined as the share of the sum of state capital in the total equity.[4] In the estimation, we focus on firms that are fully owned by the state (i.e., firms with a 100-percent state ownership ratio) and do not export in year $t - 1$. These firms are considered to be privatized in year t if the state ownership ratio is less 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.

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[2] These deflators are available at http://www.econ.kuleuven.be/public/n07057/China/.
[3] When labor productivity is used, the main results do not change.
[4] When we define the state ownership ratio as the share of the sum of state capital and collective capital, or the sum of state capital and legal capital, the main results do not change qualitatively and are similar quantitatively.

3.3   Descriptive Statistics

The sample to examine effects of privatization of non-exporting fully state-owned enterprises (hereafter, SOEs are defined as fully state-owned enterprises) 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 1. 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 2. 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   Effects of Privatization

Now, we examine whether privatized firms are more likely to export and if so, through which channels the effect of privatization arises, using the propensity score matching method described in Section 2.3. We use kernel matching for the benchmark estimation and caliper matching for a robustness check, finding that both matching methods generate quantitatively similar results. For simple presentation, we only show the results from kernel matching.

We first run a probit model to estimate how SOEs are chosen for privatization. According to the results shown in column (1) of Table 3, the TFP level has a positive effect, the number of workers have an inverted U-shaped effect, and 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, which is reasonably high for matching purposes.

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 check whether pre-privatization conditions are similar between the two groups. The results of the balancing tests shown in Table 4 indicate that although privatized SOEs and remaining SOEs are systematically different before matching, the two groups share very similar characteristics

after matching. Also, column (2) of Table 3 indicates that after matching, any covariate does not significantly affect privatization of SOEs. The results from these balancing tests indicate that the matching is successfully done.

The results for the effects of privatization are shown in Table 5. 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 2, 2.4 percent of non-exporting SOEs start exporting in the next year. According to the PSM estimation, privatization increases the probability of exporting in the year of privatization by 2.1 percentage points, exporting in the next year by 1.9, and exporting two years later by 1.7. Therefore, roughly speaking, privatization approximately doubles the probability of exporting for non-exporting SOEs.

The latter rows of Table 5 show whether effects of privatization on firm performance. First, as Jefferson and Su (2006) and Bai et al. (2009) find, the effect of privatization on productivity growth is positive, statistically significant, and quantitatively large. Privatization improves TFP by 3.3 percent in the same year and by 3.7 and 3.8 percent one and two years later, respectively. The effect of privatization on the firm size measured by the number of workers is also significantly positive and quantitatively large.

Second, privatization does not have a significant effect on the two short-term financial variables, i.e., the liquidity ratio and the leverage ratio. However, privatization lowers the ratio of long-term debts to total assets after two years. We can interpret this result on long-term debts in two different ways. On the one hand, lowering long-term debts may imply that privatized SOEs face tighter credit constraints and thus cannot borrow as much as they could do if they would remained state-owned. On the other hand, lowering long-term debts may imply that privatized SOEs improve financial conditions by purposefully reducing debts. Since the negative effect on long-term debts emerges two years after privatization, it is more likely that the latter is the case. In other words, if privatized SOEs face tighter credit constraints so that they have to reduce the amount of debts, the negative effect should emerge immediately after the privatization. Therefore, we conclude that privatized SOEs purposefully reduce their long-term debts to achieve healthier financial conditions.

4.2   Sources of Effects of Privatization on Export

The next question is through which channels the effect of privatization on exporting decisions arises. To answer this question, we first consider determinants of export. Du and Girma (2007) show that Chinese firms are more likely to export when they are more productive, larger, and

less credit-constrained (i.e., with more bank loans). To confirm and quantify these effects using our data, we employ three-stage least squares (3SLS) estimation in which instrumental variables are lagged regressors and the state ownership ratio so that we can correct for possible biases in estimation of determinants of exporting decisions due to endogeneity.

