KATHOLIEKE UNIVERSITEIT
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Faculty of Economics and
Applied Economics

Department of Economics

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by

Hang GAO
Johannes VAN BIESEBROECK

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DISCUSSION
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Effects of Deregulation and Vertical Unbundling on the Performance of China's Electricity Generation Sector*

Hang Gao
K.U.Leuven

Johannes Van Biesebroeck†
K.U.Leuven and CEPR

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Abstract

We study whether the 2002 deregulation and vertical unbundling of the Chinese electricity sector has boosted productivity in the generation segment of the industry. Controlling explicitly for sources of price-heterogeneity across firms and for firm-fixed effects, we find deregulation to be associated with a reduction in labor input and material use of 6 and 4 percent, respectively. This effect only appears two years after the reforms, is robust to alternative ways of identifying restructured firms, and to the nonrandom selection of restructured firms using a matching estimator. Input use of new state-owned firms that start operations two years into the reform period does not differ significantly anymore from input use of private sector entrants.

JEL codes: L5, L9, O4

Keywords: Productivity, regulation

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†E-mail: hang.gao@econ.kuleuven.be and jo.vanbiesebroeck@econ.kuleuven.be.

1 Introduction

We study whether the 2002 deregulation and vertical unbundling of the Chinese electricity sector has boosted productivity in the generation segment of the industry. Efficient operation of electricity plants is particularly important in China for several reasons. The manufacturing sector accounts for an unusually large share of the economy and it requires reliable and ever increasing amounts of electric power. Half of all provinces have suffered rolling brownouts in the last decade because of local capacity shortages and refusals of producers to generate electricity when generating costs surpassed regulated prices. Total electricity demand continues to grow and new capacity is constantly added. Operating the existing capacity more efficiently would slow down the rate of new capacity additions and allow for more careful planning. Given China's status as the world's largest emitter of greenhouse gasses, it is an issue of global importance.

As in most countries, deregulation of the electricity sector is an ongoing process—see Joskow (1997) for an overview of the U.S. experience. Landmark changes for China occurred in 1985, when investment in electricity generation by local governments, foreign companies, and private investors was first permitted, and in 1997, when the modern enterprise system replaced the former central planning system. We study the effects of the latest policy changes which were published as a blueprint for electricity reform in 2002. Competition was introduced by dismantling the former vertically integrated public utility into competitive generation and distribution enterprises. Next, a market-oriented price mechanism would be established through competitive electricity wholesale markets and the relaxation of coal price regulations. While the restructuring was implemented immediately, much of the pricing reforms were suspended or scaled back subsequently.

Several aspects of electricity sector deregulation have received attention in the literature. Most notably, the abuse of market power has been faulted as the main culprit for the failure of the California regulatory regime in the summer of 2000 by Borenstein, Bushnell, and Wolak (2002) and Joskow and Kahn (2002). In contrast, cross-country studies like Hattori and Tsutsui (2004) have documented the ability of deregulation to temper price growth. Newbery and Pollitt (1997) conduct a cost-benefit analysis of the reforms in the United Kingdom taking into account generator efficiency gains, but also effects on emissions, pricing, and investment incentives. Knittel (2002) reports evidence of efficiency increases at generation plants associated with the diffusion of incentive regulation.

The main purpose of the vertical unbundling and divestiture of generation assets in China was to improve the production efficiency of electricity generation. Several studies have attempted to quantify ex-post operating efficiency gains from similar restructuring episodes in other countries. Newbery and Pollitt (1997) find that the move from a state-owned monopoly to a privatized, competitive generation market in the United Kingdom was accompanied by significant labor force reductions. This represents both restructuring and privatization effects. Hiebert (2002) provides the first econometric plant-level evidence studying the effect of U.S. restructuring over the 1988–1997 period using a stochastic frontier production function. He finds a substantial improvement in generation plant efficiency for coal plants in states that restructured in or before 1996, but no effect for gas-fired plants in the same states nor for policies enacted in 1997.

A shortcoming of these earlier studies is that productivity is assumed orthogonal to input choices. Following the path-breaking work of Olley and Pakes (1996), more recent studies have taken endogeneity of productivity more seriously. In particular, Fabrizio, Rose, and Wolfram (2007) first estimate productivity using the control function approach and then apply a difference-in-differences method to uncover the effect of regulatory restructuring on U.S. electricity generation efficiency.[1] Their results suggest that labor and non-fuel efficiency of investor-owned utilities (IOUs) in the states that passed restructuring legislation increased by 3 to 5 percent relative to IOUs in non-restructuring states, and by 6 to 12 percent relative to municipal and federal plants insulated from restructuring incentives. They find little improvement in fuel efficiency.[2]

Two studies that rely on country-level information and exploit the differential timing of reforms across countries also find evidence for efficiency gains with deregulation. Steiner (2000) studies OECD countries and finds that while changes in legal rules only translate slowly into changes in conduct, unbundling of generation and introducing private ownership has a positive and significant impact on most performance measures. Zhang, Parker, and Kirkpatrick (2008) study reforms in 51 developing countries. They find favorable effects on service penetration, capacity expansion, labor efficiency, and prices for industrial users.

A final important point of comparison is Du, Mao, and Shi (2008), which is the first study the evaluate the combined effect of reforms in 1997 and 2002 on the production efficiency of China's electricity generation industry. They closely follow

[1] They stress that failing to recognize that shocks to input productivity may induce firms to adjust targeted output leads to an upward bias in the estimated efficiency effects by almost a factor of two in some cases.

[2] Bushnell and Wolfram (2005) focus specifically on the effect of divestitures on fuel efficiency. They find a 2% gain for divested plants, but a similar efficiency gain for plants that remained under utility ownership but now faced incentive regulation. Hence, their conclusion that changes in incentives rather than ownership were the main drivers of fuel efficiency improvement.

the estimation approach in Fabrizio et al. (2007) and find large efficiency gains of 29 percent in labor input and 35 percent in non-fuel materials for the plants divested from the former Ministry of Electricity Power (MEP) or the State Power Company (SPC) relative to other firms, but no evidence of greater efficiency in fuel use. These effects are cumulatively over a nine year period between two census years, they only observe two cross sections of firms in 1995 and 2004, and translate to 2.9 and 3.4 percent efficiency gain annually.

To study the electricity industry in China we face a number of complicating factors and we advance the existing literate on three counts: (i) using more comprehensive and more recent data, (ii) controlling explicitly for firm-heterogeneity in coal and output prices, and (iii) verifying the robustness of estimates to alternative ways of defining deregulated entities. We briefly elaborate on each contribution.

First, we observe the universe of all Chinese fossil-fired electricity generation companies from 1998 to 2007.[3] As we can track firms over time in a panel, we can include fixed effects to capture an important dimension of firm-heterogeneity and only need instrumental variables to control for the remaining endogeneity.[4] Moreover, observing several years in the post-reform period allows us to pay particular attention to the timing of the effects as adjustments are likely to take some time in a capital intensive sector. We also use matching techniques from the treatment evaluation literature to control for the possibility of non-random treatment of firms.

Second, the continued existence of regulated prices, not only for the final product but most importantly for the coal input, makes the productivity growth calculations a delicate undertaking (Wang, 2007). The complicated system of generation tariffs and the dual track pricing on the coal market lead to output and input prices that differ across firms. This missing data problem is addressed by relying on institutional details of the electricity and coal markets that suggest firm-age, size, location, and legal ownership structure as key explanatory factors for the price differences at the firm level.[5]

[3]Our sample only contains firms that are state-owned or with sales above five million RMB, but any generation plant easily satisfies the second criteria.

[4]Du et al. (2008) used provincial thermal power output as a proxy for plant-level electricity output and provincial fuel usage as a proxy for plant-level fuel input.

[5]In order to stimulate electricity generation, the government guaranteed a healthy profit to firms active in this industry. The electricity tariff structure is rather inscrutable, but at a regional level the prices are fixed (Wang, 2007). More importantly, the price of 'electricity coal' is still being tightly controlled and kept 30-40 Yuan per tonne below the market rate, which determines the coal price for other uses than electricity generation and has largely been deregulated. Such underpricing creates scarcity and large state-owned generation plants have preferential access to the regulated coal coming from state-owned mines (Mathys, 2011).

