Barriers to Entry and Regional Economic Growth in China*

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Abstract

Labor productivity in manufacturing differs starkly across regions in China. We document that productivity, wages, and start-up rates of non-state firms have nevertheless experienced rapid regional convergence after 1995. To analyze these patterns, we construct a Hopenhayn (1992) model that incorporates location-specific capital wedges, output wedges, and entry barriers. Using Chinese Industry Census data we estimate these wedges and examine their role in explaining differences in performance and growth across prefectures. Entry barriers explain most of the differences. We investigate the empirical covariates of these entry barriers and find that changes in barriers are causally related to changes in the size of the state sector: a smaller state sector leads to lower entry barriers.

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Keywords: Chinese economic growth; SOEs; firm entry; entry barriers; capital wedges; output wedges; SOE reform.

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1 Introduction

Since the onset of economic reform in the late 1970s, China has gone from one of the poorest economies in the world to being a middle-income country. The main source of this growth has been the expansion of the non-state sector (Zhu (2012)), especially in manufacturing. While the non-state sector has experienced rapid expansion at the national level, the growth has been highly uneven with significant differences across regions and localities. By the mid-1990s, this was reflected in sizeable local differences in productivity, wages, and the number and size of non-state enterprises (NSOE). Subsequently, differences between localities in the non-state sector began to disappear and from the mid-1990s China experienced a remarkably rapid economic convergence between localities, not only in value added per worker in non-state firms, but also in TFP, capital per worker, and wage rates. Moreover, the performance of non-state firms improved significantly more in areas where employment in state-owned firms declined faster.

The purpose of this paper is to examine the initial dispersion and subsequent convergence in the performance of the non-state sector through the lens of a macroeconomic model where the distribution and selection of firms matter for productive efficiency. In particular, we use this framework as an accounting device to determine which factors drove the initial dispersion across locations and the subsequent convergence. The theoretical framework is motivated by the empirical observation that the creation and selection of new firms in China's non-state sector have been the most important source of productivity and output growth in the manufacturing sector (Brandt, Van Biesebroeck and Zhang (2012)).

A number of factors might be responsible for differences in new firm creation and growth between regions, including human capital differences and distortions, taxes, and subsidies imposed by local governments (cf. Huang (2003)). To quantify the role of various channels we construct a Hopenhayn (1992) model, extended to allow three distortions. Following Hsieh and Klenow (2009) we allow for capital and output wedges. These wedges are prefecture-specific. In addition, we introduce a novel entry barrier which may differ across locations. This entry barrier takes the form of a probability that potential entrepreneurs will be allowed to operate firms. We solve the general equilibrium model analytically and show that the model aggregates. Namely, the underlying wedges can be derived using data on average wage rates and aggregate allocations of output, capital, and employment in a prefecture. Thus, by construction these wedges can account for the observed aggregate allocations in a given prefecture.

We measure the theoretical wedges using firm-level data from the Chinese Industrial Census (CIC) for 1995, 2004, and 2008. We construct data on value added, employment, capital, and average wage rates for each prefecture in China by aggregating the firm-level data. Focusing on aggregate allocations and distortions at the prefectural level – as opposed to the firm level – makes the analysis robust to measurement error at the firm level. To our knowledge our paper is the first to quantify distortions driving regional growth in China. The CIC data have some clear advantages. First, national account data are not available at the prefectural level. Second, the CIC data allow us to study theoretical predictions about the number of firm entrants and the firm size distribution since these data cover the entire manufacturing industry, not only large firms.

We use this framework to explore what factors/wedges are most important for accounting for the aggregate differences and convergence across prefectures in China. We find that the entry barrier is the main driver of both the initial (1995) dispersion and the subsequent convergence in wages and TFP across locations in China. Thus, the influence of capital or output market distortions, which in the Chinese context have also been identified as important (Hsieh and Klenow (2009); Song, Storesletten and Zilibotti (2011)) seems to be secondary for explaining the regional convergence of the non-state sectors in China. Instead, we conclude that local variations in the entry barriers are responsible for the regional economic heterogeneity and convergence.

We study the measured entry barriers in greater detail and show, in the spirit of Cheremukhin, Golosov, Guriev and Tsyvinski (2017, 2015), that these theoretical distortions can be tied to auxiliary empirical evidence for distortions. In particular, our measured entry barriers match up closely with measures of the formal costs of starting a business in China reported in the "Doing Business in China 2008" report by the World Bank (2008) for provincial capitals in China. This provides valuable external validation for our estimates.

Then, using prefecture-level information − beyond data on the aggregate allocations in the non-state manufacturing sector − we investigate the empirical drivers of the wedges. We are able to link systematically the size of these entry barriers and their changes to the size of China's state-owned enterprise (SOE) sector, and to several variables reflecting local fiscal capacity. In the mid-1990s, entry barriers were sizeably larger in localities with a larger SOE presence. In almost every dimension – the rate of start-up of new firms, size of firms, TFP, and wages – we find that new NSOE firms are weaker where SOEs are more dominant. However, after the mid-1990s the fortune turned to the better for prefectures which originally had a large state sector: on average, output per worker, TFP, wages, and capital per worker in non-state firms grew faster in these prefectures than elsewhere. This process is related to the fact that these same locations experienced large reductions in entry barriers because of large reductions in state employment.

Our results on the effect of the state sector are robust to potential concerns about endogeneity, omitted variables, and measurement error. We address such concerns with a Bartik (1991) instrumental variable approach. In a major policy change in 1997, the Chinese government allowed SOEs to be crowded out by non-state firms in some but not all industrial sectors. Interacting the initial local sectoral distribution of SOEs with the industry-specific decline in SOE employment at the national level predicts very accurately the reduction in local SOE employment. Using the 1995 SOE distribution as a Bartik instrument we find that the 1995-2004 reduction in SOE employment is systematically related to the reduction in entry barriers: larger predicted declines in SOE employment are associated with larger reductions in entry barriers. To study the link between entry barriers and SOE employment in the cross-section we apply an alternative instrumental variable approach that exploits various lagged instruments. The results confirm the findings using the Bartik instrument.

To motivate the empirical link between observed entry barriers and the size of the SOE sector, we develop a simple political economy model of local governments' incentives to influence the three wedges. In the model, local authorities face pressure to protect state-owned firms. Since non-state firms compete for resources with SOEs, the government uses these wedges to distort NSOEs' behavior in order to help SOEs. If local cadres care about the profits of local entrepreneurs, then restricting NSOE entry provides the best trade-off between ensuring that SOEs remain sufficiently competitive and supporting the NSOE profits.

Finally, we extend the benchmark model to allow for firm-specific capital and output wedges as in Restuccia and Rogerson (2008). We reestimate the model and find that the entry barriers continue to account for most of the regional convergence in wages and TFP. Moreover, the entry barriers estimated from the extended model are highly correlated with those of the benchmark model.

Our paper makes a number of contributions. First, we provide an analytical framework that can be used as an accounting device to identify distortions that inhibit or stimulate growth in a development context. Second, we use this framework to provide new insights for understanding growth dynamics in China. We identify new firm behavior and the removal of barriers to entry as the main driver of regional wages and TFP growth. Third, we document an important set of new empirical facts on regional economic development in China, emphasizing the strong convergence in wages, TFP, labor productivity, and capital per worker across regions after the mid-1990s. Fourth, we study the empirical determinants of the prefecture-specific barriers to entry. Our Bartik IV approach reveals a novel and important channel: SOEs cause larger entry barriers for non-state firms. This finding points to an important additional benefit of the reforms of the state-owned sector of the late 1990s: as SOEs were scaled back, the entry barriers for private firms came down. This in turn paved the way for the subsequent rapid economic growth.

Our paper builds on and contributes to several literatures. There exists an extensive literature analyzing China's economic rise. At the aggregate level, productivity growth has been identified as the most important source of Chinese growth (Zhu (2012)). Especially important in this context is the rapid growth in TFP in the non-agricultural sector (Cheremukhin et al. (2015)), and within non-agriculture, the non-state sector (Brandt and Zhu, 2010). Cheremukhin et al. (2015) attribute 62% of the growth in GDP in China between 1978-2012 to productivity growth in non-agriculture. The downsizing of the state sector in the mid-to-late 1990s and a reallocation of resources to the non-state sector has been singled out in the literature for its role in the improvement of TFP in the non-agricultural sector (Hsieh and Klenow (2009), Song et. al. (2011), and Song and Hsieh (2015)). But even more important to TFP growth in non-agriculture is the contribution of new and better non-state firms, which enjoyed a significant premium over incumbent firms (Brandt et. al., 2012 and 2017). Our analysis explicitly links these gains in the manufacturing sector to the fall in local entry barriers for new firms−which we find is also tied to restructuring of the state sector−and the ensuing convergence in productivity across regions.

Our model and results are in line with this literature. In particular, we find that TFP is the primary driver of both growth and convergence. Moreover, aggregate TFP is determined by the establishment of new non-state firms, and this process is the main focus of our theory. Our message is that local TFP growth can be linked to changes in local entry barriers for new firms.

Our paper builds on the literature using wedge analysis to infer sources of distortion for understanding economic growth (see e.g. Chari, Kehoe and McGrattan (2007) and, in a developing economy context, Cheremukhin et al. (2015, 2017)). A large literature emphasizes distortions and misallocation of resources for understanding cross-country differences in economic development (see e.g. Restuccia and Rogerson (2008) and Hsieh and Klenow (2009)). This literature identifies a number of distortions that may be important, including implicit taxes on capital, labor, and output. In the Chinese context the literature has emphasized both capital market distortions (Hsieh and Klenow (2009), Song et al. (2011), Brandt and Zhu, 2010)) and labor market distortions (Tombe and Zhu (2019)). We incorporate such distortions in our analysis. Similar to Barseghyan and DiCecio (2011), we also emphasize the role of entry barriers for new firms in accounting for TFP differences, although we focus on regional convergence in China while they focus on dispersion across countries.

Finally, our paper contributes to the large macroeconomic literature studying growth and convergence across countries and regions (Barro and Sala-i-Martin (1991); Mankiw, Romer and Weil (1992); Gennaioli, La Porta, Lopez De Silanes and Shleifer (2014)).

The rest of the paper is organized as follows. Section 2 documents the empirical economic development across more than 300 prefectures. Section 3 lays out a version of the Hopenhayn (1992) model extended to incorporate a novel entry barrier. Section 4 uses the entry barrier model to measure the distortions across prefectures and Section 4 quantifies the role of wedges for growth and convergence. Section 5 studies the empirical drivers of the prefecture-specific measured entry barriers while Section 7 studies an extension of the model that allows for firm heterogeneity in wedges. Section 8 concludes.

2 Empirical Evidence

2.1 Data description

Chinese Industrial Census. Our main data source is the 1995, 2004, and the 2008 Chinese Industrial Census (CIC) carried out by China's National Bureau of Statistics (NBS).[1] The CIC covers all of the manufacturing sector[2] and provides firm-level data on gross output, value added, employment, the gross capital stock, depreciation, total wage bill, as well as information on firm year of establishment, ownership type, and main sector of business. For these three years, we have firm-level records on 0.53, 1.37 and 2.08 million firms, respectively.[3]

In order to make these data comparable across the three census years, we have addressed a number of issues related to changes that occurred in China's industrial classification system, ownership categories, and prefecture boundaries. We draw on concordances described in Brandt et al. (2012) for ownership types and industrial sectors, and extend the concordance on prefecture boundaries in Baum-Snow, Brandt, Henderson, Turner and Zhang (2017) to cover all prefectures. We also utilize deflators developed by Brandt et al. (2012) for the purposes of constructing real measures of industrial output and estimates of the real capital stock (see Appendix A).

Using the CIC data on firm type by ownership, we identify non-state-owned firms as all firms other than those listed as state-owned, state solely-funded limited liability companies, or shareholding companies. We have experimented with alternative definitions of NSOEs. In general, our results – both in the cross-section and over time – are robust to these alternative definitions.[4]

2.2 Regional dispersion and convergence

We start by documenting the initial dispersion and subsequent convergence across locations in a set of key economic variables: the average wage per worker, aggregate value added per worker, aggregate capital per worker, and aggregate TFP, all measured at the prefecture level.

Table 1: Dispersion and Rates of Convergence.

	labor productivity $\frac{Y}{N}$	wage rate w	capital per worker $\frac{K}{N}$	Aggr. TFP Z
Annualized rate of β-convergence 1995-2004	7.8%	7.6%	11.7%	3.7%
Annualized rate of β-convergence 1995-2008	9.2%	5.7%	11.9%	5.1%
90-10 ratio in 1995	4.4	2.3	3.5	3.1
90-10 ratio in 2004	2.8	1.8	3.1	3.2
90-10 ratio in 2008	3.1	3.0	4.2	2.9

Notes: The table reports the annualized rates of β-convergence and the cross-sectional dispersion in aggregate outcomes across prefectures in China. The annualized β-convergence coefficient between times t_0 and t_0+T for variable x is estimated from the regression $\left(\frac{1}{T}\right)\ln\left(\frac{x_{p,t_0+T}}{x_{p,t_0}}\right) = a - \frac{1}{T}\left(1-e^{-\beta T}\right)\ln(x_{p,t_0}) + \varepsilon_{pt_0}$, where ε_{pt_0} is an error term.

Figure 1 and Table 1 document the dispersion across prefectures in 1995 and the dynamics of the aggregate variables between 1995 and 2004. Each panel is a scatter plot of the level of a variable (on a log scale on the x-axis) against the growth in the variable over the 1995-2004 period. Figure B-1 in Appendix B documents the corresponding statistics for the 1995-2008 period.

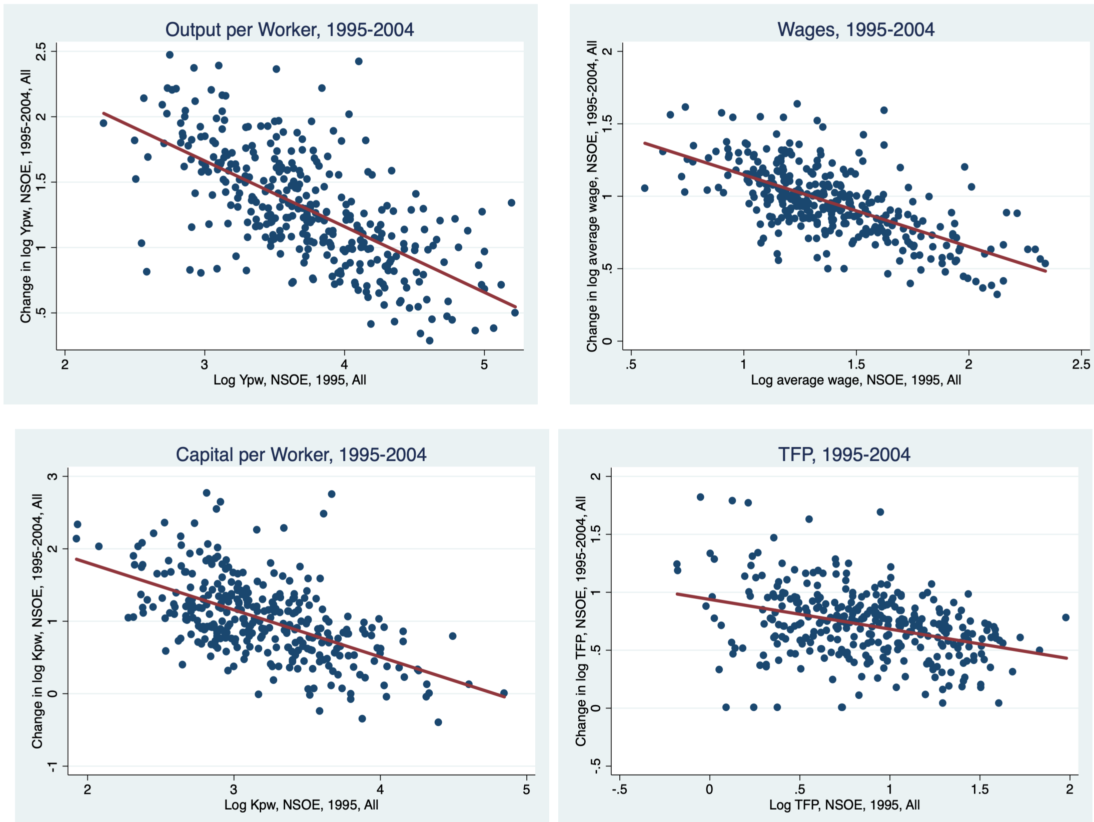
[1] We also draw on firm-level data collected by the NBS for 1992 on all independent accounting units (0.39 million), which covers a slightly smaller subset of firms than the CIC.

[2] The 2004 and 2008 CIC also provide data for the service sector, but unfortunately similar information was not reported in 1995.

[3] The firm-level records are not exhaustive, but cover upwards of 90 percent of industrial activity.

[4] See Appendix A for details.

Figure 1: Convergence in the NSOE sector, 1995-2004.



Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line.

The first observation is that there is large dispersion across prefectures in these aggregate outcomes. Second, there was substantial conditional convergence in labor productivity between 1995 and 2004. The negative slope of the regression line in the top left panel implies that the growth in labor productivity over this period was larger in prefectures with low initial labor productivity. The annualized rate of β-convergence was around 8% (cf. Table 1), implying that it only takes about 9 years to cut the difference in labor productivity between any two prefectures by half.

To put this rate of convergence in perspective, we compare it with the rate of convergence in GDP per capita across regions in other countries. Using data from 1,528 regions from 83 countries, Gennaioli et al. (2014) document that the speed of per capita income convergence within countries has been about 2 percent per year. This is similar to what Barro and Sala-i-Martin (1991) found for convergence across states in the US (1880-1988). By this metric, regional convergence in labor productivity in China's non-state manufacturing sectors was exceptional, both from a historical and international perspective.

Consider now the dynamics of wages per worker in non-state manufacturing, documented in the top right panel of Figure 1. Wages per worker in the non-state sector measure the annual average earnings per worker. The 1995 dispersion across prefectures in average wage per worker is large, albeit less dispersed than labor productivity. The annualized rate of β-convergence was also large—at 7.6 percent—suggesting that it took about 9 years to reduce average wage differences between two prefectures by half.

The regional dispersion in capital per worker in non-state manufacturing firms is documented in the bottom left panel of Figure 1. Both the initial 1995 regional dispersion and the rate of convergence are large, with an annualized β-convergence of close to 12 percent.

The bottom right panel of Figure 1 documents the convergence and dispersion across prefectures in aggregate TFP. We calculate the growth in aggregate TFP as the weighted average growth in the Solow residual across industries. In particular, we calculate the Solow residual in each industry, Z_j, in line with standard growth accounting. Namely, Z_j is the residual from a sector-specific production function that takes total capital and labor (used in the industry and prefecture) as inputs; $Y = ZF(K, N)$, where Y, K, and N are, respectively, production, capital, and labor in industry j. This assumes a constant returns to scale Cobb-Douglas production function, $Y = ZK^{\alpha_j}N^{1-\alpha_j}$, with industry-specific shares (α_j). The Solow residual is computed for each (2-digit) industry in a prefecture in a particular year. In order to compute the growth in aggregate TFP for a particular prefecture j, we first compute the change in $\ln Z_j$ for each (2-digit) industry and then aggregate up the industry-specific growth rates using as weights the relative share – averaged across 1995 and 2004 – of the composite input $K^{\alpha}N^{1-\alpha}$ of each industry in that prefecture. The 1995 dispersion in aggregate TFP across prefectures is very large, with a 90-10 ratio of 3.1. Aggregate TFP exhibits regional annualized convergence between 1995 and 2004 of about 4 percent.[5]

Finally, we note that between 1995 and 2004 the β-convergence is so strong that even the cross-sectional dispersion in all variables except TFP fell: Table 1 shows that the dispersion is lower in 2004 than in 1995 for all variables except TFP, indicating σ-convergence across prefectures. However, after 2004 the overall cross-sectional dispersion in wages and capital per worker increased while it remained fairly constant for TFP and labor productivity. In the presence of shocks, the dispersion in, for example, productivity can increase even if there is conditional convergence (see Barro and Sala-i-Martin (1991) for a discussion).

Co-movements between TFP, wages, and new firm entry. Table 2 documents the correlation matrix in levels and growth for these variables. We define the *entry rate* Γ of new private firms in a prefecture as the share of employment in *new* NSOE firms, i.e., firms established during the last two years, relative to total employment in manufacturing in the prefecture. We interpret this statistic as a measure of firm entry. As is clear from Table 2, all variables are positively correlated. This holds true in the cross section in 1995 as well as in changes over the 1995-2004 or 1995-2008 periods. This positive correlation will be important for identifying the key forces driving the differential performance of the non-state sectors across prefectures.

2.3 The size of state sector and non-state sector performance

Earlier research documents that the growth of non-state firms is correlated with the presence of SOEs. Brandt, Hsieh and Zhu (2008), for example, show that across provinces the growth in non-state nonagricultural output is negatively correlated with the initial size of the state sector, measured by the 1978 share of aggregate value added in non-agriculture produced by state-owned firms. Motivated by this evidence, we now document that in our data the performance of non-state firms is also related to the size of the state sector.

We start by analyzing the performance of non-state firms (NSOE) in the 1995 cross section of prefectures. We continue to focus on average wages, labor productivity, capital per worker, and TFP for non-state firms. We also document entry rates of new NSOEs for each prefecture. In

[5] We show below that the convergence results are robust to adding a set of controls, including province-level fixed effects. When controls are included, the rate of convergence is larger for TFP and slightly lower for wages, output per worker, and capital per worker. See Section 6.2 for details.

Table 2: Comovements in Wages, TFP, and Firm Entry.

	1995				1995-2004			1995-2008		
	$\ln W$	$\ln TFP$	$\ln \Gamma$		$\Delta \ln W$	$\Delta \ln TFP$	$\Delta \ln \Gamma$	$\Delta \ln W$	$\Delta \ln TFP$	$\Delta \ln \Gamma$
$\ln W$	1.00			$\Delta \ln W$	1.00			1.00		
$\ln TFP$	0.30	1.00		$\Delta \ln TFP$	0.25	1.00		0.20	1.00	
$\ln \Gamma$	0.26	0.39	1.00	$\Delta \ln \Gamma$	0.26	0.10	1.00	0.15	0.19	1.00

Notes: The table reports the correlations between log wages, log TFP, and log firm entry in 1995 as well as the correlations between the changes in log wages, log TFP, and log firm entry in 1995-2004 or 1995-2008.

order to illustrate the correlation with the size of the state sector, we sort prefectures according to the local preponderance of state firms. We let s_p denote the size of the state sector in prefecture p, measured by the fraction of output in manufacturing produced by state firms. Our results are essentially the same if we use the fraction of workers employed by state firms as our measure of the size of the state sector.