Table 6 presents results from the 3SLS estimations. As shown in Du and Girma (2007), the productivity level measured by TFP and the firm size measured by the number of workers have a positive and significant effect on the export decision. In each of columns (1)-(3) of Table 4, we use as a measure of financial conditions 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 debts to the total assets. Among these three measures, the ratio of long-term debts to the total assets has a positive and significant effect, while the two variables representing short-term financial conditions have no significant effect. This finding suggests that firms which can increase their long-term debts are more likely to engage in export, since exporting may require long-term investment in product modification or marketing abroad. All of these results are consistent with the findings of Du and Girma (2007).

Combining these results with the results in Section 4.1, we can identify channels of the effect of privatization on exporting decisions and evaluate them quantitatively. First, privatization improves the TFP level by 3.28 percent on average (Table 5). Since the coefficient on the log of TFP in column (3) of Table 6 is 0.0157, privatization raises the probability of export by 0.0515 percent (= 3.28 percent * 0.0157) through productivity improvement. Second, privatization raises the probability of exporting by 0.0373 percent (= 2.81 percent * 0.0133) through increasing the firm size. Since privatization increases the probability of exporting by about 2-percent, the effects of privatization on exporting decisions through productivity and size improvement are small in size. Third, privatized SOEs decrease long-term debts 2 years after privatization, and the decline in debts further negatively affects exporting decisions. However, the negative effect of privatization on the probability of exporting through lowering debts is -0.00713 percent (= -0.61 percent * 0.0117) and again, quantitatively negligible.

In summary, although we identified positive effects of privatization on exporting decisions through improvement in productivity and size and a negative effect through lowering long-term debts in the medium run, these effects are quantitatively small or negligible. In other words, privatized SOEs are more likely to engage in export than remaining SOEs, mostly due to unobservable factors which cannot be explained by productivity, firm size, or financial conditions.

A potential candidate of such unobservable factors is changes in firms' attitude toward profits and risks. When privatized, former SOEs are not protected by the government any more

so that they may have to expand their business to survive. One way to expand it is to export to foreign markets. Moreover, privatized SOEs become more risk-taking to be profitable than SOEs and may thus initiate exporting. Todo and Sato (2011) indeed find that more risk-taking firms are more likely to export and that the effect of risk preference on exporting decisions is far more important in magnitude than the effect of productivity and firm size, using firm-level data for Japanese small and medium enterprises. Since we do not have data on firms' attitude toward profits and risks, we cannot formally test our hypotheses above. However, it should be emphasized that these psychological factors may play an important role in effects of privatization.

4.3   Long-Run Effects of Privatization

We further examine long-run effects of privatization on exporting decisions, because some new exporters expand their exports as they continue to export, while others quit exporting within a few years after starting to export. Akhmetova (2010) and Akhmetova and Mitaritonna (2010) emphasize that firms often start with a small amount of exports and expand it over time due to learning by exporting. Eaton et al. (2007) observe that more than half of all exporters in Columbia exit from export markets in the next year. Similar evidence can be seen in the case of China. Figure 3 shows the change from 2001 to 2007 in the average export intensity, or the ratio of exports to total sales, of firms which started exporting in 2001[5] and continued to export in 2007. The average export intensity was 32 percent in the first year of exporting, 2001, and increased to 38 percent in 2007. Table 7 shows top five patterns of export dynamics of firms which started exporting in 2001. This table indicates that 30 percent of new exporters in 2001 quitted exporting in the next year, 2002. The fraction of exitters decreases over time from 2001 to 2004, as Akhmetova (2010) and Akhmetova and Mitaritonna (2010) suggest. More than half of the new exporters in 2001 quitted exporting within four years. The high exit rate is also consistent with the finding of Eaton et al. (2007) for Columbia.

Given these findings, we are concerned about how long the effect of privatization on exporting decisions lasts. For example, suppose that effects from learning by exporting are small, and that privatized SOEs tend to become too aggressive towards exporting and go beyond their competitiveness in foreign markets. Then, privatized SOEs may start exporting, but many of the new exporters quit exporting within a few years since they find exporting unprofitable to them. If this is the case, privatized SOEs are more likely to engage in export in the short run after privatization but not in the long run.