Third, identifying divested subsidiaries, plants or firms, is not without ambiguity. As is well known, many ownership categories in China are hybrid forms between state and private ownership. Du et al. (2008) focused on plants previously owned by the former MEP or SPC. It is likely, though, that subsidiaries that had already transformed to listed shareholder companies or already had a diversified ownership structure in 2002 were less affected by the restructuring. For example, their management is more likely to have been stable under the reforms. At the same time, the attempts to deregulate output and input prices are likely to have influenced not only MEP and SPC subsidiaries, but all state-owned companies with preferential access to coal quotas or with influence over electricity prices. Following Zhang, Zhang, and Zhao (2001), we rely in the benchmark estimates on the legal ownership categorization in 2002 to identify restructured firms and we perform sensitivity checks using alternative definitions.

The estimates suggest a positive impact of the latest round of regulatory reforms on production efficiency, appearing two years after the publication of the blueprint in 2002. Restructured fossil fuel-fired electricity generation firms reduced labor input by 6 percent and material input (fuel and non-fuel combined) by 4 percent relative to private or foreign-owned firms, which are only indirectly affected by the reforms. These effects of deregulation on productivity growth are shown to be robust using conservative evaluation methods that control for unobserved firm-effects, nonrandom selection of restructured firms, and different definitions of restructuring.

Subsequently, we have compared the remaining differences in productivity levels by ownership category at the end of the sample period. State firms already in existence in the first year of the sample still used significantly more labor input than other firms at the end of the period, although the gap narrowed by half if they changed ownership in the meantime. Their material input use is slightly higher as well, but the difference is barely significant. Importantly, the labor and material input use of state firms that were established in 2004, two years into the reform period, is not significantly different from other firms.[6]

The remainder of the paper is organized as follows. Section 2 reviews the electricity reforms in China. Section 3 presents the empirical model and estimation strategy. Section 4 describes the data and provides summary statistics. Section 5 presents the results, inferences and robustness checks. Section 6 discusses the remaining productivity differences between state and privately owned firms, both for new and existing firms. The last section concludes.

[6] A similar analysis in the pre-reform period did find significantly higher input use by state-owned entrants.

2 Electricity sector restructuring

The history of China's electricity sector deregulation since the start of the general market-oriented economic reforms in 1979 is marked by three important restructuring episodes. Prior to 1985, China's electricity sector was treated like a natural monopoly, even more so than in many other countries. The central government determined all prices and quantities, including for coal and electricity, and had the final say in all investment decisions. In subsequent steps, new objectives were introduced: attract private investment, separate the administrative authority from the business operations, and vertically unbundle generation and transmission.

To cope with rapidly expanding demand and frequent brownouts, the reforms of 1985 started a long process of deregulation. To relieve power shortages—the excess demand amounted to 450-500 kWh or nearly 12% of total generation capacity (Wang, 2007)—and lack of capital, the government enacted the "Temporary Provision on Promoting Fund-Raising for Investment in the Electricity Sector and Implementing Different Electricity Tariffs" policy. It allowed for the first time investment in electricity generation from local governments, domestic enterprises, and foreign companies.

The transmission lines and distribution grids were still owned by the Ministry of Electricity Power (MEP). The new independent power producers (IPPs) provided electricity at tariffs set by the government and obtained regulated profits, whereas state-owned plants continued to sell at lower "plan" prices. The independent generation tariffs were reset every year based on accounting cost information and according to rate-of-return regulation principles also used in Western countries. Tariffs differed across plants of the same firm, sometimes even across units within a single plant.[7] By the end of the 1990s, more than half of all electricity was generated in non-state-owned plants (Du et al., 2008).

Following this first phase of reforms, electricity tariffs set by the government increased rapidly. The next stage of deregulation occurred in 1997 when the MEP was split up into two bodies. A new public utility, State Power Company (SPC), took over the electricity assets, including generation plants, transmission, and distribution grids. The State Economic and Trade Committee (SETC) became the new regulator and took over the administrative and decision-making functions of MEP. One of its immediate actions was to slow down the increase in electricity prices.

The third and final reform was the introduction of competition in the generation

[7] The calculation of tariffs depended on the type of company, the origin of the capital, and accounting costs. The objective was for investments by IPPs to be recovered within a fixed period, generally within 10 years.

sector by dismantling the SPC in 2002. Its generation and transmission assets were divested into five large generation corporations and two transmission companies.[8] Because the newly formed companies had exclusive areas of operation, the divesture was more of a move to decentralization than the introduction of genuine competition.[9] It did make possible the celebrated Chinese practice of local experimentation with regulation and reforms.

In the same year, an independent regulatory agency, the State Electricity Regulatory Commission (SERC) was created to supervise and establish a legal framework for the electricity market (Pittman and Zhang, 2008). The government also attempted to establish a market-oriented pricing mechanism for electricity. Five competitive regional wholesale electricity markets were scheduled to be constructed (Xu and Chen, 2006). Generators would bid into these regional wholesale markets and gain grid-accessing priority according to their bidding.[10] This arrangement would benefit more efficient generators and encourage all firms to improve their productivity.

Equally important, the government also announced to stop its policy of guiding coal prices and that it would allow a market to develop. This lead to rapid price increases for coal, while the government was reluctant to allow these cost increases immediately be reflected into electricity generation tariffs. Rising inflation was one concern, and the possibility of harming the competitiveness of the manufacturing sector a second. The pricing conflict between coal and electricity producers after 2002 exacerbated electricity shortages. The government eventually decided to postpone, or perhaps even call-off, the planned transition to market-pricing in this sector. The evolution of prices for electricity, fuel, and the general consumer price index is illustrated in Figure A.1 in the Appendix.

The episode underscores two important features of the post-reform electricity sector in China. First, improving the efficiency of generating firms is vital for their survival as fuel prices are bound to increase and their ability to pass on cost increases to customers is severely limited. Second, regulated electricity coal prices are substantially below market and firms with access to this rationed commodity will appear as being highly productive as their fuel input use is underpriced. State-owned electricity generators are more likely to have the necessary connections to access the cheap coal produced by state-owned mines. Because the group of firms for which we

[8]The five generation companies are China Huaneng Corporation, China Datang Corporation, China Huadian Corporation, China Guodian Corporation and China Power Investment Corporation, and the two grids companies are State Grid Company and China Southern Grid Company.

[9]Competition with the IPPs established in the first phase of the reforms continued unabated.

[10]The New Electricity Trading Arrangement in the U.K. was a comparable mechanism (Wang, 2007) and many other countries established similar regional power pools at the wholesale level.

want to measure the productivity growth following deregulation is likely to be most affected by fluctuations in the availability of subsidized coal, it is crucial to control explicitly for coal price heterogeneity in the analysis.[11]

While the government's policy on price competition was reversed, the last round of reform in 2002 did trigger the restructuring of state-owned firms that were subsidiaries of the former SPC. The objective of our analysis is to assess whether this reform boosted productivity in restructured fossil fuel-fired generation firms relative to the productivity evolution at other companies. The control group of IPPs was only indirectly affected by the reforms, through the competition they faced, and we will control for possible effects on electricity prices.

3 Empirical model specification

The key estimating equations to evaluate the effects on productivity are taken from Fabrizio et al. (2007) and Du et al. (2008) with a few modifications. Because in the Chinese electricity sector firms do not directly control prices and because demand is highly inelastic, a cost minimization framework is most appropriate to model short-run firm behavior. The Leontief functional form for the production function reflects the inability to substitute in the short run between fuel, on one hand, and capital and labor input, on the other hand.[12]

$$\begin{aligned} Q &= \min_{M,L} \left\{ f_1(M,\beta,\varepsilon_M), f_2(K,L,\alpha,\varepsilon_L) \right\} \\ \text{s.t.} &\quad Q \geq \bar{Q} \end{aligned} \tag{1}$$

We treat capital input K as quasi-fixed in the short run as this is predetermined before labor and material input are decided. In the Chinese case, this applies even more strongly as firms in the electricity sector have only partial control over their own investment strategy which is subject to government approval. Full details on the derivation of factor demands are in Fabrizio et al. (2007), but it is clear that in equilibrium both terms in the production function will hold with equality. A first-order Taylor approximation to any monotonically increasing $f_1(\cdot)$ function immediately produces a log-linear material demand equation:

$$\ln M_{it} = \beta_Q \ln Q_{it} + \beta_t + \beta_i + \varepsilon^M_{it} \tag{2}$$

[11] In 2004, SERC implemented the "Coal and Electricity Prices Co-move" policy to arbitrate disputes over coal and electricity pricing, but it seems that both sides and even the end users of electricity are unsatisfied with this policy (Wang, 2007).

