Identifying the Aggregate Productivity Effects of Entry and Size Restrictions: An Empirical Analysis of License Reform in India*

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December 28, 2009

Abstract

Recent studies have stressed the importance of distortions in resource allocation across heterogeneous establishments in generating large TFP effects. There is however little direct empirical evidence from actual policy experiments on the magnitude of these effects. In this paper we propose and implement a simple methodology that empirically identifies the separate effects of entry and size restrictions on aggregate TFP and apply it to an analysis of the policy of industrial licensing in India. The licensing policy was notably reformed in the mid-1980s, and we take advantage of this policy experiment to identify the effects of the licensing regime using factory-level data covering this period. Our results suggest that the reform resulted in an aggregate TFP improvement of nearly 22% in the deregulated sector, and that the reductions in entry and size restrictions stemming from the reform contributed 25% and 75% to this improvement, respectively.

JEL Codes: L5, O14, O47

1 Introduction

Industrial policy in many poor and developing countries is often characterized by a tension between the desire to encourage industrial development and a desire to control it. The many institutional and policy constraints imposed on the industrial sector in developing countries have become the focus of a very large literature. Earlier work in this area emphasized the role of policies that limited industrial growth in the aggregate, building on representative-firm models. A more recent strand of the literature emphasizes the heterogeneity of units within industries, and examines the role of distortions in resource allocation across these

*I am indebted to Giuseppe Moscarini, Peter Schott and Christopher Udry for many helpful discussions and encouragement. I am also grateful to Costas Arkolakis, Santosh Anagol, Michael Boozer, Simeon Djankov, Amit Khandelwal, Chris Ksoll, Giovanni Maggi, Siddharth Sharma, T. N. Srinivasan and seminar participants at Yale, NEUDC and EIEF, Rome for valuable comments and criticism. I am also grateful to Lori Beaman and John Gray for their advice and help with the data.

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units in generating aggregate efficiency losses. Restuccia and Rogerson (2008) develop and calibrate a theoretical model in which establishments face idiosyncratic prices, showing that the associated productivity losses can be very substantial. Hsieh and Klenow (2009) examine the role of idiosyncratic input and output taxes that distort the allocation of inputs across establishments in China and India, and attempt to identify the magnitude of the distortions. In a similar vein, Bartelsman, Haltiwanger and Scarpetta (2009) conduct a cross-country comparison to identify the extent of misallocation of resources by examining the within-industry correlation between firm-size and productivity.

Drawing on the insights of this more recent literature, this paper attempts to separately estimate the effects of multiple distortions in the manufacturing sector of India. Differently from the papers mentioned above, we focus on developing a method of identifying the effects of a very specific set of distortions that relate to a particular industrial policy, as opposed to the more generic distortions that have been considered so far. Industrial development in India affords a fascinating and fertile case study because outside the communist world, almost no other economy has placed as many and as varied a set of controls on the industrial sector. However, this very complexity of the policy environment makes it extremely difficult to assess the impact of any one control. In addition, policy changes can rarely be assumed to be exogenous. Empirical studies therefore tend to be based on particular reform episodes that can arguably be characterized as exogenous shocks, while restricting attention to some salient aspect of the reform (see for example Topalova 2005, Krishna and Mitra 1998 and Aghion et al 2008). This paper is no exception – we focus attention on the policy reform of 1985, which significantly relaxed size and entry restrictions for a subset of manufacturing industries. The policy in question was the policy of industrial licensing that required firms to obtain government permission before setting up a factory or expanding output in an existing factory. As has been argued elsewhere (see Aghion et al), the reform of 1985 was an unanticipated event, following on the heels of a political assassination.

Inferring the restrictiveness of the licensing policy is challenging because the only available data on factories in India consist of repeated cross-sections, ruling out the possibility of observing growth at the firm level. Even if we confine ourselves to inferences based on the entire distribution of firm sizes, an additional challenge is that licensing policy directly regulated size as well as entry. The removal of the policy can therefore be expected to have increased firm sizes by relaxing size restrictions, but also possibly to have reduced firm sizes due to the increased competition induced by the relaxation of entry constraints, the net effect therefore being theoretically ambiguous.

In this paper, we attempt to identify the effects of entry and size deregulation separately by focusing on a single, robust pattern in the data. Using three years of factory-level data covering a six-year period following the reform we find that in industries that were deregulated, an initial phase of factory expansion accompanied by some exit was followed by a phase of entry. We interpret this as reflecting the operation of short- and long-run effects of the reform. In the short-run, the effect of entry deregulation is likely to have been smaller than the effect of size deregulation (key to this hypothesis is the assumption that the reform was largely unanticipated), so that the short-run increase in average factory size would be due almost entirely to the reduction in size constraints. Comparing the short-run to the pre-reform situation therefore allows us to gauge the extent of reduction in size constraints, and in turn the comparison of long-run and pre-reform data allows for an estimation of the entry effect as a residual.

Because the data are not a panel at the factory level it becomes difficult to obtain credible estimates of firm-level productivity based on estimation of a production function.[1] Instead, we base our identification of aggregate productivity effects on a theoretical model of a competitive industry with heterogeneous firms, in which the license policy is modeled as imposing a cost of entry as well as a tax on output. Matching the model to the data allows us to estimate the percentage reduction in the entry costs and the output tax due to the reform, and the productivity improvements associated with each. To clarify briefly how this method is different from a production function estimation: the approach taken here avoids estimating firm level productivities, and instead uses structural assumptions to generate a mapping between the observed industry aggregates and aggregate productivity. In this broad sense, the method is similar to a growth accounting technique, the difference being that our method is based on an equilibrium model with heterogeneous firms, and therefore imposes more structure on the data. In this particular context, the extra structure helps in identifying the channels of productivity improvement, which is an advantage over both production function and growth accounting methods.

Understanding the effects of size and entry restrictions is quite important from a welfare perspective, but the effects will depend crucially on the nature of costs and the behavior of firms. As Lucas (1978) pointed out, if the technology is characterized by constant returns to scale and firms are perfectly competitive, the size distribution of firms is irrelevant, implying that restricting firm sizes is a harmless policy (as long as entry costs are zero). In more realistic settings, regulating size and entry can have welfare significance. For example, Bertrand and Kramarz (2002) in their study of French retailing find that entry regulation significantly increased market concentration and retailer prices. In the absence of data on plant-level prices, however, estimating the impact of regulations on mark-ups and consumer welfare requires imposing strong assumptions on demand and/or the type of competition. In this paper, we have confined ourselves to a setting in which firms are price-takers and regulations merely impact the allocative efficiency - if the restrictions were removed, the same amount of aggregate output could be produced more efficiently.

The policy reform of 1985 has been studied by others, notably by Aghion et al (2008), who examine how it interacted with differences in labor regulations across the states of India to produce different outcomes. Our paper differs from theirs in two important respects: firstly, in its exclusive focus on the license policy itself and secondly in the results, which being based on more disaggregated data than used by Aghion et al, turn out to indicate the operation of distinct short- and long-run effects, which in turn allows us to separate out the different aspects of the reform. The results of our analysis indicate that the license reform resulted in an aggregate productivity improvement of around 22%, of which 25% and 75% can be attributed to the relaxation of size and entry constraints, respectively. These results are also significant when placed in the context of the turnaround in TFP growth in India that began at around the time of the reform, reported elsewhere in the literature (see,

[1] This is due to the well-known econometric problem of endogeneity of inputs - in particular, the correlation between a firm's inputs and its unobserved productivity will produce inconsistent estimates of the production function (see Marschak and Andrews 1944). Controlling for unobserved productivity with fixed-effects, and/or using further structural assumptions on input choices (see for example, Olley and Pakes 1996 and Levinsohn and Petrin 2003) typically require panel data.

for example, Bosworth, Collins and Virmani 2007). More generally, these numbers are also indicative of the extent to which institutions and policy barriers are responsible for the low levels of TFP in poor countries, a point that is being increasingly stressed in the literature on cross-country income differences (see for example, Hall and Jones 1999 and Parente and Prescott 2000).