Wages, TFP, value added per worker, and capital per worker for NSOE entrants. Figure 2 documents aggregate outcomes for NSOE firms across prefectures. We report these facts for new firms – defined as firms established between 1993 and 1995 – since firm creation will be a central focus in our theory below. The figure reveals that new entrants in prefectures with a larger state presence in 1995 (high s_p prefectures) pay lower wages, have lower TFP, lower value added per worker, and less capital per worker.[6,7] On the basis of OLS regressions, the SOE output share in 1995 accounts for 12% of the variation in wages across prefectures, 40% of the variation in aggregate TFP, 39% of the variation in value added per worker, and 9% of the variation in capital per worker.

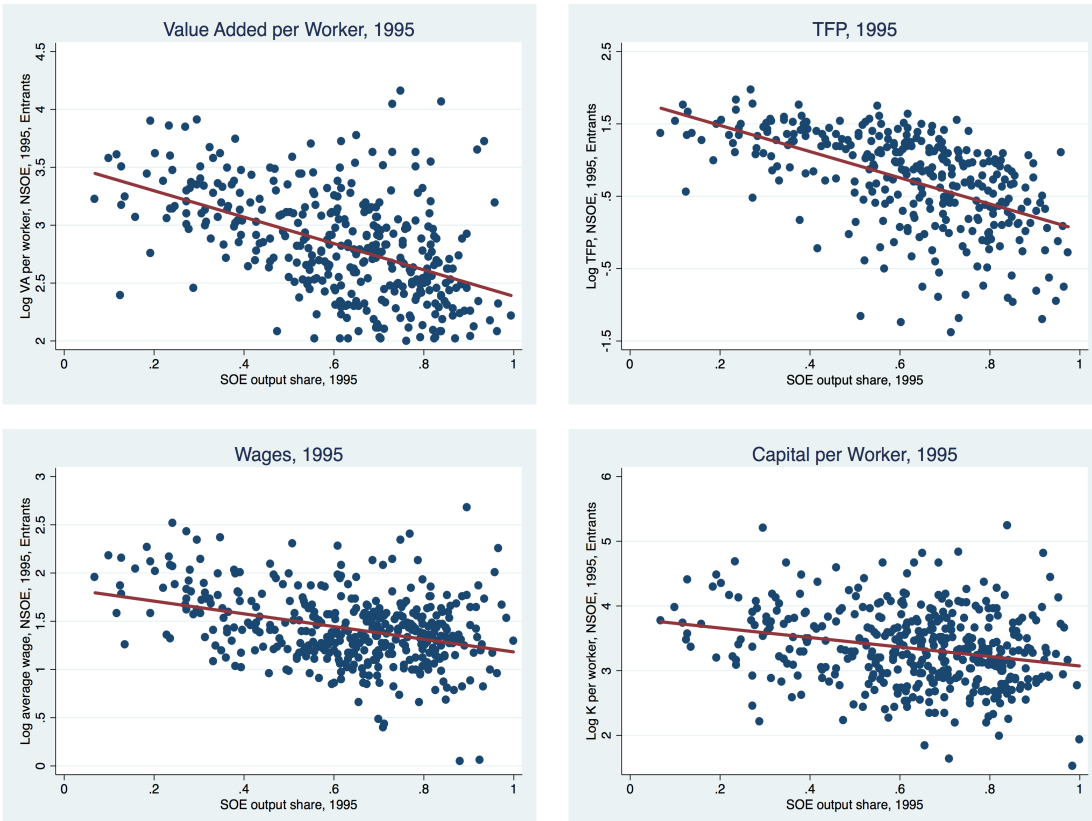
Firm entry in the NSOE sector. Prefectures with high s_p also have substantially lower entry of NSOE firms. The left panel in Figure 3 plots the number of new firms in a prefecture as a share of all new firms at the national level. Clearly, most of the new NSOE entrants were established in prefectures in which the state sector was less prominent in 1992. The right panel in Figure 3 measures 1995 employment in new NSOE firms as a fraction of total employment in that prefecture in 1992. Again, most of new NSOE employment originates in prefectures that had a low s_p in 1992.

To summarize, the 1995 CIC cross-section reveals that in prefectures with high s_p, there were relatively fewer NSOE entrants and NSOE entrants were weaker in multiple dimensions – they paid lower wages and had lower total factor productivity, lower value added per worker, and lower capital per worker. The same relationship between the size of the state sector and the performance of the non-state sector also holds in terms of observed changes between 1995 and 2004, as reported

[6]Figure B-2 in Appendix B shows that the same patterns hold up for all non-state firms, i.e., when also including incumbent NSOE firms.

[7]One concern is that the negative relationship between the size of the state sector and productivity in the non-state sectors is a product of unobserved heterogeneity at the prefecture level. State owned enterprises might be located in more "backward" prefectures where endowments of human capital are lower. In regressions in Section 6.1 we add a set of control variables, including average human capital and the employment share of agriculture. The results are robust to these controls.

Figure 2: Characteristics of NSOE Entrants in 1995.



Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line. The 1995 SOE output share in a prefecture is on the horizontal axis.

Figure 3: NSOE Firm Entry during 1993-1995.

Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line.

in Figure B-3: In prefectures with a larger decline in the share of the state sector, the non-state sector experienced a larger increase in its value added per worker, TFP, wages, and capital per worker.

In the next section, we develop a model to identify the main forces – in terms of prefecture-level wedges and distortions – that are behind the relationship between the performance of the non-state sector and the state sector.[8]

3 A Hopenhayn Model of Heterogeneous Entrepreneurs

This section lays out a theory of private (non-state) firms across locations. The main purpose is to derive predictions about aggregate firm performance in each location.

The economy consists of a set of locations. Each location is a small open economy where labor is location specific and supplied inelastically. Capital can be allocated freely across locations and commands an exogenous common rental rate $(r+\delta)$. In the main analysis we take the labor supply offered to private firms in location j, N_j, as exogenous.[9]

Firms produce a homogeneous good with decreasing returns to scale. The production function is Cobb-Douglas,

$$y_i = z_i^{1-\eta} \left(k_i^{1-\alpha} n_i^{\alpha}\right)^{\eta} , \tag{1}$$

where y_i is the firm's value added, k_i is the firm's capital stock, n_i is the firm's employment, and z_i is the firm's total factor productivity. The parameter $\eta \in (0,1)$ captures the decreasing returns. We allow the parameter $\alpha \in (0,1)$ to differ across locations, reflecting heterogeneity in the technological labor income share $\alpha\eta$. Firms pay workers a location-specific wage rate w. In addition, firms face standard distortions on output and capital given by τ_y and τ_k.[10] These wedges are common for all firms in the location.[11] There is a fixed cost ν for operating a firm. This cost is constant across all locations.

Following a long tradition in the misallocation literature, the model is static, comprising two stages: a firm entry stage and a production stage. Each location has a measure M_j of potential entrepreneurs. Each potential entrepreneur can operate one firm and this firm is endowed with a productivity z. The distribution of productivities of potential entrepreneurs is given by a p.d.f. $f(z)$. We assume that z is Pareto distributed, i.e., that $f(z) = \xi \underline{z} z^{-\xi-1}$, where $\xi > 1$, $\underline{z} \geq 1$, and $z \in \left[\underline{z}^{1/\xi}, \infty\right)$. The Pareto assumption is made for analytical convenience. We show in Section 3.3 that the upper tail of the TFP distribution in China is approximately Pareto distributed.

A key source of heterogeneity across locations is that they differ in the *effective* number of potential entrepreneurs. In particular, we assume that a location-specific fraction ψ of entrepreneurs who want to produce will not obtain a license to operate and will therefore be prevented from entering. We refer to the fraction of potential entrepreneurs who have the option to operate, $(1-\psi)$, as the *gross entry barrier*.[12] This entry barrier can be interpreted as a lottery over licenses.

[8] As we will see, these wedges and distortions are strongly correlated with the size of the state sector. Clearly, interpreting these correlations is not straightforward due to potential econometric issues such as endogeneity, measurement error, and omitted variable bias. We tackle these issues in Section 6 and show that there is indeed a causal link between the size of the state sector in a prefecture and the distortions and the performance of its non-state sector.

[9] In Section 6.3 we extend the model to incorporate a second type of firms – state firms – which compete with private firms for local workers.

[10] We interpret these wedges as implicit taxes, where these taxes are not recorded as costs and thus do not affect the measured value added y_i. Moreover, following Hsieh and Klenow (2009), we abstract from labor wedges. Our analysis differs from Hsieh and Klenow (2009), however, in that we allow wage rates to differ across locations.

[11] Section 7 extends the analysis to allow for firm-specific capital and output wedges. As we shall see, our quantitative findings and the main message of the paper remain robust to this extension.

[12] Hopenhayn (1992) proposes an alternative model of entry barriers with an infinite supply of potential entrepreneurs. Each entrepreneur who considers entering must first pay a fixed cost of obtaining a stochastic draw of

It is important that this barrier is independent of the firm's productivity. As we shall see below, this feature will induce negative selection of firms entering locations with a large ψ.

3.1 The Firm problem

We start by analyzing the production stage and then study the entry decision.

For convenience we drop the firm subscript i. Firms maximize profits and take as given the wedges and prices. The firms' objective, conditional on operating, is given by:

$$\Pi = \max_{k,n} \left\{ (1-\tau_y)\, y - wn - (1+\tau_k)\,(r+\delta)\, k \right\}. \tag{2}$$

Using the firm's first-order conditions, the optimal choices are given by,

$$\begin{aligned} y^* &= z \cdot \bar{y}\left(\tau_y, \tau_k, r, w\right) \\ k^* &= z(1-\alpha)\eta \frac{(1-\tau_y)}{(1+\tau_k)(r+\delta)} \cdot \bar{y}\left(\tau_y, \tau_k, r, w\right) \\ n^* &= z\alpha\eta \frac{(1-\tau_y)}{w} \cdot \bar{y}\left(\tau_y, \tau_k, r, w\right) \\ \Pi^* &= (1-\tau_y)\,(1-\eta)\, z \cdot \bar{y}\left(\tau_y, \tau_k, r, w\right), \end{aligned} \tag{3}$$

where

$$\bar{y}\left(\tau_y, \tau_k, r, w\right) \equiv \left[(1-\tau_y)\,\eta\right]^{\frac{\eta}{1-\eta}} \left(\frac{(1-\alpha)}{(1+\tau_k)\,(r+\delta)}\right)^{\frac{(1-\alpha)\eta}{1-\eta}} \left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}.$$

Given the vector of distortions and prices (τ_y, τ_k, r, w), there exists a cutoff $z^* = z^*\left(\tau_y, \tau_k, r, w\right)$ such that all potential entrepreneurs with $z \geq z^*$ will choose to operate firms. Given the profit function Π, this cutoff z^* is determined by the condition $\nu = (1-\tau_y)\,(1-\eta) \cdot z^* \cdot \bar{y}\left(\tau_y, \tau_k, r, w\right)$, implying

$$z^* = \frac{\nu}{(1-\tau_y)^{\frac{1}{1-\eta}}\, \eta^{\frac{\eta}{1-\eta}}\, (1-\eta)} \left(\left(\frac{(1+\tau_k)\,(r+\delta)}{1-\alpha}\right)^{1-\alpha} \left(\frac{w}{\alpha}\right)^{\alpha} \right)^{\frac{\eta}{1-\eta}}. \tag{4}$$

3.2 Equilibrium

The equilibrium allocations and prices can be determined separately for each location, given the wedges and location-specific set of workers and potential entrepreneurs (N_j, M_j). The reason is that capital is fully mobile and the locations produce a homogeneous good. Differences across locations are therefore determined entirely by differences in wedges and (N_j, M_j). Note that we take the wedges and (N_j, M_j) as exogenous and abstract from endogenous migration decisions of workers and entrepreneurs. However, we will take migration and population growth into account in the quantitative analysis insofar as they influence the realized empirical distribution of (N_j, M_j). We discuss below how endogenizing migration decisions would affect our analysis.

We now solve for the equilibrium wage w and the associated aggregate output, capital stock, and measured aggregate TFP in each location. Market clearing in the labor market requires that

$$M(1-\psi) \int_{z^*}^{\infty} n\left(z\right) f\left(z\right) dz = N, \tag{5}$$

firm TFP and the cost is incurred before the TFP is realized. The predictions of our model differ qualitatively from Hopenhayn (1992) in the effect of labor supply N_j. In Hopenhayn (1992) changes in N_j have no effects on allocations and wages. We show below that in our model an increase in labor supply will lower equilibrium wages and TFP.

where M is the number of potential entrepreneurs.

Imposing labor market clearing and optimal firm behavior (equations (3)-(4)), we can solve analytically for the equilibrium wage as a function of the distortions. We state this as a proposition.

Proposition 1 *The equilibrium wage in a location is given by*

$$\begin{aligned}\ln w &= \mu(1-\eta)\ln\left[(1-\psi)\underline{z}\frac{M}{N}\right] \\ &+\mu\xi\ln(1-\tau_y)-\mu(1-\alpha)\xi\eta\ln\left[(1+\tau_k)(r+\delta)\right]+\Omega(\alpha,\eta,\xi,\nu), \qquad (6)\end{aligned}$$

where $\mu\equiv\frac{1}{1-\eta+\xi\alpha\eta}>0$ *and the term* $\Omega(\alpha,\eta,\xi)$ *is defined in Appendix C. The equilibrium wage is falling in* N/M, τ_y , τ_k, *and* ψ.

Our empirical analysis focuses on wages, aggregate TFP, and new firm entry. We treat these objects as summary statistics that we study through the lens of our model. Section 2.2 documents regional convergence in wages and aggregate TFP where TFP is defined as the Solow residual, in line with standard growth accounting. We now calculate the Solow residual in a way that is theoretically consistent with our model.[13] Using the equilibrium wage w to calculate z^* and aggregating over firms' optimal choices allows us to determine the implied Solow residual as a function of the wedges,

$$\begin{aligned}\ln Z &= \mu\alpha\eta(1-\eta)\ln\left[(1-\psi)\underline{z}\frac{M}{N}\right]-\mu(1-\eta)\ln(1-\tau_y) \\ &+\mu(1-\eta)\left[1+(\xi-1)\alpha\eta\right]\ln\left[(1+\tau_k)(r+\delta)\right]+\hat{\Omega}(\alpha,\eta,\xi,\nu). \qquad (7)\end{aligned}$$

Note that aggregate TFP is increasing in τ_k, τ_y, and M and decreasing in ψ and N. Moreover, the term $\hat{\Omega}(\alpha,\eta,\xi,\nu)$, defined in Appendix C, does not interact with the wedges.

Let the measure of firms entering the location be denoted $\Gamma = Pr(z\geq z^*)$. The theoretical prediction for firm entry Γ is given by

$$\begin{aligned}\ln\Gamma &= \ln\left[M(1-\psi)\int_{z^*}^{\infty}\underline{z}\xi z^{-\xi-1}dz\right] \\ &= \mu(1-\eta)\ln\left[(1-\psi)\underline{z}M\right]+\mu\alpha\eta\xi\ln(N) \qquad (8)\\ &+\mu\xi\ln(1-\tau_y)-\mu\xi\eta(1-\alpha)\ln\left[(1+\tau_k)(r+\delta)\right]+\tilde{\Omega}(\alpha,\eta,\xi,\nu),\end{aligned}$$

where the term $\tilde{\Omega}(\alpha,\eta,\xi,\nu)$ is independent of the wedges (see Appendix C for derivation). It follows immediately that the number of firm entrants is falling in τ_y, τ_k, and ψ and rising in N and M.

Comparative statics on wages, TFP, and firm entry. The analytical characterizations above allow us to study how the endogenous outcomes we study in the empirical analysis – namely wage rates, firm entry, and the Solow residual – react to changes in the wedges and the N/M ratio. Consider, first, the entry barrier. A larger ψ will lower the number of potential entrants. If the productivity cutoff z^* was held constant there would be fewer entrants and less demand for labor. To clear the labor market the wage must fall in order to induce each firm to hire more workers and to attract more entrants. The TFP cutoff z^* falls in response to the lower wages: firms with lower productivity are able to operate since labor is cheaper. This induces negative selection which in turn lowers the aggregate TFP. The result is that an increase in the entry barrier ψ will lower firm entry, wages, and aggregate TFP. Thus, aggregate TFP, firm entry, and wage rates all move in the same direction in response to movements in ψ (cf. Table 3).

[13] In growth accounting, the Solow residual is calculated by imposing an aggregate Cobb-Douglas production production function with constant returns to scale with a weight equal to the empirical labor-income share on labor and a weight of one minus the labor share on capital. To calculate this object in our model, we impose a weight $\alpha\eta$ on labor and a residual weight $1-\alpha\eta$ on capital, defining the Solow residual as $\ln Z\equiv\ln Y-\alpha\eta\ln N-(1-\alpha\eta)\ln K$.

Table 3: Comparative statics of varying the wedges (τ_y, τ_k, ψ) and aggregate labor supply per potential entrepreneur, N/M.

	$(1-\tau_y)$	$(1+\tau_k)$	$(1-\psi)$	N/M
wage rate w	+	-	+	-
Solow residual Z	-	+	+	-
Entry Γ	+	-	+	+

Consider now the effect of changing the labor supply per potential entrepreneur, N/M. A larger N or a lower M requires a lower equilibrium wage in order to clear the labor market. The lower wage induces a lower TFP cutoff z^* which in turn implies both more firm entry and a lower aggregate TFP due to negative selection. Thus, a larger labor supply N and fewer potential entrepreneurs M cause lower aggregate TFP, lower wage rate, and more firm entry (cf. Table 3).

Finally, increasing τ_y and τ_k will lower profits and distort the optimal firm size and optimal use of capital in the firm, making it less attractive for potential entrepreneurs to enter. This will increase the cutoff z^* and lower firm entry and demand for workers, which in turn lowers the equilibrium wage. However, larger τ_y and τ_k will increase the Solow residual, which is our measure of aggregate TFP.

Discussion of alternative forms of heterogeneity. Our analysis assumes geographical variation in the wedges (ψ, τ_y, τ_k) and holds the parameters ν and $\underline{z}$ constant across locations. We could alternatively have allowed variation in the lower bound for the distribution of firm TFP, $\underline{z}^{1/\xi}$. Variation in $\underline{z}$ would capture prefecture-level TFP shifts affecting all firms proportionally such as for example agglomeration or human capital differences across locations. Note that the terms $(1-\psi)$ and $\underline{z}$ enter multiplicatively in equations (6)-(7). Thus, variation in $\underline{z}$ would have the same effect on wages, aggregate TFP, and firm entry as would variation in the entry barrier $(1-\psi)$. This equivalence is due to the Pareto distribution assumption: shifting the distribution of potential entrepreneurs down (i.e., lower $\underline{z}$) is equivalent to lowering the effective number of potential entrepreneurs. Our theory is therefore not designed to quantify the role of selection versus proportional TFP shifts in accounting for local growth.[14] We have focused our analysis on geographic heterogeneity in the entry barrier but recognize that heterogeneity in the distribution of entrepreneurs offers an alternative interpretation of the data. The next section provides evidence supporting the assumption that the distribution of potential entrepreneurs is similar across prefectures. We could also have assumed heterogeneity in the fixed operating cost ν instead of modeling heterogeneity in ψ. However, we find it more intuitive to let differences across locations, over and above capital and output frictions, to be captured by heterogeneity in ψ.

[14] Combes, Duranton, Gobillon, Puga and Roux (2012) propose an innovative approach to estimating the role of agglomeration versus selection and apply their methodology to study TFP differences across cities in France.

3.3 Evidence for the selection mechanism

Our theory has strong and testable implications for how the cross-sectional distribution of firm TFP differs across locations. Our maintained assumption is that potential entrepreneurs in all locations draw their productivity from the same distribution f. As a result, measured TFP can vary across locations only because of differences in the selection of entrepreneurs: a high threshold z^* induces high selection of firms and high measured aggregate TFP. This selection mechanism is the hallmark of our theory and has sharp implications for the shape of the firm-TFP distribution and how it differs across locations.

First and foremost, the lower tail of the TFP distribution should be thinner in locations with larger aggregate TFP. The reason is that entrepreneurs in a high-TFP location have a larger z^*. Therefore, high-TFP prefectures should have distributions that are left-truncated at a higher level of TFP. Second, the upper tail of the TFP distribution of a location should be invariant to its aggregate TFP. The reason is that all locations draw entrepreneurs from the same distribution f, implying an identical upper tail of the firm-TFP distribution across locations. Namely, even though there are fewer productive entrants in low-performing locations where z^* is low, the distribution of the most productive firms should, according to the model, be identical.

Consider now the TFP distributions of two locations – one with low aggregate TFP and one with a high aggregate TFP. A direct implication of the predictions above is that conditional on firm TFP being above some level z, average TFP should be very different across the high- and low-TFP locations for low z because the low-TFP location has more mass in the lower tail. However, the difference should be very small for a high z because the upper tails of the TFP distributions should be similar.

We provide a straightforward test of this prediction. We first sort all prefectures into two groups according to their aggregate TFP – one with high and one with low aggregate TFP. For each percentile s of the overall distribution of firm-specific TFP we calculate the ratio of conditional means for the group with high versus low aggregate TFP, denoted $CMR(s)$,

$$CMR(s) \equiv \frac{E\{z_H \mid z_H \geq Z(s)\}}{E\{z_L \mid z_L \geq Z(s)\}},$$

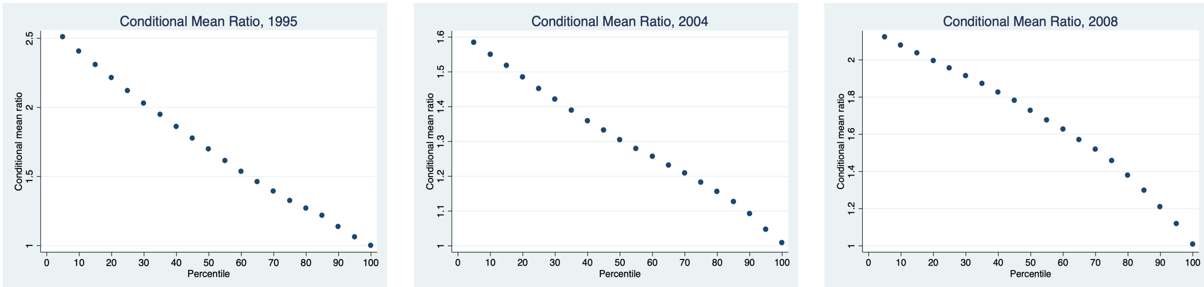
where z_H and z_L represent firm-specific TFP in the groups with high versus low aggregate TFP, respectively, and $Z(s)$ is the TFP level of percentile s in the overall TFP distribution. The key prediction is that the conditional means ratio, $CMR(s)$, should fall monotonically in s and converge to unity as s becomes large. Note that this is a general prediction of our selection-based theory which holds true for any distribution f, including the Pareto distribution we have assumed for tractability.

