University of Toronto
Department of Economics

Working Paper 652

Barriers to Entry and Regional Economic Growth in China

By Loren Brandt, Gueorgui Kambourov and Kjetil Storesletten

January 05, 2020

Barriers to Entry and Regional Economic Growth in China*

Loren Brandt† University of Toronto

Gueorgui Kambourov‡ University of Toronto

Kjetil Storesletten§ University of Oslo

December 2019

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 barriers are causally related to the size of the state sector

JEL Classification: O11, O14, O16, O40, O53, P25, R13, D22, D24, E24.

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

*This version: December 1, 2019. We thank Tommaso Porzio, Paulina Restrepo-Echavarria, Hsuan-Li Su, Aleh Tsyvinski and seminar and conference participants at ABFER Singapore, Asian Meeting of the Econometric Society, ASSA Annual Meeting, Bank of Finland Conference on China's Financial Markets and the Global Economy, Barcelona GSE Summer Forum, Beijing Conference on Business Cycles/Financial Markets/ and Monetary Policy, Beijing Conference on Deepening Economic Reforms, Beijing University 3rd International Conference on New Structural Economics, Canadian Economics Association Meeting, Cornell, Emory, FGV Sao Paulo School of Economics, Glasgow, Guelph, Hawaii, IAES Conference, IMF, Kent, Leuven, Mannheim 3rd Workshop on the Spacial Dimensions of the Labour Market, McMaster, Minneapolis Fed, MIT Sloan, Montreal Workshop on Markets with Frictions, NBER Chinese Economy Working Group, Notre Dame, Oslo, Rochester, Society for Economic Dynamics Meeting, Sichuan University, Stockholm School of Economics Conference on Industrial Upgrading and Urbanization, St. Andrews Workshop on Structural Transformation, Tsinghua, UBC, University of Toronto/Bank of Canada Conference on the Chinese Economy, Vienna Macro, Western Economics Anniversary Conference, Wharton, Wisconsin-Madison, World Bank, and Yale. Brandt, Kambourov, and Storesletten have received funding from the European Research Council under the European Union's Seventh Framework Programme (FP7/2007-2013)/ERC grant agreement n. 324085. Brandt and Kambourov acknowledge financial support from the Social Sciences and Humanities Research Council grants #435-2015-0987 and #435-2014-0815, respectively.

†University of Toronto, Department of Economics, 150 St. George St., Toronto, Ontario M5S 3G7, Canada. E-mail: brandt@chass.utoronto.ca.

‡University of Toronto, Department of Economics, 150 St. George St., Toronto, Ontario M5S 3G7, Canada. E-mail: g.kambourov@utoronto.ca.

§University of Oslo, Department of Economics, 0317 Oslo, Norway. E-mail: kjetil.storesletten@econ.uio.no.

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 macro-economic 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 changes. 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) and Melitz (2003) 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 who would like to enter will be allowed to operate. 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.

The three distortions affect the economy through different mechanisms. Increasing any of the distortions will lower the equilibrium wage rate in a prefecture through lowering the entry rate of new firms and thereby lowering the demand for workers. However, the distortions have differential effects on aggregate prefecture TFP. Larger output and capital distortions imply that only the most productive firms will choose to operate. This positive selection of entrants induces an increase in aggregate TFP. In contrast, larger entry barriers will lower the productivity threshold for entering firms due to lower equilibrium wages. This creates negative selection and, hence, lower aggregate TFP. Thus, the entry barrier is the only distortion that can cause a positive correlation between wages and aggregate TFP across locations and over time. This turns out to be a key feature of the data.

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 across prefectures in China. We find that the entry barrier is the main driver of 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 patterns.

We study the measured entry barriers in greater detail and show, in the spirit of Cheremukhin, Golosov, Guriev and Tsyvinski (2017a,b), 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 cities in China. This provides valuable external validation for our estimates. Moreover, using data on actual creation of new firms − data which we did not target when estimating the wedges − we show that firm creation is primarily explained by the entry barrier.

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 and omitted variables. 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 cadre 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. Using an IV approach, we document 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 the rise of Chinese manufacturing during the great transformation (see e.g. Brandt, Rawski and Sutton (2008b), Young (2003), Zhu (2012), and references therein). Several papers emphasize the role of the reform of the state sector in the late 1990s for understanding this growth (Hsieh and Song (2015), Song et al. (2011)).

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. (2017b,a)). 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)). Similar to Barseghyan and DiCecio (2011), we also emphasize the role of entry barriers for new firms in accounting for TFP differences, although they focus on dispersion across countries while we focus on regional convergence in China.

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)). To our knowledge ours is the first paper using wedge analysis to analyze cross-region convergence in output, wages, and TFP.

The rest of the paper is organized as follows. Section 2 empirically documents the economic development across more than 300 prefectures. Section 3 lays out a version of the Hopenhayn (1992) and Melitz (2003) model extended to incorporate a novel entry barrier. Section 4 uses the entry barrier model to measure the distortions across prefectures. Section 5 studies the empirical drivers of the prefecture-specific measured entry barriers while Section 6 studies an extension of the model that allows for firm heterogeneity in wedges. Section 7 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	10.2%	8.7%	14.4%	2.3%
Annualized rate of β-convergence 2004-2008	7.9%	2.5%	3.3%	4.6%
st.dev. of log in 1995	0.53	0.33	0.51	0.59
st.dev. of log in 2004	0.43	0.25	0.50	0.54
st.dev. of log in 2008	0.47	0.38	0.69	0.51

Notes: The table reports the annualized rates of β-convergence and the cross-sectional dispersion in aggregate outcomes across prefectures in China. Dipersion is measured as standard deviation of log. 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 2004-2008 period.

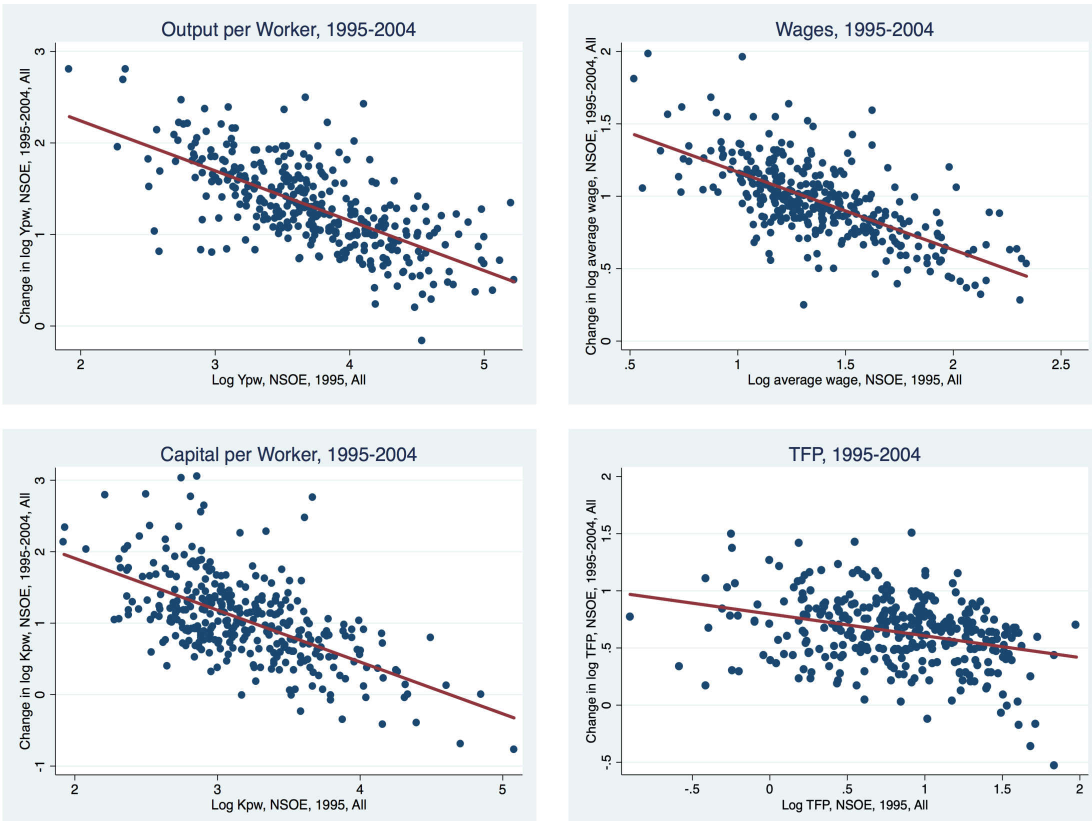
[1]We also draw on firm-level data collected by the NBS for 1992 on all independent accounting units (0.39 million), which covers a slightly smaller subset of firms than the CIC.

[2]The 2004 and 2008 CIC also provide data for the service sector, but unfortunately similar information was not collected 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. The top left panel of Figure 1 documents output per worker, which we denote labor productivity, across prefectures. In 1995 the cross-sectional dispersion in output per worker, as measured by the standard deviation of logs, was 0.53 (cf. Table 1). Second, there was substantial 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 10.2% (cf. Table 1), implying that it only takes about seven years to cut the difference in labor productivity between any two prefectures by half. The annualized rate of convergence across regions in labor productivity falls slightly to 7.9 percent over the 2004-2008 time period, suggesting that convergence remained strong after 2004.

To put the magnitude of this rate of convergence in perspective, we compare it with the rate of convergence in GDP per capita across regions in other countries. Barro and Sala-i-Martin (1991) and Sala-i-Martin (1996) document that the annualized rate of convergence has been about 2 percent across states in the US (1880-1988), across 73 regions in Europe (1950-1985), and across large industrialized countries (USA, Japan, and five Western European countries). 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 a particular firm measure the annual average earnings per worker in that firm, computed as the firm's total wage bill divided by the firm's employment. The 1995 dispersion in average wage rates is large, albeit less dispersed than labor productivity: the standard deviation of log wages was 0.33 in 1995. The annualized rate of β-convergence was also very large – 8.7 percent – suggesting that it took about eight years to reduce average wage differences between two prefectures by 50 percent. However, after 2004 the rate of regional convergence in wages falls substantially to a mere 2.5 percent.

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 14.4 percent. After 2004 the annualized rate of regional convergence in capital per worker falls to 3.3 percent.