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[5] This means that firms which did not export in 2000 but exported in 2001.

To test this hypothesis, we estimate effects of privatization in year t on exporting decisions and firm performance in year t to $t + 5$. We employ the same method as we did in Section 4.1, except for one thing. Although our sample in Section 4.1 consists of firms which existed continuously from year $t – 1$ to $t + 2$, here we include firms which existed from year $t – 1$ to $t + 5$, as well as those which existed in years $t – 1$ and t but disappeared during the period from $t + 1$ to $t + 5$. This is because when we examine long-run effects, attrition bias may be substantial. When a firm disappears from the dataset, we assume that the firm is not engaging in export. One disadvantage of this method is that since we cannot observe productivity, employment, or financial conditions of firms which are no longer in the data set, we should limit our attention to the effect of privatization on export decisions. The sample consists of 7,085 and is much smaller than the sample in Section 4.1 where 13,991 firms which existed continuously from year $t – 1$ to $t + 2$ are included.

Figure 4 summarizes the long-run effects of privatization graphically. The horizontal axis indicates years after privatization, and the vertical axis indicates the effect of privatization on the probability of export. The dot and the line segment for each year respectively represent the point estimate and the confidence interval at the 95-percent level of the effect of privatization obtained from PSM estimations as in Section 4.1. Therefore, if the lower bound of a line segment is above zero, the effect is statistically significant at the 5-percent level.

The effect of privatization on the export dummy for any year is positive and significant at the 5-percent significance level. The point estimate of the privatization effect is stable over time at around 2 percent. This finding suggests that privatization has a positive effect on exporting decisions even in the long run and implies that exporters can remain in the export market probably because they improve productivity through learning by exporting.

5.   Concluding Remarks

This paper examines whether or not privatization of Chinese state-owned enterprises (SOEs) increases the probability of exporting and if so, what channels generate such effect. Using firm-level data for the Chinese manufacturing sector for the period 2000-2007, we find that privatization has a positive effect on exporting decisions, productivity, and firm size and a negative effect on firms' long-term debt. We also find that Chinese firms are more likely to engage in export, when the productivity level, firm size, or long-term debt is larger. These two sets of results suggest that privatization has positive effects on exporting decisions through improving productivity and firm size and has a negative effect through lowering debts. However, quantitative analysis reveals that effects of privatization through these three channels are quantitatively small. Therefore, we conclude that the positive effect of privatization on

exporting decisions comes mostly from other unobservable factors, most probably changes in attitude toward profits and risks associated with privatization. However, because of data limitation, we cannot formally test this possible channel of the effect of privatization of exporting decisions. We let future studies tackle this issue.

