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# Effects of Privatization on Exporting Decisions: Firm-level Evidence from Chinese State-owned Enterprises

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This paper examines whether privatizing Chinese state-owned enterprises increases the probability of exporting and, if so, what factors generate such an effect. Using firm-level data for the Chinese manufacturing sector for the 2000–2007 period, we find that privatization positively affects a firm's productivity, size, and decision to export, whereas we find that it negatively affects the level of a firm's long-term debt. We also find that Chinese firms are more likely to export when the productivity level, firm size, or the level of long-term debt increases. Taken together, these two sets of results suggest that privatization positively affects the likelihood that a firm will export by improving productivity and increasing firm size, whereas it negatively affects such a likelihood by lowering the long-term debt level of the firm. However, a quantitative analysis reveals that the effects of privatization that occur through these three channels are only slight. Therefore, we conclude that the positive effect of privatization on the likelihood of exporting is mainly the result of unobservable factors that are most likely related to changes in attitude about the profits and risks associated with privatization.
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## INTRODUCTION

As state-owned enterprises (SOEs) in transition economies have been privatized since the 1990s, effects of privatization have been examined in depth. It is often claimed that SOEs are slower to adopt strategies to compete in a global market than are private firms because managers of SOEs have distorted incentives and are less exposed to competitive pressure (Yusuf *et al.*, 2006). Therefore, privatization of SOEs is expected to improve their performance if they are to survive in competitive markets. Empirical studies using firm-level data have in fact found that the effect of privatization of SOEs in Central and Eastern Europe on productivity is positive and statistically significant while the effect of privatization in the former Soviet Union is less clear, as surveyed in Estrin *et al.* (2009). Privatizing SOEs in China also improves productivity and increases firm size, according to Jefferson and Su (2006) and Bai *et al.* (2009).

Given a higher productivity level and higher incentive to grow, privatized SOEs may be more able and willing to enter foreign markets than the remaining SOEs. Melitz (2003) predicts that only firms with a sufficiently high productivity level and a sufficiently large size can serve foreign markets through exporting, because exporting requires initial costs of entry. Empirical studies using firm-level data mostly support this theoretical prediction: higher productivity and larger size lead to firms' higher propensity to export (see, eg, Clerides *et al.*, 1998; Bernard and Jensen, 1999, 2004; Greenaway and Kneller, 2004).

However, it is not clear whether privatization of SOEs indeed encourages exports of privatized SOEs, because privatization influences firms' credit constraints, which, in turn, affect export decisions. Du and Girma (2007) used firm-level data from China to show that Chinese firms are more likely to export when they have greater levels of long-term debt through bank loans.[1] They interpret this evidence as showing that access to credits encourages firms to export because credit enables firms to cover the initial costs of export. On the other hand, whether privatization tightens or softens the credit constraints of former SOEs is unclear. Qian and Roland (1996) posit that privatization tightens credit constraints because state ownership is frequently associated with soft budget constraints. However, Lin and Li (2008) argue that privatized SOEs require greater subsidies than SOEs when the government imposes policy burdens on such firms and that privatization therefore softens budget constraints. Lizal and Svejnar (2002) indeed find that both SOEs and private

[1] Greenaway *et al.* (2007), Du and Girma (2007), Muuls (2008), and Feenstra *et al.* (2011) also examined whether the financial conditions of firms affect the decision to export using data for non-transition economies.

firms in the Czech Republic are credit constrained. Accordingly, whether privatization of SOEs encourages export decisions by alleviating credit constraints is unclear.

Therefore, the net effect of privatization of SOEs on export decisions remains an empirical question and is of great interest. However, to the authors' best knowledge, no study has empirically examined this issue using firm-level data for transition economies, although there are several related studies. For example, Filatotchev *et al.* (2001) investigated how different types of governance and management strategies after privatization in the former Soviet Union affect export intensity. Shinkle and Kriauciunas (2010) found that the effect of firm size and age on export intensity in Central and Eastern European firms varies depending on economic institutions in the transition economies. However, these papers did not look into the effect of privatization itself. Therefore, this paper investigates whether privatized SOEs are more likely to start exporting than SOEs, using a rich firm-level data set for the Chinese manufacturing sector in the 2000–2007 period. Moreover, we investigate the three channels discussed above, productivity, firm size, and long-term debt levels, to determine the effects of privatization on the decision to export.

Employing propensity score matching (PSM) estimation to deal with endogeneity of privatization, we find that privatization leads to a larger probability of exporting, higher productivity, and larger size. Conversely, it also leads to a smaller ratio of long-term debt to total assets. Additionally, we find that Chinese firms are more likely to engage in exporting when productivity levels, firm size, or long-term debt are larger, as Du and Girma (2007) have found.