The bottom right panel of Figure 1 documents the dispersion and dynamics in aggregate TFP in each prefecture. 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. The production function is 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, Z, for a particular prefecture between 1995 and 2004, we first compute the change in $\ln Z_j$ for each (2-digit) industry in that prefecture. The industry-specific growth rates are then weighted using 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, and equal to 0.59. However, aggregate TFP exhibits regional convergence between 1995 and 2004 of 2.3 percent, and even faster convergence after 2004.[5]

Finally, we note that between 1995 and 2004 the β-convergence is so strong that even the cross-sectional dispersion in all variables fell: Table 1 shows that the dispersion is lower in 2004 than in 1995, indicating σ-convergence across prefectures. However, after 2004 the overall cross-sectional dispersion in wages, labor productivity, and capital per worker increased (while it continued to fall for TFP). 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* of new private firms in a prefecture, Γ, 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.[6] This positive correlation will be important for identifying the key forces driving the differential performance of the non-state sectors across prefectures.

[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 5.2 for details.

[6] Aggregate employment in NSOE manufacturing firms grew rapidly during the 1995-2008 period and 87 percent of the prefectures experienced positive growth. Moreover, this growth was larger in prefectures that initially paid higher wages.

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.

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 (2008a), 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.

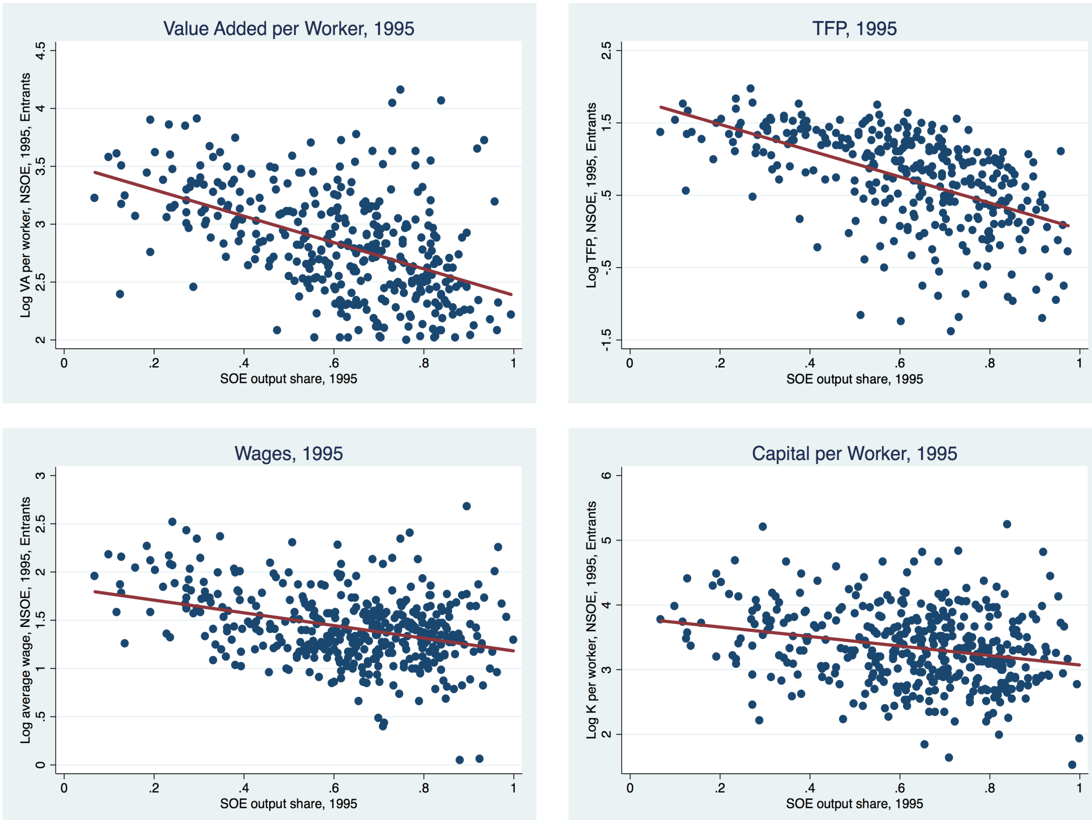
We start by analyzing the performance of non-state firms (NSOE) in the 1995 cross section of prefectures. We continue to focus on average wages, labor productivity, capital per worker, and TFP for non-state firms. We also document entry rates of new NSOEs for each prefecture. In order to illustrate the 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.[7,8] 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.

[7] 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.

[8] 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 5.1 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.

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 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.[9]

[9] 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 and

Figure 3: NSOE Firm Entry during 1993-1995.

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

3 A Hopenhayn-Melitz 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 the aggregate firm performance in each location. Since the geographic unit of measurement in the subsequent empirical work will be a prefecture, we often refer to a location as a *prefecture.*

3.1 Environment

The economy consists of a set of locations. Each location is a small open economy where labor is location specific and supplied inelastically and capital can be allocated freely across locations. In the main analysis we take the labor supply offered to private firms in location j, N_j, as exogenous and abstract from state firms.[10]

Firms produce a homogenous 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 a common rental rate $(r+\delta)$ on capital and face a location-specific wage rate w. In addition, firms face standard distortions on output and capital given by τ_y and τ_k.[11] These wedges are common for all firms in the location.[12] There is a fixed

omitted variable bias. We tackle these issues in Section 5 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.

[10] In Section 5.3 we extend the model to incorporate state firms which compete with non-state firms for workers. The purpose of that extension is to motivate the political choice of wedges.

[11] 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. The reason is that labor wedges cannot be separately identified from capital and output wedges. It is therefore convenient to normalize the labor wedge to zero. Note, however, that our analysis differs crucially from Hsieh and Klenow (2009) in that we allow wages to differ across locations.

[12] In Section 6 we extend the analysis to allow for firm-specific capital and output wedges. As we shall see, our

cost ν for operating a firm. This cost is constant across all locations.

Following Melitz (2003), the model is static, comprising two stages: a firm entry stage and a production stage. Each location has a measure M 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 $f(z) = \underline{z}\xi z^{-\xi-1}$, where $\xi > 1$, $\underline{z} \geq 1$, and $z \in \left[\underline{z}^{1/\xi}, \infty\right)$.

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*.[13] This entry barrier can be interpreted as a lottery over licenses. 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 entering firms in locations with a large ψ.

3.2 The Firm problem

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

Profit maximization. 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,

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

where

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

The entry decision. Given the vector of distortions and prices (τ_y, τ_k, r, w), there exists a cutoff $z^* = z^*(\tau_y, \tau_k, r, w)$ 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 $\nu = (1-\tau_y)(1-\eta) \cdot z^* \cdot \bar{y}(\tau_y, \tau_k, r, w)$, implying

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

quantitative findings and the main message of the paper remain robust to this extension.

[13] Hopenhayn (1992) proposes an alternative model of entry barriers. He assumes an infinite supply of potential entrepreneurs. Each entrepreneur who considers entering must first pay a fixed cost of obtaining a stochastic draw of 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. However, as we discuss below, in our model an increase in labor supply will lower equilibrium wages and TFP.

3.3 Equilibrium

We can now compute the equilibrium wage w and the associated aggregate output, capital stock, and measured aggregate TFP in the non-state sector, given a labor supply N. Without loss of generality we normalize the number of potential entrepreneurs to unity, $M = 1$.

Market clearing in the labor market requires that $(1-\psi)\int_{z^*}^{\infty} n(z) f(z)\, dz = N$. 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*

$$\begin{aligned}\ln w &= \mu(1-\eta)\ln\left[\frac{(1-\psi)\underline{z}^{\xi}}{N}\right] \\ &\quad +\mu\xi\ln(1-\tau_y) - \mu(1-\alpha)\xi\eta\ln[(1+\tau_k)(r+\delta)] + \Omega(\alpha,\eta,\xi,\nu), \qquad (5)\end{aligned}$$

where $\mu \equiv \frac{1}{1-\eta+\xi\alpha\eta} > 0$ *and*

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

The equilibrium wage is falling in N, τ_y , τ_k, *and* ψ.

The analytical characterization of the equilibrium wage in terms of the output wedge, the capital wedge, and the entry barrier, allows us to obtain sharp comparative statics, which we return to below.

Our empirical analysis focuses on wages and aggregate TFP across locations. Given the wage rate w we can calculate the equilibrium aggregate TFP in each location. We measure aggregate TFP as the Solow residual following a standard growth accounting procedure (as we did in Section 2.2). In particular, we impose an aggregate Cobb-Douglas production function with a weight $\alpha\eta$ on labor and $1-\alpha\eta$ on capital, $\ln Z \equiv \ln Y - \alpha\eta N - (1-\alpha\eta)K$. 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,

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

where

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

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

It is useful to lay out the theoretical predictions for firm entry, namely, the measure of firms entering the location, denoted $\Gamma = Pr(z \geq z^*)$. This is given by

$$\begin{aligned}\ln\Gamma &= \ln[(1-\psi)\int_{z^*}^{\infty}\underline{z}^{\xi}\xi z^{-\xi-1}dz] \\ &= \mu(1-\eta)\ln\left[(1-\psi)\underline{z}^{\xi}\right] + \mu\alpha\eta\xi\ln(N) \qquad (7)\\ &\quad +\mu\xi\ln(1-\tau_y) - \mu\xi\eta(1-\alpha)\ln[(1+\tau_k)(r+\delta)] + \tilde{\Omega}(\alpha,\eta,\xi,\nu),\end{aligned}$$

where $\tilde{\Omega}(\alpha, \eta, \xi, \nu)$ is a constant. It follows immediately that the number of firm entrants is rising in N and falling in τ_y, τ_k, and ψ.

3.4 Comparative statics on wages, TFP, and firm entry

We now summarize the comparative statics of the wedges and of labor supply on the endogenous outcomes we will study in the empirical analysis, namely wage rates, aggregate TFP, and firm entry. Consider the effect of the various wedges on the equilibrium allocations and prices. As is clear from equations (5)-(6), increasing τ_y and τ_k will lower the equilibrium wage and increase the aggregate TFP. The mechanism is that increasing these wedges will lower profits and distort the optimal size and optimal use of capital in the firm, which makes it less attractive for potential entrepreneurs to enter. This increases the TFP cutoff z^*, thereby inducing *positive selection* among entrants: only the most productive entrepreneurs will enter when there are large distortions to capital and output, i.e., when τ_y and τ_k are large. Lower entry in turn reduces the demand for labor, inducing a lower equilibrium w. The key insight is that capital wedges and output wedges cause the wage rate and the aggregate TFP to move in opposite directions.