The remainder of the paper is organized as follows: Section 2 details the reforms, Section 3 describes the data and presents some key descriptive results that motivate the rest of the analysis. Section 4 builds a theoretical model based on these results, Section 5 outlines the estimation strategy and presents the results and Section 6 concludes.

2 Industrial Reform in India

India is an interesting example of an economy in which government regulations also appear to have shaped the pattern of specialization. Several authors (notably Kochhar et al 2006) have remarked on the peculiar pattern of India's development; although industrialization has been a strong policy emphasis from the time of independence, the particular strategy of industrialization that India adopted emphasized investment in the capital goods sector as a pre-requisite to successful long-term industrialization. In practice, this was achieved by import-substitution and a rigid set of controls that regulated the flow of private investment into industries. This strategy, coupled with the subsidization of tertiary education and the boom in the demand for services in the 1990s, has created the paradox of a poor economy specializing in capital-intensive and skill-intensive sectors. As Kochhar et al point out, the restrictions on entry and capacity creation in the private sector also resulted in relatively small establishments - in 1990, the average manufacturing firm in India was more than 10 times smaller than its counterpart in the US. By 2004, the industrial sector contributed only 28% of total value added in the economy and only accounted for 18% of total employment (Bosworth and Collins 2007). This contrasts with the industrial sector in China, which contributed nearly 60% of total value added while employing about 20% of the workforce.

In the early decades the results of these policies were unspectacular, but not sufficiently alarming to engender doubt in the system of controls. However, a long period of productivity slowdown in the 1970s (see Bosworth, Collins and Virmani 2007) caused policy-makers to rethink the soundness of the regulatory regime. By this time, it was also clear that restrictive regulations in practice had become an anti-competitive tool and an expedient for bureaucratic corruption. In 1985, following the assassination of Indira Gandhi and the accession to power of her son Rajiv Gandhi, the infamous "license-raj" was partially reformed by removing a significant subset of industries from its jurisdiction. At the same time, however, the licensing restrictions on factories in other industries were also eased, as detailed below.

The system of industrial licensing controlled both entry as well as output in Indian industry by requiring each firm to obtain an official license to operate while also stipulating the output the firm was allowed to produce during the period of validity of the license (see Chari 2009). The enforcement of the output license was achieved by controlling the quantity of essential raw materials such as fuel and coal that were allocated to the firm. The official guidelines for issuing new licenses stipulated that new projects were to be rejected or approved on the basis of whether existing capacity in the industry was considered sufficient to meet projected demand. The actual granting of the licenses was subject to the vagaries of the bureaucracy, and since every project required at least a few licenses, the incentive to pursue any investment was severely limited.

The actual coverage of the policy (and its evolution) was extremely complicated: new investments that were smaller than a specified amount were exempt from licensing, unless they constituted additions to an existing factory above a separate size threshold, with these limits evolving continuously over time. In 1982, the exemption limit stood at 50 million rupees and this was further increased across the board to 150 million rupees in 1985. Furthermore, additions to capacity that did not constitute more than 25% of existing capacity were also exempted from licensing requirements, and this limit was also raised across the board to 49% in 1985. The 1985 reform also introduced a program of "re-endorsement of capacity", whereby establishments that were running above 80% capacity utilization were allowed to automatically get a license to increase production by at least a third. License policy also applied differentially to large industrial houses, defined as industrial enterprises whose assets exceeded 200 million rupees. These enterprises came under the ambit of the Monopolies and Restrictive Trade Practices (MRTP) Act, and any existing licensing exemptions did not apply to them. In 1985, the definition of large industrial houses was relaxed across the board, to include only enterprises whose assets exceeded 1000 million rupees.

The second major episode of license reform occurred in 1991, when a host of factors brought the country to the brink of a balance-of-payments crisis. 1991 also saw the assassination of Rajiv Gandhi. The new government, led by Narasimha Rao, sought for and obtained a bailout from the IMF, which, as part of the conditions of the loan, insisted on major economic reforms. A new industrial policy was established that, among other things, virtually abolished the system of licensing (retaining it in only a few industries).

Because the effects of the de-licensing of 1991 are difficult to disentangle from the effects of concurrent reforms on other fronts (notably, a dramatic tariff liberalization), we base this study on the period 1983-1991, which includes only the reform episode of 1985.

3 Data and some key descriptive results

3.1 Data

The factory-level data used in this paper are from the Annual Survey of Industries for the years 1982-83, 1984-85, 1987-88 and 1990-91. Recalling that the reform occurred in 1985, this means that we have two years of data pre-reform and two years post-reform. The Annual Survey of Industries (ASI) is a partial census covering the organized manufacturing sector. The survey samples all factories above a threshold size (this is referred to in the survey as the "census sector") while drawing a representative sample from the rest of the population (referred to as the "sample sector"). The survey's definition of the census and sample sectors was not constant over the years in this study; however, the provided sample weights allow for a valid comparison of industries across years. An important limitation of the data is that panel identifiers are not available to link factories over time (even for the census sector), so that in effect, the data are a set of cross-sections.

The data are at the 4-digit level, as per the National Industrial Classification. Reported revenues are deflated using price deflators at the 3-digit industry level constructed from the official Wholesale Price Indices (WPI) published by the Reserve Bank of India. Labor input is measured by the number of workers as well as the total number of mandays in the year at the plant, while the capital stock is obtained as the average of the opening and closing levels of fixed capital (for the financial year) reported by the plant. The source of data on de-licensing is the Handbook of Industrial Statistics (1987) that lists the industries that were deregulated in 1985. Table 1 (in the appendix) lists the industries that were deregulated in 1985.

For the empirical analysis, we construct a balanced panel of 211 4-digit industries that are present in all four years of the data, of which 41 were deregulated in 1985.

3.2 Key Descriptive Results

As we noted in Section 2, in addition to removing a set of industries from licensing requirements, the license reform in 1985 also significantly relaxed constraints across-the-board. Our focus in the empirical analysis will therefore be to identify the differential effects of license reform on the set of industries that were completely freed from licensing. Hereafter, we will refer to these industries as 'deregulated' and the remaining industries as 'underegulated', with the understanding that this is not a literal description of the reform.

Table 2 shows that in 1983 (prior to the reform), the industries that were chosen for deregulation constituted around 16% of total employment and fixed capital and 24% of total output in the formal manufacturing sector. In terms of industry-level characteristics, such as aggregate output, employment, capital intensity and plant size (output per plant), the deregulated industries were not significantly different from the remaining set of industries, although the former had a significantly greater number of establishments. As we will see shortly, the similarity in pre-reform levels also carries over when we look at pre-reform trends in the same variables. This matches well with a similar finding by Aghion et al. This gives us some measure of confidence that post-reform differences between these two sets of industries can be attributed to the reform itself..