These results suggest that privatization has positive effects on the decision to export by improving productivity and increasing firm size, whereas it has a negative effect on the decision to export by lowering long-term debt. However, a quantitative analysis reveals that the effects of privatization through these three channels are only slight. Therefore, we conclude that the positive effect of privatization on the decision to export is primarily the result of other unobservable factors, such as changes in attitudes about the profits and risks associated with privatization. In other words, exposed to larger competitive pressure of markets, privatized SOEs are more willing to serve foreign markets through exporting. This is in line with the findings of Filatotchev *et al.* (2001) for the former Soviet Union and Shinkle and Kriauciunas (2010) for Central and Eastern Europe that governance and management strategies and institutions in the economy after privatization largely affect decisions to export.

The contribution of this paper is twofold. First, as we mentioned above, this is the first attempt to examine the effects of privatization of SOEs in transition economies on exporting decisions using firm-level data. Second, our results provide important policy implications for contemporary China. Our results suggest that, by privatizing SOEs, China can penetrate the global market and grow more rapidly. Elliott and Zhou (2013) show that, although exporting SOEs tend to be more productive than other exporters, non-exporting SOEs tend to be less productive than other non-exporters. Therefore, privatizing non-exporting SOEs in particular can contribute to higher growth of the Chinese economy, through increased productivity and propensity to serve foreign markets.

## EMPIRICAL METHODOLOGY

### Empirical framework

When we estimate the effects of privatization on SOEs, we encounter endogeneity problems because SOEs are not randomly chosen for privatization. To correct for biases related to endogeneity, we employ the PSM technique developed by Rosenbaum and Rubin (1983).[2]

In the PSM estimation, we identify the average effect of treatment on the treated (ATT), that is, the average effect of privatization on the decision to export, productivity, firm size, and financial conditions. Let $D_{it}$ be a dummy variable that represents 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 $D_{it}$. Thus, ATT can be defined as:

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

where $X_{i,t-1}$ denotes the characteristics of firm $i$ in year $t-1$. Thus, ATT is the average difference between the outcomes of privatized SOEs and their counterfactual outcomes had they not been privatized.

To identify ATT, we must assume 'strong ignorability', that is, unconfoundedness and overlap (Rosenbaum and Rubin, 1983). The unconfoundedness assumption is given by:

$$Y(1), Y(0) \amalg D \mid X \tag{2}$$

which implies that, given a set of observable characteristics $X$, the potential outcomes are independent of the treatment (privatization) assignment. The overlap assumption is given by:

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

which ensures a positive probability of privatization and non-privatization.

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

Under the strong ignorability assumption, Rosenbaum and Rubin (1983) show that the potential outcomes are also independent of the treatment, which is conditional on the probability that the firm is privatized (or the propensity score $P(X)$). Thus, the ATT term in equation 1 becomes:

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

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 either matched with a remaining SOE that has a similar propensity score or with the weighted average of the remaining SOEs (using their propensity scores to construct the weights). The second term, the expected outcome of privatized SOEs had they not been privatized, can then be estimated by the average outcome of the matched remaining SOEs.

When panel data are available, as in the case of this paper, a difference-in-difference (DID) PSM estimator of the ATT can be employed, as proposed by Heckman *et al.* (1997, 1998), in which we examine the effect on the change in the outcome measure. An advantage of using 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 that do not use DID. Formally, the DID − PSM estimator is defined as:

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

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

**Practical procedure**

To present the effects of privatization on initiating exports more clearly, we focus on firms that are fully owned by the state and did not export in year $t-1$. Thus, our treatment group consists of SOEs that are fully state-owned, did not export in year $t-1$, and were privatized in year $t$, whereas our control group consists of SOEs that are fully state-owned, did not export in year $t-1$, and remained state-owned in year $t$.

To obtain the DID − PSM estimator showing the impact of the privatization of non-exporting SOEs in our data set, 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) and include the following: the log of total factor productivity (TFP); the log of the number of workers; the liquidity ratio (defined as firms' current assets minus their 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 square term of the log of the number of workers, and the square term of firms' 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. We impose a common support condition in both methods to satisfy the overlap assumption (equation 3); we drop observations in the treatment group whose propensity score is either higher than the maximum score or lower than the minimum score from the control group. In the case of caliper matching, each observation in the treatment group is matched with the control observation that has the closest propensity score to the treated observation's score within the maximum score distance, that is, the caliper. In our 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. For the weight function $W$ in equation 5, we use the Epanechnikov kernel function and set the bandwidth to 0.06.

Following Arnold and Javorcik (2009), we match the treatment observations with control observations from the same year. In the case of evaluating the effects of job training programs, Heckman *et al.* (1997) found that matching estimates perform well when both participants and non-participants reside in the same local labor market. Therefore, they argue that geographic mismatches should be avoided in the matching estimation. In this paper, temporal mismatches may be more substantial than geographic mismatches because the data of this paper cover an 8-year period (as explained in detail later), and SOEs were privatized throughout this period. Therefore, a time restriction is imposed in this study.