While τ_y and τ_k have similar qualitative effects on wages, aggregate TFP, and firm entry, they have different effects on the aggregate labor income share and on the aggregate capital-wage-bill ratio. This is what identifies these wedges. We return to this below.

Consider now varying 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).

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

	$(1-\tau_y)$	$(1+\tau_k)$	$(1-\psi)$	N
wage rate w	+	-	+	-
Solow residual Z	-	+	+	-
Entry Γ	+	-	+	+
Labor income share wN/Y	+	0	0	0
Wage bill/capital ratio wN/K	0	+	0	0
$\frac{Y}{N}$	+	-	+	-

Finally, consider the comparative statics for changing labor supply N. A larger N 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 causes lower aggregate TFP, lower wage rate, and more firm entry (cf. Table 3).

3.5 Heterogeneity in other parameters

In our analysis we hold the parameters $(\eta, \xi, \nu, \underline{z})$ constant across locations. We could in principle have allowed geographical variation in any of these parameters. There are several reasons why we have ignored such variation.

Consider first ξ, the Pareto parameter for f, the distribution of firm-specific TFP. Recall that all firm selection rests on variation in the TFP cutoff z^*. From eq. (3) the firm size is linear in z so the firm-size distribution inherits the distribution of z above z^*. Due to heterogeneity in z^* driven, for example, by the other wedges, the lower tail of the firm-size distribution might differ across locations. However, so long as the f distributions share the same parameter ξ across locations, the upper tail of the firm-size distribution and the firm TFP distribution should be identical across locations. Namely, even though there are few productive entrants in low-performing locations (where z^* is low), the distribution of the most productive firms should, according to the model, be identical.

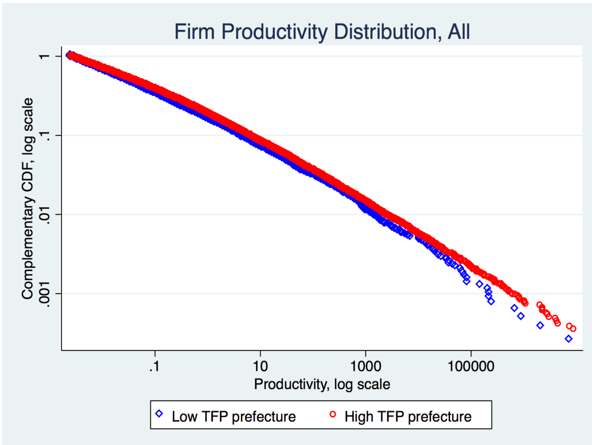
To investigate this theoretical implication we sort the prefectures according to their aggregate TFP. High-TFP (low-TFP) prefectures have aggregate TFP above (below) the median aggregate TFP in 1995. For each group we plot the distribution of firm-specific TFP conditional on their TFP being above the 90th percentile in the overall distribution.

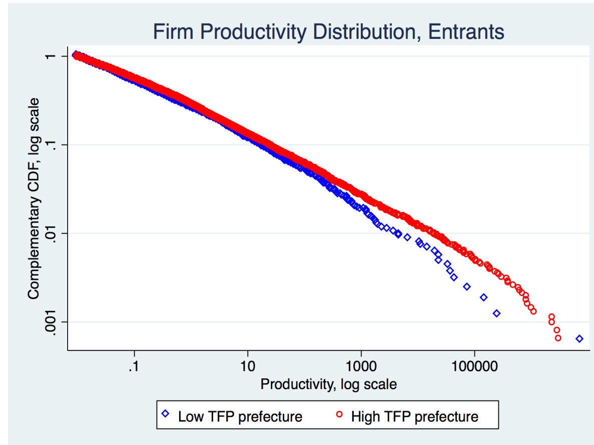
We first conduct this analysis for all firms. We then repeat the analysis for new firms only. i.e., for the subset of firms established after 1993. Figure 4 plots in log scales the complementary cumulative distribution functions for z in low and high TFP prefectures, respectively, for firms in the top 10% of the overall productivity distribution. The two distributions are remarkably 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. We conclude that it is plausible to abstract from geographical heterogeneity in the Pareto parameter ξ. The implied Pareto parameter for the firm size distribution in our sample of non-state manufacturing firms in China is 1.055, which is remarkably similar to the corresponding Pareto tail value that Axtell (2001) reports for the United States, 1.06.

The model assumes that the parameter capturing the fixed operating cost ν is identical across locations. We could alternatively assume heterogeneity in ν instead of modeling heterogeneity in ψ. However, we find it more intuitive to let differences in locations, over and above capital and output frictions, to be captured by heterogeneity in ψ.

Finally, consider the lower bound for the distribution of firm TFP, $\underline{z}^{\xi}$. Note that the terms $(1-\psi)$ and $\underline{z}^{\xi}$ enter multiplicatively in equations (5)-(6). Thus, variation in $\underline{z}^{\xi}$ 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 $\underline{z}^{\xi}$) is equivalent to lowering the effective number of potential entrepreneurs. We prefer to restrict our analysis to geographic heterogeneity in the entry barrier (as opposed to heterogeneity in the distribution of entrepreneurs) because we find it – in the Chinese context – more natural to envision differences across prefectures in government policies rather than differences in the distribution of potential entrepreneurs.

Figure 4: The Truncated Distribution of $\ln z$ for Prefectures Low and High Aggregate TFP in 1995.





Notes: All prefectures are separated into two groups based on their aggregate Solow residual. The figure plots the the complementary cumulative distribution function for the entire firm productivity distribution in 1995, conditional on firm TFP being in the top 10% of the firm TFP distribution. The results for the sample of all firms (entrants) are on the left (right) panel.

4 Measuring the Wedges

We now use the benchmark model to estimate the wedges using data from the Industrial Census. This exercise is in the spirit of Chari et al. (2007) and Hsieh and Klenow (2009). The purpose is to study the drivers of the correlation structure and the regional convergence of economic performance documented in Section 2. Recall from Table 3 that the entry barrier is the only wedge that on its own would give rise to positive correlation between wages, aggregate TFP, and firm entry, as documented above. We shall argue below that the entry barrier emerges as the quantitatively most salient factor in accounting for the changes over time and, hence, the convergence in wages and aggregate TFP that motivated our analysis.

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:

$$\begin{aligned} 1-\tau_y &= \frac{1}{\alpha\eta}\frac{w_i n_i}{y_i}\ , \\ 1+\tau_k &= \frac{1-\alpha}{\alpha}\cdot\frac{w_i n_i}{(r+\delta)\,k_i}\ . \end{aligned} \tag{8}$$

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.[14] In Section 6 we extend the analysis to allow firm-specific wedges.

[14]Bils, Klenow and Ruane (2017) argue that measuring distortions at the firm level 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, very few firms can be linked over time in our data because of changes in the assignment of firm IDs.

Using equation (8) 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 weighted average labor-income share for each firm in that prefecture, weighted by the firm's relative value added:

$$\Delta_p^y = \sum_{j=1}^{J} \left(\frac{1}{\alpha_j \eta} \sum_{i\in(j,p)} \frac{w_i n_i}{y_i} \frac{y_i}{Y_{j,p}} \right) \frac{Y_{j,p}}{Y_p}, \tag{9}$$

where $\alpha_j \eta$ is the technological labor share of industry j. We take these shares from Hsieh and Klenow (2009).

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. Then:

$$\Delta_p^k = \sum_{j=1}^{J} \left(\frac{1-\alpha_j}{\alpha_j} \sum_{i\in(j,p)} \frac{w_i n_i}{k_i} \frac{k_i}{K_{j,p}} \right) \frac{K_{j,p}}{K_p}. \tag{10}$$

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).[15] 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). Figure 5 plots the results for the gross output and gross capital wedges in each prefecture in 1995.[16]

The left panel in Figure 5 shows the gross output wedge in 1995 for each prefecture as a function of the 1995 SOE output share in that prefecture. The gross output wedge is increasing in s, where s denotes the 1995 SOE output share in the prefecture. This implies that in 1995 non-SOE firms in some of the high-s prefectures are receiving subsidies while non-SOE firms in the low s prefectures are being significantly taxed. The right panel in Figure 5 shows the gross capital wedge in 1995 for each prefecture as a function of the 1995 SOE output share in that prefecture. The gross capital wedge 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 Section 2. The slightly increasing capital wedge is a force for wages to fall in s. However, the effect of τ_y dominates. This suggests that the entry barrier is crucial for accounting for the cross-sectional patterns in the data.

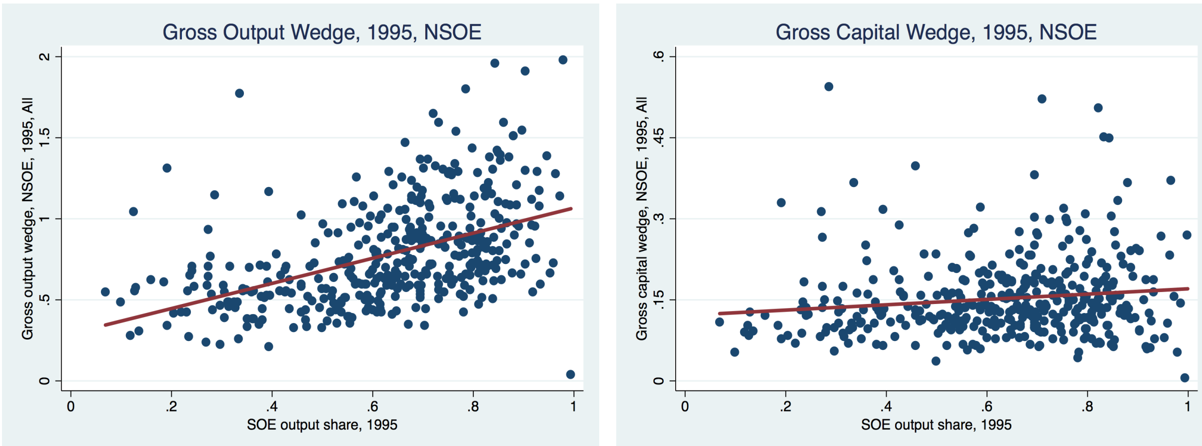
4.2 Log gross entry barrier, $\ln(1-\psi)$

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 (5), we derive an analytical

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

[16] Appendix B shows the results for 2004 and 2008. We compute the gross output and gross capital wedges using all firms in a given cross-section. The results based on the sample of new entrants are very similar and are provided in Appendix B.