Figure 4 reports the results. To mitigate the potential problem of measurement error we have trimmed the top and bottom 1% of firms in all industries. Moreover, firm TFP is defined relative to the average TFP in the industry at the national level in that year. The figures show that the ratio of conditional means is monotone decreasing and converges to unity.[15] We view this evidence as support for the joint hypothesis that TFP differences across prefectures is driven by selection and that the distribution of potential entrepreneurs is similar across locations.

Finally, to illustrate that the upper tail of the TFP distributions is very similar for prefectures with high versus low aggregate TFP, we plot for each group the entire empirical distribution of firm-specific TFP conditional on their TFP being above the 95th percentile in the overall distribution. Figure 5 plots in log scales the complementary cumulative distribution functions for z in low and

[15] Note that if we had not done any trimming, the conditional mean ratio $CMR(s)$ would still be falling in s although it would not fall all the way to unity, as seen in Figure B-4.

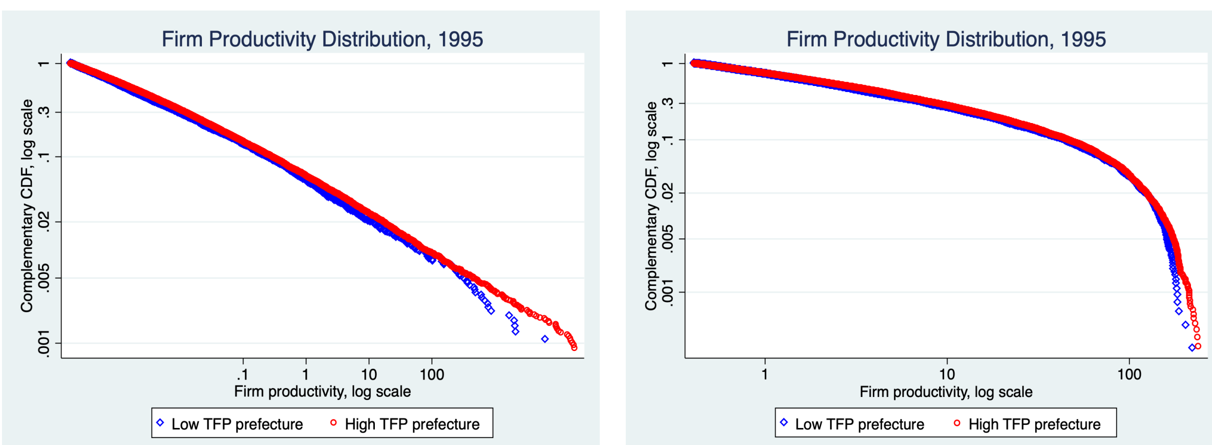
Figure 4: Conditional Mean Ratio: 1995, 2004, and 2008



Notes: In each year, the top and bottom 1% of firms in each industry, in terms of productivity, are trimmed.

high TFP prefectures, respectively, for firms in the top 5% of the overall productivity distribution. The left panel is for the full sample (no trimming) while the right panel is for the sample where we trim the top and bottom 1% of all firms in each industry. The upper tails of the TFP distributions for high-TFP and low-TFP prefectures are strikingly similar, consistent with our assumption that the distributions are the same and also consistent with the mechanism in our model through which the wedges affect the lower tail of the firm size distribution but not the upper tail. Figures B-5 and B-6 shows that this pattern is observed in all sample years. The fact that the graphs in Figure 5 are approximately linear suggests that the upper tail of the empirical firm TFP distribution is close to being Pareto distributed. The Pareto parameter for the firm-size distribution in our sample of non-state manufacturing firms in China is 1.05, which is similar to the corresponding Pareto tail value that Axtell (2001) reports for the United States, 1.06.

Figure 5: Upper Tail of the Distribution of $\ln z$ for Prefectures with Low and High Aggregate TFP



Notes: All prefectures are separated into two groups based on their aggregate Solow residual. The figure plots the complementary cumulative distribution function for the entire firm productivity distribution in 1995, conditional on firm TFP being in the top 5% of the firm TFP distribution. The left panel is for the full sample (no trimming) and the right panel is for the sample where we trim the top and bottom 1% of all firms in each industry.

4 Measuring the Wedges

We now use the benchmark model to estimate the wedges using data from the Industrial Census. The purpose is to study the drivers of the correlation structure and the regional convergence of economic performance documented in Section 2.

4.1 Log gross output and capital wedges

Following Hsieh and Klenow (2009) we use the first-order conditions for k_i and n_i from the firm's problem (2) to identify the wedges τ_y and τ_k:

$$\begin{aligned} 1-\tau_y &= \frac{1}{\alpha\eta}\frac{w_i n_i}{y_i}\ , \\ 1+\tau_k &= \frac{1-\alpha}{\alpha}\cdot\frac{w_i n_i}{(r+\delta)\,k_i}\ . \end{aligned} \tag{9}$$

In our main analysis we abstract from dispersion in firm-specific wedges within prefectures. This choice makes the analysis robust to measurement error in the firm-level data.[16] In Section 7 we extend the analysis to allow firm-specific wedges.

Using equation (9) we compute the gross output wedge and the gross capital wedge in a given prefecture. In deriving the wedges, we take into account that the technological labor-income share differs across industries and that the industrial structure differs across prefectures. Let $Y_{j,p} = \sum_{i\in(j,p)} y_i$ be the total value added for all firms in industry j in prefecture p, and let $Y_p = \sum_{j=1}^{J} Y_{j,p}$ be the total value added in prefecture p. The gross output wedge in prefecture p, Δ_p^y, is measured as the average labor-income share for each firm in that prefecture, weighted by the firm's relative value added:

$$\Delta_p^y = \sum_{j=1}^{J}\left(\frac{1}{\alpha_j\eta}\sum_{i\in(j,p)}\frac{w_i n_i}{y_i}\frac{y_i}{Y_{j,p}}\right)\frac{Y_{j,p}}{Y_p}, \tag{10}$$

where $\alpha_j\eta$ is the technological labor share of industry j.

Similarly, the gross capital wedge in prefecture p, Δ_p^k, is computed as the weighted average wage bill per unit of capital for each firm in that prefecture, weighted by the firm's relative capital stock. Let $K_{j,p} = \sum_{i\in(j,p)} k_i$ be the total capital for all firms in industry j in prefecture p, and let $K_p = \sum_{j=1}^{J} K_{j,p}$ be the total capital in prefecture p. The gross capital wedge Δ_p^k is then:

$$\Delta_p^k = \sum_{j=1}^{J}\left(\frac{1-\alpha_j}{\alpha_j}\sum_{i\in(j,p)}\frac{w_i n_i}{k_i}\frac{k_i}{K_{j,p}}\right)\frac{K_{j,p}}{K_p}. \tag{11}$$

Finally, we calculate $\alpha\eta$ in each prefecture as the weighted average of the technological labor income shares, weighted by the value added of each industry, $\alpha\eta(p) = \sum_{j=1}^{J}(\alpha\eta)_j Y_{j,p}/Y_p$.

For each firm in the Chinese Industrial Census we have data on the wage bill ($w_i n_i$), on the firm's value added (y_i), and on the firm's capital stock (k_i).[17] We use the information on the labor shares of 2-digit industries ($\alpha_j\eta$) used in Hsieh and Klenow (2009) and a decreasing returns to scale parameter of $\eta = 0.85$ as in Restuccia and Rogerson (2008).

[16] Bils, Klenow and Ruane (2021) argue that measuring firm-level distortions is highly sensitive to measurement error in firm-level data. To address this issue they assume that the distortions are constant over time, using a balanced panel of firms. This approach is not feasible for us because our focus is precisely on changes in distortions over time. Besides, many firms cannot be linked over time in our data because of changes in the assignment of firm IDs.

[17] See Appendix A for a discussion of the procedure to construct the real capital stock at the firm level.

Table 4 reports the the mean and standard deviation of the gross wedges over time. Over time, the gross output wedge is falling and the gross capital wedge is rising. This implies that τ_y and τ_k are both rising, indicating that the average distortions are increasing for non-state manufacturing firms during this period. The increase in average τ_y is driven by the labor-income share falling over time while the increase in τ_k is driven by the wage bill per unit of capital increasing over time (see equation (10) and columns (1) and (5) of Table 4). These patterns are qualitatively consistent with Bai, Hsieh and Qian (2006). Using national accounts data they document that the labor-income share is falling for the economy as a whole over this period. They also find that the return to capital in secondary industries increases slightly between 1995 and 2004. The dispersion across prefectures in the output wedge fell substantially over time, while the capital-wedge dispersion was at the same level in 1995 and 2008 (columns (3) and (6) in Table 4).

Table 4: Moments for Wedges

	$\frac{WN}{Y}$	Output wedge τ_y		$\frac{WN}{K}$	Capital wedge τ_k		Entry barrier $\ln(1-\psi)$	
		avg. Δ_y	std$(\ln \Delta_y)$		avg. Δ_k	std$(\ln \Delta_k)$	avg. $\ln(1-\psi)$	std$(\ln(1-\psi))$
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1995	0.283	1.00	0.188	0.152	1.00	0.045	-6.29	3.11
2004	0.213	0.80	0.125	0.173	1.19	0.043	-3.92	2.77
2008	0.168	0.63	0.154	0.195	1.40	0.085	-3.48	2.12

Notes: The table reports mean and standard deviation of gross wedges and entry barriers across prefectures. The means for the gross output and capital wedges are normalized to unity in 1995. The entries in columns (1) and (4) refer to ratios calculated using aggregated data on the wage bill WN, capital K, and value added Y at the national level for non-state manufacturing firms. Data source: Chinese Industry Census.

Figure 6 reports scatter plots of the wedges against the SOE output share s in each prefecture in 1995.[18] Appendix B reports corresponding plots for 2004 and 2008. The gross output wedge (left panel of Figure 6) is increasing in s, implying that non-SOE firms in high-s prefectures tend to receive subsidies while non-SOE firms in the low-s prefectures tend to be taxed in 1995. The gross capital wedge (right panel of Figure 6) is slightly increasing with s.

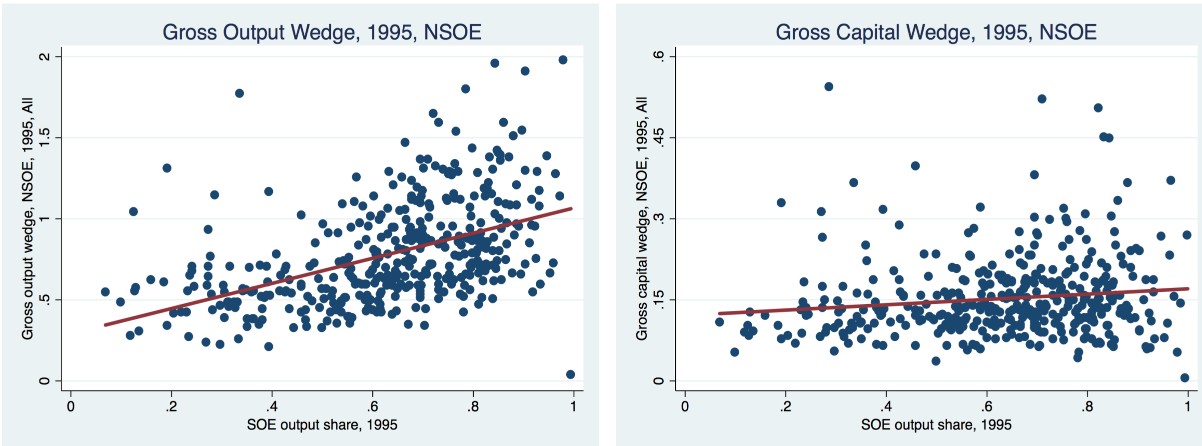
A gross output wedge, Δ_p^y, that is strongly increasing with s implies that we should observe higher wages for non-SOE firms in the high-s prefectures compared to the low-s prefectures. This pattern is the exact opposite of the empirical evidence presented in Figure 2. The fact that the capital wedge is slightly increasing in s is a force for wages to fall in s. However, the effect of τ_y dominates. This suggests that the capital and output wedges cannot on their own account for the main cross-sectional patterns of wages.

4.2 Measuring labor supply and potential entrepreneurs

Evaluating w, Z, and Γ in equations (6)-(8) requires that we measure the number of workers in the non-state sector per potential entrepreneur, N/M. This ratio could be affected by migration, structural change, and SOE employment. In particular, migration could affect both the number of manufacturing workers (through worker migration) and the number of potential entrepreneurs (through migrants starting new firms). To address this issue we present empirical data on migration

[18]We compute the wedges using data on all firms. The results based on the sample of new entrants are very similar and are provided in Appendix B.

Figure 6: Gross Output and Gross Capital Wedges, 1995, All Firms, NSOE Sector.



Notes: Each dot represents a prefecture. The left (right) panel plots the gross output (capital) wedge in the NSOE sector in 1995. The SOE output share in 1995 in each prefecture is on the horizontal axis.

of workers and entrepreneurs and lay out a model that allows us to measure N/M in light of these data.

Data from the Chinese Population Census allow us to identify migrants, i.e., individuals who are living and working in a location other than the one in which they have their Hukou (household registration), and the occupation of each individual that is working. Entrepreneur is not an occupational category, and we use Chief Executive Officer (CEO) as a proxy. We believe this is a good approximation because the owner of a firm is in most cases also the CEO. The vast majority of firms in China are controlled by an individual with more than 50% of the shares. In 2010 this number was 7 million out of 8 million firms. From now on we refer to CEOs as entrepreneurs.

Table 5 reports the shares of manufacturing workers and entrepreneurs who are migrants. Note first that migrants are prevalent, accounting for more than one fifth of the entrepreneurs and almost one third of the manufacturing workers in 2010. Moreover, the shares of migrant workers and entrepreneurs are increasing over time. Second, and most importantly, the migration rates of workers and entrepreneurs are positively correlated (see column (3) of Table 5). This shows that locations that attract many migrant workers also attract many migrant entrepreneurs.[19]

Table 5: Migrant workers and entrepreneurs

	(1)	(2)	(3)
	$m_n \equiv \frac{\text{migrant workers}}{\text{total workers}}$	$m_e \equiv \frac{\text{migrant entrepreneurs}}{\text{total entrepreneurs}}$	$corr(m_n, m_e)$
1990	0.055	0.024	0.527
2000	0.249	0.115	0.346
2010	0.304	0.220	0.340

Notes: The table reports the share of workers and entrepreneurs who are migrants. Columns (1)-(2) show the aggregate shares of migrants. Column (3) reports the correlation between the worker migrant share and the entrepreneur migrant share across counties in China. Source: China Census. See Appendix A.5 for details about measurements.

[19]Migration rates are also positively correlated with wages. Thus, both workers and entrepreneurs tend to move to locations with higher wages.

To measure M in light of migration and SOE employment, we make the following assumptions. First, we assume that every local (non-migrant) manufacturing worker is a potential entrepreneur, i.e., $M_D = N_D$. Second, we observe that the migrant share is lower for entrepreneurs than for workers. This suggests that migrants may have a lower entrepreneurial potential than locals have. To capture this effect we assume that the effective number of potential migrant entrepreneurs is given by $M_I = \kappa N_I$, where let κ denotes the entrepreneurial potential of migrants. We allow κ to vary across prefectures. Assuming that non-migrant and migrant entrepreneurs draw from the same distribution of potential TFP, the effective number of potential entrepreneurs is given by $M = M_D + M_I$. Finally, we measure the number of workers employed by non-state (NSOE) firms, N, as the total number of manufacturing workers net of SOE employment N_S, i.e., $N = N_D + N_I - N_S$.[20]

The effective number of non-state workers per potential entrepreneur, N/M, can then be expressed as follows (see Appendix A for a derivation),

$$\frac{N}{M} = \underbrace{\left(1 - \frac{N_{SOE}}{N_D + N_I}\right)}_{\text{NSOE share}} \underbrace{\frac{1 - m_e}{1 - m_n}}_{\text{non-migrant ratio}} \tag{12}$$

Namely, the ratio N/M – which we measure for each prefecture and each year – is the product of the share of manufacturing workers employed in the NSOE sector times the non-migrant ratio $(1 - m_e)/(1 - m_n)$, i.e., the share of locals among the entrepreneurs over the share of locals among the workers.

4.3 Log gross entry barrier, $\ln(1 - \psi)$

The theoretical framework outlined in Section 3 allows us to measure the entry barrier for each prefecture. Using the expression for the equilibrium wage in a prefecture (6), we derive an analytical expression for the log gross entry barrier in a prefecture:

$$\begin{aligned} \ln(1 - \psi_p) &= \frac{1 - \eta + \xi\alpha\eta}{1 - \eta} \ln w_p - \frac{\xi}{1 - \eta} \ln \Delta_p^y + \frac{\xi\eta(1 - \alpha)}{1 - \eta} \ln \Delta_p^k \\ &\quad + \ln(N_p/M_p) + \bar{\Omega}(\alpha(p), \eta, \xi, \underline{z}, \nu), \end{aligned} \tag{13}$$

where $\bar{\Omega} = -\left(\frac{1-\eta+\xi\alpha\eta}{1-\eta}\right) \ln \alpha - \frac{(1-\alpha)\xi\eta}{(1-\eta)} \ln(1-\alpha) - \ln \underline{z} + (1-\xi)\ln(1-\eta) + \ln\left(\frac{\xi-1}{\xi}\nu^{\xi-1}\right) - \left(\frac{1-\eta+\xi\eta}{1-\eta}\right) \ln \eta$. The barrier ψ_p can then be identified using data on the average wage in each prefecture w_p, combined with our measures of Δ_p^y, Δ_p^k, and $\alpha(p)$.

The remaining parameters, which are common across all prefectures, are chosen as follows. The Pareto parameter ξ is obtained by exploiting the theoretical implication that the upper tail of the firm TFP distribution is the same in all prefectures. The Pareto assumption implies that $E(z|z \geq z^*)/z^* = \xi/(\xi - 1)$. Focusing on the 5% most productive firms implies $\xi = 1.05$.[21]

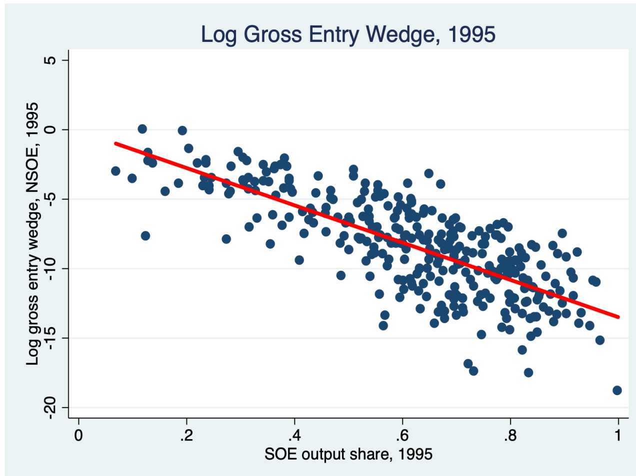
Finally, using equation (13), we compute the log gross entry barrier $\ln(1 - \psi_p)$ for all prefectures in the economy. Table 4 reports the moments for the entry barriers. The average $\ln(1 - \psi)$ is increasing and the dispersion across prefectures is falling over time. This indicates that the average and the dispersion in the barrier ψ are falling over time.

[20] Here, we introduce the class of SOE workers merely for measurement purposes. In Section 6.3 we study a model extension that incorporates SOE workers explicitly in the labor market.

[21] We normalize the fixed cost of operating a firm, ν, so that the smallest optimal size for a firm with TFP at the threshold $z = z^*$ is one worker: $n^*(z^*) = 1$. Moreover, we set the lower bound for the distribution of potential TFP, $\underline{z}$, so that the prefecture with the lowest barrier is normalized to zero ($\psi_p = 0$).

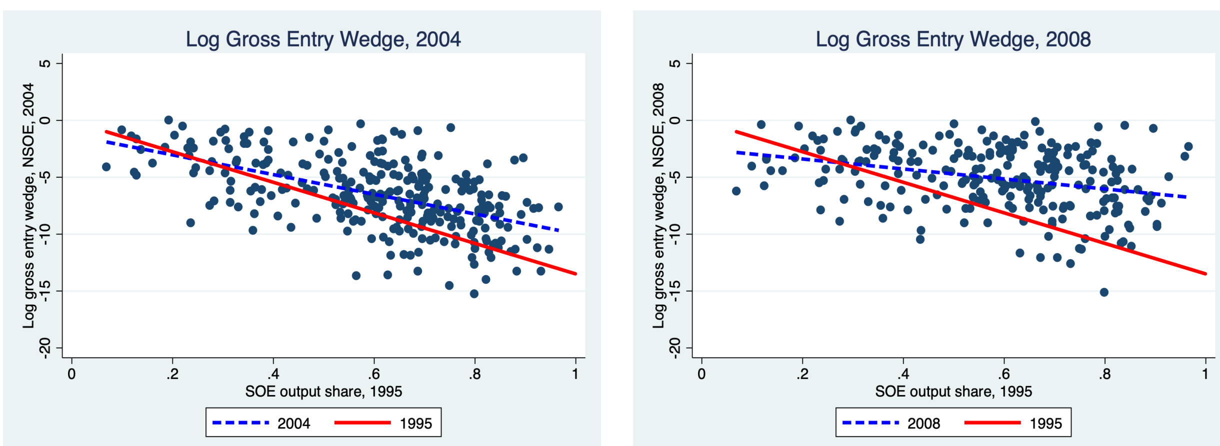
Figures 7 and 8 plot ψ against the 1995 output share of SOE firms s. The figures reveal a strong negative relationship between the entry barrier $1 - \psi$ and the SOE share s: a higher barrier ψ is associated with a larger s. This relationship becomes less pronounced over time.