After the matching, the treatment and control groups should have similar characteristics before privatization. To check whether this is the case, we employ two types of balancing tests. First, a simple $t$-test is used to examine whether the mean of each covariate differs between the treatment and control groups after matching. In addition, following Girma and Görg (2007), Hotelling's $T$-squared test is performed to jointly test the equality of the means between the two groups for all covariates. Second, we run a probit model using the post-matching sample and compare the significance of the coefficients and pseudo-$R^2$ with those obtained from the probit estimation using the pre-matching sample. These tests were proposed by Sianesi (2004). If matching is successful, then the post-matching probit should have no explanatory power such that the pseudo-$R^2$ should be low and the estimated coefficients should to be close to zero.

Because the treatment and control groups pass the balancing tests, we compute the DID − PSM estimator using equation 5. To take advantage of the panel data for this paper, the length of years between treatment and the impact evaluation (*s* in equation 5) is set at 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 because the multiple steps in PSM estimation, such as the estimation of propensity scores and the use of matching procedures, lead to more variation in the PSM estimators than in standard estimators with only one step.

## DATA

### Privatization in China

SOEs have gradually been privatized since the early 1990s in China, as the central government grew concerned about their considerable debt. In 1995, the central government endorsed the policy of ‘retain the large, release the small’ (*Zhua Da Fang Xiao*). Since then, small- and medium-sized SOEs are more likely to have been privatized than larger SOEs because of the economic and strategic importance of the large-sized SOEs. 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 can occur in a variety of ways, including through reorganizations, mergers and takeovers, leasing and management contracts, and conversion to shareholding companies. Certain SOEs are completely privatized, whereas others are only partially privatized. Certain partially privatized SOEs remain under government control after privatization.

Figure 1 illustrates the changes in the number of Chinese firms by export and ownership status from 2000 to 2007 over the entire sample of 1,361,776 firm-year observations. In 2000, 44% of all firms were SOEs, that is, firms with a state-ownership ratio of 50% or more, but that share declined drastically to 6.1% in 2007. Thus, our sample period witnessed drastic privatization. Approximately 20% of firms were exporters, and the percentage of exporters that were SOEs declined from 13.6% in 2000 to 1.7% in 2007. In stark contrast, the number of private exporters increased from 19,500 in 2000 to 53,900 in 2007.

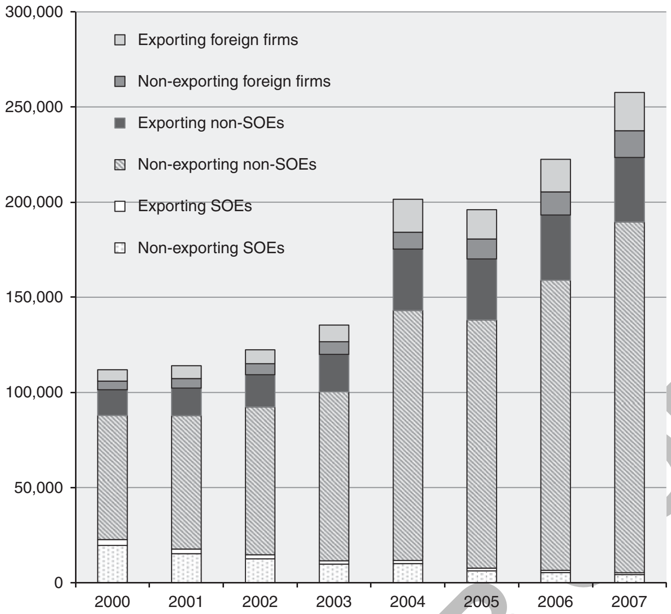


**Figure 1:** Number of firms by export status and by state ownership from 2000 to 2007
*Note*: SOEs in this figure are defined as firms with a state ownership (share of state capital) of 50% or more.



## Description of the data set

The data utilized in this paper are from 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; responding to the survey is compulsory.

Our sample covers the 2000–2007 period. We dropped firms from the sample for which the reported sales, exports, or book value of fixed assets were negative; we also dropped firms with exports that exceeded their sales in any year. We constructed the real values of outputs, inputs, and capital stocks using the industry-level deflators constructed by Brandt *et al.* (2011).[3]

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

[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.

**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 $t+3$.

The state-ownership ratio of a firm is defined as the share of the sum of state capital in the total equity of an enterprise.[5] In this estimation, we focus on firms that were fully owned by the state (ie, firms with a 100% state-ownership ratio) that did not export in year $t-1$. These firms are considered privatized in year $t$ if the state-ownership ratio was less than 50% in years $t$, $t+1$, and $t+2$.

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

## Descriptive statistics

The sample we use to examine the effects of privatizing non-exporting fully state-owned enterprises (hereafter, SOEs are defined as fully state-owned enterprises) consists of 13,991 firm-year observations. Because we define privatization in year $t$ as being privatized in year $t$ and continuing to be privatized up to year $t+2$ and use the variables in year $t-1$ as independent variables, we focus on privatization from 2001 to 2005, although the entire data set covers the period 2000–2007. Table 1 presents the number of firms by privatization, export status, and year. This table shows that 9%–18% of incumbent SOEs were privatized each year. Among the privatized SOEs, some firms began exporting immediately after privatization. The number of

[5] When we define the state-ownership ratio as the share of the sum of state capital and collective capital or as the sum of state capital and legal capital, the main results do not change qualitatively and are similar quantitatively.

