Productivity Dispersion: Misallocation or Adjustment Frictions?

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Abstract

Recent research maintains that the observed variation in productivity across firms reflects resource misallocation and concludes that large GDP gains may be obtained from market-liberalizing polices. Our theoretical analysis examines the impact on productivity dispersion of reallocation frictions in the form of costs of entry, operation, and restructuring, and shows that reforms reducing these frictions may raise dispersion of productivity across firms. The model does not imply a negative relationship between aggregate productivity and productivity dispersion. Our empirical analysis focuses on episodes of liberalizing policy reforms in the U.S. and six East European transition economies. Deregulation of U.S. telecommunications equipment manufacturing is associated with increased, not reduced, productivity dispersion, and every transition economy in our sample shows a sharp rise in dispersion after liberalization. Productivity dispersion under central planning is similar to that in the U.S., and it rises faster in countries adopting faster paces of liberalization. Lagged productivity dispersion predicts higher future productivity growth. The analysis suggests there is no simple relationship between the policy environment and productivity dispersion.

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1. Introduction

Recent research examining business-level data has documented a robust regularity of substantial productivity dispersion even within narrowly defined industries (Bartelsman and Doms 2000; Syverson 2011). In an early study, for instance, Syverson (2004) reports productivity at the 75^{th} percentile nearly double that at the 25^{th} percentile for labor productivity and about 50 percent higher for total factor productivity within four-digit manufacturing industries in the U.S. Findings such as these pose puzzles for economists, though they may be less surprising to non-economists who have not been weaned on models of representative agents and frictionless competition.

One source of the puzzlement is the apparent inefficiency of productivity dispersion. If the same technology is available to all producers, then improving the productivity of poor performers can raise aggregate output. But what prevents productivity equalization? One way to explain persistent dispersion is through the presence of idiosyncratic taxes on output or inputs that change the marginal conditions in different ways for different participants. For instance, a firm facing a higher output tax will ceteris paribus produce less than one facing a lower output tax. Unifying tax rates would eliminate this source of distortion in production decisions, and market reforms that "level the playing field" can thus raise aggregate output. Restuccia and Rogerson (2008) and Hsieh and Klenow (2009) develop models of what we call "static distortions," and the latter provide an empirical analysis of the potential aggregate gains from reducing productivity dispersion in China and India to the level in the U.S.

A different set of factors influencing productivity dispersion, which we focus on in this paper, is analyzed in models of industry dynamics going back to Jovanovic (1982), Hopenhayn (1992), and Ericson and Pakes (1995). These factors comprise various types of reallocation frictions, including sunk costs of entry, fixed costs of operating, and costs of investment with stochastic outcomes. The magnitudes of these costs are partially a function of technological considerations, but they are also affected by policies that change entry barriers, bankruptcy costs, bailout possibilities, and access to finance. By contrast with the static distortions, however, policies to reduce such frictions need not decrease productivity dispersion, as we show in our theoretical analysis below. Indeed, they may result in higher levels of dispersion, especially in the short run. The reason is that while lower frictions tend to strengthen selection mechanisms by raising the threshold productivity for firm survival, they also encourage experimentation that raises dispersion. One type of experimentation is entry, when an entrepreneur receives a draw from a productivity distribution, as in Jovanovic (1982) and Hopenhayn (1992). A second type is investing in or restructuring an incumbent firm, involving a draw from another distribution, one with a higher mean than the firm's pre-investment or pre-restructuring productivity but also with a non-trivial variance and possibly including reduced productivity as an outcome. For each type of reallocation friction, our model shows that reducing the friction may raise experimentation. The result, because of the randomness of outcomes, is a new productivity distribution, potentially one with increased dispersion.

Our empirical analysis uses firm-level data to provide evidence of the effects of reducing reallocation frictions in some wide-ranging settings. First, we consider the U.S. telecommunications equipment sector, which was gradually deregulated between the late 1960s and early 1980s. Our analysis builds on Olley and Pakes' (1996) study of this sector, but they do not examine productivity dispersion, which is our focus here. Second, we analyze the evolution of productivity dispersion in six economies undergoing a transition from central planning to some form of liberalized market economy, but with widely varying paces of reform. Using U.S. productivity dispersion as a benchmark, as in Hsieh and Klenow (2009), we track productivity dispersion in the manufacturing sectors back into the planning period and as they liberalize, some of them very quickly (with a "big bang" of reforms) and others more slowly (the "gradualists").