We now turn to the key descriptive results that motivate the identification strategy used in this paper. We attempt to identify the effects of the reform on key variables, netting out the effect of other confounding factors. The regression model may be written as:

$$\log y_{it} = \alpha + \beta_t + \gamma_t.1[Dereg_i = 1] + \delta_i + \varepsilon_{it}$$

where y_{it} represents the outcome of interest for industry i at time t; $Dereg_i$ is an indicator for whether industry i was deregulated, β_t captures a year fixed-effect, γ_t captures a differential year-effect for deregulated industries, δ_i captures an industry-specific effect and ε_{it} represents an industry-time specific shock. The industry-level fixed-effects are intended to control for unobservable (fixed) industry characteristics that may be correlated with deregulation. Because it will prove to be more convenient to interpret, we implement this specification in first-differences, i.e. we will look at changes in y over time and relate these changes to deregulation.

The effects of the reform on industry-level outcomes are therefore obtained from a regression of the following form:

$$\begin{aligned}\Delta \log y_{it} &= \beta_1.Dereg_i.Year1985 + \beta_2.Dereg_i.Yr1988 + \beta_3.Dereg_i.Yr1991 \\ &\quad +\gamma_1.Yr1985 + \gamma_2.Yr1988 + \gamma_3.Yr1991 + \epsilon_{it}\end{aligned}$$

where Δ represents a one-period difference operator, $Yr1985$, $Yr1988$ and $Yr1988$ represent dummies for the years 1985, 1988 and 1991 (because the regression is in differences, the first period in the data, 1983, is dropped). In this specification, β_1 captures the differential change in outcome (relative to control industries) for deregulated industries for the period 1983-1985 (pre-reform); β_2 and β_3 capture the differential change in outcome for deregulated industries for the periods 1985-1988 and 1988-1991, respectively. To assess the robustness of the results, we also report the results including 3-digit industry fixed-effects (corresponding to time-trends at this level in the non-differenced data). Finally, to adjust the standard errors for common shocks across industries in each period, we cluster the standard errors at the 2-digit industry-year level.

Table 3 shows the results of running this regression using industry-level number of factories, total output, employment (in terms of workers as well as mandays) and capital respectively. In Columns I, III, V, VII and IX we show the results from the base specification, while Columns II, IV, VI, VIII and X show the results once we also control for 2-digit industry fixed effects (this will capture the effect of industry-specific time-trends in y). In both specifications we cluster the errors at the 2-digit industry-year level, in order to allow for correlated demand/cost shocks at this level.

The pattern is interesting: Prior to the reform, both sets of industries appear to be declining slightly, in terms of number of factories, output, employment and capital, and there is little evidence of any difference in pre-reform trends for the two groups. Both groups of industries experience a significant amount of exit and expansion in aggregate output, employment and capital in the immediate post-reform period 1985-88, with the deregulated industries experiencing a greater expansion, as can be seen from the coefficient β_2 in these regressions. In the following period 1988-1991 the underegulated industries begin to decline again in output, employment and capital and show little entry or exit; in contrast the deregulated industries experience a large amount of entry and a small amount of expansion in aggregate terms (but a significantly large amount in relative terms).

These results appear (for the most part) to be consistent with the details of the policy reform. The across-the-board easing of licensing constraints appears to have resulted in a rationalization of all industries, with the expansion of productive factories and the exit of less productive ones. The abolition of constraints on size in the deregulated industries is evident in the differential expansion of these industries in the period 1985-88. But the considerable subsequent amount of entry and expansion (the latter in relative terms only) for the deregulated industries is indicative of the easing of entry constraints for these industries, and strongly suggests the operation of short- and long-run effects of license reform. In particular, it seems plausible that the short-run response primarily reflects the effect of easing size constraints, with the entry margin only responding with a lag. In turn, this suggests a strategy for separately estimating the size and entry constraints inherent in the licensing policy.

In the next section, we develop a simple estimable model that will allow us to assess the extent of reduction in the output and entry constraints, and the associated aggregate productivity improvements.

4 Theoretical Model

The goal of this section is to develop a model that can (a) reproduce the observed pattern of response in the data and (b) be estimated using the data at hand. Because we cannot infer firm dynamics in the data, we have focused on drawing inferences from changes in aggregate quantities. To relate these observed dynamics theoretically to changes in size and entry constraints, our model must be amenable to easy aggregation. To get around the difficulties of constructing a dynamic model with heterogeneous firms that exhibits simple aggregate dynamics, we instead construct a quasi-dynamic model along the lines of Melitz and Ottaviano (2008). In such a model, the equilibrium is static, but the comparative statics can be artificially decomposed into short-run and long-run elements. As we will see, this model is estimable in a fairly transparent way.

4.1 Setup

We first describe the demand side of the model. There are M manufacturing sector goods (industries), each of which is homogeneous. The inverse demand function for good i is assumed linear with no cross-price effects:

$$p_i = a_i - b_i q_i$$

where p_i is the price of the i-th good relative to the wage (the numeraire in this model).[2]

The manufacturing sector uses both labor and capital in production. Input markets are assumed to be perfectly competitive. Furthermore, all inputs are assumed to be in perfectly elastic supply, an admittedly extreme assumption for inputs other than labor in India, but one that considerably simplifies the analysis by allowing us to abstract away from general equilibrium effects of the reform working through changes in factor prices. However, this assumption is stronger than required: as long as the industries that were de-licensed had the same capital-intensity as the remaining industries (and this turns out to be true in the data), the estimation will still capture the relative productivity effects of the reform. This is established more formally in Appendix A.

We describe below the production side and the equilibrium conditions that obtain for a representative industry. This simplifies the notation by allowing us to drop industry subscripts.

Firms in the representative industry produce a homogeneous good and are price-takers. There is an unbounded pool of potential entrants. These potential entrants are ex-ante identical; however, they learn their respective (constant) marginal costs of production once they have paid a (sunk) cost of entry equivalent to f units of labor, which corresponds to the cost of obtaining a license to enter an industry. This entry cost may be thought of as

[2] Because we will estimate TFP effects instead of welfare effects, we omit writing out the representative consumer's utility function and maximization problem.

including bribe costs, time costs of applying for a license etc. The marginal cost x (measured in units of labor) is assumed to be a random drawing from a distribution with cdf $G(.)$ and a finite mean. There are no fixed costs of production, so that a firm's marginal cost may be thought of as the inverse of its level of TFP (we make this explicit shortly). The assumption that inputs are in perfectly elastic supply implies that the distribution of marginal costs, $G(.)$, is stable.[3]

After paying the cost of entry and learning its marginal cost, x, an entrant can decide whether to stay and produce or to exit. If it chooses to produce, it must also obtain an output license to produce its desired level of output. We model the cost of obtaining an output license as being paid in units of labor and being a quadratic function of output. The total marginal cost for a firm that has drawn a technological marginal cost of production x is therefore given by:

$$c = x + \theta y$$

where θy represents the cost of obtaining a license to produce an additional unit of output, given the current level of output, y.