**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 ($t$) | 0.024 | 0.152 | 0 | 1 |
| Export dummy ($t+1$) | 0.032 | 0.177 | 0 | 1 |
| Export dummy ($t+2$) | 0.036 | 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 ($t-1$) | 1.190 | 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.

observations declines over time because the number of incumbent SOEs declined as SOEs were privatized.

Summary statistics for the key variables in the sample are shown in Table 2. The second, third, and fourth rows indicate that 2.4% of non-exporting SOEs in year $t-1$ were exporters in year $t$, 3.2% were exporters in year $t+1$, and 3.6% were exporters in year $t+2$. Note that some of the exporters remained SOEs whereas others were privatized.

## RESULTS

### Effects of privatization

We now use the PSM method described in the Empirical Methodology section to examine whether privatized firms are more likely to export and, if so, through which channels the effects of privatization arise. We use kernel matching for the benchmark estimation and caliper matching for a robustness check, and we find that both matching methods generate quantitatively similar results. To simplify the 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 has an inverted U-shaped effect, and firm age has a U-shaped effect. The liquidity ratio has a negative effect, which implies that firms with larger net current assets are less likely to be privatized. The pseudo-$R^2$ 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 the remaining non-exporting SOEs within

**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 | −0.077 | −0.061 |
| | (0.065) | (0.129) |
| Log of the number of workers | 0.857*** | 0.231 |
| | (0.078) | (0.155) |
| Log of the number of workers squared | −0.073*** | −0.022 |
| | (0.007) | (0.014) |
| Ratio of long-term debt to total asset | −0.119 | −0.036 |
| | (0.108) | (0.221) |
| Liquidity ratio | −0.163*** | 0.047 |
| | (0.055) | (0.110) |
| Age | −0.042*** | −0.003 |
| | (0.004) | (0.00688) |
| Age squared | (5.18e−04)*** | 3.66e−05 |
| | (7.09e−05) | (1.43e−04) |
| Log of total exports in the region | 0.042*** | 0.004 |
| | (0.011) | (0.018) |
| 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-$R^2$ | 0.118 | 0.005 |

*Notes*: Standard errors are in parentheses. 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.
*, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.

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 the remaining SOEs are systematically different before matching, the two groups share similar characteristics after matching. Column (2) of Table 3 also indicates that no covariate significantly affects the privatization of SOEs after matching. The results from these balancing tests indicate that the matching was successful.

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 begin exporting after privatization, and the effect is quantitatively large. As shown in Table 2, 2.4% of non-exporting SOEs begin 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, of exporting in the year after privatization by 1.9%, and of exporting 2 years after

**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 | 0.070 |
| 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 assets | 0.088 | 0.099 | 2.85*** | 0.089 | 0.090 | 0.182 |
| Liquidity ratio | −0.033 | −0.011 | 2.99*** | −0.032 | −0.036 | 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 (ie, former SOEs privatized in year *t*) and the control group (ie, remaining SOEs) using *t*-tests.
*, **, and *** denote statistical significance at the 10%, 5%, and 1% levels, respectively.

**Table 5:** PSM estimation of the effects of privatization

| Outcome variable (*Y*) | Time difference | Mean for treatment group | Mean for matched control group | Difference | *t* statistics |
|---|---|---|---|---|---|
| Export dummy | $Y_t - Y_{t-1}$ | 0.050 | 0.029 | 0.021 | 3.380*** |
| | $Y_{t+1} - Y_{t-1}$ | 0.059 | 0.040 | 0.019 | 2.928*** |
| | $Y_{t+2} - Y_{t-1}$ | 0.058 | 0.042 | 0.017 | 2.433** |
| Log of TFP | $Y_t - Y_{t-1}$ | 0.035 | 0.003 | 0.033 | 4.946*** |
| | $Y_{t+1} - Y_{t-1}$ | 0.076 | 0.039 | 0.037 | 5.284*** |
| | $Y_{t+2} - Y_{t-1}$ | 0.114 | 0.076 | 0.038 | 5.084*** |
| Log of labor | $Y_t - Y_{t-1}$ | −0.029 | −0.057 | 0.028 | 2.179** |
| | $Y_{t+1} - Y_{t-1}$ | −0.076 | −0.106 | 0.031 | 1.981** |
| | $Y_{t+2} - Y_{t-1}$ | −0.101 | −0.171 | 0.070 | 4.012*** |
| Liquidity ratio | $Y_t - Y_{t-1}$ | 0.003 | 0.004 | −0.001 | 0.069 |
| | $Y_{t+1} - Y_{t-1}$ | 0.006 | 0.005 | 0.001 | 0.184 |
| | $Y_{t+2} - Y_{t-1}$ | 0.018 | 0.007 | 0.011 | 1.378 |
| Leverage ratio | $Y_t - Y_{t-1}$ | 0.004 | −0.019 | 0.023 | 1.019 |
| | $Y_{t+1} - Y_{t-1}$ | −0.015 | −0.020 | 0.005 | 0.195 |
| | $Y_{t+2} - Y_{t-1}$ | −0.049 | −0.021 | −0.028 | 1.021 |
| Long-term liability ratio | $Y_t - Y_{t-1}$ | −0.006 | −0.005 | −0.001 | 0.481 |
| | $Y_{t+1} - Y_{t-1}$ | −0.015 | −0.009 | −0.006 | 1.809 |
| | $Y_{t+2} - Y_{t-1}$ | −0.018 | −0.012 | −0.006 | 2.206** |