In each of these cases, the evidence suggests that deregulation and other reforms tend to raise, rather than to reduce, productivity dispersion. Dispersion rises throughout the deregulation of telecommunications equipment manufacturing in the U.S. Remarkably, our calculations of productivity dispersion in Soviet Russia, Soviet Ukraine, and Hungary under central planning are very similar to those for the U.S. In all the transition economies, dispersion rises with reforms, and it rises faster the quicker the reform process proceeds. Regression analysis shows that the extent of reform (as measured by an index from the European Bank of Reconstruction and Development) positively predicts productivity dispersion, and that productivity dispersion is associated with subsequent growth in aggregate productivity. These findings do not exclude an important role for selection mechanisms in truncating the left tail of the productivity distribution, but they suggest a perhaps dominant role for experimentation in widening or thickening both tails.

Some caveats are in order. Our theoretical model implies that overall productivity dispersion may either rise or fall in response to policies that reduce reallocation frictions, although it does predict that dispersion rises among new entrants to the industry when entry frictions fall. The model does not include idiosyncratic taxation, which if reduced would by itself have the effect of lowering dispersion, as shown by Restuccia and Rogerson (2008) and Hsieh and Klenow (2009). But most economic reforms, such as deregulation and liberalization, have effects on frictions as well as on static distortions. Indeed, the purposes of many reform programs are expressed in terms of reducing entry barriers, hardening budget constraints, and increasing access to capital – implying that studies of productivity dispersion may benefit from taking such frictions into account. Our model is also very simple in assuming price-taking behavior, but rather than investigating more complicated market environments our point is only to demonstrate the possibility of different changes in productivity dispersion resulting from reduced frictions.

A major caveat about our data is that, like most data sets, ours permits us to measure only revenue – not quantity – based total factor productivity (but see Foster et al. 2008). This implies that we cannot distinguish changes in pricing or markup behavior from changes in technical efficiency, which would be a particular problem if we were trying to measure potential aggregate efficiency gains from reduced dispersion. Our broader point, however, is that higher dispersion may reflect greater experimentation and therefore higher future growth, so that productivity dispersion need not measure misallocation.

The next section provides our theoretical analysis, Section 3 describes the data, and Section 4 lays out the empirical results. Section 5 provides a brief conclusion. Proofs of theoretical propositions are contained in an Appendix.

2. Theoretical Analysis

We focus on the role of certain institutional factors, represented by three types of adjustment frictions: cost of entry, fixed cost of operation, and cost of investment or restructuring. The cost of entry includes all sunk costs of starting a business. The fixed cost embeds all costs of operation that recur on a periodic basis and do not change with the scale of the business. Finally, the cost of restructuring entails all sunk costs associated with investments intended to raise efficiency. These costs typically vary over time and across industries and countries. For instance, deregulation in an industry or transition to a market economy from a planned one can make both starting a business or expanding it easier, by lowering entry and restructuring costs. Similarly, changes in the regulatory environment, such as more stringent requirements for quality, tighter environmental compliance, or harder budget constraints, can lead to higher fixed costs of operating a business. Major technological innovations in an industry can also result in substantial changes in the costs of entry, exit and restructuring, leading to reallocation. [2]

The model is based on the elements in widely-used dynamic models of competitive industries, such as those of Jovanovic (1982) and Hopenhayn (1992). A key element in these models is a stochastic process for firm productivity that evolves exogenously and drives firm entry, growth, and exit. We add the ability of firms to invest in potential improvements in their productivity, as in Jovanovic and MacDonald (1994) and Ericson and Pakes (1995). In addition, we highlight the contribution of pre-entry heterogeneity among potential entrants to productivity dispersion. Such pre-entry heterogeneity in productivity is absent in the models of Jovanovic (1982) and Hopenhayn (1992), but is featured in recent models of entrepreneurship with selection processes at the entry stage, such as that of Nocke (2006).[3]

Our approach differs from one strand of research that explores the role of "static" distortions, such as taxes and subsidies, varying in proportion with output or variable inputs.[4] There are no such distortions in the current model, which for simplicity also abstracts from adjustment costs applicable to production inputs and any other frictions. [5] We focus instead on the productivity distribution among firms as a function of entry, fixed, and restructuring costs. The aggregate productivity and the dispersion of productivity depend on the magnitudes of these three costs. For