Regarding the specific parametrization chosen, the quadratic representation of output tax is a better approximation to reality than a linear tax in this context, since the actual constraints imposed by the license policy were highly non-linear (as explained in Section 2) - large firms were subject to significantly greater restrictions when it came to expanding output. From a purely modeling perspective, the assumption of a quadratic (as opposed to linear) non-production cost is required in this model to obtain a non-degenerate firm-size distribution.[4] An alternative modeling strategy would be to rely on decreasing returns to scale in production to obtain a non-degenerate size distribution (as in Lucas 1978), while modeling the license cost as a linear output tax. A different issue regarding our specification of output tax is that in reality the tax was asymmetric - the license policy did not impose any costs on output reductions. However, because asymmetric costs cannot be introduced in a static equilibrium model without specifying arbitrary initial conditions, we prefer to stick with the symmetric cost specification. We can only caution that this point be kept in mind when interpreting the results.

Each firm maximizes its profit given the price, p, and this determines its output, revenue and profit:

$$\text{Output: } y(x) = \frac{p-x}{\theta} \tag{4.1}$$

$$\text{Revenue: } r(x) = p\left(\frac{p-x}{\theta}\right) \tag{4.2}$$

$$\text{Profit: } \pi(x) = \frac{(p-x)^2}{2\theta} \tag{4.3}$$

For any given price, p, there exists a cutoff (technological) marginal cost, x^*, such that an entrant who has drawn x^* will be indifferent between entering and staying out. Since the cost of obtaining an output license is a continuous function of output, it follows that $x^* = p$

[3] The model can therefore equivalently be rewritten so that firms draw a level of TFP, instead of a marginal cost.

[4] This is because a linear non-production cost combined with linear production costs would result in the most productive firm taking over the market.

and that this marginal firm produces zero output, i.e. $y(x^*) = 0$. We can therefore rewrite Eqns (4.1), (4.2) and (4.3) in terms of this cutoff cost:

$$y(x) = \frac{x^* - x}{\theta}$$

$$r(x) = x^*(\frac{x^* - x}{\theta})$$

$$\pi(x) = \frac{(x^* - x)^2}{2\theta}$$

4.2 Solving for the equilibrium

There is entry into the industry until the cost of entry equals the ex-ante expected profit. The cutoff cost, x^*, and hence the equilibrium price, are determined by the free entry condition:

$$\int_0^{x^*} \pi(x)dG(x) = f \tag{4.4}$$

This also pins down average output, revenue and profit in the industry. The cutoff marginal cost, x^*, implies a cutoff productivity, i.e. the productivity of the least efficient firm that survives. Specifically, under the maintained assumption of constant returns to scale technology, the cutoff productivity, denoted by ϕ^*, can be written as:

$$\phi^* = \frac{\Psi}{x^*}$$

where Ψ is a function of the input price vector (which is fixed under the assumption that inputs are in perfectly elastic supply) and the coefficients of the inputs in the production function. Denoting the prior distribution of productivities by $H(.)$,we can define the (unweighted) mean productivity of the industry:

$$\bar{\phi} = \int_{\phi^*}^{\infty} \phi d\tilde{H}(\phi)$$

where $\tilde{H}(.)$ denotes the distribution of productivities of the surviving firms. We can also define an output-weighted productivity index for the industry:

$$I_p = \frac{1}{\bar{y}} \int_{\phi^*}^{\infty} \phi y(\phi) d\tilde{H}(\phi)$$

Finally, the mass of firms, N, is determined by the equality of supply and demand:

$$p = x^* = a - bN \int_0^{x^*} y(x) d\tilde{G}(x)$$

where $\tilde{G}(x)$ is the conditional cost distribution of surviving firms. To solve explicitly for these equilibrium quantities, we assume, as in Melitz and Ottaviano (2005), that the prior distribution $G(.)$ of cost draws has the following cdf:

$$G(x) = (\frac{x}{x_m})^k$$

where the support of the distribution is $[0, x_m]$ and the parameter k is assumed to be greater than 1.[5] This is equivalent to assuming that the distribution of TFP (the inverse of the marginal cost, x), whose cdf we have called $H(.)$, is that of a Pareto-distributed random variable. The Pareto distribution is analytically convenient, while also being an empirically good approximation to productivity distributions. A useful property of the $G(.)$ distribution is that it preserves its form after a right-truncation. This implies that the conditional distribution of surviving firms inherits a cdf of the same form:

$$\tilde{G}(x) = (\frac{x}{x^*})^k$$

We now solve for x^* from Eqn (4.4):

$$x^* = [(k+1)(k+2)x_m^k \theta f]^{\frac{1}{k+2}}$$

Industry level averages of output, revenue and profit are therefore given by:

$$\bar{y} = \frac{x^*}{\theta(k+1)}$$

$$\bar{r} = \frac{(x^*)^2}{\theta(k+1)}$$

$$\bar{\pi} = \frac{(x^*)^2}{\theta(k+1)(k+2)}$$

The unweighted and output-weighted productivity measures can be shown to be linear in ϕ^*:

$$\bar{\phi} = \frac{k}{k-1}\phi^*$$

$$I_p = \frac{k-1}{k+1}\phi^*$$

[5] This is required for the model to have an equilibrium, essentially because expected profits are unbounded if k is less than 1.

We may also be interested in obtaining a measure of industry productivity that incorporates the resources used up in paying license costs. To this end, we define an index of average industry cost that includes the output license tax:

$$I_p^L = \frac{1}{\bar{y}} \int_0^{x^*} (x + \frac{\theta}{2} y) y d\tilde{G}(x)$$

We can show that this index is also related in a linear way to the cutoff x^*:

$$I_p^L = \frac{k+1}{k+2} x^*$$

This is convenient because it implies that aggregate productivity moves one-for-one with the cutoff productivity, ϕ^*, and hence the cutoff marginal cost, x^*, so that it is sufficient to identify changes in the latter in order to estimate productivity effects. It is also easily seen that the cutoff (and hence the average) productivity increases in θ and f. Intuitively, larger entry barriers insulate incumbent firms from competition and allow unproductive firms to survive, while limits to expansion prevent the more productive firms from expanding and driving out the inefficient ones. We expect therefore that a reform that reduces θ and f should improve productivity and raise total output.

The equilibrium number of firms and total output are given by:

$$N = \frac{\theta(k+1)}{b} \frac{a - x^*}{x^*}$$

$$Q = \frac{(a - x^*)}{b}$$

How do the number of firms and average industry output respond to changes in θ and f? Propositions 1 and 2 below state the relevant effects:

Proposition 1: Average output, revenue and profit per firm ($\bar{y}, \bar{r}$ and $\bar{\pi}$ respectively) are increasing in f and decreasing in θ.

Proposition 2: The number of firms N is decreasing in f and increasing in θ.

These results are in line with intuition: a reduction in entry costs encourages entry and thereby results in a larger number of firms and smaller firm sizes, whereas a reduction in the output license cost[6] allows more productive firms to expand and thereby increases firm size while reducing the number of firms required to serve the market. The long-run effect of de-licensing can therefore be summarized as follows: (a) total output and average productivity rise, (b) the effect on average output, revenue, profit and the number of firms is ambiguous, since the changes in entry and output license costs have opposite effects on these quantities.

[6]Note that the resource cost of obtaining an output license for the average firm is given by $\frac{\theta}{2}\bar{y}^2$. *Holding fixed this average firm size*, the change in the license cost due to the reform is given by the change in θ, so that hereafter we will refer to the latter as the change in the output license cost.

4.3 Deriving the short-run response

We follow Melitz and Ottaviano (2008) in introducing the concept of a short-run equilibrium in this framework. The short-run is defined to be a time-frame in which there is no entry. There is a fixed set of incumbent firms which react to the policy shock by expanding, contracting or shutting down.