*Note*: 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% levels, respectively.

privatization by 1.7%. Therefore, privatization roughly doubles the probability of exporting for non-exporting SOEs.

The latter rows of Table 5 show the effects of privatization on firm performance. First, the effect of privatization on productivity growth is positive, statistically significant, and quantitatively large, as Jefferson and Su (2006) and Bai *et al.* (2009) found for China and Estrin *et al.* (2009) surveyed for Central and Eastern Europe. Privatization improves TFP by 3.3% in the same year and by 3.7% and 3.8% 1 and 2 years later, respectively. The effect of privatization on firm size, as 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, that is, the liquidity ratio and the leverage ratio. However, privatization lowers the ratio of long-term debts to total assets after 2 years. We can interpret this result regarding long-term debt in two different ways. On the one hand, lowering long-term debt may imply that privatized SOEs face tighter credit constraints and therefore cannot borrow as much as they could have had they remained state-owned. On the other hand, lowering long-term debt may imply that privatized SOEs improve their financial positions by intentionally reducing debt. Because the negative effect on long-term debt emerges 2 years after privatization, it is more likely that the latter is true; if tighter credit constraints led privatized SOEs to reduce their debts, the negative effect should emerge immediately after privatization. Therefore, we conclude that privatized SOEs purposefully reduce their long-term debt to achieve healthier financial conditions.



## Sources of the effects of privatization on export

The next question concerns the channels through which the effects of privatization on decisions to export are felt. To address this issue, we first consider the determinants of being an exporter. Du and Girma (2007) show that Chinese firms are more likely to export when they are more productive, larger, and less credit-constrained (ie, with more bank debt). To confirm and quantify these effects using our data, we employ a generalized method of moments (GMM) estimation in which our instrumental variables are lagged regressors and the state-ownership ratio. This estimation allows us to correct for possible biases that might arise because of endogeneity when estimating the determinants of exporting decisions.

Table 6 presents results from the GMM estimations. As shown in Du and Girma (2007), the productivity level (as measured by TFP) and the firm size (as measured by the number of workers) have positive and significant effects on the decision to export. In columns (1)–(3) of Table 4, we use the liquidity ratio (the ratio of current assets minus current debt to total assets), the

**Table 6:** Determinants of export (dependent variable: export dummy)

| | (1) GMM | (2) GMM | (3) GMM |
|---|---|---|---|
| Export dummy ($t$−1) | 0.894*** | 0.894*** | 0.894*** |
| | (0.00149) | (0.00149) | (0.00149) |
| Log of TFP ($t$−1) | 0.0156*** | 0.0159*** | 0.0162*** |
| | (0.00260) | (0.00257) | (0.00251) |
| Log of labor ($t$−1) | 0.0134*** | 0.0133*** | 0.0132*** |
| | (0.000444) | (0.000441) | (0.000438) |
| Age | −4.08e−04*** | −4.08e−04*** | −4.16e−04*** |
| | (3.23e−05) | (3.23e−05) | (3.24e−05) |
| Log of exports in the region ($t$−1) | 3.87e−03*** | 3.87e−03*** | 3.91e−03*** |
| | (2.68e−04) | (2.68e−04) | (2.69e−04) |
| Liquidity ratio ($t$−1) | 9.22e−04 | | |
| | (1.88e−03) | | |
| Leverage ratio ($t$−1) | | 1.93e−04 | |
| | | (8.38e−04) | |
| Ratio of long-term debt to assets ($t$−1) | | | 0.0108** |
| | | | (4.98e−3) |
| Industry dummies | Yes | Yes | Yes |
| Region dummies | Yes | Yes | Yes |
| Year dummies | Yes | Yes | Yes |
| Number of observations | 600,204 | 600,016 | 600,204 |
| $R^2$ | 0.652 | 0.652 | 0.652 |

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



leverage ratio (the ratio of current debt to current assets), or the ratio of long-term debt to total assets as a measure of financial conditions. Among these three measures, the ratio of long-term debt to total assets has a positive and significant effect, whereas the two variables that represent short-term financial conditions have no significant effects. This finding suggests that firms that can increase their long-term debt are more likely to engage in export because exporting may require long-term investment in product modification and/or foreign marketing. All of these results are consistent with the findings of Du and Girma (2007).