[2] See, e.g., Collard-Wexler and De Loecker (2015) for an analysis of reallocation induced by the introduction of mini-mills in the U.S. steel industry.
[3] Hopenhayn (1992) allows for the initial post-entry distribution of productivity for new entrants to differ from that of the incumbents, but like many other models of industry dynamics does not otherwise allow pre-entry heterogeneity.
[4] See, among others, Hsieh and Klenow (2009), Guner et al. (2008), Restuccia and Rogerson (2008), Midrigan and Xu (2014), and Bartelsman, Haltiwanger and Scarpetta (2013). The last paper features a fixed factor in production (overhead labor), and the distortions to revenue interact with this factor in the firm's choice of labor, but there is no analysis of the effect of a change in the cost of this fixed factor itself.
[5] For analysis of dynamic inputs and adjustment costs, see Asker, Collard-Wexler, and De Loecker (2014) and Butters (2016).

instance, in a world with no fixed costs of operation, all firms, even the least productive ones, survive, leading to a broader range of productivity levels than would prevail when fixed costs are high. Similarly, lower entry costs may allow less productive firms to enter. In general, entry costs can not only protect incumbent firms but also determine the range of productivity for potential entrants that ultimately enter – the process of selection among potential entrants. The model thus provides an alternative framework to interpret the persistent differences in aggregate productivity and its dispersion both across economies, and over time within an economy. Clearly, both the "static" type of distortions studied extensively in the literature and the institutional factors exemplified by the three costs considered here are important for understanding the connection between productivity dispersion and misallocation. Without a better grasp of how both sets of factors may evolve over time or differ across economies, interpreting the causes and consequences of productivity dispersion in an economy may be misleading, as the effects of these frictions operate alongside the effects of static distortions.

Firms in the model are price-takers in output and input markets, and they face perfectly elastic demand, as in Jovanovic (1982) and Hopenhayn (1992). In alternative models featuring monopolistic competition, such as that of Hsieh and Klenow (2009), firms face imperfectly elastic demand, and a firm's price is a fixed markup over its marginal cost, which is inversely proportional to its physical total factor productivity (TFPQ). In this setup, firms with higher TFPQ produce higher quantities, but charge lower prices. When there are no "static" distortions, this mechanism leads to exact equalization across firms of total factor productivity measured by revenue (TFPR), and hence, to no dispersion in TFPR. This is not the case in perfectly competitive models with decreasing returns and price-taking firms, including the current model, where TFPR differs across firms even in the absence of any static distortions.[6] In these models, under general conditions there will be a non-trivial level of TFPR dispersion that is positively associated with TFPQ dispersion. However, this TFPR variation is benign; absent any distortions, it is not associated with allocative inefficiencies.

The connection between institutional environment and the productivity of firms has been studied most frequently in the context of entry conditions across countries.[7] Part of this literature focuses on the connection between entry costs and aggregate productivity. In competitive models similar to the one considered here, higher entry costs can lead to lower aggregate productivity. Empirical evidence indicates that countries with lower entry costs tend to have higher output per worker. To our knowledge, however, no study so far has investigated how entry, fixed, and restructuring costs simultaneously affect productivity dispersion. Because economies and industries exhibit considerable heterogeneity in the magnitudes of these costs, the model carries predictions for the cross-sectional and time-series variation in productivity dispersion.

[6] Foster et al. (2016) show that one need only change Hsieh and Klenow's (2009) constant returns to scale assumption to non-constant returns to scale to generate TFPR dispersion from sources other than static distortions, namely shocks to TFPQ and demand.
[7] See, e.g., Barseghyan and Diceccio (2009), Nicoletti and Scarpetta (2003), and Boedo and Mukoyama (2009).

2.1 The Model

The model builds on the basic framework of Hopenhayn (1992). Consider an industry where a large number (a continuum) of firms produce a homogeneous good. Time is discrete, and firms can survive for multiple periods. Firms take output price, p, and input prices as given. The industry is also a price-taker in input markets.[8] The demand for the good is summarized by $D(p)$, a bounded and downward sloping function.[9] Firms are heterogeneous with respect to their physical productivity (TFPQ), denoted by the random variable $\theta \in [0,1]$, which evolves independently over time and across firms. There is a fixed cost, c_f, of operating in the industry, which is avoidable only if the firm exits the industry. Each period incumbent firms have the option to pay a restructuring cost, c_r, to achieve potential improvements in θ, similar to the mechanisms of investment in research and exploration by firms in Jovanovic and MacDonald (1994) and Ericson and Pakes (1995).