Figure 7: Log Gross Entry Barriers, $\ln(1 - \psi)$, 1995.



Notes: Each dot represents a prefecture. The figure plots the log gross entry barrier in the NSOE sector in 1995, and the solid red line is the fitted regression line. The 1995 SOE output share in a prefecture is on the horizontal axis.

Figure 8: Log Gross Entry Barriers, $\ln(1 - \psi)$, 2004 and 2008.



Notes: Each dot represents a prefecture. The left (right) panel plots the log gross entry barrier in the NSOE sector in 2004 (2008), and the dotted blue line is the corresponding fitted regression line. The solid red line is the fitted regression line for the log gross entry barrier in 1995. The 1995 SOE output share in a prefecture is on the horizontal axis.

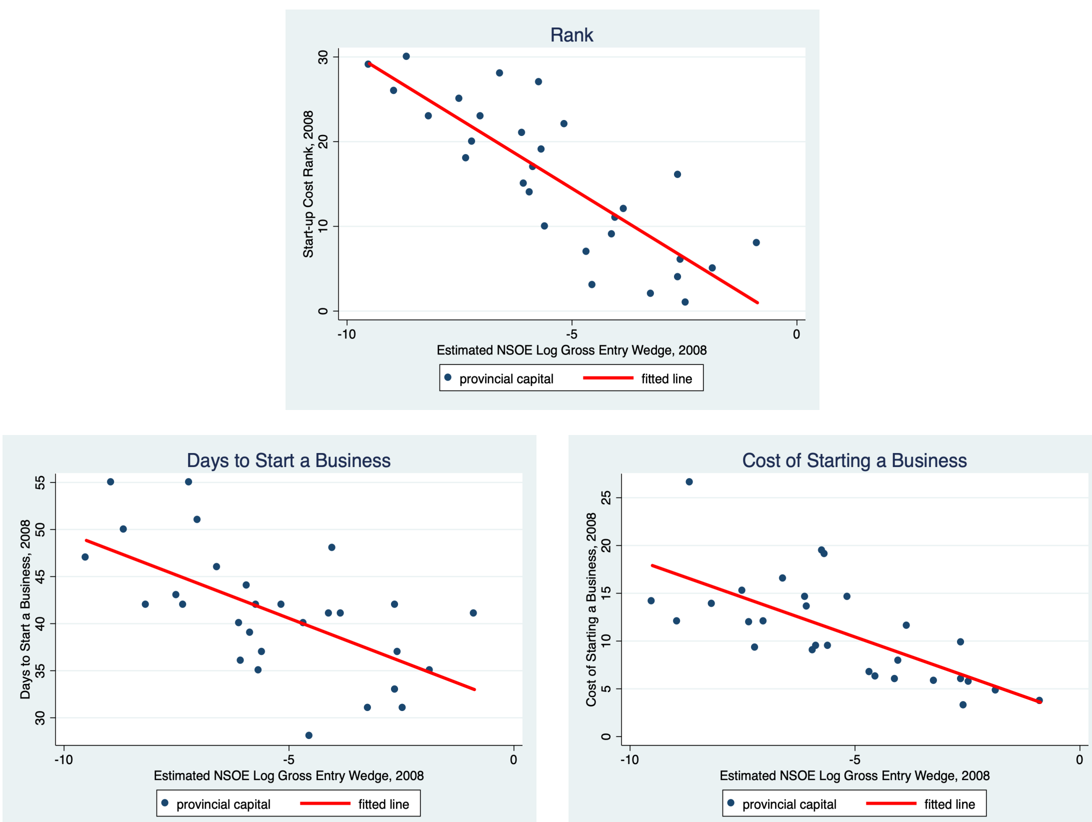
4.4 External validation of the entry barriers and wedges

We now provide external validation that our imputed entry barriers capture actual barriers to entry for private firms. To this end, we first relate our measures to those of the World Bank's "Doing

Business in China 2008" report. We then study how our measured wedges and barriers vary with the empirical entry rates of new firms.

The 2008 costs of starting a business in China. The "Doing Business in China 2008" report produced by the World Bank (2008) provides various measures of the extent to which government regulations affect private business activity. The report outlines differences in various regulations in the capital cities of 26 Chinese provinces and 4 centrally administered municipalities. We focus on the following reported indicators on how easy it is to start a business: (i) a rank computed in the report based on all available information on how easy it is to start a business; (ii) the number of days it usually takes to start a business; and (iii) the cost of starting a business, as a percent of GDP per capita. The results, reported in Figure 9, indicate that in localities where the measured entry barriers in our analysis in 2008 are higher are also the localities where the report finds high costs of starting a business. The correlations of our entry barrier $\ln(1-\psi)$ with each of the World Bank's three measures of start-up costs are respectively -0.77, -0.55, and -0.64, implying that the correlation with ψ is positive for all measures. These results provide valuable external validation for our estimates.

Figure 9: "Doing Business in China" and Entry Barriers, 2008.



Notes: Each dot represents a provincial capital city or a centrally administered municipality. Each panel shows a scatter plot of the estimated log gross entry barrier $\ln(1-\psi)$ against a World Bank measure of the cost of doing business in China in 2008: rank (top panel), days to start a business (bottom left panel), and cost of starting a business (bottom right panel). The solid red line is the fitted regression line.

Entry rates and wedges. To provide an auxiliary verification of our selection-based theory with distortions, we now study the predictions for how the three wedges should influence firm entry. As discussed in Section 3.2, increases in ψ, τ_y, and τ_k should all contribute to reduced firm entry (cf. equation (8)). However, since we did not target firm entry rates when estimating the wedges (ψ, τ_y, τ_k), the model-implied values for Γ will not necessarily be consistent with the empirical patterns for firm entry. It follows that data on entry provide an auxiliary verification of the model.

To measure the entry rate we define the rate of entry of private firms in prefecture p, $\Gamma^e_{p,t}$, as the share of employment in *new* NSOE firms – established during the last two years – relative to employment in all firms.[22] We define as new those firms that were started in year t, $t-1$, or in $t-2$. The top panel in Table 6 reports the results from the following regression in levels:

$$\ln \Gamma^e_{p,t} = \beta_0 + \beta_1 \ln(1-\tau_{y,p,t}) + \beta_2 \ln[(1+\tau_{k,p,t})(r+\delta)] + \beta_3 \ln(1-\psi_{p,t}) + \epsilon_{p,t},$$

while the bottom panel reports the results from a corresponding regression in growth rates:

$$\Delta \ln \Gamma^e_{p,t} = \gamma_0 + \gamma_1 \Delta \ln(1-\tau_{y,p,t}) + \gamma_2 \Delta \ln[(1+\tau_{k,p,t})(r+\delta)] + \gamma_3 \Delta \ln(1-\psi_{p,t}) + \epsilon_{p,t}.$$

Equation (8) predicts $\beta_1 > 0$, $\beta_2 < 0$, $\beta_3 > 0$, $\gamma_1 > 0$, $\gamma_2 < 0$, and $\gamma_3 > 0$. As is clear from the table, the data on entry rates are consistent with the predictions of the model, both in levels and in growth rates. In particular, entry barriers (higher ψ) significantly lower entry rates ($\beta_3 > 0$ and $\gamma_3 > 0$). The effects of changes in ψ, τ_y, and τ_k on entry vary over time. However, the effects of changes are quantitatively large: a one standard deviation change in $\ln(1-\psi)$ between 1995 and 2008 induces around a 30% change in the entry rate. The effects of changes in the capital and output wedges are of a similar magnitude. We interpret this as an external validation of our model and a confirmation of the mechanism through which the measured wedges and entry barriers influence the economy.

5 Accounting for convergence in TFP and wages

Our paper aims to explain the strong regional convergence in aggregate TFP and wages rates documented in Section 2. In this section we use our model as a measurement device to account for the convergence. The model features five possible factors causing changes over time in aggregate allocations and prices in a prefecture: the three wedges, the ratio of workers per potential entrepreneur, and the prefecture-specific production function (i.e., the weight on labor supply in the production function, $\alpha\eta$).[23] The production function matters because changes in industrial structure – for example, growth in the relative preponderance of labor-intensive industries – can have consequences for the labor market and wages.

Table 7 reports the annualized rate of β-convergence for aggregate TFP and wages under various counterfactual model scenarios. The first row reports the annualized 1995-2004 and 1995-2008 rates of convergence in TFP and wages.[24] To decompose the overall rates of convergence into each of the five possible sources of change, we use equations (6) and (7) to compute counterfactual rates of

[22] Our empirical measure of new firm entry differs slightly from the notion of entry in our static theoretical model, where all firms in principle would be entrants. However, the empirical measure of entry is consistent with a straightforward extension of our model to a standard dynamic Hopenhayn model incorporating firm survival and exit.

[23] Our calculations take the N/M ratio as exogenous. We discuss below how our results would change if the N/M ratio were endogenous and could respond to changes in the wedges.

[24] The numbers for the overall convergence are slightly different from the numbers reported in Table 1. There are two reasons for this. First, the calculations in Table 1 are based on data from the largest sample of prefectures for which we can calculate aggregate productivity (more than 300 prefectures). The calculations underlying Table 7

Table 6: The Firm Entry Rate and Barriers in 1995, 2004, and 2008.

	$\ln(1-\tau^y)$		$\ln(1+\tau^k)$		$\ln(1-\psi)$	
	β_1	$1sd$	β_2	$1sd$	β_3	$1sd$
1995	0.256*	10.6%	-0.216**	-9.2%	0.098***	33.9%
2004	0.101	3.4%	-0.052	-1.9%	0.032**	10.1%
2008	0.143	6.74%	-0.127	-6.8%	0.013	3.6 %
	$\Delta\ln(1-\tau^y)$		$\Delta\ln(1+\tau^k)$		$\Delta\ln(1-\psi)$	
	γ_1	$1sd$	γ_2	$1sd$	γ_3	$1sd$
1995-2004	0.689***	23.0%	-0.508***	-21.6%	0.116 ***	28.7%
1995-2008	0.504***	26.7%	-0.569***	-32.8%	0.103 ***	31.2%

Notes: The table reports the results from a regression of log gross entry rates on log gross output, capital, and entry rates in 1995, 2004, and 2008. The table also reports the percentage change in the log entry rate as a result of a one standard deviation in the variable. *** – statistically significant at 1%; ** – statistically significant at 5%; * – statistically significant at 10%.

convergence for prefecture-specific TFP and wages in 2004 (2008) under the assumption that only one factor changed between 1995 and 2004 (2008) and all other factors remained at their 1995 level. The five prefecture-specific factors we consider are (i) the entry barrier $1-\psi$; (ii) the capital wedge τ_k; (iii) the output wedge τ_y; (iv) changes in the average labor share; and (v) the effective number of workers per potential entrepreneur N/M. The effect of N/M can be further decomposed into changes in (a) the migration of workers m_n, (b) the migration of potential entrepreneurs m_e, and (c) NSOE workers per potential entrepreneur.

The main message from the first two columns of Table 7 is that changes in the entry barrier account for essentially all the convergence in TFP – the entire convergence for 1995-2004 and 85% of the overall 1995-2008 convergence. Thus, if the only change after 1995 had been the estimated changes in the entry barriers, the model would have accounted for the lion's share of convergence. This reflects the fact that the dispersion in entry barriers fell sharply over time, and more so in areas with low initial TFP and wages.

The second most important factor is the output wedge, accounting for about one quarter of

require additional data on migration, needed to measure N/M. This reduces the sample by around 15%. Second, since the model does not explicitly incorporate different industrial sectors, prefecture TFP is computed using equation (7) and an averaged prefecture-level labor share. In contrast, the TFP measures reported in Table 1 are obtained by averaging over TFP measures computed at the sectoral level in each prefecture. This leads to slight differences in the reported convergence in TFP, although the overall patterns remain unchanged.

Table 7: Annual Rate of Convergence in TFP and Wages: 1995-2004 and 1995-2008.

		TFP		Wages	
Change in		1995-2004	1995-2008	1995-2004	1995-2008
all		0.041	0.059	0.081	0.066
$(1-\psi)$		0.042	0.050	0.032	0.044
$(1+\tau_k)$		-0.005	-0.003	0.014	0.019
$(1-\tau_y)$		0.011	0.015	-0.004	-0.020
$\alpha\eta$		-0.004	-0.004	0.025	0.024
N/M		-0.001	-0.001	0.003	-0.001
-	m_N	0.002	0.002	0.011	0.009
-	m_e	-0.001	-0.001	-0.002	-0.004
-	NSOE share	-0.003	-0.002	-0.005	-0.005

Notes: The table reports the annual rate of convergence in TFP and wages across prefectures for the 1995-2004 and 1995-2008 time periods. The β-convergence coefficient for prefectures p between times t_0 and t_0+T is estimated from the regression $\left(\frac{1}{T}\right)\ln\left(\frac{y_{p,t_0+T}}{y_{p,t_0}}\right) = a - \left(\frac{1-e^{-\beta T}}{T}\right)\ln(y_{p,t_0}) + u_{pt_0,t_0+T}$, where u_{pt_0,t_0+T} represents an average of error terms, $u_{p,t}$, between times t_0 and t_0+T. Each row in the table reports what the convergence in TFP and wages would have been had only one of the listed variables changed. The row "all" allows all factors to change and captures by construction the estimated empirical convergence rate.

the convergence in aggregate TFP. This is because the dispersion across prefectures in the output wedge fell and τ_y increased more in areas where aggregate TFP was initially low.[25] The other factors play only minor quantitative roles in accounting for the convergence in aggregate TFP.

The findings for wages (the last two columns of Table 7) echo the results for aggregate TFP: the entry barrier emerges as the main factor driving the convergence of wages, accounting for two-thirds of the convergence over the 1995-2008 period and for somewhat less in the 1995-2004 period. Changes in the capital wedge and in the technological labor share also contribute to explaining parts of the convergence in wages, although these factors are quantitatively less important than the entry barrier. Note that while changes in the output wedge could explain some of the convergence in aggregate TFP, this factor contributes negatively to the convergence in wages. This reflects the fact that changes in the output wedge have opposite effects on aggregate TFP and wages (cf. Proposition 1). Recall from Table 2 that the empirical aggregate TFP and empirical wages are positively correlated, both in levels and in changes. Therefore, if the output wedge gives a positive contribution to observed convergence in TFP, it would tend to contribute negatively to observed convergence in wage rates.

[25] To see this, consider Figures B-7 and B-8 in the appendix. The figures show that the dispersion across prefectures in the measured gross output wedge $1-\tau_y$ is decreasing over time, and it is prefectures with a large SOE sector (i.e., a large SOE share) which on average experience the largest decline in $1-\tau_y$ and, hence the largest decline in implicit subsidies.

Consider, next, the role of the effective labor supply N/M, which we measure as private manufacturing employment per potential entrepreneur (N/M). Recall that changes in N/M move wages and aggregate TFP in the same direction. Table 7 documents that the measured changes in N/M are not quantitatively important in accounting for convergence in TFP and wages. This means that the changes in N/M were largely uncorrelated with changes in wages and TFP. If anything, N/M provided a slightly negative contribution to convergence, suggesting that N/M must have grown slightly more (less) in places that had relatively low (high) wages and TFP in 1995.

This conclusion does not necessarily imply that the contribution of each of the components of effective labor supply per entrepreneur is small. We now decompose the contribution of N/M into components reflecting migration of workers (m_N), migration of entrepreneurs (m_e), and the employment share of non-state firms, i.e., the share of workers in manufacturing not employed by state firms (NSOE share).

Consider, first, the role of migration of workers (m_N), holding fixed all other components, including m_e and the NSOE share. As one might expect, m_N has a large effect on convergence of wages, accounting for around 14 percent of the 1995-2008 convergence. This is because worker migration increases N/M and workers tend to migrate from low-wage to high-wage locations. The contribution to convergence of aggregate TFP is also positive, albeit quantitatively small.

Next, consider migration of entrepreneurs: m_e has a negative contribution to convergence of wages and TFP. This is because a larger m_e lowers N/M and because migration flows of entrepreneurs tend to follow migration of workers, counteracting the effect of worker migration.

Finally, consider the effect of changes in the NSOE share, i.e., the share of non-state employment in manufacturing. The contribution to convergence is negative both for TFP and wages. One reason for this is that the NSOE share increased almost everywhere but more so in the places where wages and TFP were initially low. We conclude that the reason why N/M has close to zero effect on convergence is because the effects of m_N, m_e, and NSOE share tend to neutralize each other.

Our decomposition analysis has taken the geographical distribution of workers and potential entrepreneurs as exogenous and abstracted from possible endogeneity in N and M. In particular, the decompositions in Table 7 are performed keeping all factors – including the N/M ratio – constant at their 1995 levels and changing one factor at the time. This amounts to imposing a zero elasticity of N/M to changes in the wedges. How would our analysis change if we endogenized the process of migration and manufacturing employment? General-equilibrium effects of wedges through migration decisions could in principle affect how changes in wedges propagate to allocations and prices. For example, the effect of a lower entry barrier could be mitigated if the associated increase in wages would trigger an increase in N/M.[26] We believe these potential concerns are not quantitatively important because the realized movements in N/M turn out to be quantitatively unimportant for understanding regional convergence (cf. Table 7). This suggests that the elasticity of N/M to changes in the wedges must be low, possibly because better business environments might attract both more workers and more potential entrepreneurs, rendering the N/M ratio approximately unchanged. We therefore conclude that our decomposition analysis of the effect of the wedges would be quantitatively and qualitatively robust to incorporating the endogenous effects of N/M.[27]

[26]Note that our measurements of the wedges (ψ, τ_k, τ_y) would not be affected by endogenizing N/M because the measurement equations (10)-(11) and (13) explicitly incorporate data on N/M.

[27]An alternative approach would have been to provide explicit model of migration of workers, entrepreneurs, and growth in manufacturing employment. One drawback with such an approach is that we would have to do all the analysis at the province level instead of at the prefecture level. The reason is that we have data on the province of origin for workers and entrepreneurs but not prefecture-level data on origin. We would then lose significant and valuable data on variation across prefectures within provinces. We leave such an extension for future work.

5.1 Aggregate growth

While the focus of this paper is on regional convergence, our model has implications also for aggregate growth since the wedges and migration patterns can, by construction, account for the aggregate allocations for each prefecture. We now use the model to investigate the contribution of each of the wedges for aggregate growth. Recall that we have normalized the entry barriers so that for each year t, the prefecture with the lowest entry barrier is normalized to zero, i.e., $\min_p\{\psi_{p,t}\} = 0$. Therefore, if entry barriers were falling everywhere, which seems likely, our normalization of ψ would lead us to underestimate the role of changes in entry barriers for accounting for growth.[28]

Table 8 reports the results. The most important driver of GDP growth is, as one might expect, growth in the term $\underline{z}$, which reflects general TFP growth, common to all locations. The second most important driver is the dynamics of the entry barriers, accounting for more than one-fifth of the aggregate TFP growth and about half the aggregate wage growth. The other wedges play a negative but small role for growth in wages and TFP. The output wedge has a particularly large negative contribution to wages because this wedge tended to become substantially more adverse over time. We conclude that growth in wages and output are mainly associated with TFP growth which in turn is driven by growth in $\underline{z}$ and convergence in ψ. The conclusion that aggregate GDP growth in China is primarily driven by productivity growth is in line with the finding of the broader dynamic macroeconomic literature on China, including Brandt and Zhu (2010), Cheremukhin et al. (2015), Song et al. (2011), and Tombe and Zhu (2019).

Table 8: Increase in Y, Y/N, TFP and Wages: 1995-2008.

	Y	Y/N	TFP	w
Change in log	2.27	1.80	1.19	1.16
Fraction (in percent) accounted for by:				
$(1-\psi)$	16	21	15	35
$\underline{z}$	81	102	76	160
$(1-\tau_y)$	-12	-15	9	-67
$(1+\tau_k)$	-9	-11	7	-15
N/M	24	3	-7	-13
	100	100	100	100

Notes: The table report the share of aggregate growth that can be accounted for by changes in each of the factors.

[28] A decline in entry barriers everywhere is suggested by the changing relationship between the Chinese Communist Party (CCP) and private entrepreneurs. In 2001, for example, CCP leader Jiang Zimin recommended that private entrepreneurs be allowed to join the Party. On this point, see Dickson (2003).

6 The Role of the State Sector

Section 5 established that the entry barrier is the most important factor for understanding the convergence in aggregate TFP and wages across prefectures in China. Earlier, we also presented evidence suggesting a strong positive relationship between the size of the SOE sector and the size of the entry barrier in a prefecture. So far, we have interpreted this as a mere correlation. In this section, we argue that there is a causal relationship between the size of the SOE sector and the entry barrier in a prefecture – a larger SOE sector in a prefecture is associated with a larger entry barrier in the cross section, and prefectures that experienced larger declines in their SOE sector shares also saw larger decreases in their entry barriers. As a consequence, the size of the state sector, through its effect on entry barriers, should influence wages, value-added per worker, TFP, and capital per worker in the non-state sector.

6.1 Entry barriers and the size of the state sector

6.1.1 The state sector in the cross-section

We start by studying the empirical covariates for the entry barrier in the cross section. Utilizing data for 1995, 2004 and 2008, we estimate equation (14) in the cross section, where $\ln(1-\psi)_{p,t}$ is the log gross entry barrier in prefecture p in year t, $S_{p,t}$ is the employment share of the state sector in prefecture p in year t, $X_{p,t}$ is a vector of prefecture characteristics that might also influence entry barriers, and $\epsilon_{p,t}$ is an idiosyncratic error term:

$$\ln(1-\psi)_{p,t} = \beta_0 + \beta \cdot S_{p,t} + X_{p,t}\gamma' + \epsilon_{p,t}\,. \tag{14}$$

Using data from the 1990 Census, we control for prefecture-level differences in educational attainment, labor force participation, and the share of workers in agriculture. In addition, for 1995 we have information on the profitability of SOE firms in each prefecture, as well as fiscal revenue per government worker in each prefecture. For 2004, we also have fiscal data, but do not have information from the enterprise census on profitability. Since the number of government workers is determined exogenously – set by a centrally determined policy rule as a percentage of the registered population – differences in fiscal revenue per worker must largely reflect differences on the revenue side. Effects of these variables on the entry wedge could be working through a number of alternative channels. In prefectures where SOEs were less profitable, local governments may have been more concerned about competition from non-state firms that could have reduced SOE profitability. Fewer rents in the SOEs may have also made local officials more predatory towards the non-state sector. More fiscal resources, some of which came from SOEs, may have had the opposite effect on cadre behavior towards private firms, and made it easier for local governments to make complementary investments to support the state sector.