To derive the short-run response to the policy shock, we reason as follows. Suppose the industry is initially in the pre-reform steady-state: the cutoff marginal cost is x^* and the number of firms operating is N^*. Denote the distribution of productivities of these firms by $\bar{G}(.)$. Note that the support of this distribution is $[0, x^*]$. Suppose now that there is a policy shock that reduces θ from θ_0 to θ_1 and f from f_0 to f_1. Since there is no entry in the short run, the change in the entry cost can have no short-run impact. However, the change in θ changes the optimal quantities for the incumbent firms. The fall in θ allows everyone to produce more, but the resulting fall in prices must cause some of the less-productive firms to suspend operation.

A new short-run cutoff cost obtains due to the exit of some relatively unproductive firms: we denote this by x^S. The new number of firms, denoted by N^S, is the set of incumbents whose cost is less than x_s:

$$\begin{aligned} N^S &= N^* \bar{G}(x^S) \\ &= N^* (\frac{x^S}{x^*})^k \end{aligned} \tag{4.5}$$

This relation, together with the market-clearing condition, determines x^S and hence N^S in terms of x^* and N^*. Proposition 3 verifies that in the short-run, average output increases in response to the policy shock.

Proposition 3: If $\theta_1 < \theta_0$, then $\bar{y}^S > \bar{y}^*$

Note that we can still write the average output in this short-run equilibrium in terms of the new cutoff cost:

$$\bar{y}^S = \frac{x^S}{\theta_1(k+1)}$$

Hence,

$$\frac{\bar{y}^S}{\bar{y}^*} = \frac{x^S}{x^*}\frac{\theta_0}{\theta_1} = (\frac{N^S}{N^*})^{1/k}\frac{\theta_0}{\theta_1} \tag{4.6}$$

This equation relates the short-run change in average establishment size to an unanticipated change in θ and f. The interpretation of this relation is that the change in average establishment output is due to the direct effect of the change in expansion costs (θ) as well as the indirect effect of this change in θ on the equilibrium market price, the latter being proxied by the change in the number of establishments. We will make use of this equation in the estimation.

4.4 Moving to the long-run equilibrium

We now derive the change in average output when we move from the short-run to the long-run equilibrium. Super-scripting all post-reform long-run equilibrium values by L, we can write:

$$\bar{y}^L = \frac{x^L}{\theta_1(k+1)}$$

Further, we know that

$$\frac{\bar{y}^L}{\bar{y}^*} = \frac{x^L}{x^*}\frac{\theta_0}{\theta_1} = [\frac{\theta_1}{\theta_0}\frac{f_1}{f_0}]^{\frac{1}{k+2}}(\frac{\theta_0}{\theta_1}) \tag{4.7}$$

where the last equality follows from the fact that $x_L = [(k+1)(k+2)x_m^k\theta_1 f_1]^{\frac{1}{k+2}}$. We will use this equation in conjunction with Eqn (4.6) to obtain an estimate of the change in entry costs.

Finally, Proposition 4 verifies that average output declines and the number of firms increases in going from the short-run to the long-run equilibrium.

Proposition 4: If $\theta_1 < \theta_0$ and $f_1 < f_0$, then $\bar{y}^L < \bar{y}^S$ and $N^L > N^S$.

The model therefore qualitatively reproduces the dynamics of establishment size observed in the data. In the next section, we use the relations implied by the model to obtain estimates of the changes in the entry and the output costs, and thereby, a decomposition of the productivity effects of the reform.

5 Empirical Specification and Results

We now turn to the exercise of using the equations of the model to obtain the implied TFP gain and its decomposition into entry cost and output tax effects. The estimation of these effects hinges on using Eqns (4.6) and (4.7), reproduced below with industry subscripts:

$$\frac{\bar{y}_i^S}{\bar{y}_i^*} = \frac{x_i^S}{x_i^*}\frac{\theta_0}{\theta_1} = (\frac{N_i^S}{N_i^*})^{1/k}\frac{\theta_0}{\theta_1} \tag{5.1}$$

$$\frac{\bar{y}_i^L}{\bar{y}_i^S} = [\frac{\theta_1}{\theta_0}\frac{f_1}{f_0}]^{\frac{1}{k+2}}(\frac{\theta_0}{\theta_1}) \tag{5.2}$$

where we will assume for the purposes of estimation that the changes in θ and f are common to all deregulated industries. Because the reform also eased size constraints for factories in other industries, we will attempt to identify the differential change in θ for the deregulated industries. We will also allow for the value of k to differ across the two sets of industries.

We first consider Eqn (5.1). In principle, we could implement this equation as a regression using only the deregulated industries, in which case the intercept term would be expected to pick up the change in θ. To ensure that the estimated intercept is not picking up an economy-wide trend in productivity or an economy-wide change in θ, we instead include all the industries and implement a flexible version of this equation as a simple regression using data on the short-run period 1985-1988:

$$\log(\frac{\bar{y}_i^{1988}}{\bar{y}_i^{1985}}) = \alpha + \beta.Dereg_i + \gamma_1\log(\frac{N_i^{1988}}{N_i^{1985}}) + \gamma_2\log(\frac{N_i^{1988}}{N_i^{1985}}).Dereg_i + \xi_i \tag{5.3}$$

This specification allows the coefficient on the change in the number of plants to be different for deregulated and control industries as would be the case if there was no change in price for the underegulated industries or if the value of k differs across the two sets of industries. The coefficient β in this regression will pick up the differential change in θ for the deregulated industries. Additionally, we note that the coefficients γ_1 and γ_2 must sum to $1/k_D$ where k_D is the value of k for the deregulated industries. Finally, we include fixed effects for the 2-digit industry that industry i belongs to, in order to pick up any common trends at this level.

Turning to Eqn (5.2), we can cast as this as an estimable equation as follows:

$$\log(\frac{\bar{y}_i^{1991}}{\bar{y}_i^{1988}}) = \alpha' + \beta'.Dereg_i + \nu_i \tag{5.4}$$

where we once again include the control group in order to pick up any overall change in f and θ and 2-digit industry fixed effects to capture common trends.[7] This time, the regression only uses the long-run period corresponding to the two years 1985 and 1991. In this regression, the coefficient β' will capture the term on the RHS of Eqn (5.2), which involves changes in both f and θ. To obtain a measure of the change in f alone, we must combine the coefficient β' with the coefficient β from the first regression. Specifically, it is easy to see that:

$$\beta' = (\frac{f_1}{f_0})^{\frac{1}{k+2}} \beta^{\frac{k+1}{k+2}} \tag{5.5}$$

which we can use to calculate f_1/f_0 (using the estimated value of k from the previous regression.

Table 4 shows the coefficient estimates from the two regressions above, in columns 1 and 2 respectively. The coefficient on the indicator for deregulated industries in the regression in Eqn (5.3) indicates a relative change in θ of roughly 22%. Although γ_1 is small and insignificant, γ_2 is large, reflecting a strong association between size changes and exit in the deregulated industries. The implied value of k for these industries is about 2.2, which is fairly close to the average k value of 1.8 reported in Del Gatto, Mion and Ottaviano (2006).