Figure 1: Number of Firms by Export Status and by State Ownership from 2000 to 2007
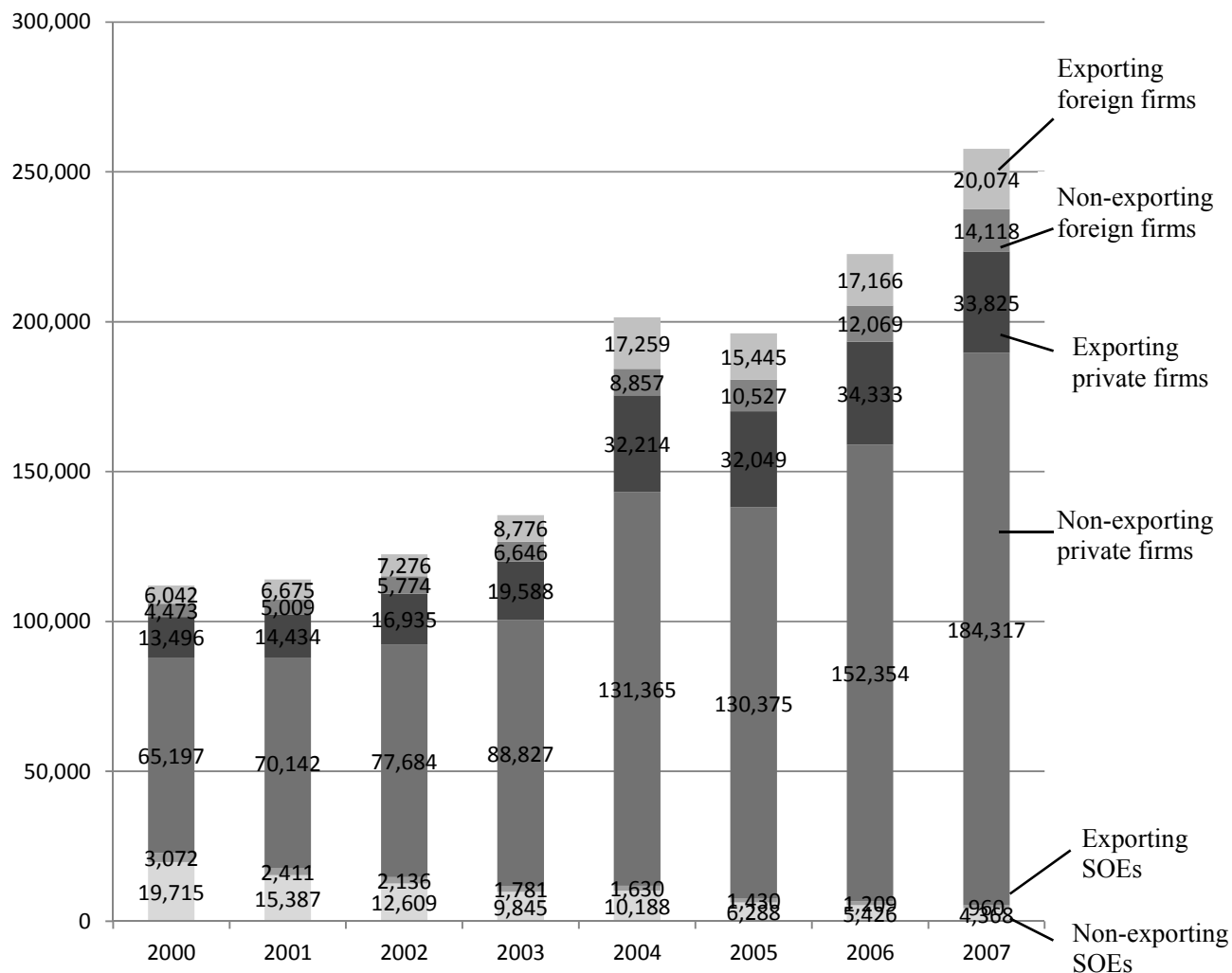
Note: SOEs in this figure are defined as firms with a state ownership (share of state capital) of 50 percent or more.