Estimation of equation (14) may be affected by the endogeneity of the state sector as well as measurement issues. Changes in ownership codes between 1995 and 2004 make identification of state ownership and control less precise in later years. Shareholding companies become an important new ownership type in 2004, a majority of which are state-owned and controlled. Our data only identify shareholding companies and do not separately identify those under state control, leading to a likely overestimate in the size of the state sector when all shareholding companies are classified as state-owned.

To address these concerns, we use a set of alternative IVs. IV_{lag} uses as an instrument the lagged value, $S_{p,t-1}$, of the SOE employment share of prefecture p, where the lagged value refers to the SOE employment share in prefecture p, observed in the previous Chinese Industrial Census (CIC).

The next two instruments exploit information on the size of the state sector in 1978, which itself heavily reflects historical factors exogenous to prefectures such as the Third Front policies under the CCP in the 1960s and early 1970s and the Kuomointang (KMT) shift of industrial capacity inland (see Naughton (1988) and Brandt, Ma and Rawski (2017)). Reflecting these policies, coastal provinces had less manufacturing activity per capita and also a smaller role of the state sector in manufacturing than the interior provinces when reforms began in the late 1970s. We construct the IV_{1978} instrument using the sample of firms in the 1995 Census that were established in or before 1978, and compute an SOE employment share for prefecture p. Because of limited firm exit between 1978 and 1995, this provides a good measure of the size of the state sector in 1978. Finally, we run the analysis at the province level and construct the IV_{prov} instrument at the province level using 1978 provincial data on SOE output shares in industry.

We report the cross-sectional results in Table 9. In the individual cross sections for 1995, 2004 and 2008, the OLS coefficient on the size of the state sector is consistently negative and highly significant, and declines slightly over time. These results suggest that prefectures with the largest (smallest) state sectors had the highest (lowest) entry barriers. Consider now the IV regressions. In all first-stage versions of the regressions the instrument is highly significant and the R^2 is high. Compared to the OLS, the IV results suggest larger effects on the size of the state sector on measured entry barriers, with the differences most pronounced in 2004, the year in which measurement issues are likely most pronounced. This is consistent with a standard attenuation bias of the OLS due to measurement error. For 1995 we also find that entry barriers were lower in prefectures in which the state sector was more profitable, and lower in prefectures in which fiscal revenue per government worker was larger. For 2004 we do not have information on SOE profitability, but find that fiscal revenue continues to be important.

6.1.2 Changes in the size of the state sector

A potential concern for the cross-sectional results in Table 9 is that our estimates of the effect of the state sector remain contaminated by the effect of unobserved heterogeneity. There are several additional solutions. In order to eliminate any time-invariant fixed effects at the prefecture level that might be correlated with $S_{p,t}$, we can exploit the panel dimension of the data and estimate Equation (14) in first differences, or Equation (15),

$$\Delta \ln(1-\psi)_{p,t} = \beta_0 + \beta \cdot \Delta S_{p,t} + \Delta X_{p,t}\gamma' + \Delta\epsilon_{p,t}\,. \tag{15}$$

Conditional on prefecture fixed effects, changes in the share of SOEs in a prefecture may still be potentially endogenous: Unobserved shocks may affect both the share of the state sector in a prefecture and entry barriers. We also cannot rule out here the influence of measurement error related to changes in ownership codes or the possibility of reverse causality, namely, that changes in entry barriers influence the employment and output of SOEs.

A Bartik instrument. To address these concerns, we take advantage of the major 1997 policy reform embedded in China's Ninth Five-Year plan to restructure the state sector. The program was to close down loss-making state-owned firms under the slogan "Grasp the Large, Let Go of the Small" (*Zhuada Fangxiao*).[29] In addition to reducing the size of the state sector in terms of the number of firms and workers, a major objective of this reform was to concentrate state industry

[29] Some firms that were SOEs in 1995 were privatized as a consequence of this policy. However, these are minor compared to the number of *de novo* private firms that were established before 2004.

Table 9: The Entry Barrier in 1995, 2004, and 2008.

	$\ln(1-\psi)$	*OLS*	IV_{lag}	IV_{1978}	IV_{prov}
1995	e^{soe}	-12.42***	-15.36***	-13.63***	-14.44***
		(1.20)	(1.52)	(1.52)	(5.63)
	$\ln FREV$	1.32***	0.88*	1.14**	0.73
		(0.45)	(0.47)	(0.47)	(1.52)
	$\ln PROF^{soe}$	0.24*	0.24*	0.24*	0.19
		(0.14)	(0.14)	(0.14)	(0.40)
	Controls	Yes	Yes	Yes	Yes
First stage:					
	IV coefficient		0.69***	0.95***	0.89***
	st. error		(0.04)	(0.06)	(0.18)
	R^2		0.74	0.73	0.80
2004	e^{soe}	-8.61***	-12.30***	-14.44***	-18.75***
		(1.22)	(1.94)	(2.19)	(5.45)
	$\ln FREV$	1.48***	1.08**	0.84*	0.80
		(0.44)	(0.48)	(0.50)	(0.80)
	Controls	Yes	Yes	Yes	Yes
First stage:					
	IV coefficient		0.62***	0.70***	0.85***
	st. error		(0.05)	(0.07)	(0.22)
	R^2		0.54	0.48	0.64
2008	e^{soe}	-6.79***	-7.97***	-8.35***	-4.85
		(1.14)	(1.32)	(1.94)	(3.43)
	Controls	Yes	Yes	Yes	Yes
First stage:					
	IV coefficient		0.89***	0.70***	1.04**
	st. error		(0.03)	(0.06)	(0.27)
	R^2		0.78	0.45	0.49

Notes: The table reports the OLS and IV results from a regression of the log gross entry barrier on the SOE employment share (e^{soe}), fiscal revenues per government worker ($FREV$), and SOE profitability ($PROF^{soe}$) in a prefecture in 1995, 2004, and 2008. Controls include average educational attainment, agricultural employment share, and labor force participation rate in 1990. e^{soe} available in all years, $FREV$ − in 1995 and 2004, and $PROF^{soe}$ in 1995. Standard errors are in parentheses. *** − statistically significant at 1%; ** − statistically significant at 5%; * − statistically significant at 10%.

activity in sectors identified as strategic or pillar. Typically, these were more capital and skill-labor intensive sectors that were often upstream in the value chain.

We construct Bartik (1991) instruments for the changes in local SOE employment by using national-level data on the changes between 1995 and 2004 in SOE employment at the sector level.[30] A weighted average of changes at the national level should be a good predictor of prefecture-level changes in SOE employment, where the weights are the share of total SOE employment in a prefecture in 1995 in each sector k. The instrument we use scales the predicted percentage change in SOE employment by the share of SOEs in total manufacturing employment in 1995. This allows the impact of the policy change to be larger where the state sector is initially more prominent.[31]

The key identifying assumption of this instrument is that the composition of local employment in the state-sector is orthogonal to the error term in Equation (15). This is plausible in light of our earlier discussion that local SOE employment and industrial composition of SOE firms in 1995 were largely a product of central government policies before 1978 and arguably random across locations. Note also that the 1997 reform was a national policy and arguably exogenous to each prefecture.

Table 10: Change in the Entry Barrier and the SOE Sector, 1995-2004.

$\Delta \ln(1-\psi)$	*OLS*	*OLS*	IV_{Bartik}	IV_{Bartik}
ΔS	-2.06*	-1.82	-8.99**	-10.28***
	(1.16)	(1.16)	(2.63)	(2.91)
$\Delta \ln FREV$		1.24***		0.76*
		(0.40)		(0.46)
First stage:				
IV coefficient			0.62***	0.65***
st. error			(0.07)	(0.08)
R^2			0.22	0.25

Notes: The table reports the OLS and IV results from a regression of the change in the log gross entry barrier on the change in the SOE employment share (e^{soe}) and in the log fiscal revenues per government worker ($\ln FREV$) in a prefecture between 1995 and 2004. Standard errors are in parentheses. *** − statistically significant at 1%; ** − statistically significant at 5%; * − statistically significant at 10%.

In Table 10, we report the results from the fixed effects regression using the data for 1995 and

[30] Since we do not have a similarly good IV for the changes in the size of the state sector between 2004 and 2008, we limit our analysis to the changes between 1995 and 2004.

[31] Formally, the instrument is constructed as follows. Let the level of SOE employment at the national level in sector k and time t be given by $E^{soe}_{k,t}$. The national growth in SOE employment in sector k can then be expressed as $\mu^{soe}_k \equiv E^{soe}_{k,2004}/E^{soe}_{k,1995} - 1$. Moreover, let $E^{soe}_{k,t,p}$ denote the SOE employment in sector k in prefecture p in period t.The weights for constructing the Bartik instrument are the share of SOE employment accounted for by SOE firms in sector k in 1995, i.e., $\phi_{p,k} \equiv E^{soe}_{k,1995,p}/\sum_j E^{soe}_{j,1995,p}$, where these shares sum to unity, $\sum_k \phi_{p,k} = 1$. Finally, and let $E_{t,p}$ denote total manufacturing employment (SOE plus NSOE) and let $S_{t,p} \equiv \sum_j E^{soe}_{j,t,p}/E_{t,p}$ denote the SOE share of total manufacturing employment in period t and prefecture p. Our instrument for the change in SOE employment between 1995-2004 measured relative to the state sector's share of total manufacturing employment in prefecture p is given by:

$$IV_p = S_{t,p} \sum_k \phi_{p,k} * \mu^{soe}_k .$$

2004.[32] Results for the simple first differences reported in columns (1) and (2) continue to indicate that the entry barriers fell more in areas where state employment declined. However, the magnitude of the effect is significantly smaller – only one third to one quarter – than that suggested by results in Table 9. Columns (3) and (4) report Bartik instrument results, with first-stage results reported in the lower panel. Changes at the national level in sector level SOE employment are a very good predictor of changes in the share of the SOEs by prefecture. The IV coefficient on the size of the state sector is also significantly larger than the OLS FE estimates in (1) and (2), and the magnitude of the coefficient is more in line with our estimates from the cross-section. These estimates imply that the size of the state sector has a causal and economically significant negative effect on entry barriers at the prefecture level. Moreover, in prefectures where the SOE employment is predicted to fall more between 1995 and 2004, entry wedges experience an even faster decline.

6.2 The state sector and regional convergence

The previous section established that employment changes in state-owned firms cause lower entry barriers. Motivated by this evidence, we now revisit our empirical results in Table 1 on regional convergence and investigate how changes in state employment affect the observed growth rates of wages, output per worker, capital per worker, and TFP in NSOE firms. Due to the endogeneity concerns discussed above, we address this question by applying our Bartik instrument.[33] We include in the regression the initial level of $\log(x)$ in 1995 and province-level fixed effects.

Table 11 contains two important results. First, the results on regional β convergence are robust to including instrumented changes in SOE employment and province fixed effects as explanatory variables. When these controls are included, the rate of convergence is larger for TFP and is slightly smaller for wages, output per worker, and capital per worker – all relative to the results in Table 1.[34] Second, and more importantly, the instrumented changes in SOE employment have strong and significant effects on regional growth rates of wages, value-added per worker, TFP, and capital per worker of non-state firms.[35] This establishes a causal role of SOE employment for NSOE firm performance. Namely, in prefectures where the SOE employment is predicted to fall more between 1995 and 2004, the prefecture experiences faster growth in wages, output per worker, capital per worker, and aggregate TFP.

To illustrate the quantitative magnitudes of the results in Table 11, we evaluate the effect of a one standard deviation change of the instrument, a decline of about 9 percent of total manufacturing employment. In this case, the 1995-2004 *annualized* growth in wages is predicted to increase by 0.6 percent, output per worker by 1.2 percent, aggregate TFP by 0.4 percent, and capital per worker by 1.8 percent. These effects are economically significant.

[32] The analysis in this section has focused on the entry barrier. However, state employment could matter also for the other wedges. In Table B-2 we report results for the effect of the Bartik-instrumented change in SOE employment on the gross output and capital wedges. That table shows that a larger predicted reduction in SOE employment causes a reduction in τ_k and an increase in τ_y. All effects are significant.

[33] Simply including the observed changes in SOE employment as an explanatory variable in the growth regressions would be subject to the same potential endogeneity and reverse causality issues as the entry-barrier regressions in Section 6.1. Potential general-equilibrium effects of SOE layoffs on non-state firms would, if anything, run against our subsequent findings. For example, if laid-off state workers were to put downward pressure on wages in private firms, then more SOE layoffs should be associated with lower wage growth in private firms.

[34] Once the controls are added, the annualized rate of convergence increases to 4.6% for TFP, up from 3.7% in Table 1. For wages, output per worker, and capital per worker the rates of convergence fall to 5.5%, 6.6%, and 8.4%, respectively.

[35] The results for aggregate TFP growth are robust to alternative weighting schemes for calculating TFP growth. In Table A-1 in Appendix A.4 we show that the coefficient on the Bartik instrument is robust to using instead the relative share of value added Y of each industry in that prefecture and to basing the weights entirely on 1995.

Table 11: NSOE growth and the role of the state sector, 1995-2004

	$\Delta \ln w$	$\Delta \ln(\text{VA}/N)$	$\Delta \ln TFP$	$\Delta \ln(K/N)$
$\Delta \hat{S}_p$	-0.61***	-1.24***	-0.33	-1.90***
	(0.18)	(0.30)	(0.22)	(0.40)
$\ln x_{1995}$	-0.36***	-0.47***	-0.34***	-0.51***
	(0.05)	(0.05)	(0.04)	(0.07)
Province F.E.	Yes	Yes	Yes	Yes
First stage				
β_{IV}	0.83	0.85	0.87	0.86
s.e.	(0.09)	(0.09)	(0.09)	(0.09)
adj. R^2	0.38	0.36	0.36	0.36

Notes: The table reports the 1995-2004 growth in average wages, value added (VA) per worker, TFP, and log capital per worker for NSOE firms across prefectures. The role of the state sector growth is instrumented by IV_b. Notes: Standard errors are in parentheses. *** – statistically significant at 1%; ** – statistically significant at 5%; * – statistically significant at 10%.

Based on this evidence, we conclude that the 1997 SOE reform was a major contributor to growth and regional convergence for private sector in China. Moreover, our results in Section 6.1 suggest that a key mechanism for this was though the effect of SOEs on the entry barriers.

6.3 A political economy model of wedges

Why would the size of the state sector matter for the wedges and barriers facing non-state firms? To address this issue, this section provides a version of the benchmark model extended to incorporate the presence of SOEs alongside private firms. The purpose of the extension is to develop a simple political economy model for the determination of the wedges that can provide a theoretical motivation for the causal relationship between the observed entry barriers and the size of the SOE sector that we documented in Tables 9 and 10. We emphasize the important role played by local cadres for explaining this link.

We assume that there is a unit measure of potential SOEs with the same production function as NSOEs, eq. (1). For simplicity we abstract from wedges on output and capital for SOEs (i.e., $\tau_y^{SOE} = \tau_k^{SOE} = 0$). We model the labor market the same way as Song et al. (2011), where the SOEs hire workers in competition with the NSOE sector.[36] Following the analysis in Section 3, the aggregate labor demand of SOEs is then given by

$$\Lambda_{SOE} = \frac{\xi \underline{z}}{\xi - 1} \left(\frac{1-\eta}{\nu} \right)^{\xi - 1} \left(\frac{(1-\alpha)\eta}{r+\delta} \right)^{\xi \frac{(1-\alpha)\eta}{1-\eta}} \left(\frac{\alpha \eta}{w} \right)^{1 + \xi \frac{\alpha\eta}{1-\eta}} .$$

We assume that the three wedges for private firms, (ψ, τ_y, τ_k), are set by the local government in the prefecture. We label the decision maker as the *local cadre.* We impose two constraints on the

[36] For simplicity we assume that SOEs and NSOEs pay the same wages. Forcing SOEs to pay an exogenous wage premium for workers would not affect the qualitative results. The key assumption is that SOEs compete with private firms for some factor in short supply, be it workers, high-skilled workers, managers, land, or other input factors.

wedges. First, they must be non-negative.[37] Second, local cadres set the wedges to ensure that the equilibrium state employment in the prefecture meets an exogenous target $\Lambda_{SOE} = \bar{\Lambda}_{SOE}$, which is set by higher levels of government and can differ between prefectures.[38]

Note that an increase in any of the wedges will increase SOE employment. The cadre therefore faces a trade off between the various wedges when meeting the hiring requirement. To see this, note that market clearing requires that non-state labor demand is $N = 1 - \bar{\Lambda}_{SOE}$ (where aggregate labor supply has been normalized to unity). Substituting NSOE labor demand and the equilibrium wage rate into this market-clearing condition yields a condition linking the wedges to the hiring requirement,

$$\frac{1-\bar{\Lambda}_{SOE}}{\bar{\Lambda}_{SOE}} \frac{1}{(1-\psi)} = (1-\tau_y)^{\frac{\xi}{1-\eta}} \left(\frac{1}{1+\tau_k}\right)^{\xi \frac{(1-\alpha)\eta}{1-\eta}} . \tag{16}$$

It follows that the (target) state employment $\bar{\Lambda}_{SOE}$ is increasing in each of the wedges, (ψ, τ_k, τ_y). The reason is that an increase in any of the wedges lowers NSOE demand for workers and, hence, equilibrium wages. This affects SOE employment along both the extensive and the intensive margin: with lower wages less efficient SOE firms can operate (i.e., more SOE entry), and the lower wages make it optimal for each SOE firm to hire more workers.

We focus on the case where $\bar{\Lambda}_{SOE} > 1/2$ to ensure that the SOE employment constraint is relevant in the sense that state firms need to be favored relative to non-state firms in order to satisfy the SOE hiring constraint.

Consider now the objective of the local cadre. We assume that the cadre wants to maximize profits for an entrepreneur, conditional on obtaining a licence and their TFP, z. This captures the notion of *crony capitalism*, i.e., that the cadre may want to help a friend (crony) who is a potential NSOE entrepreneur (see e.g. Bai, Hsieh and Song (2018) for a motivation for this assumption), but that the cadre has limited policy instruments for achieving this goal. On the one hand, the cadre can subsidize the entrepreneur by choosing low capital or output wedges (although all firms will benefit from these subsidies). On the other hand, the cadre can restrict entry for anonymous potential entrepreneurs by setting a large ψ, while at the same time guaranteeing that their entrepreneur friend will be allowed to operate.

Conditional on operating the firm the entrepreneur's profits $-$ net of the implicit taxes on capital and output $-$ are given by:

$$\Pi(z) = \frac{z}{1-\psi} \frac{1-\bar{\Lambda}_{SOE}}{\left(\bar{\Lambda}_{SOE}\right)^{\frac{1-\eta}{\xi\alpha\eta+1-\eta}}} \cdot \frac{1-\eta}{1+\mu} \left(\frac{\xi \underline{z}}{\xi-1} \left(\frac{1-\eta}{\nu}\right)^{\xi-1} \left(\frac{(1-\alpha)\eta}{r+\delta}\right)^{\xi\frac{(1-\alpha)\eta}{1-\eta}} \right)^{\frac{1-\eta}{\xi\alpha\eta+1-\eta}} .$$

The profits $\Pi(z)$ are increasing in the entry barrier. The reason is that entrepreneurial talent is a scarce resource, and with fewer potential entrepreneurs the profits are higher conditional on z. However, note that the entrepreneur's expected profits are independent of the output and capital wedges. This is due to the fact that profits $\Pi(z)$ can be expressed as a function z, ψ, and the right-hand side of equation (16). Thus, conditional on ψ and $\bar{\Lambda}_{SOE}$, any combination of (τ_k, τ_y) that satisfies equation (16) will give rise to the same profits. A lower τ_k will therefore have to be offset by a higher τ_y in order to satisfy the hiring constraint, rendering profits invariant.

Under these assumptions about the local cadre's problem, we find that the optimal way to satisfy the hiring requirement is to set the capital and output wedges to zero and set ψ so as to

[37] The constraint $\psi \geq 0$ is natural. The constraints $\tau_y \geq 0$ and $\tau_k \geq 0$ can be motivated by limited government funds ruling out outright subsidies.

[38] See, for example, Brandt and Zhu (2000) and Wang (2017) for possible political economy motivations for such a requirement on state employment.

satisfy equation (16). This implies a high correlation between SOE employment Λ_{SOE} and entry barriers ψ. We state this result as a proposition (proved in the text).

Proposition 2 *The constrained optimal choice of wedges* (ψ, τ_y, τ_k) *is to set* $\tau_k = \tau_y = 0$ *and* $\psi > 0$. *Moreover, an exogenous increase in* $\bar{\Lambda}_{SOE}$ *implies a larger entry barrier* ψ.

In the empirical analysis in Section 6.1 we introduced two Instrumental Variables for $\bar{\Lambda}_{SOE}$: lagged SOE employment and the Bartik instrument. First, the central and provincial governments may want the local government to maintain the current level of SOE employment, thereby upholding the legacy of the state sector. In this case the historical level of state employment in the prefecture should be expected to influence $\bar{\Lambda}_{SOE}$. Second, the 1997 SOE reform, which was imposed by the central government, involved large-scale reductions in state employment in industries deemed to be non-strategic from the point of view of national security. We interpret this as an exogenous reduction in $\bar{\Lambda}_{SOE}$. The implication of Proposition 2 and our choice of Instrumental Variables for the hiring constraint $\bar{\Lambda}_{SOE}$ is that the entry barrier should be larger in areas with historically higher levels of state employment and should fall more in areas where state employment was more significantly scaled back after 1997. This is consistent with the results in Tables 9 and 10.