[12] This specification was first used in Van Biesebroeck (2003) to capture the inability to substitute between intermediate inputs and other production factors in automobile assembly plants.

M comprising both fuel and non-fuel expenditures,[13] and the error term ε^M captures measurement error and factors that affect a firm's material efficiency. We include time and firm-fixed effects to soak up some of the productivity heterogeneity. Because we only observe output and material input in value terms, we face a missing data problem for input and output prices. We discuss and address it below.

Capital and labor input are assumed to be substitutable to some extent, and the $f_2(\cdot)$ function is modeled using the Cobb-Douglas functional form: $K^{\alpha_K}L^{\alpha_L}e^{\varepsilon_L}$. Assuming that capital is exogenous but varying over time, the main job of the plant manager is to choose labor input each period to minimize the total wage bill, while satisfying the output constraint. Rewriting the first order condition, produces the following labor demand equation

$$L_{it} \quad = \quad (\lambda_{it}\alpha_L Q_{it})/W_{it}.$$

λ_{it} is the Lagrange multiplier associated with the output constraint and W_{it} is the wage level. In logarithms, the intrinsically unobservable term $\ln(\lambda_{it}\alpha_L)$ appears in the demand equation. The shadow price of output λ_{it} will vary with the available capital, the market environment, and firm-specific conditions, such the ownership of other generation plants by the same firm. As proxies for it, we include the time-varying capital stock, time and firm-fixed effects.

$$\ln L_{it} \quad = \quad \ln Q_{it} - \ln W_{it} + \alpha_K \ln K_{it} + \alpha_i + \alpha_t + \varepsilon^L_{it} \tag{3}$$

ε^L_{it} contains measurement error, productivity shocks, and remaining time-varying firm-specific variation in the shadow price of output. Problems of simultaneity between output and the error term are discussed below.

In contrast with Fabrizio et al. (2007), we include the capital stock explicitly in the labor demand equation rather than using 'plant-epoch' fixed effect for periods in between large investments. Production capacity is more frequently adjusted at the firm level (our unit of analysis) and the rapid growth in the Chinese electricity sector leads to frequent capacity additions also at existing facilities.[14]

To estimate the two equations (2) and (3), we need to address the simultaneity of productivity shocks and input choices, the missing price bias, and the nonrandom selection of restructured firms.

[13]Our data only contains information on total intermediate material inputs, which combines 'fuel material input' and 'industrial intermediate inputs'. The latter consists of transportation cost, repairs and storage cost, and intermediate service inputs, such as interest expenditure, advertising, insurance, education and travel cost. The assumption that total material input is proportional to output is often made to justify estimating a production function in value added terms—see Van Biesebroeck (2007) for a discussion.

[14]At the firm level, average fixed assets more than doubled between 1998 and 2007.

3.1 Simultaneity

Some of the assumptions regarding timing of decisions and realization of shocks have not been spelled out here in detail, but the interested reader is referred to Van Biesebroeck (2003) and Fabrizio, et al. (2007). The key simultaneity issue is that productivity shocks that affect the factor demands might be correlated with output. Managers could adjust the amount of electricity to produce after they observe the idiosyncratic shocks to labor and input productivity. This problem is likely to be more severe in the material than in the labor equation, as material input responds more directly to demand fluctuations. Adjusting the employment decision that was made based on demand expectation is not always easy, especially in a highly regulated sector with a lot of state ownership.

Fabrizio et al. (2007) and Du et al. (2008) have used aggregate electricity demand at the state or provincial level as an instrument for plant-level output. For this approach to be effective, firm-level electricity revenue and province-level electricity consumption need to be positively correlated. Over the sample period, the association is quite weak: the correlation is only 0.067 even pooling all firm-year observations. There are several reasons for this, some of them unique to the Chinese situation.

First, the correlation is diminished by heterogeneity in electricity prices which is to a large extent outside the firms' control. Second, many provinces are large and contain regions where firms are de facto local monopolies, only partially affected by province-wide demand fluctuations. Third, provincial demand and production is often unbalanced. For example, under the 'West Development Strategy' the government launched a vast project to transport electricity from western provinces to the developed coastal area. Weak instruments will lead to large standard errors and in finite samples to inconsistent and biased estimators (Verbeek, 2008).

An important advantage of our analysis over Du et al. (2008) is the availability of a relatively long (ten year) panel, allowing us to include firm-fixed effects. These soak up much of the heterogeneity in the cross-firm dimension which reduces the simultaneity problem in the quasi-differenced equation. We follow the Arellano and Bond (1991) approach and use lagged values of output as instrument for the changes in output. The strong correlation of firm-level output over time is indicative of limited changes in generation capacity between consecutive years and stable long-term contracts to sell electricity to the distribution grid. Both factors diminish the firms' ability and incentive to respond to short-term productivity fluctuations.[15]

[15] These same features of electricity markets also weaken the possibility of selection bias (Olley and Pakes, 1996). Firms are unlikely to base exit decisions on idiosyncratic productivity shocks

3.2 Missing price bias

As mentioned in the derivation of equations (2) and (3), we ideally would like to use physical quantities for inputs and output, but these are unavailable for material input (coal) and electricity production. As a result, prices are omitted variables that enter the error term and cause inconsistent estimates as they are certain to influence the levels of physical input and output.[16]

To address this problem, Du et al. (2008) replaced the dependent variable in the material demand equation—plant-level fuel input—with provincial fuel use. The similar replacement of firm-level electricity output with provincial electricity demand to take care of the simultaneity issue in the previous section also resolves the missing output price problem. Using the aggregate variable as an instrument instead of a proxy would be a better approach, but that would be impossible for the material input. The weak correlation between firm-level and provincial variables in our sample makes this approach problematic anyway.

An alternative solution would be to include deflated aggregate sales as an additional control in the input demand equations. Klette and Griliches (1996) have shown that this absorbs all output price heterogeneity if the industry can be characterized as monopolistically competitive with constant elasticity of substitution demand. In the Chinese electricity industry, it is implausible that firm-level price changes are always proportional to province-wide demand changes. They are the result of market power by regional monopolies and of variations over time (as market prices fluctuate) in the value of political connections that give some firms preferential access to subsidized coal or higher electricity tariffs.

Two factors temper the missing price problem on the output side. The constraints on electricity price increases after 2002 mentioned earlier limited the scope for differential price paths across firms, while the firm-fixed effects that we include capture much of the cross-sectional heterogeneity. State-owned companies sell most of their electricity at pre-determined transfer prices to the distribution networks. The IPPs sell at differentiated tariffs which are largely determined by firms' ownership type and the date and size of initial investment, although accounting information does enter the rate of return calculations to some extent.

On the input side, the dual-track price system provides some firms with access to regulated subsidized power coal. Other firms, especially newly entered IPPs, have

as many are local monopolies and aggregate electricity demand grew enormously over the sample period.

[16] There have even been instances of plants deliberately choosing to stop production to avoid losses when coal prices soared on the private market and regulated electricity prices remained constant.

to pay much higher market prices. Some evidence suggests that the price differences between the regulated an unregulated markets have fluctuated over time (Wang, 2007). The value of individual firms' preferential access to subsidized coal must therefore also fluctuate over time.

In a case study of the Chinese power coal market, Mathys (2011) investigates which variables—firm characteristics and market features—have the most explanatory power for price differentials and access to regulated power coal. The most important determinant is found to be a firm's location which determines ease of access to domestic mines, to imported coal, and to the congested transportation infrastructure. Provincial dummies capture much of these location effects. They also control for residual variation in output prices, because relative price differences across localities tend to be stable, while there is some variation in provincial price regulation over time. Firm age, size, and ownership are found to matter as well, as they help predict historical plan allocations and bargaining power.[17]

In the estimating equations, we use industry deflators for electricity output and coal input. We substitute out the differences between firm-specific prices and the average using a second order polynomial in all the price-control variables. Because the last three variables vary over time, the location controls are not collinear with the firm-fixed effects.