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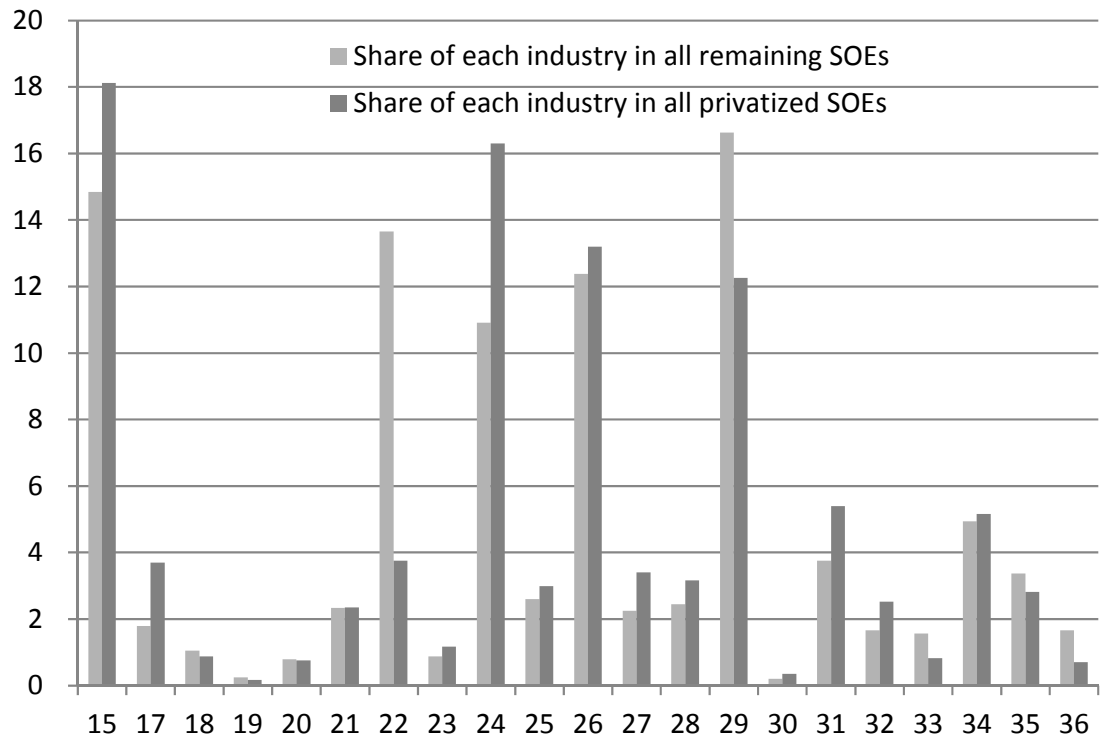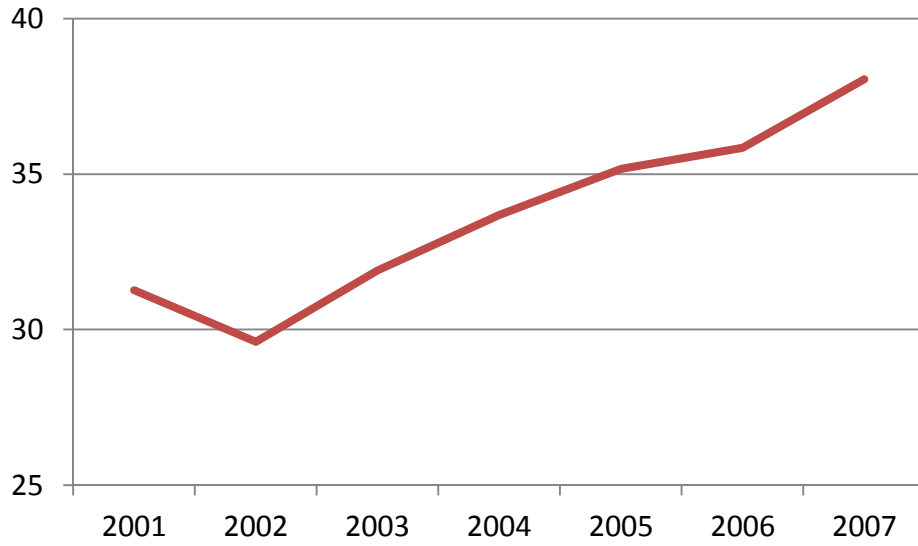Figure 2: Differences across Industries

Figure 3: Average Percentage of Exports to Total Sales for Firms
Which Started Exporting in 2001 and Continued to Export in 2007

Note: This figure is based on the whole sample, including private firms.