6.3.1 Discussion: Why would SOEs matter for new private start-ups?

The political economy model of the determination of wedges above assumes that the local government faces pressure to meet an exogenous target for state employment, $\bar{N}_{SOE}$. We motivate this assumption as follows. Local officials, e.g. party secretaries and mayors, are appointed by higher levels of government and are tasked with multiple objectives. Much of the focus in the literature – see e.g. Li and Zhou (2005) and Xu (2011) – is on the high-powered incentives local leaders have to promote economic growth, but equally important through the nomenklatura system is their role in supporting state-owned enterprises. The performance of SOEs is important for Communist Party and for officials at all levels. Indeed, state-owned firms themselves have multiple mandates. As a major source of employment in the cities, SOEs have been perceived as policy instruments for maintaining social stability, especially during economic downturns (Wang (2017)). Local cadres are beneficiaries of the success of SOEs in meeting the objectives of higher levels of government and of the Communist Party. SOEs are also potentially important sources of local government revenue and rents for local officials, often in the form of valuable jobs for family members and relatives as well as through highly lucrative business relationships with these same firms.

A key premise in the political economy model of this section is that local government has access to policy instruments that may suppress the entry of private firms, and that local cadre often apply such policies, especially in areas where the state sector is prevalent. Market liberalization and easier entry for new private firms arguably pose threats to the position of the SOEs through pressures in the product market, and more importantly, through the competition for local scarce factors. Thus, by mitigating the growth of private firms, local cadre can prevent the flight of the most capable managers and workers (and other scarce factors) from the SOEs to the private sector. Whiting (2006) documents that local officials erect various forms of barriers to entry and argues that the motivation for engaging in such behavior is that they seek to protect firms owned by local governments. This behavior manifests itself in the form of making it more difficult to obtain access to land, electricity and other scare intermediate inputs, over which local governments have some discretion and control. In addition, in newly emerging sectors, ministries have often restricted entry by issuing few licences and by allocating these licenses to SOEs (Huang (2003)). More generally, local cadre can use their discretion over granting business licenses and influence over access to critical inputs to enrich family and friends in their networks, and thus themselves.

Barriers to entry in environments in which SOEs are dominant also take more indirect forms. Suppliers to state-owned firm must typically go through a lengthy certification process. On paper, this certification is to ensure that the supplier has the capabilities to meet the requirements laid out by the SOE. However, in practice the purpose of this process is to limit the access to act as a supplier to the SOEs to firms linked through personal networks either to officials in the state sector or local government (Interviews, 2017).

7 Extension: Heterogeneity of Wedges across Firms

In the model we have analyzed so far we assumed that the capital and output wedges were the same for all firms in a prefecture. In this section we extend our benchmark model to allow capital and output wedges to be firm-specific. Namely, we assume that there is heterogeneity in τ_{ik} and τ_{iy} across firms not only across locations but also across firms within each prefecture. We maintain the assumption that all prefectures have the same distribution f of potential z. However, due to selection in participation, there will be, in equilibrium, a correlation between z and wedges among firms that choose to operate.

Each potential entrepreneur can observe both her potential TFP, z_i, and her potential wedges, $\{\tau_{ik}, \tau_{iy}\}$, before deciding to enter. As we shall see, the entry decision of the potential entrepreneur depends on the entrepreneur's realized wedges $\{\tau_{ik}, \tau_{iy}\}$. Therefore, the equilibrium distribution of observed TFP will be correlated with the wedges, even though the distribution of potential TFP is, by assumption, independent of the wedges. In order to ensure that the problem is analytically tractable we assume that the distribution of *potential* wedges is jointly log-normal across firms in each prefecture. Denote the density function as $g(\tau_k, \tau_y)$, and let the moments be given by:

$$\begin{aligned} E\left(\ln(1+\tau_k)\right) &= \ln\left(1+\bar{\tau}_k\right) - \frac{\sigma_k}{2} \\ E\left(\ln(1-\tau_y)\right) &= \ln\left(1-\bar{\tau}_y\right) - \frac{\sigma_y}{2} \\ var\left(\ln(1+\tau_k)\right) &= \sigma_k \\ var\left(\ln(1-\tau_y)\right) &= \sigma_y \\ cov\left(\ln(1+\tau_k), \ln(1-\tau_y)\right) &= \sigma_{ky}. \end{aligned} \tag{17}$$

Note that the dispersion in wedges are mean-preserving spreads, implying that $E\left((1+\tau_k)\right) = 1+\bar{\tau}_k$ and $E\left((1-\tau_y)\right) = 1-\bar{\tau}_y$. Moreover, this extended model nests our benchmark model when $\sigma_k = \sigma_y = \sigma_{ky} = 0$.

Conditional on the individual state $s_i = \{z_i, \tau_{ik}, \tau_{iy}\}$, the optimal firm choices are still given by equations (3)-(4). Note in particular that the cutoff threshold $z^*(\tau_{ik}, \tau_{iy}, r, w)$ now differs across firms. Given the distributional assumptions it is possible to solve analytically for the wage that clears the labor market and for the associated aggregate Solow residual. We summarize these results in the following proposition.

Proposition 3 *Suppose there is within-prefecture heterogeneity in capital and output wedges. The equilibrium wage in a location is then given by*

$$\begin{aligned} \ln w &= \mu(1-\eta)\ln\left[(1-\psi)\underline{z}\frac{M}{N}\right] \\ &\quad +\mu\xi\ln\left(1-\bar{\tau}^y\right) - \mu\xi\eta\left(1-\alpha\right)\ln\left(\left(1+\bar{\tau}^k\right)(r+\delta)\right) + \Omega \\ &\quad +\mu\xi\left(\frac{\xi}{1-\eta}-1\right)\frac{\sigma_y}{2} + \mu\xi\eta\left(1-\alpha\right)\left(\frac{\xi\eta\left(1-\alpha\right)}{1-\eta}+1\right)\frac{\sigma_k}{2} - \mu\xi^2\left(1-\alpha\right)\frac{\eta}{1-\eta}\sigma_{ky}, \end{aligned} \tag{18}$$

Moreover, the Solow residual is given by

$$\begin{aligned}\ln Z &= \mu\alpha\eta(1-\eta)\ln\left[(1-\psi)\underline{z}\frac{M}{N}\right] - \mu(1-\eta)\ln(1-\bar{\tau}_y) \\ &+\mu(1-\eta)(1-\alpha\eta(1-\xi))\ln(1+\bar{\tau}_k) + \hat{\Omega} \\ &-(\xi-(1-\eta))\left(\frac{1}{1-\eta}+\mu\right)\frac{\sigma_y}{2} \\ &-[2(1-\eta)(1-\alpha\eta)+\alpha\xi\eta(2-\eta(1+\alpha))]\frac{\mu(1-\eta+(1-\alpha)\xi\eta)}{(1-\eta)}\frac{\sigma_k}{2} \\ &+\left((1-\eta)(\eta(1-\alpha)+(\xi-1)\alpha\eta+1)+\alpha\eta^2(1-\alpha)\xi\right)\frac{\mu\xi}{1-\eta}\sigma_{ky}.\end{aligned} \tag{19}$$

The Solow residual is falling in ψ, N/M, σ_y, σ_k, and $-\sigma_{ky}$, while it is increasing in $\bar{\tau}_k$ and $\bar{\tau}_y$. The equilibrium wage is falling in ψ, N/M, $\bar{\tau}_k$, and $\bar{\tau}_y$, while it is increasing in σ_y, σ_k, and $-\sigma_{ky}$.

Note first that if there is no heterogeneity in wedges (i.e., $\sigma_y = \sigma_k = \sigma_{ky} = 0$), then the equilibrium wage rate and Solow residual will be equal to their counterparts in the model without cross-sectional dispersion (cf. eq. (6) and (7)). Therefore, the qualitative effects of ψ, $\bar{\tau}_k$, and $\bar{\tau}_y$ are the same as before (see Proposition 3).

Consider now the effect of the second moments. As is clear from equation (18), a mean-preserving spread of the wedges – represented by an increase in the variance of $\ln(1+\tau_k)$ or $\ln(1-\tau_y)$ – will increase the wage rate. Similarly, an increase in the correlation between τ_k and τ_y, i.e., a smaller covariance σ_{ky}, will also increase the wage rate. The reason is that larger dispersion in firm-specific wedges and a tighter link between τ_k and τ_y, will increase aggregate labor demand. On the one hand, the large firms will become larger, which obviously increases demand for workers. On the other hand, while the small firms become smaller (or drop out), this will not lower much the demand for workers since they already hired few workers.

Consider now the expression for the aggregate TFP in a prefecture (equation (19)). Note that the comparative statics of the second moments on aggregate TFP are the opposite of those on the wage rate. Namely, the aggregate TFP will *fall* in response to a mean-preserving spread in capital and output wedges (i.e., increases in the variances of $\ln(1+\tau_k)$ and $\ln(1-\tau_y)$). Moreover, TFP will also fall in response to a higher correlation between τ_k and τ_y (i.e., a smaller σ_{ky}).[39] The reason is negative selection: low-TFP firms with capital and output subsidies (i.e., negative τ_k and τ_y) will be large while high-TFP firms with large capital and output wedges will be small or maybe even induced to drop out.

We conclude that the comparative statics for the cross-sectional dispersion in τ_k and τ_y (i.e., comparative statics of $\{\sigma_y, \sigma_k, -\sigma_{ky}\}$) are qualitatively similar to the comparative statics for the prefecture-specific gross output wedge $1-\bar{\tau_y}$ which we listed in Table 3. In particular, changes in the dispersion have opposite effects on wages and aggregate TFP.

We now revisit measurement of the wedges when incorporating cross-sectional dispersion in output and capital wedges. To this end, we must identify the wedges while taking into account the equilibrium distribution of observed allocations and wedges. Proposition 4 outlines a strategy for estimating the entry barriers based on the first and second moments of the *observed* wedges.

Proposition 4 *The parameters of the joint log-normal distribution of potential wedges, $\{\bar{\tau}_k, \bar{\tau}_y, \sigma_k, \sigma_k, \sigma_{ky}\}$, can be identified by the following cross-sectional first and second moments for observed wedges.*

[39] The comparative statics for the Solow residual echo the theoretical finding of Hsieh and Klenow (2009). However, they did not study the comparative statics on the equilibrium wage rate.

$$\frac{std\left\{(1+\tau_k)(r+\delta)|z\geq z^*\right\}}{E\left\{(1+\tau_k)(r+\delta)|z\geq z^*\right\}} = \sqrt{\exp(\sigma_k)-1} \tag{20}$$

$$\frac{std\left\{(1-\tau_y)|z\geq z^*\right\}}{E\left\{(1-\tau_y)|z\geq z^*\right\}} = \sqrt{\exp(\sigma_y)-1} \tag{21}$$

$$\frac{cov\left\{(1+\tau_k)(r+\delta),(1-\tau_y)|z\geq z^*\right\}}{E\left\{(1-\tau_y)|z\geq z^*\right\}\cdot E\left\{(1+\tau_k)(r+\delta)|z\geq z^*\right\}} = \exp(-\sigma_{ky})-1 \tag{22}$$

$$E\left\{(1-\tau_y)|z\geq z^*\right\} = \exp\left(\ln(1-\bar{\tau}_y)+2\frac{\xi}{1-\eta}\frac{\sigma_y}{2}-\left(\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\sigma_{ky}\right) \tag{23}$$

$$E\left\{(1+\tau_k)(r+\delta)|z\geq z^*\right\} = \exp\left(\ln[(1+\bar{\tau}_k)(r+\delta)]+2\xi(1-\alpha)\frac{\eta}{1-\eta}\frac{\sigma_k}{2}-\frac{\xi}{1-\eta}\sigma_{ky}\right) \tag{24}$$

The proposition implies that even though there is selection in which firms choose to enter (namely, firms with low τ_k and τ_y will be more likely to enter) the prefecture-specific moments for the distribution of wedges can still be identified by using a suitable empirical strategy. In particular, equations (20)-(24) show that the prefecture-specific means $\bar{\tau}_y$ and $\bar{\tau}_k$ and the variance-covariance matrix of the wedges can be identified using the coefficient of variation of the observed firms, i.e., the firms that were selected to enter.

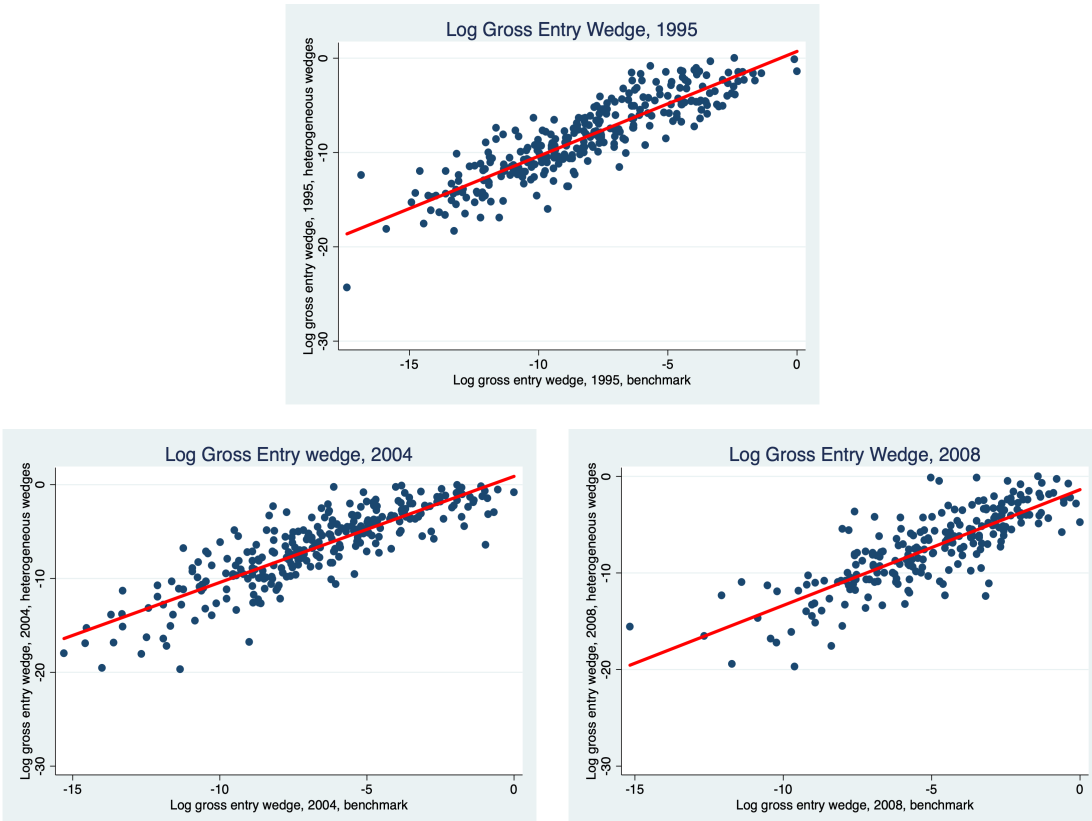
Given the prefecture-specific moments $\{\bar{\tau}_y, \bar{\tau}_k, \sigma_y, \sigma_k, \sigma_{ky}\}$ and the wage rate w, we can identify the entry barrier ψ by inverting equation (18), as we did in Section 4. Several results are worth pointing out. First, the entry barriers in the heterogeneous-wedge model are highly correlated with the entry barriers in the benchmark model. Figure 10 plots them $-$ for 1995, 2004, and 2008 $-$ against the entry barriers in the benchmark model. The correlation is high: 0.88 in 1995, 0.86 in 2004, and 0.82 in 2008. Moreover, the entry barriers decline over time and tend to be higher in prefectures with a high SOE output share. Second, when accounting for convergence in wages and TFP over time, as presented in Table B-3, the entry barriers continue to account for a large share of the convergence. Overall, the dispersion in the capital wedges has no effect on wage and TFP convergence while the dispersion in the output wedges and the covariance between the output and capital wedges affect only the convergence in wages. Finally, as presented in Table B-4, a decline in a prefecture's SOE share over time leads to a decline in its entry barrier, with both the OLS and Bartik instruments estimates being large and negative, although the Bartik instrument results are estimated with less precision.[40]

8 Conclusion

This paper studies regional economic growth in China. Using firm-level data from the Chinese Industrial Census, we document that China experienced a remarkable regional convergence in wages, TFP, productivity, and capital per worker in non-state manufacturing firms at the prefecture level during the period 1995 to 2008. The main aim of the paper is to analyze the factors behind the initial dispersion and subsequent regional convergence in wages and TFP. To this end we propose a tractable version of the Hopenhayn (1992) model of firm heterogeneity and new firm creation, extended to incorporate three distortions: standard capital and output wedges, common to all firms in a prefecture, and a novel entry barrier. The general equilibrium model is solved analytically. It features endogenous aggregate TFP and allows us to measure the three wedges using data on aggregate allocations for wages, output, employment, and capital.

[40]Due to what we perceive as severe measurement error in the data, we drop the top and bottom 15% of the firms in terms of output and capital wedges in each prefecture. Thus, although the results from the heterogeneous-wedge model are insightful, we consider our benchmark model as our preferred choice.

Figure 10: Log Gross Entry Barriers, Benchmark Model and Model with Wedge Heterogeneity.



Notes: Each dot represents a prefecture. The graphs plot the log gross entry barriers in the benchmark model and in the model with heterogeneous wedges in 1995, 2004, and 2008. The solid red line is the fitted regression line.

Using the model as an accounting device, we exploit the aggregate prefecture-level data to measure these distortions for each prefecture. We document that entry barriers are salient in accounting for the regional dispersion and subsequent convergence in China after 1995. In contrast, the capital and output wedges play only a limited role in explaining the empirical regional convergence. Given the preponderance of the entry barriers in accounting for economic performance, we investigate the empirical drivers of these distortions. Based on a Bartik instrumental variable approach exploiting the major 1997 SOE reform that resulted in a large decline in the role of state-owned firms in many industries, we find that the presence of state-owned firms was a source of larger entry barriers for non-state firms. We provide a political economy model of distortions to motivate the empirical link between SOEs and entry barriers for non-state firms.

Our analysis has made a number of simplifying assumptions, often dictated by data limitations. To minimize the role of measurement error we focus on prefecture-level distortions and abstract from firm-level distortions within a prefecture. However, our main findings turn out to be robust to allowing firm-level dispersion in capital and output wedges. A natural next step would be to extend our static analysis to a dynamic setting. Such an extension would yield rich theoretical predictions for new firm creation and the distribution of firms that could help discriminate between alternative sources of local growth and convergence such as entry barriers, agglomeration, human capital, input prices, and exposure to trade and competition.

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Online Appendix for
"Barriers to Entry and
Regional Economic Growth in China"

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A Data

A.1 Dataset

Our main data source is the 1995, 2004, and the 2008 Chinese Industrial Census (CIC) carried out by China's National Bureau of Statistics (NBS).[41] The CIC covers all of the manufacturing sector[42] and provides rich firm-level data on gross output, value added, employment, the gross capital stock, depreciation, total wages, as well as information on firm year of establishment, ownership type, and main sector of business. For these three years, we have firm-level records on 0.53, 1.37 and 2.08 million firms, respectively.[43]

In order to make these data comparable across the three census years, we needed to address a number of issues related to changes that occurred in China's industrial classification system, ownership categories, and prefecture boundaries. We draw on concordances described in Brandt et al. (2012) for ownership types and industrial sectors, and extend the concordance on prefecture boundaries in Baum-Snow et al. (2017) to cover all prefectures. We also utilize deflators developed by Brandt and Rawski (2008) for the purposes of constructing real measures of industrial output, and estimates of the real capital stock.

A.2 Defining non-state-owned enterprises

The NBS provides a detailed breakdown of firm type by ownership for firms in the CIC. In 1995, there are 12 ownership categories, of which one covers state-owned firms. On the basis of the slightly more detailed classification in use in 2004 (and 2008), we define state owned to include firms listed as state-owned, state solely-funded limited liability companies, and shareholding companies. Shareholding companies during this period are largely state-controlled, but a subset of these firms is not. Non-state-owned enterprises are then defined as all enterprises that are not state-owned. A stricter definition of state-owned would exclude the shareholding companies. In addition, for each firm we have a breakdown of equity in the firm between state, collective, private, legal person, and foreign. Alternative definitions of SOE and NSOE ownership can be constructed on the basis of these variables, as well as using a combination of the categorical ownership variables and data on ownership equity. The latter information is especially helpful for identifying state-controlled shareholding companies.

[41]We also draw on firm-level data for 1992 on all independent accounting units (0.39 million), which covers a slightly smaller subset of firms than the census and has information on a smaller set of variables.

[42]The 2004 and 2008 Census also provide data for the service sector, but unfortunately similar information was not collected in 1995.

[43]The firm-level records are not exhaustive, but cover in upwards of 90 percent of industrial activity.

A.3 Constructing real capital in 1995, 2004, and 2008

We construct measures of real capital using a procedure similar to the one in Brandt et al. (2012) and Hsieh and Song (2015). The source of the measurement problem is that firms do not report the real capital stock. Instead, they report the value of their accumulated fixed investments at original purchase prices.

To estimate the capital stock in the year when we have data (say year T), we first estimate the capital stock of each firm in the year it was established (say year $T - n$). The identifying assumption is that the firm's capital grew at the same rate as aggregate capital in the firm's 2-digit industry and province cell. The annualized growth rate in nominal capital for each industry-province cell is then estimated using NBS data.

We then use the same aggregate capital growth rates to estimate the accumulated nominal investments for each year the firm has existed. The difference in accumulated nominal investments between year $t - 1$ and t represents the nominal investment in year t. The nominal investments are then deflated using the capital price deflator from Brandt and Rawski (2008).

Finally, given the imputed real investments sequence and the initial capital stock, we calculate the real capital stock in year T assuming an annual depreciation rate of 9%.

A.4 Alternative definitions of prefecture-specific TFP growth

In Table 11 we calculated prefecture-specific aggregate TFP growth as a weighted average of industry-specific TFP growth, where the weight of each industry was the industry's relative share of value added, averaged across 1995 and 2004. We now show that these results are robust to alternative weighting schemes for calculating TFP growth.

We consider alternative weights along two dimensions: using relative value added versus relative inputs (Y versus $K^{\alpha}N^{(1-\alpha)}$) and using 1995 versus an average of 1995 and 2004. This leads to four cases, where the benchmark is case (4), reported in Table 11.

1) Use the relative share of Y in 1995.
2) Use the relative share of Y, averaged across 1995 and 2004.
3) Use the relative share of $K^{\alpha}N^{(1-\alpha)}$ in 1995.
4) Use the relative share of $K^{\alpha}N^{(1-\alpha)}$ averaged across 1995 and 2004.