Column 2 reports the results of estimating the regression in Eqn (5.4). Relative to the control industries, there is little to no differential change in size over this long period, indicating that the earlier increase in average factory size was largely undone by the subsequent entry of new firms. Using the relation in Eqn (5.5), we estimate the implied change in entry costs for the deregulated industries to be around 70% (specifically, taking the overall change in size from 1985-1991 to be exactly zero gives an estimate of 70.4%).

Taking these estimates together, we can say something about the aggregate productivity improvement due to the license reform. We first recall that this aggregate improvement is equivalent to the change in cutoff cost x^*, which can be related to the changes in θ and f as follows:

$$\frac{x^L}{x^*} = (\frac{\theta_1}{\theta_2}\frac{f_1}{f_2})^{\frac{1}{k+2}}$$

[7] Because this regression uses only one time period 1985-88, we cannot include 4-digit fixed effects. We also avoid including 3-digit fixed effects in this regression, because many 3-digit industries in the data comprise only one 4-digit industry, so that the amount of variation absorbed by 3-digit effects is extremely high.

Doing the calculation yields an aggregate TFP improvement (over the entire period) of 21.5%, of which 25% and 75% can be attributed to the easing of size and entry constraints respectively, indicating a 5-6% TFP improvement associated with relaxing size restrictions and a 15-16% TFP improvement associated with the lowering of entry costs.

The estimates obtained above reflect the TFP improvement due to a reallocation of resources from low-TFP to high-TFP producers. Because these estimates are based on the structural relations between distortionary costs, average firm-size and aggregate productivity, they necessarily constitute somewhat indirect evidence of reallocation. We now attempt to find some direct evidence of resource reallocation following the reform. To do so, we employ the decomposition of industry-level productivity suggested by Olley and Pakes (1996).[8] The essential idea is to examine the correlation between size and productivity in the data - if the reform induces some productivity-enhancing reallocation, then this should show up as an increase in this correlation. The method is based on the following algebraic identity:

$$p_{it} = \bar{p}_{it} + \sum_{j \in i} (s_{ijt} - \bar{s}_{it})(p_{ijt} - \bar{p}_{it}) \tag{5.6}$$

where p_{it} is a share-weighted average of (either labor or multi-factor) plant-level productivity, with the aggregation being over plants in industry i; s_{ijt} is the plant-level weight (usually either output or employment shares of plant j in imndustry i); $\bar{p}_{it}$ and $\bar{s}_{it}$ are the simple averages of p_{ijt} and s_{ijt}. The second term on the right-hand side above captures the covariance between size and productivity within the industry.[9]

Because we do not have suitable deflators for intermediate inputs nor a plant-level panel, we implement the Olley-Pakes decomposition using the logarithm of plant-level revenue labor productivity (instead of calculating plant-level TFP), with employment-shares being used as weights. Our measure of labor productivity is real output per worker (the results using mandays and value-added are very similar).[10] We perform the decomposition separately for each 4-digit industry in the data. Table 5 summarizes the results of the decomposition, reporting the value of each term in the decomposition for each of the two sets of industries, for the years 1985-1991. As in Bartelsman, Haltiwanger and Scarpetta (2009), we use industry output shares in 1991 (because we think 1991 is likely to represent the least distorted sample observation in our data) as weights to aggregate these measures for each set of industries.

The decomposition reveals that the covariance between size and productivity in deregulated industries registered a large increase over the period 1985-1988, but did not change

[8] The Olley-Pakes decomposition is a static one, and starting with Baily, Hulten and Campbell (1992) various dynamic versions of it have been proposed to identify channels of productivity improvement. These dynamic versions typically distinguish between the contributions to aggregate productivity of surviving, entering and exiting firms. Because our data are not a panel, we are unable to implement such a decomposition.

[9] Algebraically, the O-P decomposition represents a re-arrangement of a familiar identity from statistics: the sample covariance between two variables can be written as the average of the product minus the product of the averages.

[10] We view the results using this measure of productivity as merely indicative, but we note that under certain conditions, this measure of productivity is indeed closely correlated with TFP. For example, if there are no plant-level distortions or frictions, both input and output markets are perfectly competitive and the technology exhibits constant returns to scale, then the ratio of labor productivities between two plants within an industry is simply the ratio of their TFP levels.

significantly thereafter. In contrast, the covariance term registered a small and steady rate of growth for the underegulated industries over the entire period 1985-1991. At the end of the sample period, the covariance term stood at 0.26 for the underegulated industries and at 1.03 for the deregulated industries. The decomposition also indicates that for deregulated industries, the increase in industry productivity over the period 1985-88 was largely due to a reallocation of labor across plants, while the increase in productivity over the later period 1988-91 was due primarily to an increase in productivity of the average plant.

It may appear surprising that the large amount of observed net entry in the period 1988-1991 for the deregulated industries (see Section 3.2) does not appear to have resulted in any significant increase in the covariance between employment shares and labor productivity. However, this finding is in fact not inconsistent with our previous results: of the two aspects of the reform, namely reduction in entry costs and reduction in size restrictions, only the latter directly distorts the allocation of resources across surviving firms. Corresponding to the reduction in size restrictions we do indeed observe an increase in the covariance term over the period 1985-88. If we believe that the net entry over the subsequent period reflects the reduction in entry barriers, the associated reallocation would imply an increase in average productivity (which is what we observe) but would not necessarily register in the covariance term of the O-P decomposition. This point is readily seen when we consider the relation between size (in terms of output) and TFP in our theoretical model. The output of a firm with cost x is given by:

$$y(x) = \frac{x^*}{\theta} - \frac{x}{\theta}$$

There is thus a linear relationship between output and x, the strength of which is determined only by the extent of the size-restriction θ. Changes in the cost of entry only impact x^*, without changing the correlation between size and productivity. In short, the results using the decomposition of aggregate productivity are exactly consistent with our interpretation of the descriptive results in Section 3: the initial post-reform effect appears to reflect the reallocation due to the removal of size constraints, while the subsequent effect appears to be due to a relaxation of entry restrictions.

Overall, our findings suggest that license policy had created and preserved a fairly distorted allocation of resources. The point estimate of the reallocational TFP effect of size-restrictions in our paper (about 5%) is however well within the range of potential TFP improvements that emerge in the simulations by Restuccia and Rogerson (2008). Their simulations consider the effects of idiosyncratic output taxes at the plant-level: if the taxes are uncorrelated with firm productivity (the best-case scenario), the TFP reduction achieved by a 10% tax on ouput randomly applied to 80% of the firms (and balanced by subsidies to the remaining firms) will generate an aggregate TFP reduction of 5%. When the taxes are correlated with firm productivity, a case which is closer to our setting, the same experiment yields a TFP reduction of as much as 16%.

Our results on the TFP effects of entry costs are surprisingly large (around 15-16%). While the prevalence of entry costs around the world is well documented (see for example, Djankov et al 2002), credible estimates of the TFP effects of entry costs are hard to come by in the literature. We may mention Barseghyan and DiCecio (2009), who calibrate a theoretical model to US data and using actual measures of entry cost across countries, estimate that a one *percentage point* reduction in entry cost (with entry cost being calculated as a fraction of per capita GDP) implies a 0.14 percent increase in aggregate TFP. Although this estimate is not directly comparable with ours, it is notable for its magnitude - entry costs can clearly have very dramatic effects, a point that also appears to emerge from our analysis.