Combining these results with the results in the section 'Effects of privatization', we can identify the channels through which privatization affects exporting decisions and evaluate them quantitatively. First, privatization improves the level of TFP by an average of 3.28% (Table 5). Because the coefficient on the log of TFP in column (3) of Table 6 is 0.0162, privatization increases the likelihood of exporting by 0.0531% (=3.28% × 0.0162) through productivity improvement. Second, privatization raises the probability of exporting by 0.0371% (=2.81% × 0.0132) by increasing firm size.

Because privatization increases the probability of exporting by approximately 2%, the effects of privatization on exporting decisions through productivity and size improvement are small in size. Third, privatized SOEs decrease long-term debt 2 years after privatization, and the decline in debt levels affects the decision to export negatively. However, the negative effect of privatization by lowering debt levels on the probability of exporting is only $-0.00659\%$ ($= -0.61\% \times 0.0108$) and quantitatively negligible.

In sum, although we identified certain positive effects of privatization on exporting decisions through improved productivity and increased size and a negative effect through lower long-term debt levels in the medium run, these effects are quantitatively small or negligible. Thus, privatized SOEs are more likely to engage in exporting than firms that remain SOEs mainly because of unobservable factors that are not systematically related to productivity, firm size, or financial conditions.

A potential unobservable explanation for such effects may be that firms change their attitudes about profits and risks. After privatization, former SOEs are no longer protected by the government and are exposed to competitive markets, therefore they may have to expand their business to survive. One way to expand their business is to export to foreign markets. Moreover, privatized SOEs must take more risks than SOEs if they are to be profitable, which may also lead to exporting. Using firm-level data for Japanese small- and medium-sized enterprises, Todo and Sato (2011) found that more risk-loving firms are more likely to export and that the effect of risk preference on exporting decisions is far greater in magnitude than the effects of productivity and firm size. Because we do not have data on firm attitudes toward profits and risks, we cannot formally test this hypothesis. However, it should be emphasized that these psychological factors might play an important role in determining the effects of privatization. This conclusion is in line with the findings of Filatotchev *et al.* (2001) for the former Soviet Union and Shinkle and Kriauciunas (2010) for Central and Eastern Europe that governance and management strategies and institutions in the economy after privatization largely affect decisions to export.



### Long-run effects of privatization

We further examine the long-run effects of privatization on exporting decisions because some new exporters expand and continue to export, whereas others stop exporting within a few years of beginning. Akhmetova (2010) and Akhmetova and Mitaritonna (2012) emphasize that firms often begin with a small volume of exports and expand this amount over time because they learn by exporting. Eaton *et al.* (2007) observe that more than half of all exporters in Colombia exit export markets within 1 year. Similar evidence can be observed for China. Figure 2 shows the change from 2001 to 2007 in the

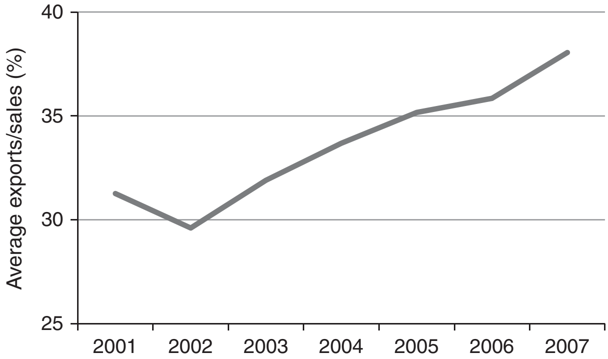


**Figure 2:** Average percentage of exports out of total sales for firms that began exporting in 2001 and continued to export in 2007
*Note*: This figure is based on the entire sample, including private firms, from the firm-level data for China.

**Table 7:** Top five patterns of export dynamics of 2,358 firms that began 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 that began exporting in 2001 are defined as firms that did not export in 2000 but exported in 2001.



average export intensity, or the ratio of exports to total sales, of firms that began exporting in 2001[6] and continued to export in 2007. The average export intensity was 32% in the first year of exporting (2001), which increased to 38% by 2007. Table 7 shows the top five patterns of export dynamics for firms that began exporting in 2001. This table indicates that 30% of new exporters in 2001 stopped exporting the next year (2002). Although the fraction of firms exiting the export market decreased over time from 2001 to 2004, as Akhmetova (2010) and Akhmetova and Mitaritonna (2012) suggest, more than half of the new exporters in 2001 had stopped exporting within 4 years. The high exit rate is consistent with the findings of Eaton *et al.* (2007) for Colombia.

Given these findings, we are concerned about the duration of the effect of privatization on exporting decisions. For example, let us suppose that the effects from learning by exporting are small and that privatized SOEs tend to export too aggressively because they overestimate their competitiveness in

[6] This refers to firms that did not export in 2000 but exported in 2001.

foreign markets. Thus, privatized SOEs may begin exporting, but many of the new exporters quit exporting within a few years because exporting is not profitable. In this case, privatized SOEs may be more likely to engage in exporting after privatization in the short run but not in the long run.