There is a large mass, $N > 0$, of potential entrants.[10] Each entrant can pay a sunk entry cost of c_e to enter. Before entry, the heterogeneity of post-entry productivity among potential entrants is summarized by a distribution $G(\phi)$. The parameter $\phi \in [0,1]$ describes a potential entrant's prior belief about its productivity in the first period following entry. The initial productivity of an entrant is revealed after the entrant incurs the entry cost c_e, and it is a draw from a continuous distribution with c.d.f. $H_e(\theta|\phi)$ and the associated density h_e. $H_e(\theta|\phi)$ is strictly decreasing in ϕ. In other words, a higher prior represents a better idea or blue-print, deeper industry knowledge, more experience, better location, or a superior managerial talent, all of which spawn a better post-entry productivity distribution, in a first order stochastic dominance (f.o.s.d.) sense.[11] New entrants do not have the option to restructure in their first period, but can do so in subsequent periods. An incumbent firm that does not restructure receives its next period productivity draw, θ', from a continuous distribution with c.d.f. $H_n(\theta'|\theta)$ and density h_n. Incumbents that have chosen to restructure, on the other hand, obtain a productivity draw from another continuous distribution with c.d.f. $H_r(\theta'|\theta)$ and density h_r.

A number of assumptions are imposed on the various processes that govern the evolution of θ. First, as in Hopenhayn (1992), H_n and H_r are both strictly decreasing in θ. In other words, the next period's productivity draw for an incumbent is higher, in an f.o.s.d. sense, when its current productivity is higher. Furthermore, restructuring results in a productivity that is on average higher than the firm's initial productivity, i.e. $E_r[\theta'|\theta] > \theta$. In addition, a restructuring incumbent obtains better outcomes, in an f.o.s.d. sense, than in the case of non-restructuring. This amounts to

[8] This assumption can be relaxed. For instance, the unit cost of labor can depend on the extent of employment in the industry. However, such additions do not alter the main messages of this theoretical section.

[9] The demand function also satisfies $\lim_{p\to\infty} D(p) = 0$. As price becomes arbitrarily large, demand vanishes, ensuring that firm profits remain bounded even at very large prices.

[10] N is assumed to be sufficiently large so that even for very low entry costs, entry cannot exhaust the mass of potential entrants. The mass of entering firms is then determined by the type of the marginal entrant, as detailed below.

[11] This formulation of the heterogeneity in entrant population differs from that in Hopenhayn (1992), where all potential entrants are ex-ante identical and they all draw from the same productivity distribution upon entry. Ex-ante entrant heterogeneity is also a feature of some models of entrepreneurship, such as Nocke (2006).

the restriction that $H_r(\theta'|\theta) < H_n(\theta'|\theta)$ for any θ', which implies $E_r[\theta'|\theta] > E_n[\theta'|\theta]$. In addition, all distributions satisfy the property that a firm can at some point receive an arbitrarily small productivity draw with positive probability, regardless of its current type, in some period in the future.[12] Note that both the entry and the restructuring processes allow for the possibility that a firm ends up with a productivity below its initial type – productivity improvements are not guaranteed.[13]

The timing of events in a period is as follows. At the beginning of the period, there is an initial mass, M, of firms and a measure of productivity across firms, μ, such that $\mu(\theta)$ gives the total mass of firms with productivity at most θ, and $\mu(1) = M$. Given this initial configuration, potential entrants decide whether to enter, and those that enter observe their productivity drawn from H_e. At the same time, incumbents decide whether to restructure or exit. All continuing incumbents then receive their productivity draws, either from H_r or H_n, depending on their decisions on whether to restructure. Production then takes place, output price is determined to clear the market, and the period ends.