Figure 4: Long-Run Effects of Privatization

Years after privatization
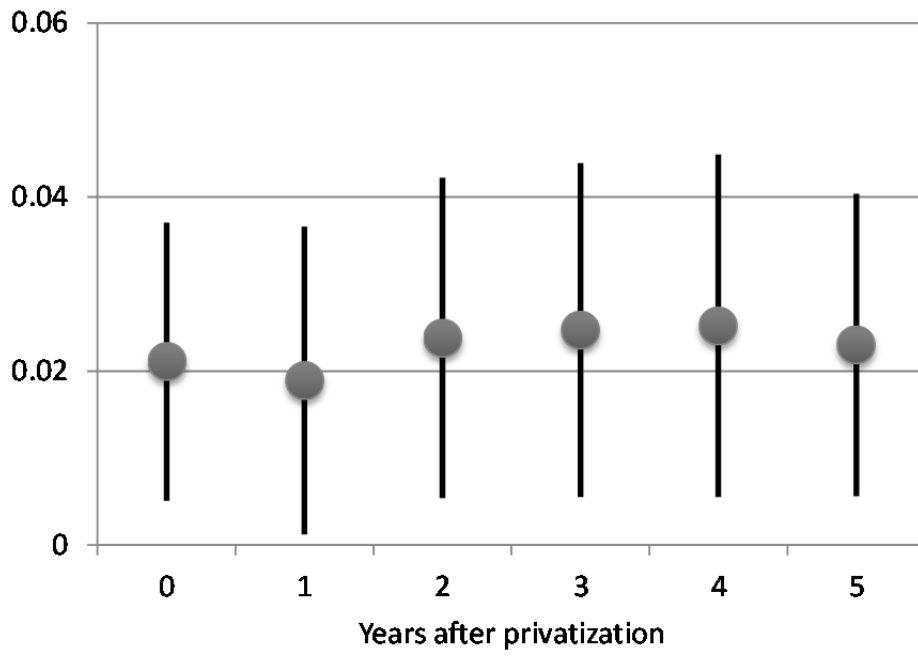
Table 1: Number of Non-exporting SOEs by Subsequent Privatization and Export Status

Year t	SOEs in years t-1 and t				SOEs in years t-1 and privatized in t				Total
	Non-exporters in t		Exporters in t		Non-exporters in t		Exporters in t		
2001	3,644	(88.4)	56	(1.4)	405	(9.8)	18	(0.4)	4,123
2002	2,661	(90.2)	35	(1.2)	245	(8.3)	9	(0.3)	2,950
2003	2,223	(88.4)	29	(1.2)	260	(10.3)	3	(0.1)	2,515
2004	1,956	(81.1)	51	(2.1)	380	(15.7)	26	(1.1)	2,413
2005	1,555	(78.1)	76	(3.8)	331	(16.6)	28	(1.4)	1,990
					Total number of observations				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 2: Summary Statistics for non-exporting SOEs

Variable	Mean	Standard Deviation	Minimum	Maximum
Privatization dummy (*t*)	0.122	0.327	0	1
Export dummy (*t*)	0.0237	0.152	0	1
Export dummy (*t*+1)	0.0324	0.177	0	1
Export dummy (*t*+2)	0.0364	0.187	0	1
Log of TFP (*t*-1)	1.163	0.374	-0.589	2.61
Log of number of workers (*t*-1)	5.096	1.358	0.000	10.428
Liquidity ratio (*t*-1)	-0.013	0.283	-1.112	0.794
Leverage ratio (*t*-1)	1.19	0.826	0	8.00
Long-term liability ratio (*t*-1)	0.098	0.143	0	0.811
Age	26.0	16.1	0	54
Log of regional exports (*t*-1)	16.0	1.78	6.22	20.0

Notes: The sample consists of firms which are fully state-owned and are not exporting in year $t - 1$. The number of observations is 13,991