Figure 5: 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.

expression for the log gross entry barrier in a prefecture:

$$\begin{aligned}\ln(1-\psi_p) &= \frac{1-\eta+\xi\alpha\eta}{1-\eta}\ln w_p - \frac{\xi}{1-\eta}\ln\Delta_p^y + \frac{\xi\eta(1-\alpha)}{1-\eta}\ln\Delta_p^k \\ &\quad + \ln N_p + \bar{\Omega}(\alpha(p),\eta,\xi,\underline{z},\nu)\,, \end{aligned} \tag{11}$$

where $\bar{\Omega}$ is a constant.[17] 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 $\alpha(p)$. Remember that we have normalized the number of potential entrepreneurs to unity. We interpret this as assuming that the number of potential entrepreneurs is proportional to total employment in manufacturing in the prefecture. Consequently, we measure non-SOE employment in year t, N_t, as the fraction of workers in manufacturing employed in the non-SOE sector relative to total manufacturing employment in 1995 in that prefecture.

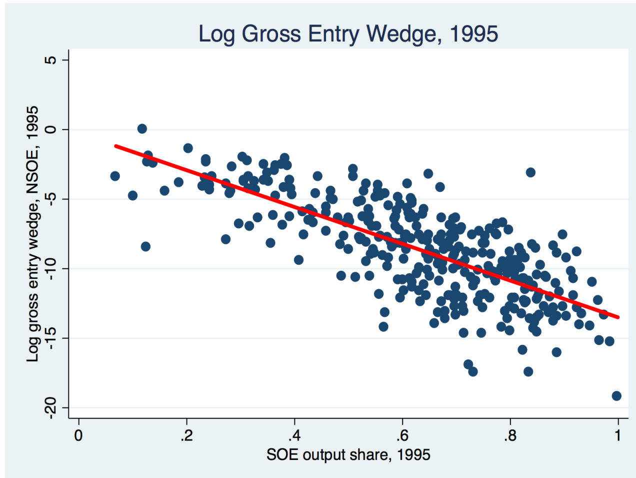
The remaining parameters, which are common across all prefectures, are chosen as follows. The Pareto parameter ξ is obtained by exploiting the theoretical implication that the upper tail of the firm TFP distribution is the same in all prefectures. The Pareto assumption implies that $E(z|z\geq z^*)/z^* = \xi/(\xi-1)$. Focusing on the 30% most productive firms implies $\xi = 1.05$.[18]

Finally, using equation (11), we compute the log gross entry barrier $\ln(1-\psi_p)$ for all prefectures in the economy. Figures 6 and 7 present the results for each prefecture in 1995, 2004, and 2008. The figures reveal a strong negative relationship between the entry barrier $1-\psi$ and the 1995 output share of SOE firms s: a higher barrier ψ is associated with a larger s. In 1995 s explains 51% of the variance in $\ln(1-\psi)$. Moreover, over time there is some convergence in ψ across prefectures.

[17] $\bar{\Omega} = \ln[\frac{\xi-1}{\xi}\nu^{\xi-1}] - (\frac{1-\eta+\xi\alpha\eta}{1-\eta})\ln\alpha\eta - \frac{(1-\alpha)\xi\eta}{(1-\eta)}\ln(1-\alpha) + (1-\xi)\ln(1-\eta)$.

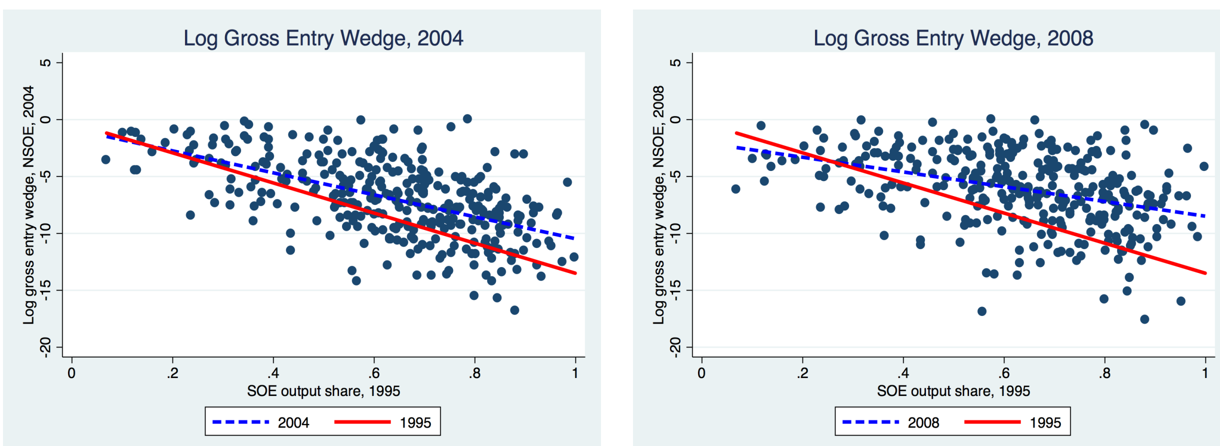
[18] The parameters ν and $\underline{z}$ do not matter for the wedges beyond normalizing the average level because these parameters do not interact with any of the wedges in equations (5), (6), (7), (9), and (10). We normalize the fixed cost of operating a firm, ν, so that the smallest optimal size for a firm with TFP at the threshold $z = z^*$ is one worker: $n^*(z^*) = 1$. Moreover, we normalize the lower bound for the distribution of potential TFP, $\underline{z}^{\xi}$, so that all potential entrepreneurs get a licence in a location without barriers, i.e., when $\psi = 0$.

Figure 6: 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 7: 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.

4.3 Accounting for convergence in TFP and wages

One of the objectives of the paper is to explain the strong regional convergence in aggregate TFP and wages rates that we documented in Section 2. We use our model as a measurement device to account for the convergence. According to the model there are five possible sources of changes over time in aggregate allocations and prices in a prefecture: changes in the three wedges, growth in labor supply (i.e., increased employment in non-SOE manufacturing), and changes in the prefecture-specific production function (i.e., the weight on labor supply in the production function, $\alpha\eta$). The latter is motivated by the fact that changes in industrial structure − for example, growth in the relative preponderance of labor-intensive industries − can be expected to induce changes in the aggregate labor intensity.[19]

Table 4 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 2004-2008 rates of convergence in TFP and wages. Note that while the convergence numbers for wages are identical to those reported in Table 1, the convergence rates for TFP are slightly higher. This is due to the fact that the model does not explicitly incorporate different industrial sectors and thus prefecture TFP is computed using equation (6) with an averaged prefecture-level labor share. In contrast, the TFP measures reported in Table 1 were obtained by averaging over TFP measures computed at the sectoral level in each prefecture. This leads to slight differences in the computed convergence in TFP, although the overall patterns remain unchanged.

To decompose the overall rates of convergence into each of the five possible sources of change, we use equations (5) and (6) to compute aggregate TFP and wages in 2004 if the only prefecture-specific change between 1995 and 2004 was in: (i) the average labor share; (ii) available labor force; (iii) capital wedge; (iv) output wedge; and (v) entry barrier. We then repeat the exercise for the 2004-2008 period. The main message from the first two columns of Table 4 is that changes in the entry barrier account for the lion's share of the convergence in aggregate TFP: if the only change between 1995 and 2004 had been the estimated change in the entry barrier, the annual rate of convergence would have been 3.5%, accounting for more than 92% of the overall 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. We will return to this point in Section 5.

The other factors play a smaller quantitative role in accounting for the convergence in aggregate TFP. The second most important factor is the output wedge, accounting for about 26% of 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.[20]

The findings for wages echo the results for aggregate TFP: the entry barrier emerges as the main explanatory factor for the convergence of wages, accounting for a large fraction of the convergence over the entire 1995-2008 period. Changes in the capital wedge and in the technological labor share also contribute to explaining parts of the convergence in wages, however these factors play quantitatively smaller roles 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, the output wedge cannot have a positive contribution to observed convergence in TFP without at the same time contributing negatively to observed convergence in wage rates.

Interestingly, changes in labor supply play only a minor quantitative role in accounting for the convergence in wages and TFP, despite its potential to move wages and aggregate TFP in the same direction (cf. Table 3). On net, the growth in private manufacturing employment was slightly larger in prefectures with initially high wages. This explains why employment changes – which incorporates migration – account for a positive albeit small share of the convergence in wages.

[19]See also the discussion in Section 3.5 motivating why the remaining parameters of the model are held constant across prefectures.

[20]To see this, consider Figures B-4 and B-5 in the appendix. As is clear from the figures, 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.

More generally, the contribution is quantitatively small because private employment increased almost everywhere and not just in the places where wages and TFP were initially high.

Table 4: Annual Rate of Convergence in TFP and Wages: 1995-2004 and 2004-2008.