3.3 Treatment identification

The 2002 reforms affected all firms indirectly as the industry restructured. Firms that were divested from the state-owned integrated public utility were impacted much more because their ownership and often their management changed, in addition to the regulatory framework governing their actions. We identify firms that were purely state-owned in 2002 as 'treated'. These are selected when the variable 'legal structure of Chinese company ownership' in 2002 indicates either state-owned company (type 110) or state-solely-funded corporation (type 151). The $STATE_0$ dummy is set to unity for this group of firms and to zero for other firms—private, foreign, collective, and mixed-ownership—that make up the control group.

As we need to include firm fixed-effects in the estimation to help control for unobservable productivity and price differences, we cannot identify cross-sectional differences in the level of input demand for restructured firms. We can identify whether the factor demands for restructured and control firms evolve differently

[17] The legacy of the planned period allows state-owned firms to access 'in-plan' power coal prices guided by National Development and Reform Commission and, equally important, guaranteed transportation by the Ministry of Railways.

over time by interacting the constant $STATE_0$ dummy with a set of time dummies. The coefficients of interest, λ_τ and μ_τ in the labor and material demand equations (4) and (5) below, are normalized to zero in the first year of the sample (1998).

Introducing both the price and restructuring controls in the factor demands leads to the following two estimating equations:

$$\begin{aligned}
\ln(EMPLOY_{it}) &= \alpha_i + \alpha_R \ln(REVENUE_{it}) + \alpha_P' X_{it} + \alpha_W \ln(WAGE_{it}) \qquad (4) \\
&+ \sum_{\tau=2}^{T} \Big(\alpha_\tau + \lambda_\tau STATE_{0i}\Big) \cdot I_{[YEAR_{it}=\tau]} + \varepsilon_{it}^{L} \\
\ln(MATERIAL_{it}) &= \beta_i + \beta_R \ln(REVENUE_{it}) + \beta_P' X_{it} \qquad (5) \\
&+ \sum_{\tau=2}^{T} \Big(\beta_\tau + \mu_\tau STATE_{0i}\Big) \cdot I_{[YEAR_{it}=\tau]} + \varepsilon_{it}^{M}
\end{aligned}$$

The set of α_i and β_i coefficients are firm-fixed effects; α_τ and β_τ are time effects faced by all firms. The electricity output is replaced by $REVENUE$ and a second order polynomial in the price-control variables ($\alpha_P' X$) as discussed earlier. The coefficients that capture the price heterogeneity in the two equations are allowed to differ as $\alpha_P' X$ only controls for firm-specific electricity prices, while $\beta_P' X$ additionally controls for coal price differences. This polynomial also absorbed the capital control in the labor equation. We follow Fabrizio et al. (2007) and Du et al. (2008) by relaxing the functional form of the labor equation to allow non-unity coefficients on revenue and wages. To control for any remaining simultaneity between firm-revenue and the error terms we instrument it using lagged revenue.

This is a traditional difference-in-differences set-up. Negative λ_τ and μ_τ coefficient estimates after 2002 would indicate larger efficiency gains for restructured than for control firms, each time relative to their own initial situation. Bertrand, Duflo, and Mullainathan (2004) argue that serial correlation may underestimate the standard error of the treatment effects and hence overstate significance levels in conventional difference-in-differences. To validate the results, we collapse the panel into a pre-reform and post-reform period, which should produce consistent standard errors if the serial correlation does not span long time periods. We also implement the randomized inference approach suggested by Bertrand et al. (2004).

4 Data and summary statistics

The firm-level data we use is collected through annual surveys by China's National Bureau of Statistics. The sample runs from 1998 to 2007 and includes all firms in the fossil fuel-fired electricity generation sector (CIC code 4411) that are either state-owned or non-state firms with sales above 5 million RMB.[18] Because the threshold is far below minimum efficient scale in the sector, in practice the sample includes the universe of generation firms. We observe the usual census variables, such as operational measures of inputs and outputs, as well as location, year of creation, and ownership type.

As mentioned earlier, electricity revenues ($REVENUE$) and intermediate input expenditures ($MATERIAL$) are observed in value terms, not in physical quantities. The latter contains the expenditures on both fuel and non-fuel material inputs. The $WAGE$ variable is defined as the total labor compensation, wage and nonwage expenditures, over total employment ($EMPLOYMENT$).[19]

Four sets of variables are included in the polynomials to control for electricity and coal prices. Firm size is measured by total fixed assets, firm age is calculated from the year of creation, the percentage of capital that is state-owned captures the ownership structure, and provincial dummies capture location and transportation conditions. The uninteracted provincial effect is absorbed by the firm-fixed effects, but the interactions with the other three variables do vary over time.

Table 1 reports for all firms active in 2002 summary statistics for the different firm characteristics, listing the state-owned 'treated' firms separately from the control group. The statistics indicate that average revenue is 17 percent higher for state-owned firms, but the difference is not significantly different from zero. They do employ a lot more workers and use more fixed assets, but only the first difference is significant. The higher employment translates into a much higher wage bill as a fraction of revenue, while both groups of firms spent approximately 73 percent of revenue on fuel and other material inputs. As expected, state-owned firms tend to be a lot older and the average share of their capital that is owned by the state is four times as high.

$\Rightarrow$ Insert Table 1 here $\Leftarrow$

[18]This amounts to approximately $US 600,000 during the sample period.

[19]For the summary statistics, revenues are deflated to 1998 with the ex-factory price index for the electricity industry, inputs with the fuel and energy purchase power index, and the labor remuneration with the consumer price index. The deflators do not influence the estimates as the model contains a full set of time dummies.

5 Results

5.1 Benchmark estimates

The impact of the 2002 reforms on input use is estimated using equations (4) and (5). Panel (a) in Figure 1 plots the coefficients and 95% confidence intervals for the year dummies in the labor demand equation. The blue markers (solid line) are for the treatment firms (state-owned in 2002) and the red markers (dashed line) are for the control group (IPPs). Labor efficiency increases slightly for both groups in the initial years and up until 2004 there is no discernable difference. In later years, the improvement for restructured firms accelerates and a statistically significant gap opens up. This evolution diminishes the employment difference between the two groups that was apparent from the summary statistics. By the end of the sample in 2007, the 0.076 log-points difference implies that labor input in restructured firms has decreased by an additional 6.8 percent compared to the 7.0 percent decline for other firms.