An incumbent firm uses labor, l, and capital, k, to produce output, $q = \theta f(k, l)$, where f is a production function that exhibits decreasing returns to scale and is strictly concave in its arguments. The firm's profit maximization problem in a period is

$$\max_{l,k} \pi(k, l; \theta, p) = p\theta f(k, l) - wl - rk,$$

where $w > 0$ is the wage and $r > 0$ is the rental rate of capital. Given the setup so far, a firm's profit function, denoted by $\tilde{\pi}(\theta) \equiv \tilde{\pi}(p, w, r; \theta)$, is strictly increasing in θ and p. A firm's output, $\tilde{q}(\theta) = \tilde{q}(p, w, r; \theta)$, is also strictly increasing in θ and p. Furthermore, we assume that $\tilde{\pi}(\theta)$ is multiplicatively separable in θ and prices (p, w, r).[14]

In a stationary environment, a firm's value can be written as

$$V(\theta) = \tilde{\pi}(\theta) - c_f + \beta \max\{0, E_r[V(\theta')|\theta] - c_r, E_n[V(\theta')|\theta]\}. \quad (1)$$

The firm obtains its maximum profit in the current period given its type. In the next period, its value depends on whether it exits, restructures, or does nothing. The value from exit is normalized to zero. The expected value from no restructuring, $E_n[V(\theta')|\theta]$, and restructuring, $E_r[V(\theta')|\theta]$, are given, respectively, by

[12] This assumption ensures that each firm faces a positive probability of exit and limits the life span of firms, allowing for continuing exit in stationary equilibrium Technically, this requires that for any θ, there exists some t, such $H^t(\varepsilon|\theta)$ is strictly positive for any given $\varepsilon > 0$, where H^t denotes the t −period ahead distribution of productivity for the firm. Note that H^t is generated from successive draws from the distributions H_i $i = r, n$, depending on whether the firm restructures in a given period.

[13] The assumptions on the nature of restructuring embed some of the assumptions imposed in some of the earlier models of innovation and learning from others, such as Jovanovic and MacDonald (1994). These common assumptions include (i) restructuring does not guarantee an improvement in productivity, and (ii) a better distribution of productivity cannot be achieved for free (the cost of restructuring is strictly positive). See the discussion in p. 29-30 in Jovanovic and MacDonald (1994).

[14] This would be the case, for instance, if f is a Cobb-Douglas type production function. This assumption is one way to ensure that the stationary equilibrium of the model is unique, if it exists – see also condition U2 in Hopenhayn (1992).

$$E_i[V(\theta')|\theta] = \int_0^1 V(\theta')h_i(\theta'|\theta)d\theta, \text{for } i = n, r.$$

Note that the expected value depends on the productivity draws from the distribution H_n in the case of no restructuring, and on the draws from H_r in the case of restructuring. Under the assumptions made so far, a unique function $V(\theta)$ as defined in (1) exists, and it is also strictly increasing in θ.[15]

Consider now the exit decision. A firm exits when

$$\max\{E_r[V(\theta')|\theta] - c_r, E_n[V(\theta')|\theta]\} \leq 0. \tag{2}$$

When a positive mass of firms exit, (2) holds with equality. Because the left hand side of (2) is strictly increasing in θ, the exit threshold, x, is unique, and all firms with $\theta \leq x$ exit.[16]

Next, turn to the entry decision. Free entry implies that the expected value of a potential entrant satisfies

$$\int_0^1 V(\theta)h_e(\theta|\phi)d\theta \leq c_e. \tag{3}$$

The expected value of entry is strictly increasing in ϕ by the properties of H_e and V. If there is positive entry, (3) holds with equality. In the case of positive entry, the marginal entrant's prior, ϕ_e, satisfies (3) with equality and all potential entrants with $\phi \geq \phi_e > 0$ enter. The mass of entering firms is then given by $(1 - G(\phi_e))N$.

Finally, consider the restructuring decision. An incumbent firm invests in restructuring if

$$E_r[V(\theta')|\theta] - c_r \geq \max\{0, E_n[V(\theta')|\theta]\}. \tag{4}$$

That is, the net benefit from restructuring exceeds the benefit the firm can obtain by exiting or not restructuring. To understand the nature of restructuring, note that the gross benefit from restructuring versus no restructuring,

$$B(\theta) = E_r[V(\theta')|\theta] - E_n[V(\theta')|\theta], \tag{5}$$

can in general be a non-monotonic function of θ, even though the two components of $B(\theta)$ are both monotonic in θ.[17] There could therefore be multiple restructuring thresholds. To impose some structure, consider the case where B is strictly decreasing. This case would apply, for instance, if H_r decreases in θ, but at a rate lower than H_n does.[18] Under this case, the gross benefit from restructuring diminishes as productivity increases, i.e. $B' < 0$. This assumption is maintained for the rest of the model.[19]

There is a positive mass of firms restructuring if (4) holds for some $\theta_r > 0$. The marginal

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[15] These results follow from the dynamic programming arguments in Stokey and Lucas (1989).