	TFP		Wages	
Change in	1995-2004	2004-2008	1995-2004	2004-2008
all	0.038	0.063	0.087	0.025
$\alpha\eta$	-0.003	-0.006	0.015	0.007
N	0.001	-0.001	0.006	-0.009
$(1+\tau^k)$	-0.006	0.003	0.007	0.008
$(1-\tau^y)$	0.010	0.024	-0.003	-0.040
$(1-\psi)$	0.035	0.040	0.034	0.052

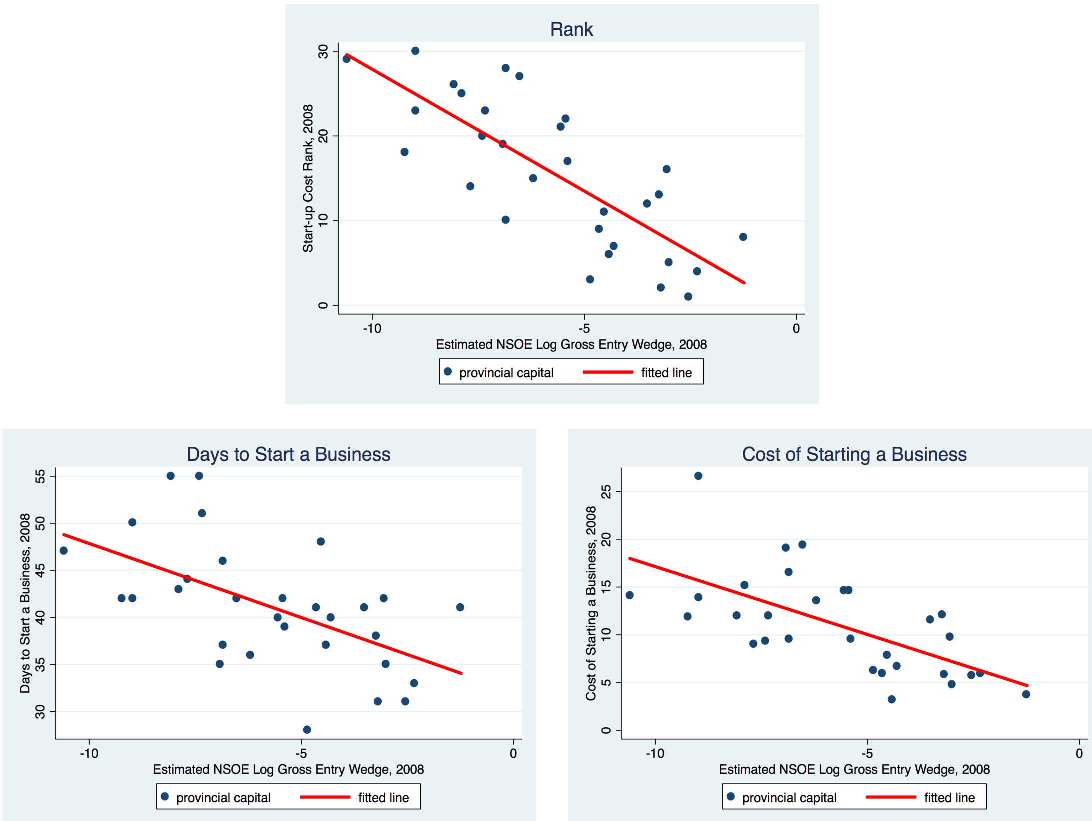
Notes: The table reports the annual rate of convergence in TFP and wages across prefectures for the 1995-2004 and 2004-2008 time periods. The β-convergence coefficient for prefectures p between times t_0 and t_0+T is estimated from the regression $\left(\frac{1}{T}\right)\ln\left(\frac{y_{p,t_0+T}}{y_{p,t_0}}\right)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

4.4 External validation of the entry barriers

Given the salience attributed to the entry barriers in Section 4.3, we now provide external validation that our imputed entry barriers capture actual barriers to entry for private firms. To this end, we perform two exercises: (1) relate our measures to those of the World Bank; and (2) study the implications of the wedges and barriers for 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 activity affects 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 8, 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 8: “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. The benchmark model has predictions for how the three wedges should influence firm entry. As discussed in Section 4.2, increases in ψ, τ_y, and τ_k should all contribute to reduced firm entry (cf. equation (7)). Since we did not target firm entry rates when estimating the wedges (ψ, τ_y, τ_k), the model will not necessarily be consistent with the empirical patterns for firm entry. It follows that data on entry provide an auxiliary test on 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.[21] We define as new those firms that were started in year t, $t-1$, or in $t-2$. The top panel in Table 5 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 the same regression in growth rates:

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

[21] 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.

Table 5: 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	0.216*	9.8%	-0.194***	-9.3%	0.091***	31.5%
2004	0.351**	12.5%	-0.202**	-8.5%	0.051***	17.4%
2008	0.180**	9.7%	-0.151**	-9.3%	0.022***	7.4
	$\Delta\ln(1-\tau^y)$		$\Delta\ln(1+\tau^k)$		$\Delta\ln(1-\psi)$	
	γ_1	$1sd$	γ_2	$1sd$	γ_3	$1sd$
1995-2004	0.477**	16.5%	-0.533***	-25.5%	0.078 ***	20.1%
2004-2008	0.059	3.1%	-0.102	-5.3%	0.003	0.7%

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%.

Equation (7) 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 $\gamma_3 > 0$). Moreover, the effect of changes in the entry barrier is quantitatively large: a one standard deviation change in $\ln(1-\psi)$ induces a 35% change in the entry rate. Also capital and output wedges influence entry rates in the predicted direction. We interpret this as external validation of our model and a confirmation of the mechanism through which the measured entry barriers influence the economy. The results also corroborate the finding that the entry barrier is quantitatively important.

5 The Role of the State Sector

Section 4 established that the entry barrier is the most important factor for understanding the dispersion and the dynamics of 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 barriers 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 barriers in a prefecture – a larger SOE sector in a prefecture is associated with a larger entry barriers 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.

5.1 Entry barriers and the size of the state sector

5.1.1 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 (12) 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:

$$\ln(1-\psi)_{p,t} = \beta_0 + \beta \cdot S_{p,t} + X_{p,t}\gamma' + \epsilon_{p,t}\,. \tag{12}$$

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.

Because of potential concerns of endogeneity in the share of the state sector, we also estimate equation (12) using 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 6. 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.

Table 6: The Entry Barrier in 1995, 2004, and 2008.

	$\ln(1-\psi)$	OLS	IV_{lag}	IV_{1978}	IV_{prov}
1995	e^{soe}	-12.42***	-15.13***	-13.56***	-14.55***
		(1.13)	(1.42)	(1.43)	(4.26)
	$\ln FREV$	1.30***	0.89**	1.13**	0.16
		(0.44)	(0.46)	(0.45)	(1.24)
	$\ln PROF^{soe}$	0.23*	0.23*	0.23*	-0.29
		(0.14)	(0.14)	(0.14)	(0.29)
	Controls	Yes	Yes	Yes	Yes
First stage:					
	IV coefficient		0.70***	0.97***	0.99***
	st. error		(0.04)	(0.06)	(0.20)
	R^2		0.74	0.73	0.75
2004	e^{soe}	-10.43***	-14.77***	-17.12***	-22.46***
		(1.24)	(2.01)	(2.30)	(6.00)
	$\ln FREV$	1.18***	0.77*	0.55	0.39
		(0.44)	(0.47)	(0.50)	(0.98)
	Controls	Yes	Yes	Yes	Yes
First stage:					
	IV coefficient		0.59***	0.66***	0.86***
	st. error		(0.05)	(0.05)	(0.24)
	R^2		0.51	0.45	0.57
2008	e^{soe}	-7.31***	-9.09***	-11.10***	-15.75***
		(0.99)	(1.15)	(1.68)	(2.83)
	Controls	Yes	Yes	Yes	Yes
First stage:					
	IV coefficient		0.87***	0.80***	1.09**
	st. error		(0.03)	(0.06)	(0.30)
	R^2		0.76	0.41	0.43

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%.

Compared to OLS, the IV results suggest slightly larger effects of the size of the state sector on measured entry barriers, and less attenuation in these effects over time. 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.

5.1.2 Changes in the size of the state sector

A potential concern for the cross-sectional results in Table 6 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 (12) in first differences, or Equation (13),

$$\Delta \ln(1-\psi)_{p,t} = \beta_0 + \beta \cdot \Delta S_{p,t} + \Delta X_{p,t}\gamma' + \Delta\epsilon_{p,t} \,. \tag{13}$$

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 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).[22] 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 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.[23] 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.[24]

[22]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.

[23]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.

[24]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:

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

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 (13). 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 7: Change in the Entry Barrier and the SOE Sector, 1995-2004.

$\Delta\ln(1-\psi)$	*OLS*	*OLS*	IV_{Bartik}	IV_{Bartik}
ΔS	-2.92***	-2.12*	-5.19**	-5.68**
	(0.99)	(1.17)	(2.21)	(2.41)
$\Delta\ln FREV$		1.16***		0.87**
		(0.37)		(0.41)
First stage:				
IV coefficient			0.66***	0.70***
st. error			(0.07)	(0.08)
R^2			0.20	0.29

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 7, we report the results from the fixed effects regression using the data for 1995 and 2004.[25] 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 6. 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 coefficients is now about half that of our estimates from the cross-sections. 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.

5.2 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

[25] The analysis in this section has foced on the entry barrier. However, state employment could matter also for the other wedges. In Table B-1 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.

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.[26] We include in the regression the initial level of $\log(x)$ in 1995 and province-level fixed effects.

Table 8: NSOE growth and the role of the state sector, 1995-2004

	$\Delta \ln w$	$\Delta \ln(VA/N)$	$\Delta \ln TFP$	$\Delta \ln(K/N)$
$\Delta \hat{S}_p$	-0.64***	-1.31***	-0.39*	-1.99***
	(0.18)	(0.32)	(0.23)	(0.42)
$\ln x_{1995}$	-0.39***	-0.45***	-0.34***	-0.53***
	(0.04)	(0.05)	(0.04)	(0.07)
Province F.E.	Yes	Yes	Yes	Yes
First stage				
β_{IV}	0.80	0.80	0.86	0.80
s.e.	(0.09)	(0.09)	(0.09)	(0.09)
adj. R^2	0.36	0.34	0.38	0.34

The table reports the 1995-2004 growth in average wages, 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%.

Table 8 contain 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.[27] 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.[28] 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 8, 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.

[26] 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 5.1. Potential general-equilibrium effects of SOE layoffs following 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.

[27] Once the controls are added, the annualized rate of convergence increases to 4.6% for TFP, up from 2.3% 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.

[28] 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.

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 5.1 suggest that a key mechanism for this was though the effect of SOEs on the entry barriers.

5.3 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 6 and 7. 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.[29] 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 wedges. First, they must be non-negative.[30] 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.[31]

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,

$$\frac{1-\bar{\Lambda}_{SOE}}{\bar{\Lambda}_{SOE}} \frac{1}{(1-\psi)} = (1-\tau_y)^{\frac{\xi}{1-\eta}} \left(\frac{1}{1+\tau_k}\right)^{\xi\frac{(1-\alpha)\eta}{1-\eta}}. \tag{14}$$

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.

[29] 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.

[30] 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.

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

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 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 (14). Thus, conditional on ψ and $\bar{\Lambda}_{SOE}$, any combination of (τ_k, τ_y) that satisfies equation (14) 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 satisfy equation (14). This implies a high correlation between SOE employment Λ_{SOE} and entry barriers ψ. We state this result as a formal remark.

Remark 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 Section 5.1 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 Remark 2 and our choice of instruments 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 1998. This is consistent with the IV results in Tables 6 and 7.

5.3.1 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 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).