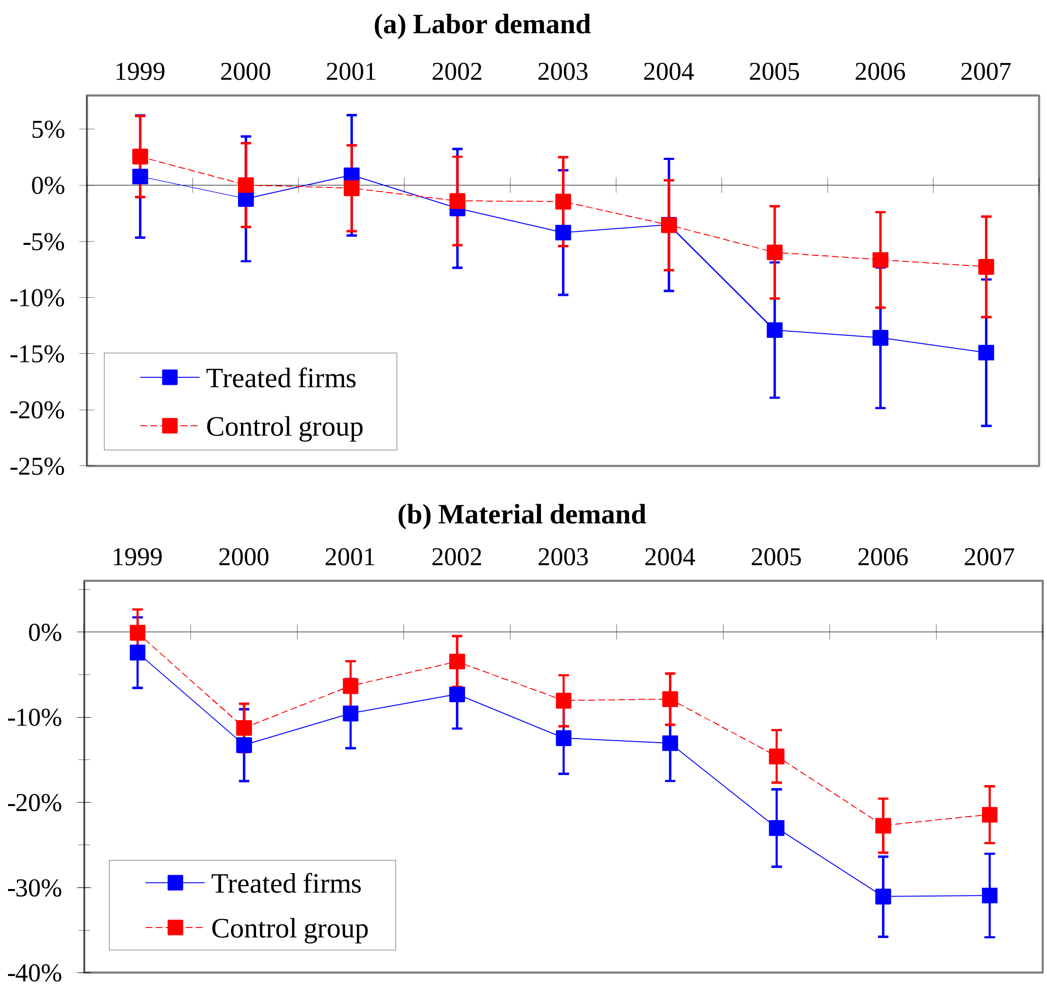
The figure in panel (b) illustrates a similar evolution for material inputs. While the average control firm lowers its material input by 19.3 percent between 1998 and 2007, the average restructured firm lowers it by 26.6 percent, again a difference of 7.3 percent. Because the gap between restructured and other firms now appears a few years earlier and the coefficients are estimated more precisely, the divergence shows up even more clearly.

The results in Figure 1 are estimated with a basic fixed effects regression, but the results are extremely similar using the lagged instrument for output. Full estimates with and without instrumenting are reported in Table A.1 in the Appendix. With the instrumental variables estimation we lose one year of observations and the time variables are measured relative to the 1999 situation. These estimates suggest an even more pronounced difference for labor input, with the point estimate on the difference in 2007 now rising to -0.102 with a t-statistic of 2.92. The differences are somewhat smaller for the reduction in material input, but they are precisely estimated. In each of the last three years, the t-statistic on the group difference is around 2.5.

As argued by Bertrand et al. (2004), serial correlation in the error term can lead to an overestimate on the coefficients in conventional difference-in-differences. The residuals of the material equation passed the Breusch–Godfrey test for autocorrelation, but the labor equation residuals did show mild serial correlation with a one year lag.[20]

[20]The p-values of Breusch–Godfrey test on material input equation are 0.698 for the fixed effects

Figure 1: Evolution of factor demands for firms in the treatment and control groups



One solution is to collapse the data into two periods with a $POST$ dummy indicating the post-reform period. The full set of time-varying restructuring effects $\sum_{\tau=2}^{T} \lambda_\tau STATE_{0i} \cdot I_{[YEAR_{it}=\tau]}$ are replaced with a single term $\lambda_P \cdot STATE_{0i} \cdot POST_t$ that switches on in the post-reform period. This more parsimonious specification makes it also easier to study the restructuring effect under alternative estimation assumptions.[21]

Bergh (1997) suggests that it can take several years for performance effects of a divesture to fully materialize. Most changes in ownership type in our data set occur

model and 0.976 if we instrument. For the employment equation, the corresponding numbers are 0.185 and 0.152.

[21] In subsequent regressions on the full sample, we maintain the full set of time effects α_τ and β_τ to still control flexibly for effects that all firms are subject to, such as changes in the market environment.

between 2002 and 2004. Moreover, it took some time to resolve the uncertainty regarding the proposed reforms of price setting in electricity and coal. As a result, many firms might not have responded to the new industry structure right away. To investigate lagged effects, we estimate the model with two alternative post-reform dummies: *POST2002* and *POST2004*.

⇒ Insert Table 2 here ⇐

Table 2 reports the estimates of the labor and material input equations. Firm-fixed effects are included and the revenue is instrumented to control for potential simultaneity with productivity levels and shocks. An uninteracted post-reform dummy controls for the decline in factor use that is shared by restructured and control firms alike. This combines the effects of general technological change and the adjustment to the new competitive environment by all firms.

The last coefficient in each column—the interaction between the restructuring dummy and the post-reform dummy—measures the efficiency gain that is only experienced by former state-owned firms. Using the two alternative reform periods, they are estimated to experience a reduction of 3.1 or 6.6 percent in labor use and a reduction of 2.5 or 4.3 percent in material input, which includes fuel. The higher effects for both factors in the limited 2004–2007 period are significant at the 1% level, the others at the 5% or 10% level. The other coefficient estimates, positive on revenue in each equation and close to unity in the material equation and negative on the wage rate, are in line with results in Fabrizio et al. (2007) and Du et al. (2008).

5.2 Verifying robustness

To further alleviate concerns about serial correlation of the residuals and the asymptotic approximation used to obtain standard errors, we follow Du et al. (2008) and implement the randomization inference method of Bertrand et al. (2004). This is a group jacknife approach, where we first run the FE-IV regression on half of the firms in the sample. In the following regressions we randomly replace ten of the firms with ten randomly drawn firms from the half of the sample that was not used initially. We can repeat this procedure a 110 times and obtain as many different estimates for the restructuring effect. A similar analysis is performed replacing twenty or fifty firms at a time, which allows 56 and 23 regressions respectively.

In Table 3 we report the average and standard deviation of the coefficient on the $STATE_{0i} \cdot POST2004_t$ interaction term.[22] In the first line we report the benchmark

[22] We do not report any of the other coefficient estimates as they hardly change. (Full results are available upon request.)

estimates from Table 2. The estimates hardly change using the randomized inference, suggesting that autocorrelation is not a serous issue for the two-period approach and that the large sample approximation to the standard errors is accurate. The average point estimate on the restructuring effect is at most reduced by one eight and it always remains significant at the 1% level.

$\Rightarrow$ Insert Table 3 here $\Leftarrow$

The difference-in-differences setup controls for unobservable differences across firms that are constant over time. The summary statistics illustrated that the treated state-owned firms differ from other firms in a number of observable ways, for example they are much older. If we suspect that not only the performance potential of firms, but also the probability of being restructured is related to observable characteristics, we should control for nonrandom selection into treatment. The treatment evaluation literature has developed several approaches, see Imbens and Wooldridge (2009) for an overview, and we implement two estimators.

To make sure that treated firms are benchmarked to comparable control firms, we can adjust for the imbalance in their covariates. We first match each treated firm to its nearest neighbor using the Mahalanobis distance measure and subsequently estimate the factor demand equations on the restricted sample of matched firms. This has the advantage that the regression does not need to fit control firms with covariates that are substantially different from treated firms.

An alternative approach is to first estimate a probit regression of treatment status on the set of covariates using data from the pre-reform period. The propensity score $\hat{p}(X)$, the predicted value from the probit regression, is then used to construct weights for the factor input regressions. In practice, we estimate separate regressions for treated and control firms in the second step. Observations in the first group are weighted by the inverse predicted probability of treatment and in the second group by the inverse probability of non-treatment.[23] The difference in the coefficient of the $POST_t$ dummy in the two regressions is the difference-in-differences estimate of the restructuring impact. Imbens and Wooldridge (2009: 38–40) argue that this estimator achieves double robustness to misspecification of both the regression and the treatment selection models.[24]

[23] Weighting the treated population by the inverse of the probability recovers the expectation of the unconditional response under treatment: $E[STATE_{0i}\Delta Y_i/p(X_i)] = E[\Delta Y_i(\omega = 1)]$, where Y is the performance measure of interest and $\omega = 1$ indicates treatment. Similarly, for the control firms it holds that $E[\{(1 - STATE_{0i})\Delta Y_i\}/\{1 - p(X_i)\}] = E[\Delta Y_i(\omega = 0)]$.

[24] See also Scharfstein, Rotnitzky, and Robins (1999) who argue the benefits of this approach to control for non-randomly missing observations.