Table 3: Probit Estimation

Dependent variable: privatization dummy

	(1) Before matching	(2) After kernel matching
Log of TFP	0.474***	0.221
	(0.161)	(0.332)
Log of TFP squared	-0.0771	-0.0606
	(0.0646)	(0.129)
Log of the number of workers	0.857***	0.231
	(0.0777)	(0.155)
Log of the number of workers squared	-0.0734***	-0.0219
	(0.00724)	(0.0144)
Ratio of long-term debt to total asset	-0.119	-0.0359
	(0.108)	(0.221)
Liquidity ratio	-0.163***	0.0468
	(0.0551)	(0.110)
Age	-0.0422***	-0.00346
	(0.00369)	(0.00688)
Age squared	0.000518***	0.0000366
	(7.09e-05)	(0.000143)
Log of total exports in the region	0.0415***	0.00365
	(0.0112)	(0.0177)
Industry dummies	Yes	No
Region dummies	Yes	No
Year dummies	Yes	Yes
Number of Observations	13,991	3,398
log likelihood	-4575.9	-2344.2
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 4: Balancing Tests

Covariate	Sample before matching			Sample after kernel matching		
	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	0.0989	2.85***	0.0886	0.0895	0.182
Liquidity ratio	-0.0326	-0.0107	2.99***	-0.0317	-0.0360	0.447
Age	19.3	27.0	18.6***	19.4	19.6	0.445
Age squared	635	977	15.3***	638	643	0.192
Log of total exports in the region	16.3	16.0	6.86***	16.3	16.4	1.39
N	1,705	12,296		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 5: PSM Estimation of Effects of Privatization

Outcome variable (Y)	Time difference	Mean for treatment group	Mean for matched control group	Difference	t statistics
Export dummy	$Y_t - Y_{t-1}$	0.0495	0.0287	0.0208	3.3798***
	$Y_{t+1} - Y_{t-1}$	0.0589	0.0398	0.0191	2.9283***
	$Y_{t+2} - Y_{t-1}$	0.0583	0.0418	0.0165	2.4329**
Log of TFP	$Y_t - Y_{t-1}$	0.0353	0.0025	0.0328	4.9459***
	$Y_{t+1} - Y_{t-1}$	0.0764	0.0390	0.0374	5.2844***
	$Y_{t+2} - Y_{t-1}$	0.1136	0.0759	0.0378	5.0841***
Log of labor	$Y_t - Y_{t-1}$	-0.0285	-0.0566	0.0281	2.1785**
	$Y_{t+1} - Y_{t-1}$	-0.0757	-0.1062	0.0306	1.9810**
	$Y_{t+2} - Y_{t-1}$	-0.1007	-0.1708	0.0700	4.0117***
Liquidity ratio	$Y_t - Y_{t-1}$	0.0032	0.0037	-0.0005	0.0693
	$Y_{t+1} - Y_{t-1}$	0.0062	0.0048	0.0013	0.1836
	$Y_{t+2} - Y_{t-1}$	0.0176	0.0071	0.0105	1.3781
Leverage ratio	$Y_t - Y_{t-1}$	0.0037	-0.0190	0.0226	1.0187
	$Y_{t+1} - Y_{t-1}$	-0.0154	-0.0204	0.0050	0.1950
	$Y_{t+2} - Y_{t-1}$	-0.0488	-0.0211	-0.0277	1.0206
Long-term liability ratio	$Y_t - Y_{t-1}$	-0.0059	-0.0045	-0.0014	0.4805
	$Y_{t+1} - Y_{t-1}$	-0.0145	-0.0085	-0.0059	1.8089
	$Y_{t+2} - Y_{t-1}$	-0.0179	-0.0118	-0.0061	2.2056**

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.

Table 6: 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 7: Top 5 Patterns of Export Dynamics of 2,358 Firms Which Started Exporting in 2001

Pattern of export (√ = export)							Number of firms	%
2001	2002	2003	2004	2005	2006	2007		
√							722	30.6
√	√	√	√	√	√	√	457	19.4
√	√						266	11.3
√	√	√					220	9.33
√	√	√	√				75	3.18

Note: Firms which started exporting in 2001 are defined as firms which did not export in 2000 but exported in 2001.

References

Akhmetova Z., 2010. Firm experimentation in new markets. mimeo, available at https://sites.google.com/site/zhanarakh/home/research.