The results are reported in Table A-1. $\ln TFP_{1995}$ is computed consistently with the given specification.

Table A-1: IV_b and alternative definitions of weighted average TFP growth

	$\Delta \ln TFP$ (1)	$\Delta \ln TFP$ (2)	$\Delta \ln TFP$ (3)	$\Delta \ln TFP$ (4)
IV_b	-0.19	-0.36	-0.34	-0.33
	(0.24)	(0.24)	(0.23)	(0.22)
$\ln TFP_{1995}$	-0.24***	-0.26***	-0.30***	-0.34***
	(0.04)	(0.04)	(0.04)	(0.04)
Province F.E.	Yes	Yes	Yes	Yes

Notes: Standard errors are in parentheses. *** − statistically significant at 1%; ** − statistically significant at 5%; * − statistically significant at 10%.

A.5 Details about measurements of migration

This appendix describes details about how the migration rates for workers and entrepreneurs are measured. We use data from the 1990, 2000, and 2010 China Census, as well as data from the 2005 1% Population Survey. 1. Workers are those age 16+ that have a job. Both entrepreneurs and workers are included. 2. Out of prefecture migrants are those who work and live outside of hukou prefecture; out of country are those who moved outside of their hukou county. 3. In 1990, migrants are defined as those who worked and lived outside for 12+ months. In 2000, this was reduced to 6 months. As result, we likely underestimate the number of migrants in 1990. 4. For all years other than 1990, entrepreneurs are identified on basis of the 3-digit occupation in the census. 5. We report two estimates for entrepreneurs for 1990. The narrow defintion includes only heads of private firms. The broader definition includes cadres and leaders of enterprises and institutions.

Given these assumptions in Section 4, we use the migration rates for workers and entrepreneurs to identify κ for each prefecture,

$$\begin{aligned} m_e &\equiv \frac{M_I}{M} = \frac{M_I}{M_D + M_I} = \frac{\kappa N_I}{N_D + \kappa N_I} = \frac{\kappa N_I}{N_D + N_I - (1-\kappa) N_I} = \frac{\kappa \frac{N_I}{N_D + N_I}}{1 - (1-\kappa)\frac{N_I}{N_D+N_I}} \\ &= \frac{\kappa m_N}{1 - (1-\kappa) m_N}, \end{aligned}$$

where m_e and m_N denote the share of migrants for entrepreneurs and workers, respectively. κ is then identified by

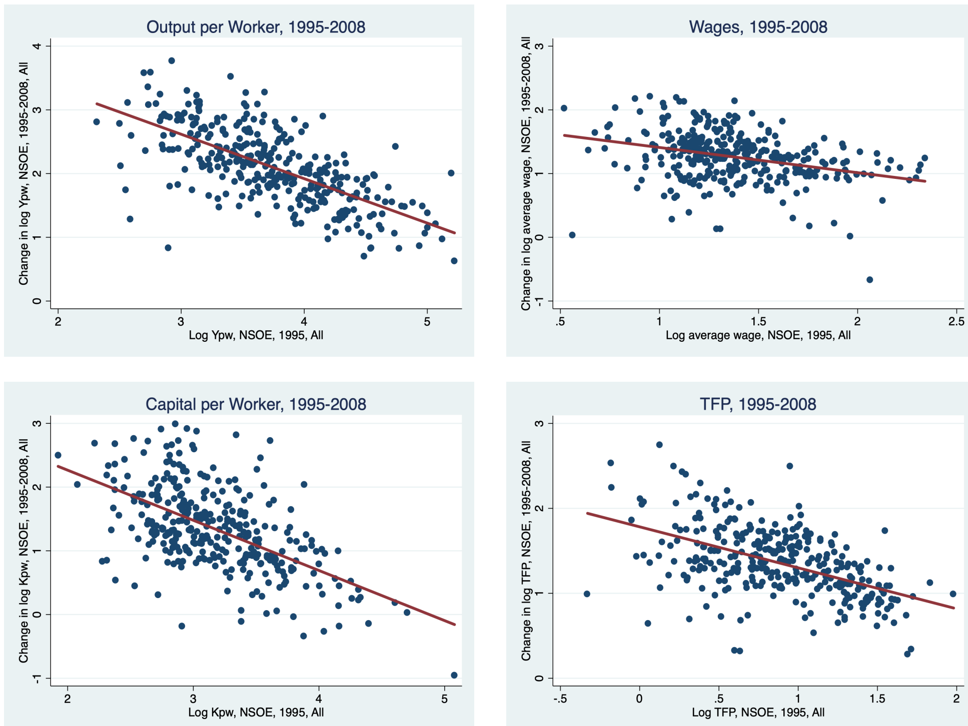
$$\kappa = \frac{m_N - 1}{m_N} \frac{m_e}{m_e - 1}.$$

It follows that the number of NSOE workers per potential entrepreneur, N/M, is given by

$$\begin{aligned} \frac{N}{M} &= \frac{M_I}{M} \frac{N_D + N_I - N_{SOE}}{M_I} = m_e \frac{N_D + N_I}{\kappa N_I} \frac{N_D + N_I - N_{SOE}}{N_D + N_I} \\ &= \frac{m_e}{\kappa} \frac{1}{m_N} \left(1 - \frac{N_{SOE}}{N_D + N_I}\right) \\ &= \underbrace{\frac{1 - m_e}{1 - m_N}}_{\text{native ratio}} \underbrace{\left(1 - \frac{N_{SOE}}{N_D + N_I}\right)}_{\text{NSOE share}} \end{aligned}$$

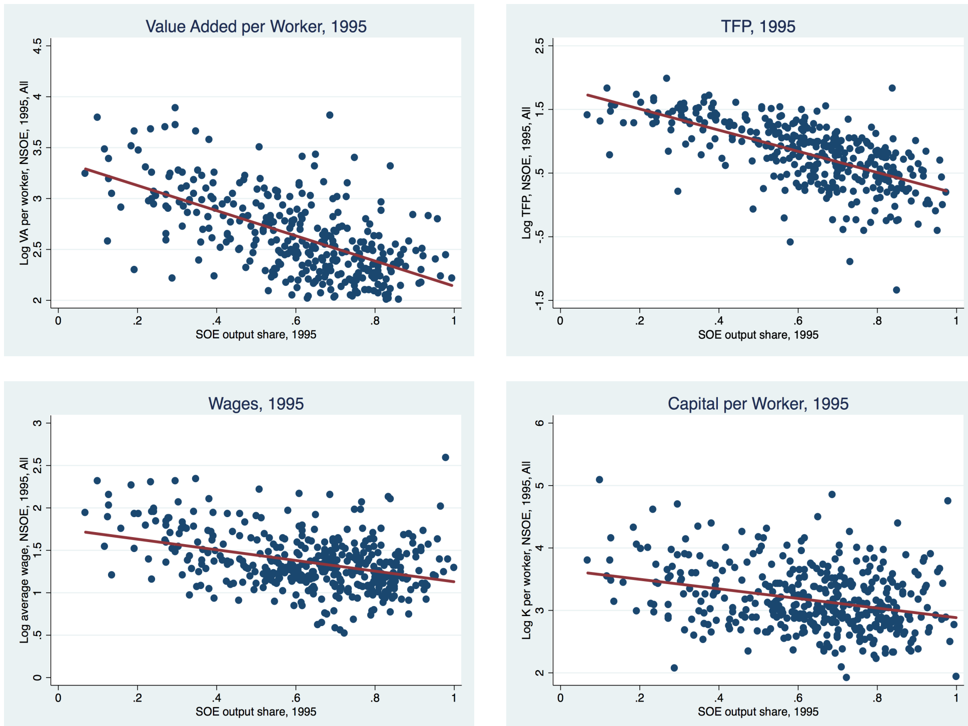
B Figures and Tables

Figure B-1: Convergence in the NSOE sector, 1995-2008.



Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line.

Figure B-2: Characteristics of NSOE Firms in 1995.



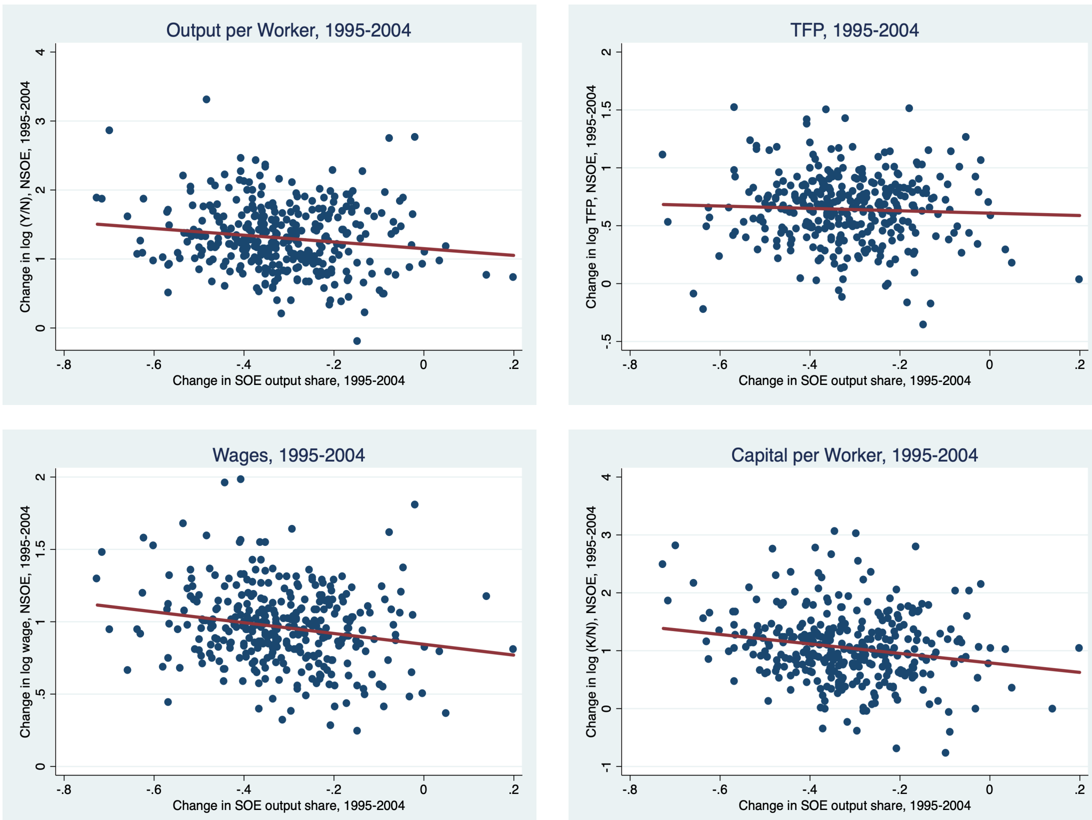
Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line. The 1995 SOE output share in a prefecture is on the horizontal axis.

Table B-1: Dispersion of Left Tail of TFP Distribution

	Low aggregate TFP		High aggregate TFP	
	Firm TFP below		Firm TFP below	
	25th percentile	50th percentile	25th percentile	50th percentile
1995	3.65	4.58	3.54	4.41
2004	3.85	4.74	3.83	4.63
2008	2.84	3.52	2.76	3.37

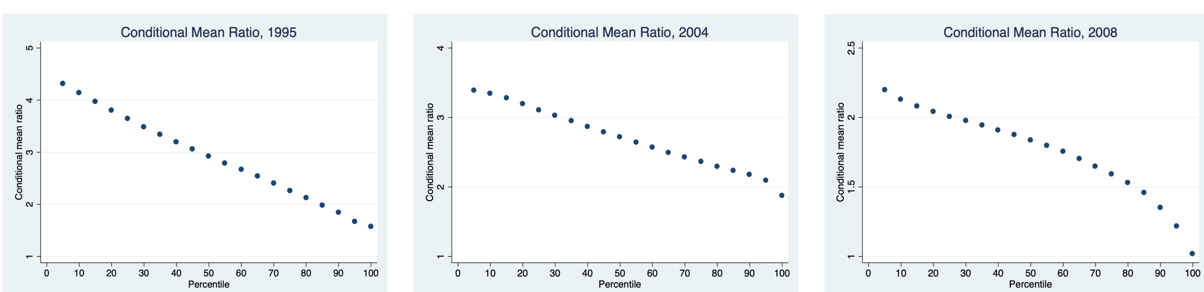
Notes: The table reports the standard deviation of log firm-specific TFP conditional on being below the 25th and 50th percentile of the national firm TFP distribution. This calculation is reported for two groups of prefectures, one group with aggregate TFP below that of the median prefecture and one group with aggregate TFP larger than or equal to that of the median prefecture.

Figure B-3: NSOE Performance and Changes in SOE Shares,1995-2004.



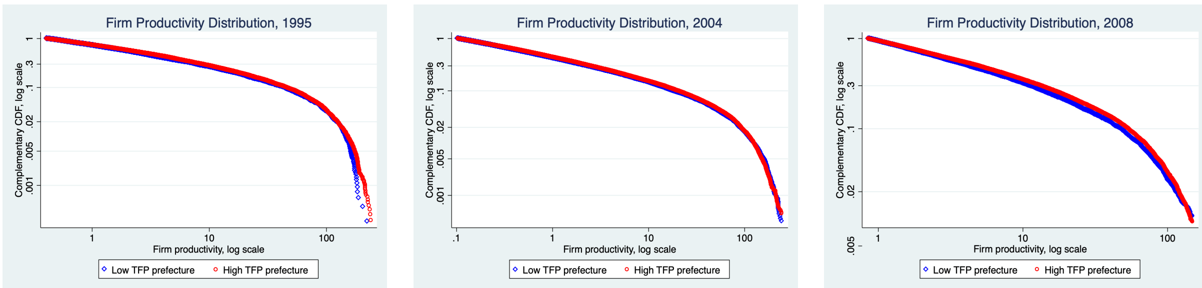
Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line. The 1995-2004 change in SOE output share in a prefecture is on the horizontal axis.

Figure B-4: Conditional Mean Ratio: 1995, 2004, and 2008



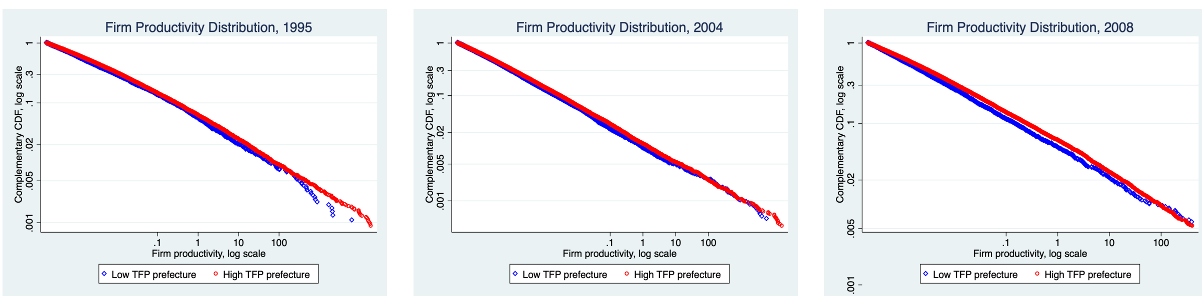
Notes: The figure describes the case when no firms from the top and the bottom of the productivity distribution are trimmed.

Figure B-5: Upper Tail of the Distribution of $\ln z$ for Prefectures with Low and High Aggregate TFP



Notes: All prefectures are separated into two groups based on their aggregate Solow residual. The figure plots the complementary cumulative distribution function for the entire firm productivity distribution in 1995, 2004, and 2008, conditional on firm TFP being in the top 5% of the firm TFP distribution in a given year. Firms in the top and bottom 1% of all firms in each industry are trimmed.

Figure B-6: Upper Tail of the Distribution of $\ln z$ for Prefectures with Low and High Aggregate TFP



Notes: All prefectures are separated into two groups based on their aggregate Solow residual. The figure plots the complementary cumulative distribution function for the entire firm productivity distribution in 1995, 2004, and 2008, conditional on firm TFP being in the top 5% of the firm TFP distribution in a given year.

Table B-2: Explaining the 1995-2004 changes in wedges.

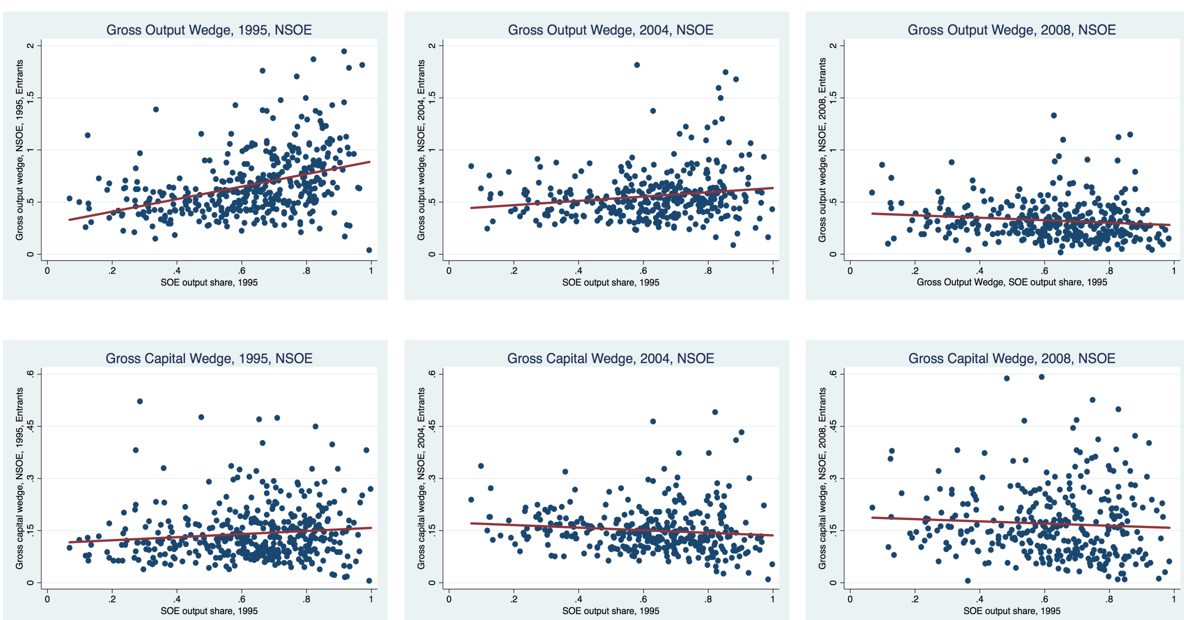
	$\Delta\ln(1-\psi)$	$\Delta\ln(1+\tau^k)$	$\Delta\ln(1-\tau^y)$
$\Delta\hat{S}_p$	-10.28***	1.37***	1.06***
	(2.91)	(0.51)	(0.41)
$\Delta\ln FREV$	0.76*	0.10	-0.03*
	(0.46)	(0.08)	(0.07)

Notes: The table reports the IV results from regression of the change in the log gross wedges on the (instrumented) change in the SOE employment share (e^{soe}) and the log fiscal revenue per government worker ($\ln FREV$) in a prefecture between 1995 and 2004. Standard errors are in parentheses. *** – statistically significant at 1%; ** – statistically significant at 5%; * – statistically significant at 10%.

Figure B-7: Gross Output and Gross Capital Wedges, 2004 and 2008, All Firms, NSOE.

Notes: Each dot represents a prefecture. The panels plot the gross output and gross capital wedges for all firms in the NSOE sector in 2004 and 2008. The SOE output share in 1995 in each prefecture is on the horizontal axis.

Figure B-8: Gross Output and Gross Capital Wedges, 1995, 2004 and 2008, Entrants, NSOE.



Notes: Each dot represents a prefecture. The panels plot the gross output and gross capital wedges for new firms in the NSOE sector in 1995, 2004, and 2008. The SOE output share in 1995 in each prefecture is on the horizontal axis.

Table B-3: Annual Rate of Convergence in TFP and Wages: 1995-2004 and 1995-2008, Heterogeneous-Wedges Model.

	TFP		Wages	
Change in	1995-2004	1995-2008	1995-2004	1995-2008
all	0.041	0.046	0.083	0.061
$\alpha\eta$	-0.001	-0.002	0.031	0.028
$\frac{N}{M}$	-0.002	-0.001	0.003	-0.001
m_N	0.000	0.001	0.011	0.010
m_e	0.000	0.000	-0.003	-0.004
$NSOE$	-0.002	-0.001	-0.005	-0.006
$(1+\bar{\tau_k})$	-0.002	-0.002	0.022	0.053
$(1-\bar{\tau_y})$	0.011	0.012	-0.032	-0.064
$(1-\psi)$	0.032	0.042	0.024	0.050
σ_k	-0.001	-0.001	-0.001	0.001
σ_y	-0.001	-0.001	0.013	0.028
σ_{ky}	-0.001	0.002	0.014	0.007

Notes: The table reports the annual rate of convergence in TFP and wages across prefectures for the 1995-2004 and 2004-2008 time periods. The β-convergence coefficient for prefectures p between times t_0 and t_0+T is estimated from the regression $\left(\frac{1}{T}\right)\ln\left(\frac{y_{p,t_0+T}}{y_{p,t_0}}\right) = a - \left(\frac{1-e^{-\beta T}}{T}\right)\ln(y_{p,t_0}) + u_{pt_0,t_0+T}$, where u_{pt_0,t_0+T} represents an average of error terms, $u_{p,t}$, between times t_0 and t_0+T. The table reports what convergence in TFP and wages had only one of the listed variables changed.

Table B-4: Change in the Entry Wedge, 1995-2004, Heterogeneous-Wedge Model.

$\Delta \ln(1-\psi)$	*OLS*	*OLS*	IV_{Bartik}	IV_{Bartik}
Δe^{soe}	-5.94***	-6.40**	-7.69**	-6.62**
	(1.65)	(1.84)	(3.34)	(3.90)
$\Delta \ln FREV$		0.59		0.58
		(0.56)		(0.61)
First stage:				
IV coefficient			0.67***	0.66***
st. error			(0.07)	(0.08)
R^2			0.24	0.24

Notes: The table reports the OLS and IV results from a regression of the change in the log gross entry wedge on the changes in SOE employment share (e^{soe}) and log fiscal revenues per government worker ($\ln FREV$) in a prefecture between 1995 and 2004. Standard errors are in parentheses. *** – statistically significant at 1%; ** – statistically significant at 5%; * – statistically significant at 10%.