To test this hypothesis, we estimate the effects of privatization in year $t$ on exporting decisions and firm performance from year $t$ to year $t+5$. We employ the same PSM estimation that we used in the section 'Effects of privatization' except for one difference. Although our sample in the section 'Effects of privatization' consists of firms that existed continuously from year $t-1$ to $t+2$, we now include firms that existed from year $t-1$ to year $t+5$, in addition to those that existed in year $t-1$ and year $t$ but disappeared during the period from year $t+1$ to year $t+5$. This inclusion is because attrition bias may be substantial when we examine long-run effects. When a firm disappears from the data set, we assume that the firm is not engaging in export. One disadvantage of this method is that we cannot observe the productivity, size of the labor force, or financial conditions of firms that are no longer in the data set. As such, we should limit our attention to the effects of privatization on export decisions. The sample consists of 7,085 firms, which is much smaller than the sample in the section 'Effects of privatization' in which 13,991 firms that existed continuously from year $t-1$ to year $t+2$ were included.

Figure 3 summarizes the long-run effects of privatization graphically. The horizontal axis indicates the number of years after privatization and the

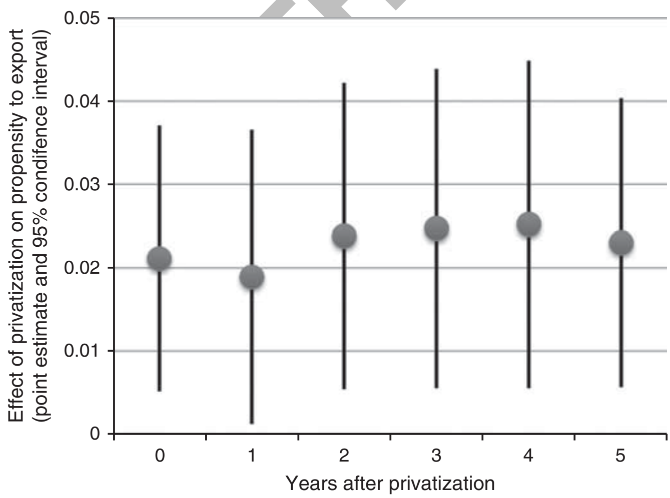


**Figure 3:** Long-run effects of privatization
*Note*: This figure is based on propensity score matching estimations of the effect of privatization on the propensity to export 0–5 years after the privatization.

vertical axis indicates the effect of privatization on the probability of export. The dot and the line segments for each year represent the point estimate and the 95% confidence interval, respectively, of the effects of privatization obtained from the PSM estimations, as in the section ‘Effects of privatization’. Therefore, if the lower bound of a line segment is above zero, the effect will be statistically significant at the 5% level.

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

## CONCLUDING REMARKS

This paper examines whether the privatization of Chinese SOEs increases the probability of exporting and, if so, what channels generate such an effect. Using firm-level data for the Chinese manufacturing sector for the 2000–2007 period, we find that privatization has a positive effect on decisions to export, productivity, and firm size. Conversely, privatization has a negative effect on the level of firms’ long-term debt. We also find that Chinese firms are more likely to engage in exporting when productivity levels, firm size, or long-term debt are larger. These two sets of results suggest that privatization has positive effects on exporting decisions because it improves productivity and firm size but has a negative effect by lowering long-term debt levels. However, a quantitative analysis reveals that the effects of privatization through these three channels are quantitatively small. Therefore, we conclude that the positive effect of privatization on exporting decisions is mainly the result of other unobservable factors that are most likely related to changes in attitudes about the profits and risks associated with privatization. However, because of data limitations, we cannot formally test whether this effect of privatization impacts exporting decisions. We will let future studies address this matter.

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## REFERENCES

Akhmetova, Z. 2010: Firm experimentation in new markets. Mimeo, available at http://www.freit.org/EITI/2011/SubmittedPapersG/Zhanar_Akhmetova1.pdf (accessed 21 March 2014).

Akhmetova, Z and Mitaritonna, C. 2012: Structural estimation of the model of firm experimentation. Mimeo, available at https://economics.adelaide.edu.au/research/seminars/2012_aug_24_Akmetova.pdf (accessed 21 March 2014).

Arnold, JM and Javorcik, BS. 2009: Gifted kids or pushy parents? Foreign acquisitions and plant performance in Indonesia. *Journal of International Economics* 79(1): 42–53.

Bai, C-E, Lu, J and Tao, Z. 2009: How does privatization work in China? *Journal of Comparative Economics* 37(3): 453–470.

Bernard, AB and Jensen, JB. 1999: Exceptional exporter performance: Cause, effect, or both? *Journal of International Economics* 47(1): 1–25.