6 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:

$$\begin{aligned}
E(\ln(1+\tau_k)) &= \ln(1+\bar{\tau}_k) - \frac{\sigma_k}{2} \\
E(\ln(1-\tau_y)) &= \ln(1-\bar{\tau}_y) - \frac{\sigma_y}{2} \\
var(\ln(1+\tau_k)) &= \sigma_k \\
var(\ln(1-\tau_y)) &= \sigma_y \\
cov(\ln(1+\tau_k), \ln(1-\tau_y)) &= \sigma_{ky}.
\end{aligned} \tag{15}$$

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

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.[32]

Proposition 3 *The equilibrium wage rate in the economy that has within-prefecture heterogeneity in capital and output wedges is given by*

$$\begin{aligned}
\ln w &= \mu(1-\eta)\ln\left[\frac{(1-\psi)\underline{z}^{\xi}}{N}\right] \\
&\quad + \mu\xi\ln(1-\bar{\tau}^y) - \mu\xi\eta(1-\alpha)\ln\left(\left(1+\bar{\tau}^k\right)(r+\delta)\right) + \Omega \\
&\quad + \mu\xi\left(\frac{\xi}{1-\eta} - 1\right)\frac{\sigma_y}{2} + \mu\xi\eta(1-\alpha)\left(\frac{\xi\eta(1-\alpha)}{1-\eta} + 1\right)\frac{\sigma_k}{2} - \mu\xi^2(1-\alpha)\frac{\eta}{1-\eta}\sigma_{ky},
\end{aligned} \tag{16}$$

Moreover, the Solow residual is given by,

$$\begin{aligned}
\ln Z &= \mu\alpha\eta(1-\eta)\ln\left[\frac{1-\psi}{N}\underline{z}^{\xi}\right] - \mu(1-\eta)\ln(1-\bar{\tau}_y) \\
&\quad + \mu(1-\eta)(1-\alpha\eta(1-\xi))\ln(1+\bar{\tau}_k) + \hat{\Omega} \\
&\quad - (\xi-(1-\eta))\left(\frac{1}{1-\eta} + \mu\right)\frac{\sigma_y}{2} \\
&\quad - [2(1-\eta)(1-\alpha\eta) + \alpha\xi\eta(2-\eta(1+\alpha))]\frac{\mu(1-\eta+(1-\alpha)\xi\eta)}{(1-\eta)}\frac{\sigma_k}{2} \\
&\quad + \left((1-\eta)(\eta(1-\alpha) + (\xi-1)\alpha\eta + 1) + \alpha\eta^2(1-\alpha)\xi\right)\frac{\mu\xi}{1-\eta}\sigma_{ky}.
\end{aligned} \tag{17}$$

The Solow residual is falling in ψ, 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, $\bar{\tau}_k$, 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

[32] See Appendix C for details.

cross-sectional dispersion (cf. eq. (5) and (6)). 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 (16), 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 (17)). 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}).[33] 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.[34]

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.*

$$\frac{std\left\{(1+\tau_k)(r+\delta)\,|z\geq z^*\right\}}{E\left\{(1+\tau_k)(r+\delta)\,|z\geq z^*\right\}} = \sqrt{\exp(\sigma_k)-1} \tag{18}$$

$$\frac{std\left\{(1-\tau_y)\,|z\geq z^*\right\}}{E\left\{(1-\tau_y)\,|z\geq z^*\right\}} = \sqrt{\exp(\sigma_y)-1} \tag{19}$$

$$\frac{cov\left\{(1+\tau_k)(r+\delta),(1-\tau_y)\,|z\geq z^*\right\}}{E\left\{(1-\tau_y)\,|z\geq z^*\right\}\cdot E\left\{(1+\tau_k)(r+\delta)\,|z\geq z^*\right\}} = \exp(-\sigma_{ky})-1 \tag{20}$$

$$E\left\{(1-\tau_y)\,|z\geq z^*\right\} = \exp\left(\ln(1-\bar{\tau}_y)+2\frac{\xi}{1-\eta}\frac{\sigma_y}{2}-\left(\xi(1-\alpha)\frac{\eta}{1-\eta}\right)\sigma_{ky}\right) \tag{21}$$

$$E\left\{(1+\tau_k)(r+\delta)\,|z\geq z^*\right\} = \exp\left(\ln[(1+\bar{\tau}_k)(r+\delta)]+2\xi(1-\alpha)\frac{\eta}{1-\eta}\frac{\sigma_k}{2}-\frac{\xi}{1-\eta}\sigma_{ky}\right) \tag{22}$$

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 (18)-(22) show that the prefecture-specific means $\bar{\tau}_y$ and $\bar{\tau}_k$ and the variance-covariance

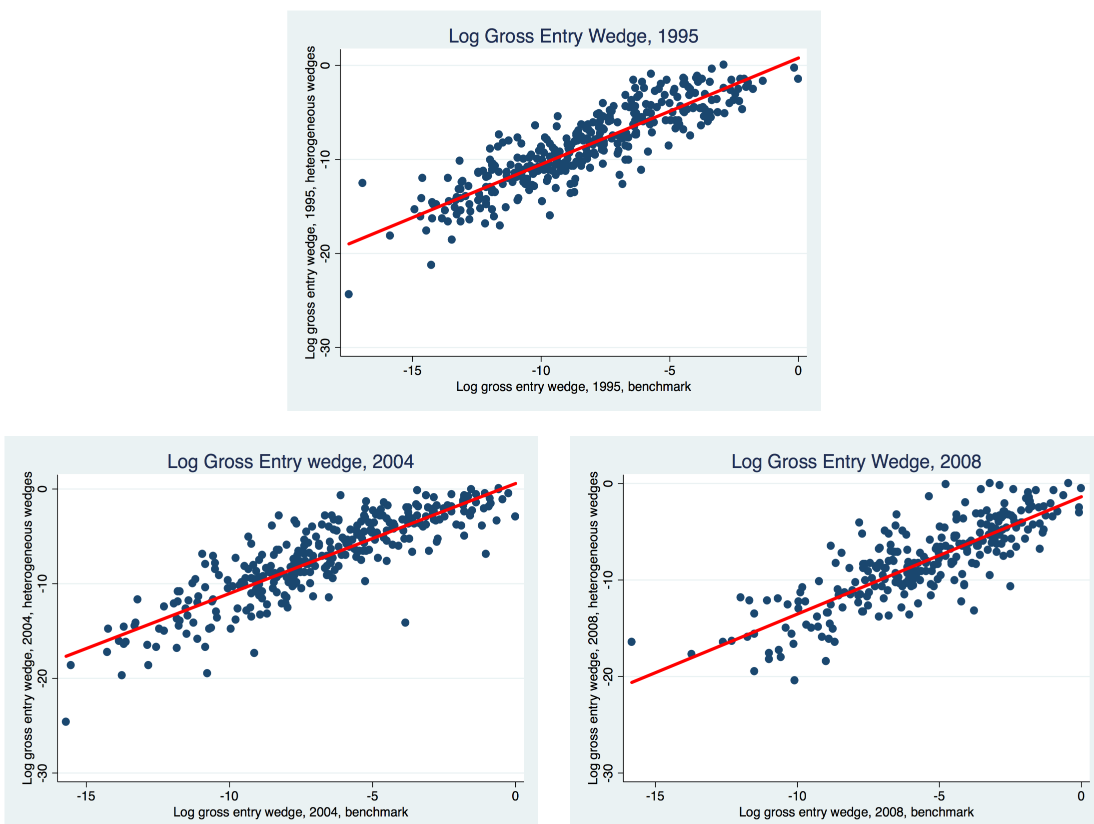
[33]The comparative statics for the Solow residual echo the theoretical finding of Hsieh and Klenow (2009). However, they did not study the comparative statics on the equilibrium wage rate.

[34]See Appendix C for details.

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 (16), 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 9 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-2, 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 affects only the convergence in wages. The covariance between the output and capital wedges, however, affects both the convergence in wages and TFP. Finally, as presented in Table B-3, 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.[35]

Figure 9: 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.

[35] 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.

7 Conclusion

This paper studies regional economic growth in China. Using firm-level data from the Chinese Industrial Census, we construct prefecture-level aggregate data for manufacturing. We document that China experienced a remarkable regional convergence in wages, TFP, productivity, and capital per worker in non-state manufacturing firms 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.

Using the model as an accounting device, we then 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. In contrast, the capital and output wedges play only a limited role in explaining the empirical regional convergence.

Finally, given the preponderance of the entry barriers in accounting for economic performance, we investigate the empirical drivers of these distortions. We find that the presence of state-owned firms gives rise to larger entry barriers for non-state firms. Moreover, based on a Bartik instrumental variable approach exploiting the major 1997 SOE reform that resulted in a decline in the role of state-owned firms in many industries, we argue that the presence of state firms has had a causal effect on increasing the 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 assumption, often dictated by data limitations. For example, to minimize the role of measurement error we have focused on prefecture-level distortions and abstracted 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. Following a standard assumption in the misallocation literature, we have assumed a Cobb-Douglas production function on the firm level, with capital and labor as the only inputs. We do not have data on input prices. This precludes an interesting avenue of research, investigating the potential role of heterogeneity in input prices. We leave this for future research.

We conclude that the gradual removal of entry barriers has been a major driver of aggregate growth and regional convergence in China. It follows that the 1997 SOE reform contributed to regional convergence to the extent that the decline in SOE presence contributed to scaling back the entry barriers. Moreover, our analysis provides a potential mechanism for the recent downturn in economic growth in China, namely, that the resurgence in the state sector following the Global Financial Crisis (see Lardy (2019)) may have contributed to larger entry barriers for non-state firms and, hence, lower non-state sector growth.

References

Axtell, R. L., "Zipf Distribution of U.S. Firm Sizes," *Science* 293 (2001), 1818–1820.

Bai, C.-E., C.-T. Hsieh and Z. Song, "Institutional Foundation of China's Growth," Mimeo, 2018.

Barro, R. J. and X. X. Sala-i-Martin, “Convergence Across States and Regions,” *Brookings Papers on Economic Activity* 1 (1991), 107–182.

Barseghyan, L. and R. DiCecio, “Entry Costs, Industry Structure, and Cross-Country Income and TFP Differences,” *Journal of Economic Theory* 146 (2011), 1828–1851.

Bartik, T., “Who Benefits from State and Local Economic Development Policies?,” Technical Report, W.E. Upjohn Institute, 1991.