Akhmetova Z., Mitaritonna C. 2010. Structural estimation of the model of firm experimentation. mimeo, available at https://sites.google.com/site/zhanarakh/home/research.

Arnold J.M., Javorcik B.S., 2005. Gifted kids or pushy parents? Foreign acquisitions and plant performance in Indonesia. World Bank Policy Research Working Paper, No. 3139.

Bai C.-E., Lu J., Tao Z., 2009. How does privatization work in China? Journal of Comparative Economics. 37, 453-70.

Bernard A.B., Jensen J.B., 1999. Exceptional exporter performance: Cause, effect, or both? Journal of International Economics. 47, 1-25.

Bernard A.B., Jensen J.B., 2004. Why some firms export. Review of Economics and Statistics. 86, 561-69.

Brandt L., Van Biesebroeck J., Zhang Y., 2011. Creative accounting or creative destruction? Firm-level productivity growth in Chinese manufacturing. Journal of Development Economics. 97, 339-351.

Clerides S.K., Lach S., Tybout J.R., 1998. Is learning by exporting important? Micro-dynamic evidence from Colombia, Mexico, and Morocco. Quarterly Journal of Economics. 113, 903-47.

Du J., Girma S., 2007. Finance and firm export in China. Kyklos. 60, 37-54.

Eaton J., Eslave M., Kugler M., Tybout J.R., 2007. Export dynamics in Columbia: Firm-level evidence. NBER Working Paper, No. W13531.

Feenstra R.C., Li Z., Yu M., 2011. Exports and credit constraints under incomplete information: Theory and evidence from China. NBER Working Paper, No. 16940.

Girma S., Görg H., 2007. Evaluating the foreign ownership wage premium using a difference-in-difference matching approach. Journal of International Economics. 72, 97-112.

Greenaway D., Guariglia A., Kneller R., 2007. Financial factors and exporting decisions. Journal of International Economics. 73, 377-95.

Greenaway D., Kneller R., 2004. Exporting and productivity in the United Kingdom. Oxford Review of Economic Policy. 20, 358-71.

Heckman J.J., Ichimura H., Todd P., 1997. Matching as an econometric evaluation estimator: Evidence from evaluating a job training programme. Review of Economic Studies. 64, 605-54.

Heckman J.J., Ichimura H., Todd P., 1998. Matching as an econometric evaluation estimator. Review of Economic Studies. 65, 261-94.

Hirano K., Imbens G.W., Ridder G., 2003. Efficient estimation of average treatment effects using the estimated propensity score. Econometrica. 71, 1161-89.

Jefferson G.H., Su J., 2006. Privatization and restructuring in china: Evidence from

shareholding ownership, 1995-2001. Journal of Comparative Economics. 34, 146-66.

Lin J.Y., Li Z., 2008. Policy burden, privatization and soft budget constraint. Journal of Comparative Economics. 36, 90-102.

Muuls M., 2008. Exporters and credit constraints: A firm-level approach. National Bank of Belgium Working Paper, No. 139.

Olley S.G., Pakes A., 1996. The dynamics of productivity in the telecommunications equipment industry. Econometrica. 64, 1263-97.

Qian Y., Roland G., 1996. The soft budget constraint in China. Japan and the World Economy. 8, 207-23.

Rosenbaum P.R., Rubin D.B., 1983. The central role of the propensity score in observational studies for causal effects. Biometrika. 70, 41-55.

Rubin D.B., 1980. Bias reduction using mahalanobis-metric matching. Biometrika. 36, 293-98.

Sianesi B., 2004. An evaluation of the active labour market programmes in Sweden. Review of Economics and Statistics. 86, 133-55.

Smith J.A., Todd P., 2005. Does matching overcome Lalonde's critique of nonexperimental estimators? Journal of Econometrics. 125, 305-53.

Todo Y., Sato H., 2011. Effects of CEOs' characteristics on internationalization of small and medium enterprises in Japan. RIETI Discussion Paper, No. 11-E-026.