Bernard, AB and Jensen, JB. 2004: Why some firms export. *Review of Economics and Statistics* 86(2): 561–569.

Brandt, L, Van Biesebroeck, J and Zhang, Y. 2011: Creative accounting or creative destruction? Firm-level productivity growth in Chinese manufacturing. *Journal of Development Economics* 97(2): 339–351.

Clerides, SK, Lach, S and Tybout, JR. 1998: Is learning by exporting important? Micro-dynamic evidence from Colombia, Mexico, and Morocco. *Quarterly Journal of Economics* 113(3): 903–947.

Du, J and Girma, S. 2007: Finance and firm export in China. *Kyklos* 60(1): 37–54.

Eaton, J, Eslave, M, Kugler, M and Tybout, JR. 2007: *Export dynamics in Columbia: Firm-level evidence*. NBER Working paper, no. 13531, National Bureau of Economic Research: Cambridge.

Elliott, R and Zhou, Y. 2013: State-owned enterprises, exporting and productivity in China: A stochastic dominance approach. *World Economy* 36(8): 1000–1028.

Estrin, S, Hanousek, J, Kocenda, E and Svejnar, J. 2009: The effects of privatization and ownership in transition economies. *Journal of Economic Literature* 47(3): 699–728.

Feenstra, RC, Li, Z and Yu, M. 2011: *Exports and credit constraints under incomplete information: Theory and evidence from China*. NBER Working paper, no. 16940, National Bureau of Economic Research: Cambridge.

Filatotchev, I, Dyomina, N, Wright, M and Buck, T. 2001: Effects of post-privatization governance and strategies on export intensity in the former Soviet Union. *Journal of International Business Studies* 32(4): 853–871.

Girma, S and Görg, H. 2007: Evaluating the foreign ownership wage premium using a difference-in-difference matching approach. *Journal of International Economics* 72(1): 97–112.

Greenaway, D, Guariglia, A and Kneller, R. 2007: Financial factors and exporting decisions. *Journal of International Economics* 73(2): 377–395.

Greenaway, D and Kneller, R. 2004: Exporting and productivity in the United Kingdom. *Oxford Review of Economic Policy* 20(3): 358–371.

Heckman, JJ, Ichimura, H and Todd, P. 1997: Matching as an econometric evaluation estimator: Evidence from evaluating a job training programme. *Review of Economic Studies* 64(4): 605–654.
Heckman, JJ, Ichimura, H and Todd, P. 1998: Matching as an econometric evaluation estimator. *Review of Economic Studies* 65(5): 261–294.
Hirano, K, Imbens, GW and Ridder, G. 2003: Efficient estimation of average treatment effects using the estimated propensity score. *Econometrica* 71(4): 1161–1189.
Jefferson, GH and Su, J. 2006: Privatization and restructuring in China: Evidence from shareholding ownership, 1995–2001. *Journal of Comparative Economics* 34(1): 146–166.
Lin, JY and Li, Z. 2008: Policy burden, privatization and soft budget constraint. *Journal of Comparative Economics* 36(1): 90–102.
Lizal, L and Svejnar, J. 2002: Investment, credit rationing, and the soft budget constraint: Evidence from Czech panel data. *Review of Economics and Statistics* 84(2): 353–370.
Melitz, MJ. 2003: The impact of trade on intra-industry reallocations and aggregate industry productivity. *Econometrica* 71(6): 1695–1725.
Muuls, M. 2008: *Exporters and credit constraints: A firm-level approach*. National Bank of Belgium: Brussels. Working paper, no. 139.
Olley, SG and Pakes, A. 1996: The dynamics of productivity in the telecommunications equipment industry. *Econometrica* 64(6): 1263–1297.
Qian, Y and Roland, G. 1996: The soft budget constraint in China. *Japan and the World Economy* 8(2): 207–223.
Rosenbaum, PR and Rubin, DB. 1983: The central role of the propensity score in observational studies for causal effects. *Biometrika* 70(1): 41–55.
Rubin, DB. 1980: Bias reduction using mahalanobis-metric matching. *Biometrika* 36(2): 293–298.
Shinkle, GA and Kriauciunas, AP. 2010: Institutions, size and age in transition economies: Implications for export growth. *Journal of International Business Studies* 41(2): 267–286.
Sianesi, B. 2004: An evaluation of the active labour market programmes in Sweden. *Review of Economics and Statistics* 86(1): 133–155.
Smith, JA and Todd, P. 2005: Does matching overcome Lalonde's critique of nonexperimental estimators? *Journal of Econometrics* 125(1–2): 305–353.
Todo, Y and Sato, H. 2011: *Effects of CEOs' characteristics on internationalization of small and medium enterprises in Japan*. RIETI Discussion paper, no. 11-E-026, Research Institute of Economy, Trade and Industry, Tokyo.
Yusuf, S, Nabeshima, K and Perkins, DH. 2006: *Under new ownership: Privatizing China's state-owned enterprises*. The World Bank: Washington D.C.