The estimates reported at the bottom of Table 3 use these two methodologies to control for observable differences between treatment and control firms that might effect the selection into treatment or the survival in the post-treatment environment. On the restricted sample of matched firms, the second stage regression gives almost identical coefficient estimates on the $STATE_{0i} \cdot POST2004_t$ interaction terms with only marginally elevated standard errors. The double-robust estimation with inverse probability weights again produces similar results. The difference in coefficients on the $POST2004_t$ dummy for the two groups of firms is virtually unchanged in the labor input regression. In the material input regression, the difference is estimated somewhat larger, but the standard error more than doubles.

5.3 Heterogeneity of restructuring effects

Till now we defined treated firm as those identified as state-owned in 2002 in the ownership type classification. A more narrow definition would be to define restructured firms only as those companies originally controlled by SPC, which is the approach followed by Du et al. (2008). Legal entities that are either subsidiaries of SPC prior to 2002 or that are subsidiaries of the Big Five generators created from the SPC breakup in 2002 can be identified in the data set based on their name.[25] This leaves several state-owned firms in the control group, but includes some SPC subsidiaries that already had mixed ownership in 2002 as treated. A third, broader definition of treatment is to include all firms with majority state-ownership of their capital in 2002, which is observed separately of their formal ownership type classification.

⇒ Insert Table 4 here ⇐

The estimates in Table 4 compare the estimated effects of the 2002 reform using all three alternatives to identify restructured firms. The results depend rather strongly on the definition used. While the benchmark definition suggests positive and significant productivity effects of restructuring, the evidence is much weaker using the alternative definitions. For material input the coefficients are still estimated negatively using both alternatives, but for Big 5 subsidiaries the effect is only significant at the 10% level and for firms with majority state-ownership the effect is not significant. For labor input the effect is reduced to one third for firms where the state is majority shareholder, while employment is even estimated to increase at Big 5 subsidiaries.

[25]Some firms with unrelated names are still defined as treated if a majority of their capital is controlled by one of the Big Five firms.

It turns out that the groups of treated firms identified using the three alternative definitions are rather different. Only half of the Big 5 subsidiaries were classified as state-owned in 2002 and only slightly more had the state as majority investor. Several of them had already transformed to shareholder companies before 2002 and some were even listed on the stock exchange. Most of the restructuring of these units might already have taken place before 2002. Implementing the ownership reforms often entailed internal restructuring, the introduction of corporate governance changes, and sometimes foreign partnership.

Big 5 subsidiaries tend to be the larger firms in the industry. While they make up less than one quarter of all firms, they represent more than half of the firms in the top quarter by revenue. Moreover, their employment to revenue ratio is less than two thirds of the sample average, while the same ratio is one half above the average using the other two definitions. It is not entirely surprising that these firms did not reduce their workforce aggressively after 2002. Moreover, while revenue growth after 2002 in restructured firms lags the sample average using all three definitions, the difference is a lot smaller for Big 5 subsidiaries than for former state-owned firms.

A problem with the comparison for different definitions is that the treatment group also changes. In the second panel of Table 4 we report results for the same regressions, but including only firms in the control group that are not treated according to any of the three definitions. The estimate of labor productivity gains is still largest using the first definition, but the differences are much reduced. For Big 5 subsidiaries, material use per unit of output declined by 2.8% more than for control firms which suggests that there were positive restructuring effects for them as well. None of the three material productivity estimates are now significantly different from each other.

The evolution of input use in firms with majority state-ownership suggests improved productivity in the post-reform period, but the point estimates are much smaller than in the benchmark case and never significant. Firms with majority state ownership, but not classified as state-owned are found to behave differently. The inclusion of other owners seems to have already changed these firms' operations prior to the 2002 reform. In spite of a high ownership share, the state seems to act differently in companies that are not formally classified as state-owned.[26]

In light of these different responses to the reforms, we investigate whether restructuring effects are heterogeneous across other dimensions as well. We estimate the benchmark equations on several sub-samples, splitting the original sample in

[26]The vast majority of firms classified as state-owned have the state as majority investor (308 out of 369, or 83.5%). The 139 companies with state-majority ownership but not registered as state-owned firms are driving the differences between the first and third regressions.

two according to size, location, age, and input efficiency.

The small-large and old-young distinction is made by comparing firms with the median in terms of fixed assets and date of creation. The cut-off points are a capital stock of 137.5 million RMB and start-up in 1993. Because firms vary in the number of years they are active, the total number of observations in both groups differ slightly. The division of firms according to their location in the "West" or "East" region correlates with the level of manufacturing development.[27] Finally, firms are divided in two groups according to their initial input efficiency, dividing revenue by employment or by material expenditures and comparing with the sample median.

⇒ Insert Table 5 here ⇐

The estimates of both input equations on each of the sub-samples are reported in Table 5, always using the FE-IV estimator and the $STATE_{0i} \cdot POST2004$ interaction term. The difference in responses along these dimensions mostly have the expected signs, but they are often surprisingly large.

While small firms, i.e. those with a small capital stock, only adjusted by reducing employment, large firms lowered both input uses with the largest adjustment for materials. Layoff costs and union activity are both likely to be higher at larger firms. The employment reduction by small firms, at almost 10%, is remarkably large. The same is true for the reduction in material use in restructured large plants, which is almost twice as high as the estimate for the full sample.

While there is no difference by firm location in the reduction of employment, the reduction in material use is vastly more pronounced in the East. There are several explanations for this pattern. The rapid growth of electricity demand from manufacturing is putting greater pressure on the system to increase production. The greater importance of market prices for coal, as plants are farther from the important coal mining areas in the West, could also play a role. At the same time, the higher level of economic development in the East could boost the deployment of more advanced technologies as well as more experienced management.

Not surprisingly, adjustments by older firms are much larger for both inputs: an employment reduction of 8.4% and reduction in material use by 4.6%. For young firms both input adjustments are estimated negatively, but insignificantly different from zero. These effects indicate convergence between different cohorts of firms and older state-owned firms shedding their historical burden of excessive employment.

More direct evidence of labor productivity convergence is provided in the last two columns of Table 5. Employment is reduced significantly more in restructured firms

[27] Center provinces are included in the West region.

with low initial labor productivity. This effect is notably absent when dividing the sample by material efficiency. The reduction in material input is rather uniform. The less efficient firms also tend to be smaller and older on average, but the correlation is a lot stronger for labor efficiency than for material efficiency.[28]

We also report the coefficients on the revenue variables in Table 5. The remaining variation indicates that there is still further potential for convergence, especially in labor productivity. Along each of the four dimensions, the group of firms with the largest productivity effect of reform also has the highest labor demand elasticity with respect to output. Differences are smaller for material demand, but in the one case that shows a clear difference, small firms are estimated to have a much higher elasticity.

In sum, the boost in labor productivity associated with the 2002 restructuring is concentrated in small, old, and inefficient firms. Material productivity has advanced most in large and Eastern firms, with a somewhat larger than average effect for older firms as well.

6 Current situation

Fossil fuel-fired plants in China, mainly coal-fired units, remain extremely important. At the end of 2010 they accounted for 73.4% of the total installed electricity generation capacity of 962 GW.[29] After demonstrating the important firm-level changes that the 2002 restructuring has caused, we close with a brief evaluation of the remaining discrepancies in efficiency among active firms at the end of our sample period.

To compare productivity levels we need to make two changes to the earlier empirical specification. First, the α_i firm-fixed effects are replaced by random effects that are integrated out using the GLS estimator.[30] As a result, we are no longer identifying the effects solely from changes over time within firms, but are also exploiting variations across firms. A second change is to replace the $\lambda_P \cdot STATE_{0i} \cdot POST_t$ interaction term with a simple $STATE_{0i}$ dummy, consistent with the cross-sectional identification strategy. The controls for unobservables are now less general and inference might not be as reliable, but these changes are necessary to compare productivity *levels* across firms.

[28]Of the firms with labor efficiency below the median, 71% are small and 60% are old. The comparable statistics for material efficiency are 60% and 52%.

[29]Total electric power generated in 2010 was 4,228 million-megawatt-hours. (Both statistics are from China Electricity Council: http://www.cec.org.cn/tongjixinxibu.)