6 Concluding Remarks

Drawing on the insights of the recent literature on aggregate productivity improvements via resource reallocations between firms, this paper attempts to estimate the aggregate productivity losses due to size and entry regulations in the manufacturing sector of India. We focus attention on the policy reform of 1985 in India, which significantly relaxed size and entry restrictions for a subset of manufacturing industries.

Using three years of factory-level data covering a six-year period following the reform we find that in industries that were deregulated, an initial phase of factory expansion accompanied by significant exit was followed by a phase of entry. We interpret this as reflecting the operation of short- and long-run effects of the reform. In the short-run, the effect of entry deregulation is likely to have been smaller than the effect of size deregulation, so that the short-run increase in average factory size would be due almost entirely to the reduction in size constraints. We construct a heterogeneous-firm model of industry equilibrium in which we model the restrictions imposed by the license policy and derive the short- and long-run effects of relaxing these restrictions. The model yields estimable equations that allows us to gauge the extent of reduction in size constraints from a comparison the short-run to the pre-reform situation, and in turn the comparison of long-run and pre-reform data allows for an estimation of the entry effect as a residual.

Our results indicate that the license reform had a significant impact, resulting in an aggregate productivity improvement of around 22%, of which 25% and 75% can be attributed to the relaxation of size and entry constraints, respectively. These results are also significant when placed in the context of the turnaround in TFP growth in India that began at around the time of the reform, reported elsewhere in the literature (see, for example, Bosworth, Collins and Virmani (2007). More generally, these numbers are also indicative of the extent to which institutions and policy barriers are responsible for the low levels of TFP in poor countries, a point that is being increasingly stressed in the literature on cross-country income differences.

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Appendix: Proofs and derivations

(a) We derive formally the assertion made in the text (Section 4.1) that the relative productivity effects of the reform can still be estimated when factor prices change as a result of the reform, as long as the treatment industries have the same capital intensity as the control industries. We prove this claim for the special case of Cobb-Douglas technology, but the claim is valid for any homothetic production function. The intuition is that when capital intensities are the same, changes in factor prices change marginal costs (and hence aggregate productivity) in the same way for both sets of industries, so that the relative change in cutoff marginal costs still captures the relative productivity effect of the reform on the treated industries.

Let the production function be given by:

$$Y = \phi K^{\alpha} L^{1-\alpha}$$

It can be shown that the firm's TFP, ϕ, is related to its marginal cost x as follows:

$$x = \frac{1}{\phi}\frac{1}{1-\alpha}(\frac{r}{w}\frac{1-\alpha}{\alpha})^{\alpha} \tag{A.1}$$

where r and w are the rental rate of capital and the wage rate, respectively. Eqn (A.1) also relates the cutoff marginal cost x^* to the cutoff productivity, ϕ^*:

$$x^* = \frac{1}{\phi^*}\frac{1}{1-\alpha}(\frac{r}{w}\frac{1-\alpha}{\alpha})^{\alpha}$$

We can now write:

$$\frac{x^L}{x^*} = \frac{\phi^*}{\phi^L}\frac{(r^L/w^L)^{\alpha}}{(r^*/w^*)^{\alpha}}$$

or, in logs,

$$log(\phi^L) - log(\phi^*) = \alpha[log(r^L/w^L) - log(r^*/w^*)] - [log(x^L) - log(x^*)] \tag{A.2}$$

where, as before, starred values refer to pre-reform equilibrium quantities and L superscripts long-run equilibrium values.

Assuming perfect input markets, the bracketed part of the first term in Eqn (A.2) is the same for treatment as well as control industries if they have the same capital intensities. It follows that measuring the change in the cutoff marginal cost for the treatment industries relative to control industries yields an estimate of the relative productivity effect.

(b) Derivations:

1. (Output-Weighted) Technological average cost

$$\begin{aligned} &= \frac{1}{\bar{y}}\int_0^{x^*} xyd\tilde{G}(x) = \frac{1}{\bar{y}}\int_0^{x^*} x.\frac{x^*-x}{\theta}.\frac{k}{x^{*k}}.x^{k-1}dx \\ &= \frac{k}{\theta\bar{y}x^{*k}}\int_0^{x^*} x^k(x^*-x)dx = \frac{k}{\theta\bar{y}x^{*k}}.\frac{x^{*k+2}}{(k+1)(k+2)} = \frac{k}{k+2}x^* \end{aligned}$$

2. Output-Weighted Average cost (including license cost)

$$\begin{aligned} &= \frac{1}{\bar{y}}\int_0^{x^*}(x+\frac{\theta}{2}y)yd\tilde{G}(x) = \frac{1}{\bar{y}}\int_0^{x^*} xyd\tilde{G}(x) + \frac{1}{\bar{y}}\int_0^{x^*}\frac{\theta}{2}y^2d\tilde{G}(x) \\ &= \frac{k}{k+2}x^* + \frac{\theta}{2\bar{y}}\int_0^{x^*}(\frac{x^*-x}{\theta})^2d\tilde{G}(x) = \frac{k}{k+2}x^* + \frac{\theta}{2\bar{y}}\frac{k}{x^{*k}}\int_0^{x^*}(\frac{x^*-x}{\theta})^2dx \\ &= \frac{k}{k+2}x^* + \frac{x^*}{k+2} = \frac{k+1}{k+2}x^* \end{aligned}$$

3. To show that $\frac{\partial\bar{y}}{\partial f} > 0$, $\frac{\partial\bar{y}}{\partial\theta} < 0$, $\frac{\partial N}{\partial f} < 0, \frac{\partial N}{\partial\theta}$:
Recall that $\bar{y} = \frac{x^*}{\theta(k+1)}$ and $x^* = [\theta f(k+1)(k+2)x_m]^{\frac{1}{k+2}}$. Therefore we have:
$\frac{\partial\bar{y}}{\partial f} = \frac{1}{\theta f(k+1)}\frac{x^*}{f} > 0$
$\frac{\partial\bar{y}}{\partial\theta} = -\frac{x^*}{(k+2)\theta^2} < 0$
Recall that $N = \frac{\theta(k+1)}{b}\frac{a-x^*}{x^*}$. This yields:
$\frac{\partial N}{\partial f} = -\frac{\theta(k+1)a}{bfx^*} < 0$
$\frac{\partial N}{\partial\theta} = -\frac{k+1}{b} < 0$
4. To show that $\bar{y}^S > \bar{y}^*$:

We have that $N^S = N^*(\frac{x^S}{x^*})^k$ and that $\bar{y}^S = \frac{x^S}{\theta_1(k+1)}$ and $\bar{y}^* = \frac{x^*}{\theta_0(k+1)}$. Therefore:

$$\frac{\bar{y}^S}{\bar{y}^*} = \frac{\theta_0}{\theta_1}(\frac{N^S}{N^*})^{\frac{1}{k}}$$

We want to show that this is greater than 1. Recall that:

$$N^S = \frac{\theta_1(k+1)}{b}\frac{a-x^S}{x^S}$$

and

$$N^* = \frac{\theta_0(k+1)}{b}\frac{a-x^*}{x^*}$$

Thus

$$\begin{aligned}\frac{N^S}{N^*} &= \frac{\theta_1}{\theta_0}\frac{x^*}{x^S}\frac{a-x^S}{a-x^*} \\ &= \frac{\theta_1}{\theta_0}(\frac{N^S}{N^*})^{-\frac{1}{k}}\frac{a-x^s}{a-x^*}\end{aligned}$$

Since $x^S < x^*$, and the RHS above is less than 1, it follows that

$$\begin{aligned}\frac{\theta_1}{\theta_0}(\frac{N^S}{N^*})^{-\frac{1}{k}} &< 1 \\ &\Rightarrow \frac{\theta_0}{\theta_1}(\frac{N^S}{N^*})^{\frac{1}{k}} > 1\end{aligned}$$

as was to be proved.