Baum-Snow, N., L. Brandt, V. J. Henderson, M. A. Turner and Q. Zhang, “Roads, Railroads and Decentralization of Chinese Cities,” *Review of Economics and Statistics* 99 (2017), 435–448.

Bils, M., P. J. Klenow and C. Ruane, “Misallocation or Measurement?,” Mimeo, 2017.

Brandt, L., C.-T. Hsieh and X. Zhu, “Growth and structural transformation in China,” in L. Brandt and T. Rawski, eds., *China’s Great Economic Transformation* (Cambridge, United Kingdom: Cambridge University Press, 2008a).

Brandt, L., D. Ma and T. Rawski, “Industrialization in China,” in K. O’Rourke and J. Williamson, eds., *The Spread of Modern Industry to the Periphery since 1871* (New York: Oxford University Press, 2017).

Brandt, L., T. Rawski and J. Sutton, “China’s Industrial Development,” in L. Brandt and T. Rawski, eds., *China’s Great Economic Transformation* (Cambridge, United Kingdom: Cambridge University Press, 2008b).

Brandt, L. and T. G. Rawski, eds., *China’s Great Economic Transformation* (Cambridge, United Kingdom: Cambridge University Press, 2008).

Brandt, L., J. Van Biesebroeck and Y. Zhang, “Creative Accounting or Creative Destruction? Firm-Level Productivity Growth in Chinese Manufacturing,” *Journal of Development Economics* 97 (2012), 339–351.

Brandt, L. and X. Zhu, “Redistribution in a Decentralized Economy: Growth and Inflation in China under Reform,” *Journal of Political Economy* 108 (2000), 422–439.

Chari, V. V., P. J. Kehoe and E. R. McGrattan, “Business Cycle Accounting,” *Econometrica* 75 (2007), 781–836.

Cheremukhin, A., M. Golosov, S. Guriev and A. Tsyvinski, “The Economy of People’s Republic of China from 1953,” Mimeo, 2017a.

———, “The Industrialization and Economic Development of Russia through the Lens of a Neoclassical Growth Model,” *Review of Economic Studies* (2017b), forthcoming.

Hopenhayn, H., “Entry, Exit, and Firm Dynamics in Long Run Equilibrium,,” *Econometrica* 60 (1992), 1127–50.

Hsieh, C.-T. and P. J. Klenow, “Misallocation and Manufacturing TFP in China and India,” *Quarterly Journal of Economics* 124 (2009), 1403–1448.

Hsieh, C.-T. and Z. Song, “Grasp the Large, Let Go of the Small: The Transformation of the State Sector in China,” *Brookings Papers on Economic Activity* (2015), 295–346.

Huang, Y., *Selling China: Foreign Direct Investment during the Reform Era* (Cambridge University Press, 2003).

Lardy, N. R., *The state strikes back : the end of economic reform in China?* (Washington, DC: Peterson Institute for International Economics, 2019).

Li, H. and L.-A. Zhou, "Political Turnover and Economic Performance: The Incentive Role of Personnel Control in China," *Journal of Public Economics* 89 (2005), 1743–1762.

Mankiw, N. G., D. Romer and D. N. Weil, "A Contribution to the Empirics of Economic Growth," *Quarterly Journal of Economics* 107 (1992), 407–437.

Melitz, M. J., "The Impact of Trade on Intra-Industry Reallocations and Aggregate Industry Productivity," *Econometrica* 71 (2003), 1695–1725.

Naughton, B., "The Third Front: Defence Industrialization in Chinese Interior," *The China Quarterly* 115 (1988), 351–386.

Restuccia, D. and R. Rogerson, "Policy Distortions and Aggregate Productivity with Heterogeneous Establishments," *Review of Economic Dynamics* 11 (2008), 707–720.

Sala-i-Martin, X. X., "Regional Cohesion: Evidence and Theories of Regional Growth and Convergence," *European Economic Review* 40 (1996), 1325–1352.

Song, Z., K. Storesletten and F. Zilibotti, "Growing Like China," *American Economic Review* 101 (2011), 202–241.

Tombe, T. and X. Zhu, "Trade, Migration and Productivity: A Quantitative Analysis of China," *American Economic Review* 109 (2019), 1843–72.

Wang, Y., "Is China's Rapid Growth Sustainable? A Theory of Politico-Economic Transition and State Capitalism," Working Paper, University of Oslo, 2017.

Whiting, S. H., *Power and Wealth in Rural China: The Political Economy of Institutional Change* (Cambridge University Press, 2006).

World Bank, "Doing Business in China 2008," Report, The World Bank Group, 2008.

Xu, C., "The Fundamental Institutions of China's Reforms and Development," *Journal of Economic Literature* 49 (2011), 1076–1151.

Young, A., "Gold into Base Metals: Productivity Growth in the People's Republic of China during the Reform Period," *Journal of Political Economy* 111 (2003), 1220–1261.

Zhu, X., "Understanding China's Growth: Past, Present, and Future," *Journal of Economic Perspectives* 26 (2012), 103–124.

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

Loren Brandt (University of Toronto),
Gueorgui Kambourov (University of Toronto), and Kjetil Storesletten (University of Oslo)

December 2019

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).[36] The CIC covers all of the manufacturing sector[37] 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.[38]

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.

[36] 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.

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

[38] 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 8 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 8.

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.

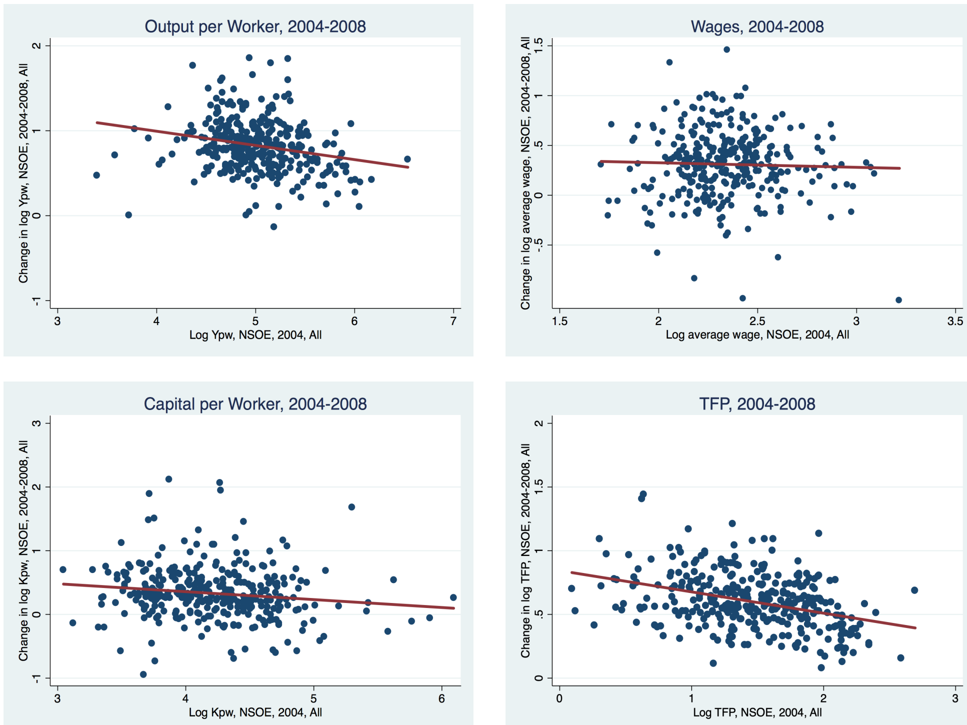
Table A-1: IV_b and alternative definitions of weighted average TFP growth

	$\Delta \ln TFP$ (1)	$\Delta \ln TFP$ (2)	$\Delta \ln TFP$ (3)	$\Delta \ln TFP$ (4)
IV_b	-0.20	-0.39*	-0.40*	-0.39*
	(0.24)	(0.24)	(0.24)	(0.23)
$\ln TFP_{1995}$	-0.23***	-0.27***	-0.27***	-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%.

B Figures and Tables

Figure B-1: Convergence in the NSOE sector, 2004-2008.



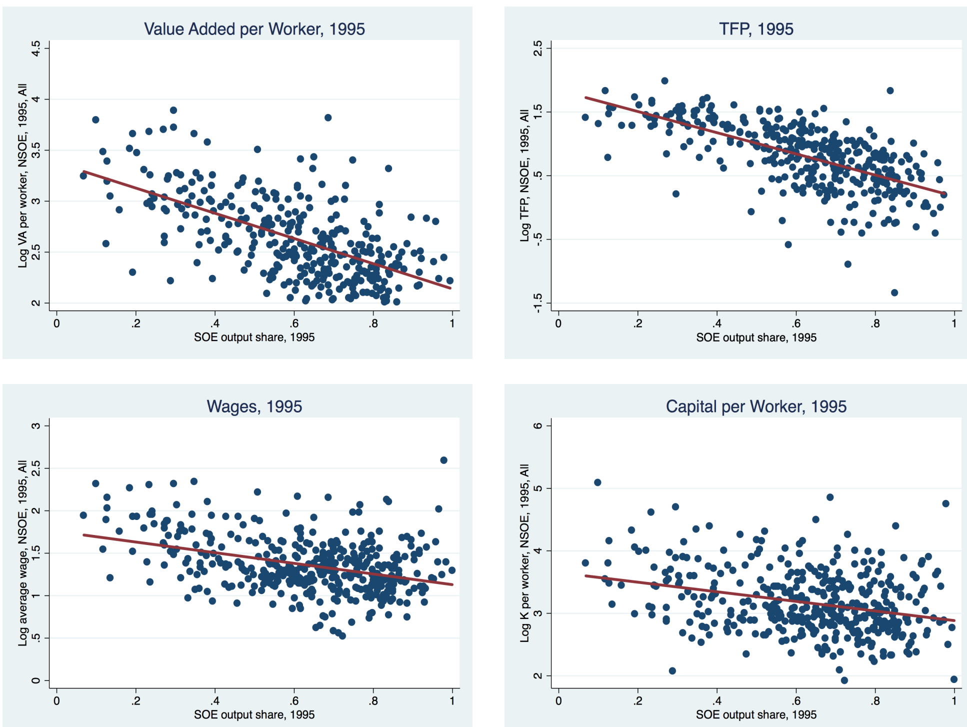
Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line.

Table B-1: Explaining the 1995-2004 changes in wedges.