[30]It provides a weighted average of the between and within estimators.

We estimate the modified input demand specifications for three situations and report coefficient estimates in Table 6. The most important results are in the last column. They characterize input use for firms created in the post-reform period—in 2004. To put them in perspective, we report in the second column comparable results in the pre-reform period—for firms created in 1998. We conduct both analyses omitting the first year for new firms as output will only reflect a partial year of operation.

⇒ Insert Table 6 here ⇐

Before turning to the results for new entrants, we report in the first column of Table 6 productivity level comparisons in the post-reform period (2005–2007) for all firms already active in 1998. The coefficients on revenue and wages are remarkably similar to those estimated using fixed effects and reported in Table 2. For firms classified as state-owned both in the first year of the sample (STATE in 1998) and in the reform period (STATE in 2004), the input difference is the sum of the two coefficients on the STATE dummies. These firms are found to still use significantly more workers in the post-reform period and the extra material input use is significant at the 10% level. For state-owned firms that had changed ownership by 2004, the labor input difference is still positive, but only halve as large: 37% instead of 74%.

The performance of firms that entered in 1998 in the pre-reform period is broadly consistent with these findings, but the standard errors are a lot higher on the smaller sample. Their extra labor input use, compared to private firm entrants, is comparable at 68%. Their extra material input use is even more pronounced, at 25%, although the significance level is again only 10%. While the reduction in excessive material input use over time occurred for all firms that were initially state-owned, the reduction in employment did require an ownership change.

The key findings are in the third column. Among new entrants in the post-reform period, state-owned firms do not use systematically more workers or material input than other firms. Both point estimates are positive, but the standard errors are extremely large. New generating capacity added in the post-reform period seems to be operated similarly in the state and private sectors of the economy, which is encouraging given the continued importance of state-owned firms in China. The absolute magnitudes of the coefficients are comparable to the 'STATE in 1998' dummies in the first column for privatized older firms, but the standard errors are a lot larger than in second column in spite of a 50% larger sample. It suggests that there is a lot of variation in input use across new entrants, but the correlation with state ownership has become very weak.

7 Conclusions

We have investigated the impact of regulatory reforms in 2002 on the performance of fossil fuel-fired electricity generation companies in China. The same differences-in-differences method used by Fabrizio et al. (2007) to estimate labor and material input efficiency in the United States underlies the analysis, but we modify the estimation to fit specific features of the Chinese situation.

Rather than using provincial proxies to deal with the missing price problem and output endogeneity, as Du et al. (2008) did, we use fixed effects to control for time invariant unobservables and instrumental variables to control for idiosyncratic firm-specific shocks. Institutional details on the operation of the electricity and the power coal markets is exploited to construct a flexible proxy for missing output and input prices. We investigate the robustness of the results to alternative treatments of econometric issues of simultaneity, serial correlation, and endogenous treatment selection.

Overall, the results strongly indicate a positive impact of deregulation and vertical unbundling on both labor and material input efficiency. We estimate that it did take several years for the effects to materialize, which explains the weaker evidence in Du et al. (2008) who only had data until 2004. The benchmark estimates suggest that in the period starting two years after the publication of the reform blueprint the average firm that was owned by the state in 2002 had reduced employment by 6 percent more than the control firms and material input by 4 percent. The magnitudes of these reform-related input reductions are plausible compared to estimated reductions in factor use experienced by all firms, which was 7.5% for employment and 22% for materials. The reform effects are notably skewed towards employment reductions.

The estimates differ to some extent if we use alternative criteria to define firms that are most directly affected by the reforms. Productivity changes are in the same direction, but less pronounced when they are identified from the broader group of firms with majority state-ownership of equity, rather than from the official state versus non-state categorization. Subsidiaries of the Big 5 generation firms that were created from the breakup of China's Power Company did not experience a comparable decline in employment, but they had already achieved a higher level of labor efficiency in 2002 and managed higher revenue growth than other state firms in the reform period. We also find that improvements in labor productivity are strongly concentrated in smaller and older firms, while material productivity improved most in large and Eastern firms.

We concluded the analysis with some cross-sectional evidence demonstrating that state-owned firms entering the industry in the post-reform period are operating as efficiently as private sector entrants. Moreover, we demonstrate that existing state-owned firms that changed ownership status did manage to reduce the labor and material productivity gap with other firms, while state-owned firms that remained state-owned throughout only shrunk the material productivity gap. Understanding the nature of competition between the new and old firms and between remaining state-owned versus restructured firms is a topic that we leave for future work. In particular, it is an open question what the advantages and disadvantages are of the ongoing addition of new generation capacity under a variety of ownership structures. The continued importance of state-owned firms in this capital-intensive sector combined with strong state control over credit provision raises the importance of validating the last findings.

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Table 1: Summary statistics by ownership category (in 2002)

	State-owned firms (st. dev.)	Control group (st. dev.)	Difference in means (st. error)
Revenue (mil. RMB)	353	303	50
	(1,008)	(636)	(51)
Employment	1149	531	617
	(3107)	(722)	(127)***
Material / Revenue	0.723	0.731	0.008
	(0.233)	(0.207)	(0.014)
Fixed Assets / Revenue	2.371	2.003	0.368
	(5.612)	(4.428)	(0.323)
Wages / Revenue	0.146	0.076	0.101
	(0.291)	(0.081)	(0.012)***
Age	22.63	10.81	11.82
	(16.59)	(9.85)	(0.83)***
State-owned capital (%)	0.828	0.214	0.613
	(0.364)	(0.362)	(0.024)***
Observations	369	654	

Note: ***, **, * denote significance at 1%, 5%, 10% level.

Table 2: Input demand equations in pre- and post-restructuring period

Dependent variable: *(N=8,387)*	ln(EMPLOYMENT)		ln(MATERIAL)	
ln(REVENUE)	0.396	0.384	0.920	0.918
	(0.025)***	(0.024)***	(0.018)***	(0.018)***
ln(WAGE)	-0.263	-0.263		
	(0.009)***	(0.009)***		
$STATE_0$*POST2002	-0.031		-0.025	
	(0.017)*		(0.013)**	
$STATE_0$*POST2004		-0.068		-0.044
		(0.018)***		(0.013)***

Note: IV-FE regression includes controls for price heterogeneity (second order polynomials in fixed assets, age, fraction of state-owned capital, and provincial dummies), firm and year fixed effects. Output is instrumented with lagged output. ***, **, * denotes significance at 1%; 5%; and 10% level.

Table 3: Verifying robustness of the restructuring effect

	ln(EMPLOYMENT)	ln(MATERIAL)
Benchmark coefficient estimate		
on $STATE_0$*POST2004	-0.068	-0.044
	(0.018)***	(0.013)***
Randomized inference		
10 (110)	-0.067	-0.039
	(0.024)***	(0.014)***
20 (56)	-0.069	-0.043
	(0.024)***	(0.015)***
50 (23)	-0.063	-0.041
	(0.026)***	(0.014)***
Mahalanobis matching	-0.066	-0.047
	(0.020)***	(0.014)***
Double-robust estimation	-0.066	-0.057
	(0.034)**	(0.034)*

Note: ***, **, * denotes significance at 1%; 5%; and 10% level.

Table 4: Alternative defintions of restructured firms

	No. of firms in 2002		ln(EMPLOY)	ln(MATERIAL)
	Treated (%)	Control		
(a) Coefficient estimate on $STATE_0$ x POST2004 DUMMY using different definitions:				
Classified as state-owned in 2002 (Benchmark)	353 (34.5%)	670	-0.068 (0.018)***	-0.044 (0.013)***
Big 5 subsidiary	228 (22.3%)	795	0.041 (0.019)**	-0.026 (0.015)*
Majority state-ownership of capital in 2002	447 (43.7%)	576	-0.022 (0.017)	-0.005 (0.013)
(b) Estimates using the same set of control firms for each definition of $STATE_0$:				
Classified as state-owned in 2002 (Benchmark)	353 (44.7%)	436	-0.044 (0.019)**	-0.038 (0.015)***
Big 5 subsidiary	228 (34.3%)	436	0.023 (0.021)	-0.028 (0.016)*
Majority state-ownership of capital in 2002	447 (50.6%)	436	-0.016 (0.018)	-0.011 (0.014)

Note: Coefficients are estimated on separate regressions using the FE-IV estimator. ***, **, * denotes significance at 1%; 5%; and 10% level.