[16] The fact that the left hand side of (2) is strictly increasing follows because both $E_r[V(\theta')|\theta]$ and $E_n[V(\theta')|\theta]$ are strictly increasing in θ, by the properties of V, H_r and H_n.

[17] Note that $E_i[V(\theta')|\theta]$ is strictly increasing in θ for $i \in \{r, n\}$ by the properties of V, H_r and H_n.

[18] This type of relationship between the two distributions would hold, if, for instance, restructuring requires learning about new technologies and such learning opportunities dwindle sufficiently fast as the firm moves further up in the productivity distribution. See, e.g. Jovanovic and MacDonald (1994) for similar discussion on how different outcomes may emerge in a model of innovation and imitation depending on the exact assumptions made on the processes for innovation and imitation.

[19] The other case, $B' \geq 0$, can also be analyzed. This case implies that more productive firms stand to gain more from restructuring. However, this case does not necessarily provide substantially different insight to the analysis of productivity dispersion.

firm type, θ_r, which engages in restructuring satisfies (4) with equality. As long as $c_r < E_r[V(\theta')|x]$, the marginal firm surviving is willing to restructure and (4) holds with equality for some $\theta_r > x$.[20] All non-exiting firms with $x < \theta \leq \theta_r$ then choose to restructure.

2.2 Equilibrium and Comparative Statics

One can define a stationary equilibrium for the model with positive entry, exit, and restructuring as follows.

Definition 1. *Given the fundamentals* $\{N, H_e, H_r, H_n, G, c_f, c_e, c_r, w, r\}$*, a stationary equilibrium with positive entry, restructuring, and exit is composed of an entry threshold* $\phi_e^* > 0$*, an exit threshold* $x^* > 0$*, a restructuring threshold* $\theta_r^* > x^*$*, a measure of firms,* μ^**, and a price* p^* *such that*

1. *Incumbent firms solve their dynamic problem to obtain the value defined by (1),*
2. ϕ_e^* *satisfies the free entry condition (3) with equality,*
3. θ_r^* *satisfies the restructuring condition (4) with equality,*
4. x^* *satisfies the exit condition (2) with equality,*
5. μ^* *satisfies, for all* $\theta \in [0,1]$*,*

$$\mu^*(\theta) = N\int_0^\theta \left(\int_{\phi_e^*}^1 h_e(z|\phi)g(\phi)d\phi\right)dz + \int_0^\theta \left(\int_{x^*}^{\theta_r^*} h_r(z|y)\mu^*(dy)\right)dz + \int_0^\theta \left(\int_{\theta_r^*}^1 h_n(z|y)\mu^*(dy)\right)dz. \tag{6}$$

6. p^* *clears the goods market*

$$D(p^*) = \int_0^1 \tilde{q}^*(\theta)\mu^*(d\theta). \tag{7}$$

If the entry cost and restructuring cost are not too high, there exists an equilibrium with positive entry, restructuring, and exit. Such an equilibrium is also unique given the assumptions so far. See Appendix A for a proof of existence and uniqueness.[21]

Consider now how the key thresholds, ϕ_e^*, x^*, and θ_r^*, change as the three parameters of interest, c_e, c_f, and c_r shift. Three types of change are considered. The first is a decline in the entry cost. Such a change can represent a removal of certain entry barriers and a relaxation of constraints on business starts. In the context of a country undergoing a transition to a market economy, a lower c_e may mean a general reduction in red tape and entry barriers. The second change is an increase in the fixed cost, c_f, brought about by, for instance, higher costs of compliance with regulations or an increase in the hardness of budget constraints. This increase may correspond to more stringent requirements for operating a business, more oversight by regulators, and fewer subsidies. The third change is a decline in the cost of restructuring, c_r. This decline may correspond to lower costs of adopting advanced technology, better business practices, lower financing costs, or in general, to reduced barriers to business expansion.

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[20] Because $E_r[V(\theta')|\theta]$ is strictly increasing in θ, $c_r < E_r[V(\theta')|x]$ holds if, for instance, $c_r < E_r[V(\theta')|0]$ – that is, the least productive firm type is willing to restructure.