	$\Delta\ln(1-\psi)$	$\Delta\ln(1+\tau^k)$	$\Delta\ln(1-\tau^y)$
$\Delta\hat{S}_p$	-5.68**	1.63***	1.05***
	(2.41)	(0.50)	(0.37)
$\Delta\ln FREV$	0.87**	0.17**	0.03*
	(0.41)	(0.08)	(0.06)

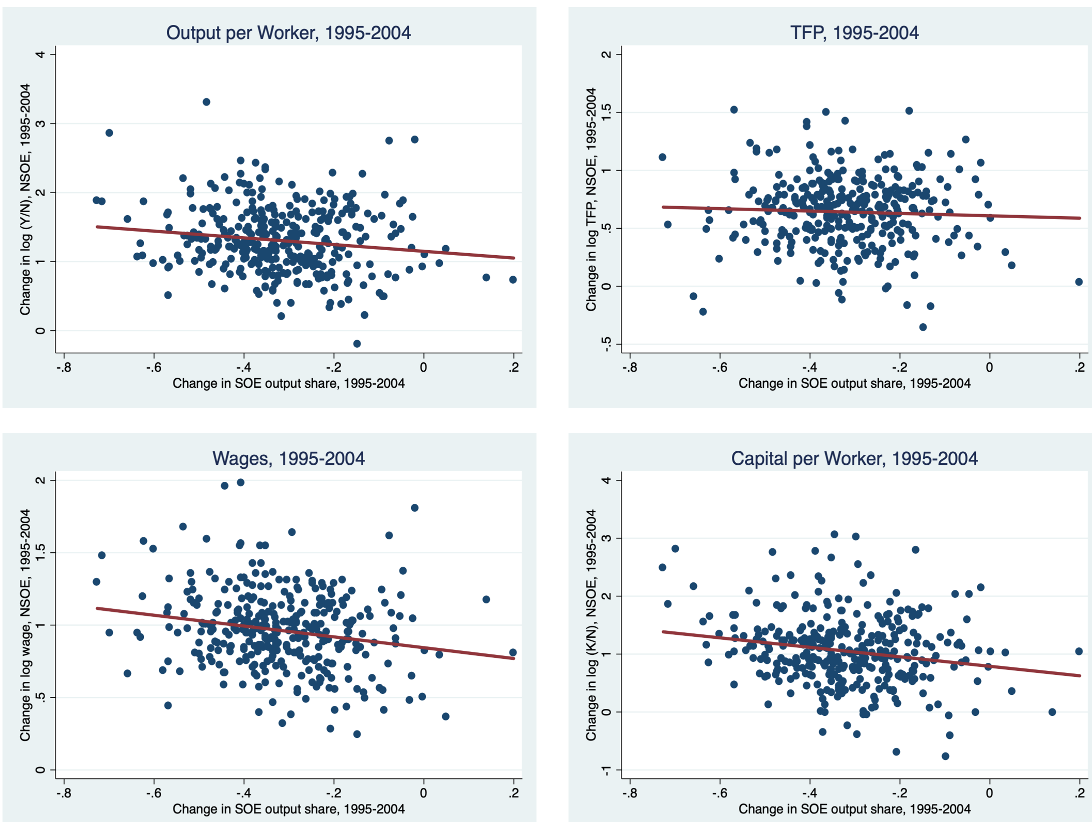
Notes: The table reports the IV results from regression of the change in the log gross wedges on the (instrumented) change in the SOE employment share (e^{soe}) and the log fiscal revenue per government worker ($\ln FREV$) in a prefecture between 1995 and 2004. Standard errors are in parentheses. ***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

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



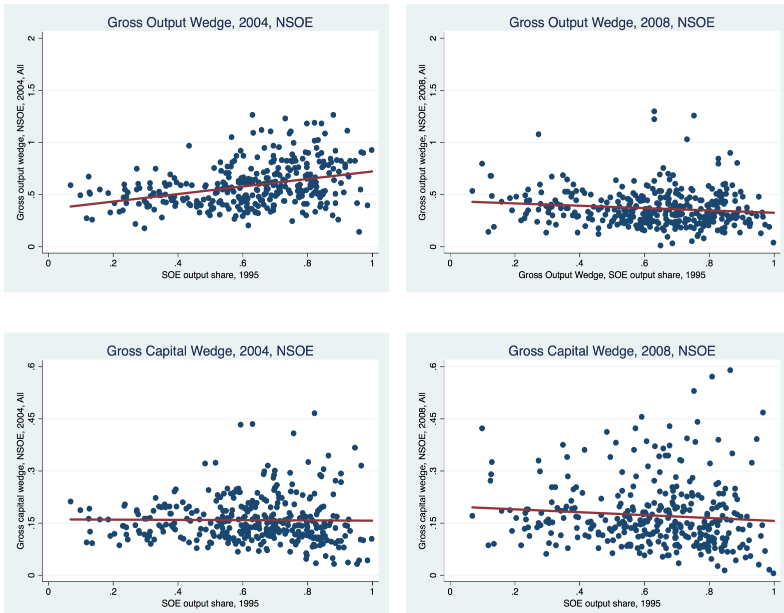
Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line. The 1995 SOE output share in a prefecture is on the horizontal axis.

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



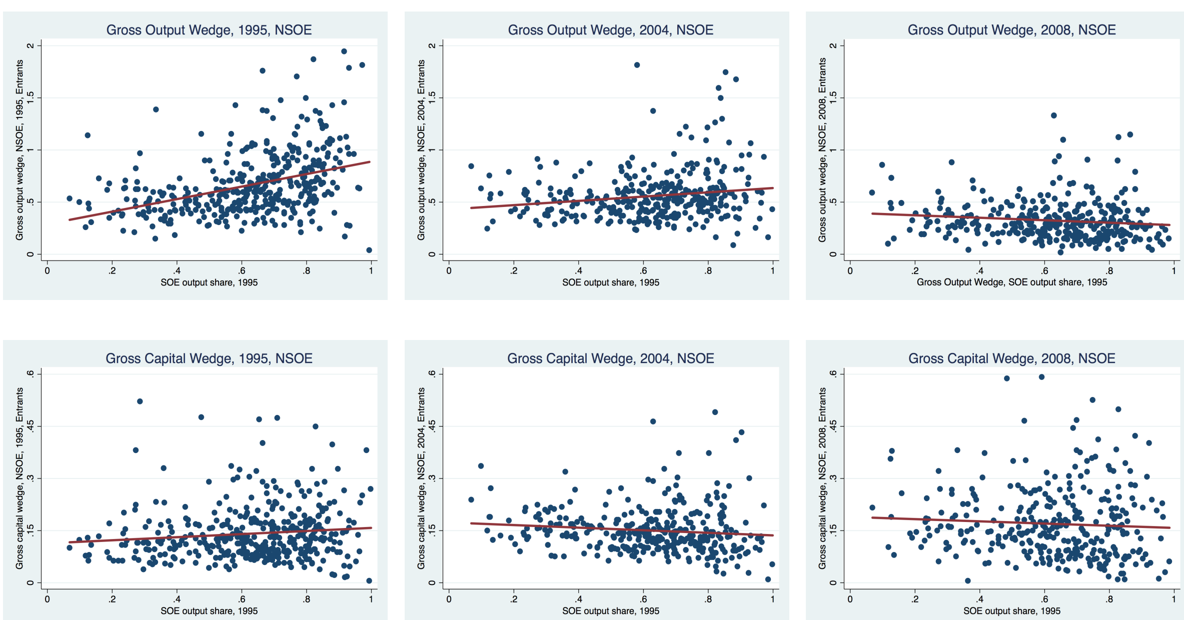
Notes: Each dot represents a prefecture, and the solid red line is the fitted regression line. The 1995-2004 change in SOE output share in a prefecture is on the horizontal axis.

Figure B-4: 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 1995, 2004, and 2008. The SOE output share in 1995 in each prefecture is on the horizontal axis.

Figure B-5: 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-2: Annual Rate of Convergence in TFP and Wages: 1995-2004 and 2004-2008, Heterogeneous-Wedges Model.

	TFP		Wages	
Change in	1995-2004	2004-2008	1995-2004	2004-2008
all	0.038	0.063	0.087	0.025
$\alpha\eta$	0.001	-0.003	0.027	0.012
N	0.001	-0.001	0.009	-0.005
$(1+\bar{\tau_k})$	-0.002	0.002	0.017	0.011
$(1-\bar{\tau_y})$	0.010	0.003	-0.021	-0.054
$(1-\psi)$	0.031	0.019	0.021	0.047
σ_k	0.000	0.000	-0.001	-0.004
σ_y	0.000	0.000	0.006	0.034
σ_{ky}	0.002	0.005	0.012	-0.002

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-3: Change in the Entry Wedge, 1995-2004, Heterogeneous-Wedge Model.

$\Delta \ln(1-\psi)$	*OLS*	*OLS*	IV_p^{ind}	IV_p^{ind}
Δe^{soe}	-6.25***	-6.61**	-4.60	-2.35
	(1.61)	(1.79)	(3.23)	(3.66)
$\Delta \ln FREV$		0.72		0.98
		(0.56)		(0.60)
First stage:				
IV coefficient			0.71***	0.69***
st. error			(0.07)	(0.08)
R^2			0.25	0.27

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 in Section 6

C.1 Proof of Proposition 3

Assume that within each prefecture there is dispersion across firms in the output and capital wedges. It is immediate that the optimal choices of each firm are still given by equations (3) and (4), although the wedges are now specific to each firm.

The joint cross-sectional distribution of the wedges is assumed to be log normal with the moments given by equation (15). 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,*

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

Proof.

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

The second equation follows from the assumption that the distribution function f (of potential z) is identical across locations. The third equation uses the fact that

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

■

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

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

From Lemma 5 the last term is given by

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

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

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

It follows immediately that the equilibrium wage is given by equation (16).

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

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

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

Now compute aggregate capital,

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

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

The Solow residual can then be calculated as

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

Substituting in the wage expression from equation (16) and simplifying yields equation (17) in the text. This concludes the proof of Proposition 3

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 $z \geq z^*\left(\tau^k,\tau^y,w\right)$. Using Lemma 5 the variance of the observed $(1-\tau^y)$ is given by

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

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 (19). 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 (18). Finally, the observed covariance is calculated as follows,

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

It follows that σ_{ky} can be identified from equation (20). This concludes the proof of Proposition 4.