C Proofs of Propositions

C.1 Proof of Propositions 1 and 3

Proposition 1 is a special case of Proposition 3 when there is no heterogeneity of wedges across firms in a prefecture ($\sigma_k = \sigma_y = \sigma_{ky} = 0$). We now proceed to prove Proposition 3. We start by providing a useful lemma.

Lemma 5 *For any constants a and b the following cross-sectional expectation holds across firms in a location,*

$$E\left\{((1+\tau_k)(r+\delta))^a (1-\tau_y)^b | z \geq z^*\right\} = \exp((b\mu_y + a\mu_k) + \left(b^2 + 2b\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2} + \left(a^2 + 2a\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\frac{\sigma_k}{2} - \left(b\xi(1-\alpha)\frac{\eta}{1-\eta} + a\frac{\xi}{1-\eta} + ab\right)\sigma_{ky})$$

Proof. Recall that the optimal choices of each firm are given by equations (3) and (4), given the firm-specific wedges. This implies,

$$\begin{aligned}
& E\left\{((1+\tau_k)(r+\delta))^a (1-\tau_y)^b | z \geq z^*\right\} \\
= & \frac{\int\int_{z^*(s)}^{\infty}\left[((1+\tau_k)(r+\delta))^a (1-\tau_y)^b\right] f(z)\, dz g(\tau_k,\tau_y)\, d\tau_k d\tau_y}{\int\int_{z^*(s)}^{\infty} f(z)\, dz g(\tau_k,\tau_y)\, d\tau_k d\tau_y} \\
= & \frac{\int((1+\tau_k)(r+\delta))^a (1-\tau_y)^b \left(\int_{z^*(s)}^{\infty} f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y}{\int\int_{z^*(s)}^{\infty} f(z)\, dz g(\tau_k,\tau_y)\, d\tau_k d\tau_y} \\
= & \frac{\exp\left(\left(\left(\frac{\xi}{1-\eta}+b\right)\mu_y + \left(a - \xi(1-\alpha)\frac{\eta}{1-\eta}\right)\mu_k\right)\right)}{\exp\left(\left(\frac{\xi}{1-\eta}\mu_y - \xi(1-\alpha)\frac{\eta}{1-\eta}\mu_k\right)\right)} \\
& \frac{\exp\left(\left(\frac{\xi}{1-\eta}+b\right)^2\frac{\sigma_y}{2} + \left(\xi(1-\alpha)\frac{\eta}{1-\eta}+a\right)^2\frac{\sigma_k}{2} - \left(\frac{\xi}{1-\eta}+b\right)\left(\xi(1-\alpha)\frac{\eta}{1-\eta}+a\right)\sigma_{ky}\right)}{\exp\left(\left(\frac{\xi}{1-\eta}\right)^2\frac{\sigma_y}{2} + \left(\xi(1-\alpha)\frac{\eta}{1-\eta}\right)^2\frac{\sigma_k}{2} - \left(\frac{\xi}{1-\eta}\right)\xi(1-\alpha)\frac{\eta}{1-\eta}\sigma_{ky}\right)} \\
= & \exp\left(b\mu_y + a\mu_k + \left(b^2 + 2b\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2} + \left(a^2 + 2a\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\frac{\sigma_k}{2} - \left(b\xi(1-\alpha)\frac{\eta}{1-\eta} + a\frac{\xi}{1-\eta} + ab\right)\sigma_{ky}\right).
\end{aligned}$$

The second equation follows from the assumption that the distribution function f (of potential z) is identical across locations. The third equation follows from the assumption that the joint cross-sectional distribution of the wedges is log normal with the moments given by equation (17) and on the fact that expected firm entry Γ given state s is given by

$$\begin{aligned}
\Gamma \equiv \int_{z^*}^{\infty} f(z)\, dz &= \int_{z^*}^{\infty} \underline{z}\xi z^{-\xi-1} dz = \underline{z}\xi\left[-\frac{1}{\xi}z^{-\xi}\right]_{z^*}^{\infty} = \underline{z}(z^*)^{-\xi} \\
&= \underline{z}\left(\frac{w}{\alpha}\right)^{-\xi\frac{\alpha\eta}{1-\eta}}\left(\frac{(1+\tau_k)(r+\delta)}{1-\alpha}\right)^{-\xi(1-\alpha)\frac{\eta}{1-\eta}}(1-\tau_y)^{\frac{\xi}{1-\eta}}\left(\nu^{-1}\eta^{\frac{\eta}{1-\eta}}(1-\eta)\right)^{\xi}.
\end{aligned}$$

■

We now solve for the equilibrium wage rate and the measured Solow residual. Given a wage w and the entrepreneurial entry decision in equation (4), the aggregate labor demand *per potential entrepreneur* in a prefecture is given by

$$\begin{aligned}
& \int\left(\int_{z^*(s)}^{\infty} n(z,s) f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
= & \int \eta^{\frac{1}{1-\eta}} (1-\tau_y)^{\frac{1}{1-\eta}} \left(\frac{(1-\alpha)}{(1+\tau_k)(r+\delta)}\right)^{\frac{(1-\alpha)\eta}{1-\eta}} \left(\frac{\alpha}{w}\right)^{1+\frac{\alpha\eta}{1-\eta}} \left(\int_{z^*(s)}^{\infty} z f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
= & \left(\frac{\alpha}{w}\right)^{1+\frac{\alpha\eta}{1-\eta}} \eta^{\frac{1}{1-\eta}} (1-\alpha)^{\frac{(1-\alpha)\eta}{1-\eta}} \int \Bigg[(1-\tau_y)^{\frac{1}{1-\eta}} \left((1+\tau_k)(r+\delta)\right)^{-\frac{(1-\alpha)\eta}{1-\eta}} \underline{z}\frac{\xi}{\xi-1}\left(\frac{w}{\alpha}\right)^{-(\xi-1)\frac{\alpha\eta}{1-\eta}} \\
& \left(\left(\frac{(1+\tau_k)(r+\delta)}{1-\alpha}\right)^{-(\xi-1)(1-\alpha)\frac{\eta}{1-\eta}} (1-\tau_y)^{\frac{\xi-1}{1-\eta}} \left(\nu^{-1}(1-\eta)\eta^{\frac{\eta}{1-\eta}}\right)^{\xi-1}\right)\Bigg] g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
= & \left(\frac{w}{\alpha}\right)^{-\frac{1-\eta+\alpha\xi\eta}{1-\eta}} \underline{z}\frac{\xi}{\xi-1}\eta^{1+\frac{\xi\eta}{1-\eta}} (1-\alpha)^{\xi\eta\frac{1-\alpha}{1-\eta}} \left(\frac{1-\eta}{\nu}\right)^{\xi-1} \\
& \cdot \int \left[(1-\tau_y)^{\frac{\xi}{1-\eta}} \left((1+\tau_k)(r+\delta)\right)^{-\xi\eta\frac{1-\alpha}{1-\eta}}\right] g(\tau_k,\tau_y)\, d\tau_k d\tau_y
\end{aligned}$$

where the second equation follows from the expression

$$\begin{aligned}
\int_{z^*}^{\infty} z f(z)\, dz &= \int_{z^*}^{\infty} \underline{z}\xi z^{-\xi} dz = \underline{z}\frac{\xi}{\xi-1}(z^*)^{-(\xi-1)} \\
&= \underline{z}\frac{\xi}{\xi-1}\left(\frac{w}{\alpha}\right)^{-(\xi-1)\frac{\alpha\eta}{1-\eta}} \left(\frac{(1+\tau_k)(r+\delta)}{1-\alpha}\right)^{-(\xi-1)(1-\alpha)\frac{\eta}{1-\eta}} (1-\tau_y)^{\frac{\xi-1}{1-\eta}} \left(\nu^{-1}(1-\eta)\eta^{\frac{\eta}{1-\eta}}\right)^{\xi-1}.
\end{aligned}$$

From Lemma 5 the last term is given by

$$\begin{aligned}
& \int \left((1-\tau_y)^{\frac{\xi}{1-\eta}} \left((1+\tau_k)(r+\delta)\right)^{-\xi\eta\frac{1-\alpha}{1-\eta}}\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
= & \exp(\mu_y)^{\frac{\xi}{1-\eta}} \exp(\mu_k)^{-\xi\eta\frac{1-\alpha}{1-\eta}} \exp\left(\left(\frac{\xi}{1-\eta}\right)^2 \frac{\sigma_y}{2} + \left(\xi\eta\frac{1-\alpha}{1-\eta}\right)^2 \frac{\sigma_k}{2} - \left(\frac{\xi}{1-\eta}\right)^2 \eta(1-\alpha)\sigma_{ky}\right).
\end{aligned}$$

Consider now the equilibrium in the labor market. The wage rate must equate labor supply N to the aggregate labor demand (in a prefecture);

$$\begin{aligned}
N &= (1-\psi) M \int\left(\int_{z^*(s)}^{\infty} n(z,s) f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= (1-\psi) M \left(\frac{w}{\alpha}\right)^{-\frac{1-\eta+\alpha\xi\eta}{1-\eta}} \underline{z}\frac{\xi}{\xi-1}\eta^{1+\frac{\xi\eta}{1-\eta}} (1-\alpha)^{\xi\eta\frac{1-\alpha}{1-\eta}} \left(\frac{1-\eta}{\nu}\right)^{\xi-1} \\
& \quad \exp\left(\frac{\xi}{1-\eta}\mu_y - \xi\eta\frac{1-\alpha}{1-\eta}\mu_k + \left(\frac{\xi}{1-\eta}\right)^2 \frac{\sigma_y}{2} + \left(\xi\eta\frac{1-\alpha}{1-\eta}\right)^2 \frac{\sigma_k}{2} - \left(\frac{\xi}{1-\eta}\right)^2 \eta(1-\alpha)\sigma_{ky}\right)
\end{aligned}$$

It follows that the equilibrium wage is given by equation (18) where the term Ω is defined as follows:

$$\Omega(\alpha,\eta,\xi,\nu) = \ln\alpha + (1-\eta)\mu \ln\left(\frac{\xi}{\xi-1}\eta^{1+\frac{\xi\eta}{1-\eta}} (1-\alpha)^{\xi\eta\frac{1-\alpha}{1-\eta}} \left(\frac{1-\eta}{\nu}\right)^{\xi-1}\right).$$

Consider now the Solow residual Z. Given the expression for w we can calculate the expressions for the prefecture-specific aggregate allocations of K , Y, and the Solow residual Z. Y is given by

$$\begin{aligned}
\frac{Y}{(1-\psi)M} &= \int\left(\int_{z^*(s)}^{\infty} y(z,s)\, f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= \int\left(\int_{z^*(s)}^{\infty}\left[z\left((1-\tau_y)\,\eta\right)^{\frac{\eta}{1-\eta}}\left(\frac{(1-\alpha)}{(1+\tau_k)(r+\delta)}\right)^{\frac{(1-\alpha)\eta}{1-\eta}}\left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}\right] f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= \eta^{\frac{\eta}{1-\eta}}(1-\alpha)^{\frac{(1-\alpha)\eta}{1-\eta}}\left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}\int\left[(1-\tau_y)^{\frac{\eta}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{-\frac{(1-\alpha)\eta}{1-\eta}}\left(\int_{z^*(s)}^{\infty} z f(z)\, dz\right)\right] g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= \eta^{\frac{\eta}{1-\eta}}(1-\alpha)^{\frac{(1-\alpha)\eta}{1-\eta}}\left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}\frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(1-\eta)^{-(1-\xi)}\eta^{-\frac{\eta(1-\xi)}{1-\eta}}(1-\alpha)^{-(1-\alpha)\frac{\eta(1-\xi)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{\frac{(1-\xi)\alpha\eta}{1-\eta}} \\
&\quad \int\left[(1-\tau_y)^{\frac{\eta}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{-\frac{(1-\alpha)\eta}{1-\eta}}(1-\tau_y)^{-\frac{1-\xi}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{(1-\alpha)(1-\xi)\frac{\eta}{1-\eta}}\right] g(\tau_k,\tau_y)\, d\tau_k d\tau_y
\end{aligned}$$

$$\begin{aligned}
&= \frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(1-\eta)^{-(1-\xi)}\eta^{\frac{\xi\eta}{1-\eta}}(1-\alpha)^{\frac{\xi\eta(1-\alpha)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{-\frac{\eta\alpha\xi}{1-\eta}} \\
&\quad \exp(\mu_y)^{\frac{\xi-1+\eta}{1-\eta}}\exp(\mu_k)^{-\xi\eta\frac{1-\alpha}{1-\eta}}\exp\left(\left(\frac{\xi-1+\eta}{1-\eta}\right)^2\frac{\sigma_y}{2}+\left(\xi\eta\frac{1-\alpha}{1-\eta}\right)^2\frac{\sigma_k}{2}-\frac{\xi\eta(1-\alpha)(\xi-1+\eta)}{(1-\eta)^2}\sigma_{ky}\right).
\end{aligned}$$

Now compute aggregate capital,

$$\begin{aligned}
\frac{K}{(1-\psi)M} &= \int\left(\int_{z^*(s)}^{\infty} k(z,s)\, f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= \int\left(\int_{z^*(s)}^{\infty}\left[z\left((1-\tau_y)\,\eta\right)^{\frac{1}{1-\eta}}\left(\frac{(1-\alpha)}{(1+\tau_k)(r+\delta)}\right)^{\frac{1-\alpha\eta}{1-\eta}}\left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}\right] f(z)\, dz\right) g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= (\eta)^{\frac{1}{1-\eta}}(1-\alpha)^{\frac{1-\alpha\eta}{1-\eta}}\left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}\int\left[(1-\tau_y)^{\frac{1}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{-\frac{1-\alpha\eta}{1-\eta}}\left(\int_{z^*(s)}^{\infty} z f(z)\, dz\right)\right] g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= (\eta)^{\frac{1}{1-\eta}}(1-\alpha)^{\frac{1-\alpha\eta}{1-\eta}}\left(\frac{\alpha}{w}\right)^{\frac{\alpha\eta}{1-\eta}}\frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(1-\eta)^{-(1-\xi)}\eta^{-\frac{\eta(1-\xi)}{1-\eta}}(1-\alpha)^{-(1-\alpha)\frac{\eta(1-\xi)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{\frac{(1-\xi)\alpha\eta}{1-\eta}} \\
&\quad \int\left[(1-\tau_y)^{\frac{1}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{-\frac{1-\alpha\eta}{1-\eta}}(1-\tau_y)^{-\frac{1-\xi}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{(1-\alpha)(1-\xi)\frac{\eta}{1-\eta}}\right] g(\tau_k,\tau_y)\, d\tau_k d\tau_y
\end{aligned}$$

$$\begin{aligned}
&= \frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(\eta)^{\frac{1-\eta+\xi\eta}{1-\eta}}(1-\eta)^{-(1-\xi)}(1-\alpha)^{\frac{1-\eta+\xi\eta(1-\alpha)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{-\frac{\xi\alpha\eta}{1-\eta}} \\
&\quad \int\left[(1-\tau_y)^{\frac{\xi}{1-\eta}}\left((1+\tau_k)(r+\delta)\right)^{\frac{\eta(1-\alpha)(1-\xi)-1+\alpha\eta}{1-\eta}}\right] g(\tau_k,\tau_y)\, d\tau_k d\tau_y \\
&= \frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(\eta)^{\frac{1-\eta+\xi\eta}{1-\eta}}(1-\eta)^{-(1-\xi)}(1-\alpha)^{\frac{1-\eta+\xi\eta(1-\alpha)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{-\frac{\xi\alpha\eta}{1-\eta}} \\
&\quad \exp(\mu_y)^{\frac{\xi}{1-\eta}}\exp(\mu_k)^{-\left(\xi\eta\frac{1-\alpha}{1-\eta}+1\right)}\exp\left(\left(\frac{\xi}{1-\eta}\right)^2\frac{\sigma_y}{2}+\left(\xi\eta\frac{1-\alpha}{1-\eta}+1\right)^2\frac{\sigma_k}{2}-\left(\frac{\xi}{1-\eta}\right)\left(\xi\eta\frac{1-\alpha}{1-\eta}+1\right)\sigma_{ky}\right)
\end{aligned}$$

The Solow residual can then be calculated as

$$\begin{aligned}
\ln Z &= \ln\left(\frac{Y}{(1-\psi)M}\right) - (1-\alpha\eta)\ln\left(\frac{K}{(1-\psi)M}\right) - \alpha\eta\ln\left(\frac{N}{(1-\psi)M}\right) \\
&= \ln\left(\frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(1-\eta)^{-(1-\xi)}\eta^{\frac{\xi\eta}{1-\eta}}(1-\alpha)^{\frac{\xi\eta(1-\alpha)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{-\frac{\eta\alpha\xi}{1-\eta}}\right) \\
&+ \frac{\xi-1+\eta}{1-\eta}\mu_y - \xi\eta\frac{1-\alpha}{1-\eta}\mu_k + \left(\frac{\xi-1+\eta}{1-\eta}\right)^2\frac{\sigma_y}{2} + \left(\xi\eta\frac{1-\alpha}{1-\eta}\right)^2\frac{\sigma_k}{2} - \frac{\xi\eta(1-\alpha)(\xi-1+\eta)}{(1-\eta)^2}\sigma_{ky} \\
&- (1-\alpha\eta)\ln\left(\frac{\xi}{\xi-1}\underline{z}(\nu)^{1-\xi}(\eta)^{\frac{1-\eta+\xi\eta}{1-\eta}}(1-\eta)^{-(1-\xi)}(1-\alpha)^{\frac{1-\eta+\xi\eta(1-\alpha)}{1-\eta}}\left(\frac{w}{\alpha}\right)^{-\frac{\xi\alpha\eta}{1-\eta}}\right) - \alpha\eta\ln\left(\frac{N}{1-\psi}\right) \\
&- (1-\alpha\eta)\left(\frac{\xi}{1-\eta}\mu_y - \left(\xi\eta\frac{1-\alpha}{1-\eta}+1\right)\mu_k + \left(\frac{\xi}{1-\eta}\right)^2\frac{\sigma_y}{2} + \left(\xi\eta\frac{1-\alpha}{1-\eta}+1\right)^2\frac{\sigma_k}{2} - \left(\frac{\xi}{1-\eta}\right)\left(\xi\eta\frac{1-\alpha}{1-\eta}+1\right)\sigma_{ky}\right).
\end{aligned}$$

Substituting in the wage expression from equation (18) and simplifying yields equation (19) in the text, where the term $\hat{\Omega}$ is defined as follows:

$$\hat{\Omega}(\alpha,\eta,\xi,\nu) = -\mu(1-\eta)[1+(\xi-1)\alpha\eta]\ln((1-\alpha)\eta) + \mu\alpha\eta(1-\eta)\ln\left(\frac{\xi}{\xi-1}\left(\frac{1-\eta}{\nu}\right)^{\xi-1}\right).$$

This concludes the proof of Proposition 3.

Finally, setting $\sigma_k = \sigma_y = \sigma_{ky} = 0$ in equations (18)-(19) yields equations (6)-(7). This concludes the proof of Proposition 1 and the derivation of equation (7).

C.2 Proof of Proposition 4

Consider now the cross-sectional moments of the observed wedges within a prefecture. Note that we are observing a truncated distribution of firms, i.e., those with $z \geq z^*(\tau^k, \tau^y, w)$. Using Lemma 5 the variance of the observed $(1-\tau^y)$ is given by

$$\begin{aligned}
& E\left\{(1-\tau_y)^2 | z \geq z^*\right\} - \left(E\left\{(1-\tau_y)|z\geq z^*\right\}\right)^2 \\
&= \exp\left(2\mu_y + 4\left(1+\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2} - \left(2\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\sigma_{ky}\right) \\
&\quad - \exp\left(2\mu_y + 2\left(1+2\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2} - 2\left(\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\sigma_{ky}\right) \\
&= \exp\left(2\left(\mu_y + \left(1+2\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2} - \xi(1-\alpha)\frac{\eta}{1-\eta}\sigma_{ky}\right)\right)(\exp(\sigma_y)-1),
\end{aligned}$$

and the mean is

$$\begin{aligned}
& E\left\{(1-\tau_y)|z\geq z^*\right\} \\
&= \exp\left(\mu_y + \left(1+2\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2} - \left(\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\sigma_{ky}\right)
\end{aligned}$$

It follows immediately that σ_y can be identified by the observed coefficient of variation, stated in equation (21). A similar argument establishes that the variance and mean of $1+\tau_k$ is given by

$$\begin{aligned}
& E\left\{((1+\tau_k)(r+\delta))^2 | z\geq z^*\right\} - \left(E\left\{(1+\tau_k)(r+\delta)|z\geq z^*\right\}\right)^2 \\
&= \exp\left(2\left(\mu_k + \left(1+2\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\frac{\sigma_k}{2} - \frac{\xi}{1-\eta}\sigma_{ky}\right)\right)(\exp(\sigma_k)-1) \\
& E\left\{(1+\tau_k)(r+\delta)|z\geq z^*\right\} \\
&= \exp\left(\mu_k + \left(1+2\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\frac{\sigma_k}{2} - \frac{\xi}{1-\eta}\sigma_{ky}\right),
\end{aligned}$$

implying that σ_k can be identified from equation (20). Finally, the observed covariance is calculated as follows,

$$
\begin{aligned}
& cov\left\{(1+\tau_k)(r+\delta),(1-\tau_y)\,|z\geq z^*\right\} \\
= \; & (\exp(-\sigma_{ky})-1)\exp\left(\mu_y+\mu_k\right. \\
& \left.+\left(1+2\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2}+\left(1+2\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\frac{\sigma_k}{2}-\left(\xi(1-\alpha)\frac{\eta}{1-\eta}+\frac{\xi}{1-\eta}\right)\sigma_{ky}\right) \\
& E\left\{((1+\tau_k)(r+\delta))(1-\tau_y)\,|z\geq z^*\right\} \\
= \; & \exp\left((\mu_y+\mu_k)\right. \\
& \left.+\left(1+2\frac{\xi}{1-\eta}\right)\frac{\sigma_y}{2}+\left(1+2\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\frac{\sigma_k}{2}-\left(\xi(1-\alpha)\frac{\eta}{1-\eta}+\frac{\xi}{1-\eta}+1\right)\sigma_{ky}\right).
\end{aligned}
$$

It follows that σ_{ky} can be identified from equation (22). This concludes the proof of Proposition 4.