Table 5: Heterogeneous restructuring effects

	Dependent variable is ln(EMPLOYMENT)							
	SIZE		LOCATION		AGE		Labor Efficiency	
	Small	Large	West	East	Old	New	Low	High
ln(REVENUE)	0.428	0.317	0.519	0.326	0.393	0.375	0.506	0.372
	(0.044)***	(0.036)***	(0.047)***	(0.028)***	(0.033)***	(0.039)***	(0.042)***	(0.040)***
ln(WAGE)	-0.308	-0.230	-0.314	-0.226	-0.264	-0.254	-0.205	-0.205
	(0.014)***	(0.012)***	(0.015)***	(0.011)***	(0.013)***	(0.012)***	(0.012)***	(0.011)***
STATE*POST04	-0.103	-0.060	-0.073	-0.070	-0.088	-0.013	-0.060	-0.026
	(0.028)***	(0.024)***	(0.026)***	(0.025)***	(0.024)***	(0.032)	(0.020)***	(0.025)

	Dependent variable is ln(EXPENDITURE)							
	SIZE		LOCATION		AGE		Material Efficiency	
	Small	Large	West	East	Old	New	Low	High
ln(REVENUE)	0.976	0.848	0.932	0.912	0.909	0.920	0.938	0.934
	(0.036)***	(0.024)***	(0.032)***	(0.023)***	(0.024)***	(0.031)***	(0.026)***	(0.027)***
STATE*POST04	-0.004	-0.079	-0.015	-0.070	-0.047	-0.024	-0.044	-0.050
	(0.024)	(0.016)***	(0.0176)	(0.021)***	(0.018)***	(0.026)	(0.019)**	(0.017)***
No. of obs.	4,144	4,243	3,787	4,600	4,284	4,103	3,963	4,424

Note: Coefficients are estimated on separate regressions using the FE-IV estimator. Standard errors in parenthesis. ***, **, * denote significance at 1%; 5%; and 10% level.

Table 6: Productivity differences in the cross-section of firms

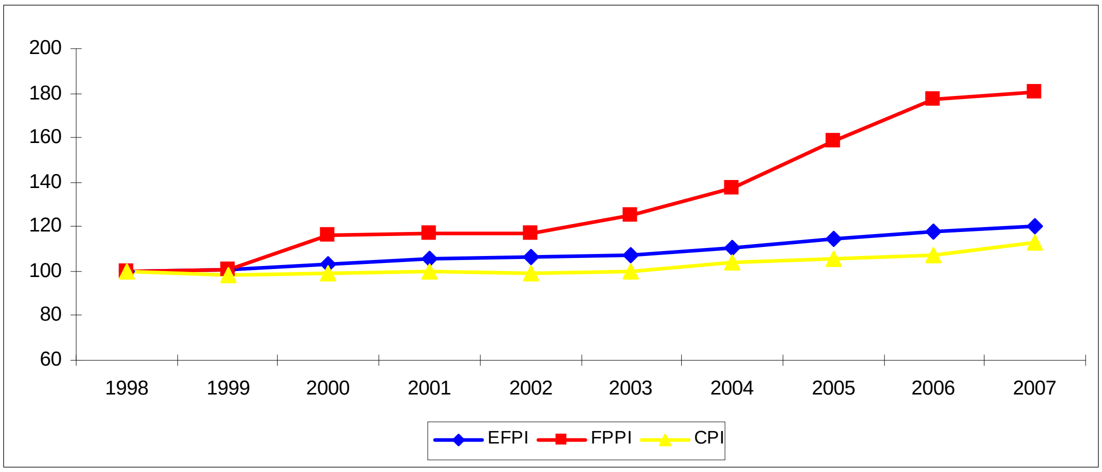
	Dependent variable is ln(EMPLOYMENT)		
Firms created:	prior to 1999	in 1998	in 2004
Observations from:	2005-2007	1999-2001	2005-2007
ln(REVENUE)	0.445	0.401	0.288
	(0.015)***	(0.084)***	(0.054)***
ln(WAGE)	-0.298	-0.196	-0.283
	(0.018)***	(0.048)***	(0.055)***
STATE in 1998	0.314	0.521	
	(0.065)***	(0.314)*	
STATE in 2004	0.240		0.294
	(0.065)***		(0.443)
No. of obs.	2,215	156	236

	Dependent variable is ln(MATERIAL)		
Firms created:	prior to 1999	in 1998	in 2004
Observations from:	2005-2007	1999-2001	2005-2007
ln(REVENUE)	0.967	0.681	0.935
	(0.024)***	(0.036)***	(0.035)***
STATE in 1998	0.044	0.211	
	(0.023)*	(0.124)*	
STATE in 2004	-0.004		0.027
	(0.022)		(0.300)
No. of obs.	2,215	156	236

Note: Coefficients are estimated using separate random effects (GLS) regressions that include year dummies. Standard errors in parenthesis. ***, **, * denotes significance at 1%; 5%; and 10% level.

Appendix

Figure A.1: Price indices for electricity (EFPI), fuel (FPPI), and inflation (CPI)



Source: National Bureau of Statistics, http://www.stats.gov.cn

Table A.1: Input demand equations with time-varying restructuring effects

Dependent Variable	ln(EMPLOYMENT)				ln(EXPENDITURE)			
	FE - basic		FE - IV		FE - basic		FE - IV	
	Obs(10831)		Obs(8387)		Obs(10831)		Obs(8387)	
Independent Variables	Coefficients	t-statistics	Coefficients	t-statistics	Coefficients	t-statistics	Coefficients	t-statistics
ln(REVENUE)	0.2371***	30.23	0.393***	16.02	0.8796***	149.87	0.9181***	49.89
ln(WAGE)	-0.2712***	-35.85	-0.2623***	-29.06				
YEAR1999	0.0256	1.39			-0.0009	-0.06		
YEAR2000	0.0003	0.01	-0.005	-0.28	-0.1125***	-7.78	-0.0907***	-6.65
YEAR2001	-0.0026	-0.13	0.0067	0.36	-0.0632***	-4.26	-0.0555***	-3.9
YEAR2002	-0.0139	-0.69	-0.0322*	-1.7	-0.0346**	-2.27	-0.0175	-1.21
YEAR2003	-0.0145	-0.72	-0.0629***	-3.18	-0.0806***	-5.27	-0.0817***	-5.41
YEAR2004	-0.0355*	-1.74	-0.0948***	-4.65	-0.0788***	-5.14	-0.0714***	-4.57
YEAR2005	-0.0597***	-2.85	-0.1178***	-5.72	-0.1458***	-9.29	-0.1333***	-8.47
YEAR2006	-0.0664***	-3.06	-0.1437***	-6.7	-0.2273***	-14.02	-0.2206***	-13.44
YEAR2007	-0.0726***	-3.18	-0.1422***	-6.3	-0.2144***	-12.62	-0.213***	-12.78
RS*YEAR1999	-0.0178	-0.53			-0.0232	-0.92		
RS*YEAR2000	-0.0123	-0.36	-0.0321	-1.02	-0.0202	-0.78	-0.0205	-0.85
RS*YEAR2001	0.0115	0.35	-0.0348	-1.09	-0.0324	-1.28	-0.0267	-1.1
RS*YEAR2002	-0.0067	-0.2	0	0	-0.0388	-1.53	-0.0232	-0.98
RS*YEAR2003	-0.0276	-0.81	-0.0215	-0.68	-0.0439*	-1.7	-0.0194	-0.81
RS*YEAR2004	0.0003	0.01	-0.004	-0.12	-0.0517*	-1.94	-0.029	-1.16
RS*YEAR2005	-0.0692**	-1.95	-0.0819**	-2.48	-0.0841***	-3.12	-0.0623***	-2.46
RS*YEAR2006	-0.0693*	-1.91	-0.0643*	-1.89	-0.0833***	-3.01	-0.0659***	-2.54
RS*YEAR2007	-0.0763**	-2.04	-0.102***	-2.92	-0.0948***	-3.32	-0.0682***	-2.55

*Note: *** denotes significance at 1%; ** denotes significance at 5%; * denotes significance at 10%*