5. To prove that $\bar{y}^L < \bar{y}^S$ and that $N^L > N^S$:

Recall that $\frac{\partial N}{\partial f}$ and $\frac{\partial N}{\partial \theta}$ are both negative. This implies that $N^L > N^*$ and therefore that $N^L > N^S$.

Further, we know that $N^L = \frac{a-x^L}{x^L}\frac{\theta_1(k+1)}{b}$ and that $N^S = \frac{a-x^S}{x^S}\frac{\theta_1(k+1)}{b}$. Thus, $N^L > N^S$ implies that $x^L < x^{S.}$.

Finally, recall that $\frac{\bar{y}^L}{\bar{y}^S} = \frac{x^L}{x^S}$ which implies that $\bar{y}^L < \bar{y}^S$.

Table 1: List of industries deregulated in 1985

Special alloy and iron castings, sponge iron and pelletisation
Steel structurals
Steam turbines
Power and distribution transformers
Power capacitors, switch gears, electrical motors
Diesel generating sets
Electronic components
Automotive ancillaries
Cycles
Industrial machinery, including rubber, printing, footwear and meat and poultry machinery
Machine tools
Agricultural implements
Plastic moulded goods
Hand tools, small tools and cutting tools
Pressure cookers, cutlery and steel furniture
Lanterns
Fuel efficient stoves
Water pumps
Industrial sewing machines
Office equipment
Surgical, industrial and scientific instruments
Drugs
Pulp and paper
Canned fruit and vegetable products
Marine products and cattle feed
Vegetable oils
Soaps, cosmetics and detergents
Leather goods
Surgical and medicinal rubber products
Glassware
Refractory and surface-lining bricks
Chinaware, pottery and sanitary ware
Tiles and graphite ceramics
Insulating boards, gypsum boards, wall boards, etc
Printing
Flour and milling industry
Chemicals

Source: *Handbook of Industrial Statistics* (1987)

Table 2: Summary statistics for deregulated and control industries in 1983

	Deregulated	*Underegulated*
Per-Industry Averages		
Output (in millions of rupees)	30.2	23.3
	(38.2)	(71.9)
Workers	17888	22538
	(23195)	(68096)
Capital (in millions of rupees)	576	696
	(907)	(3400)
No of Factories	400	254
	(809)	(604)
Aggregate figures		
Output (in millions of rupees)	1240	3970
Workers	733398	3831391
Capital (in millions of rupees)	23600	118000
No of Factories	16410	43181
Observations	41	170

Notes: Standard errors are in parentheses; "Output" is the value of total industry output in millions of 1982 rupees; "Capital" is the toal value of industry fixed capital, calculated as the average of opening and closing values of capital, in millions of 1982 rupees. Source: Annual Survey of Industries.

Table 3: Deregulation and Aggregate Industry Outcomes

		Change in logs of:									
Dependent Variables:		No of Factories		Output		Workers		Mandays		Capital	
		I	II	III	IV	V	VI	VII	VIII	IX	X
Year 85	γ_1	-0.01	-0.08*	0.04	0.11	-0.08***	-0.11**	-0.07**	-0.09	0.15***	0.09
		(0.02)	(0.04)	(0.05)	(0.15)	(0.03)	(0.05)	(0.03)	(0.06)	(0.03)	(0.15)
Year 88	γ_2	-0.28***	-0.34***	0.39**	0.46***	0.17*	0.15**	0.19*	0.17***	0.36**	0.31***
		(0.08)	(0.05)	(0.15)	(0.16)	(0.10)	(0.06)	(0.10)	(0.06)	(0.15)	(0.12)
Year 91	γ_3	0.05	-0.01	-0.31***	-0.24	-0.47***	-0.49***	-0.46***	-0.49***	-0.34***	-0.40***
		(0.06)	(0.04)	(0.09)	(0.15)	(0.05)	(0.05)	(0.06)	(0.06)	(0.07)	(0.11)
Deregulated * Year 85	β_1	0.04	0.08	0.12	0.23	0.08	0.23	0.11	0.29	0.11	0.27
		(0.05)	(0.15)	(0.07)	(0.17)	(0.06)	(0.20)	(0.07)	(0.20)	(0.06)	(0.22)
Deregulated * Year 88	β_2	0.01	0.04	0.28*	0.38**	0.15	0.30*	0.17	0.34*	0.28	0.44*
		(0.10)	(0.17)	(0.19)	(0.19)	(0.11)	(0.18)	(0.12)	(0.18)	(0.20)	(0.24)
Deregulated * Year 91	β_3	0.31***	0.35*	0.23*	0.34	0.23**	0.38	0.24**	0.41*	0.20	0.36
		(0.09)	(0.19)	(0.12)	(0.24)	(0.10)	(0.25)	(0.10)	(0.25)	(0.15)	(0.31)
3-Digit Fixed Effects?		no	yes	no	yes	no	yes	no	yes	no	yes
Number of Observations:		633	633	633	633	633	633	633	633	633	633

Notes: ***p<0.01;**p<0.05;*p<0.01; Standard errors reported in parentheses have been clustered at the 2-Digit industry-year level; The dependent variables are first-differences in logarithms of the correspondng variables; "Output" refers to the total industry output in that year; "Workers" refers to the total number of workers in the industry in that year; "Mandays" refers to the total number of mandays in the industry in that year; "Capital" refers to the average of opening and closing values of fixed capital in that year; Output and Capital have been deflated using price-deflators (with base-year 1982); "Deregulated" takes the value 1 if that industry was removed from licensing requirements in 1985;

Table 4: Estimating entry and size constraints

	Coefficient	1	2
Deregulated	β, β'	0.22***	0.03
		(0.07)	(0.13)
Δlog (No of Plants)	γ_1	0.08	
		(0.08)	
Δlog (No of Plants) * Deregulated	γ_2	0.36**	
		(0.20)	
2-Digit Fixed Effects?		yes	yes
Number of Observations		211	211

Notes: ***p<0.01;**p<0.05;*p<0.01; Standard errors reported in parentheses have been clustered at the 2-Digit industry level; the dependent variable in the regression in Column 1 is the change in the logarithm of average factory size at the industry level (measured as the output per factory) over the period 1985-88; the dependent variables in the regression in Column 2 is the change in the logarithm of average factory size at the industry level over the period 1985-91 (a two-period difference).

Table 5: Olley-Pakes Decomposition

	p_{it}	$\overline{p_{it}}$	Covariance
Deregulated			
1985	12.56	12.22	0.34
1988	13.33	12.31	1.02
1991	13.65	12.61	1.03
Underegulated			
1985	11.56	11.39	0.17
1988	11.96	11.76	0.20
1991	12.22	11.97	0.26

Notes: Reported values of the components of the Olley-Pakes decomposition represent a weighted average of values of these components for each set of industries where the weights are 1991 industry output shares. The first column reports the industry-level productivity measure, which is obtained as an employment-share weighted average of plant-level real output per worker. The second column reports the simple average of the plant-level productivities. The third column reports the within-industry covariance between plant-level productivity and employment-share.