[21] Note that, when evaluated at $\theta = 1$, equation (6) can be solved for the mass of firms in the industry, $M^* = \frac{(1-G(\phi_e^*))N}{H^*(x^*)}$, where $H^*(\theta)$ is the c.d.f. of productivity in equilibrium.

What are the effects of a decline in the entry cost c_e? A decline in the entry cost means the ex-ante expected profit required for a potential entrant to enter must now be lower for (3) to hold. Therefore, entrants with lower priors are able to enter, implying a lower entry threshold, ϕ_e^*, and hence, more entry. When input prices are fixed as assumed here, a lower entry cost (and hence higher entry) leads to lower price, which reduces the value of all firm types. Therefore, the exit threshold x^* also increases, as in Hopenhayn (1992). The higher exit threshold and lower price imply that the expected gross benefit from restructuring and no restructuring both go down. If the benefits from restructuring decline more than the benefits from no restructuring do, $B(\theta)$ in (5) will now be lower.[22] The restructuring threshold θ_r^* then decreases. In other words, the range of productivity values over which firms are willing to restructure shrinks.

Consider next the effect of an increase in the fixed cost, c_f. Starting from an equilibrium, such an increase, holding all else fixed, implies that the marginal firm type, x^*, obtains a negative value if it stays in the industry. The exit threshold must then increase to restore (2) to equality. At the same time, the expected value of potential entrants also declines, as each entrant faces a higher fixed cost and a higher likelihood of exit. The entry threshold, ϕ_e^*, increases as a result.[23] Again, if the net benefit from restructuring, $B(\theta)$, increases in response, so does the restructuring threshold, θ_r^*.

Finally, consider the effects of a lower restructuring cost, c_r. Starting at an equilibrium, a decline in c_r, all else fixed, allows more firms to restructure, as firms now need a lower expected gross benefit from restructuring to choose the option of restructuring. Thus, the restructuring threshold, θ_r^*, has to increase. As restructuring becomes easier, there is a value effect: the expected value from entry increases for each entrant type and the value of all firm types goes up, holding the output price fixed. This effect can lead to a decrease in the entry threshold, ϕ_e^*, and the exit threshold, x^*. However, because the incentives to restructure are now higher, and restructuring firms achieve a higher output in an f.o.s.d. sense than they would if they did not restructure, the total output of incumbents increases.[24] If price declines sufficiently in response, the exit threshold, x^*, and the entry threshold, ϕ_e^*, can both increase. The net effect then depends on the relative magnitudes of this price effect and the value effect. What can be said is that the exit and entry thresholds move in the same direction when c_r changes – see Proposition 1 in Appendix A.

2.3 Productivity dispersion

Now consider the main object of interest, the productivity dispersion. Note that the law of motion for the equilibrium measure, μ^*, of TFPQ is given by (6). Given any measure μ on [0,1],

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[22] Note that $E_i[V(\theta')|\theta]$ decreases as p declines for $i \in \{r, n\}$ by the fact that $V(\theta')$ is strictly increasing in p. Given the assumption that $H_r(\theta'|\theta) < H_n(\theta'|\theta)$, how much $B(\theta)$ changes as p declines depends on the rate of decline in $V(\theta')$ across different values of θ'. If a decline in price implies a higher reduction in value for more productive firms as θ' increases, then $B(\theta)$ declines.

[23] This result follows from the fact that profits of all firms types move in the same direction, by the assumed separability of the profit function, as in Hopenhayn (1992).

[24] The output of a restructuring firm is larger, on average, than the firm's initial output because $\tilde{q}(\theta)$ is strictly increasing in θ and $E_r[\theta'|\theta] > \theta$.

one can define the following linear operators

$$(\mathcal{L}_e^*\mu)(\theta) = \int_{\phi_e^*}^1 h_e\,(\theta|z)\mu(dz), (\mathcal{L}_r^*\mu)(\theta) = \int_{x^*}^{\theta_r^*} h_r\,(\theta|z)\mu(dz), (\mathcal{L}_n^*\mu)(\theta) = \int_{\theta_r^*}^1 h_n\,(\theta|z)\mu(dz) \tag{8}$$

Using these operators, (6) can be written as

$$\mu^*(\theta) = N(\textstyle\sum_{k=0}^{\infty}(\mathcal{L}_r^* + \mathcal{L}_n^*)^k)(\mathcal{L}_e^* g)(\theta), \tag{9}$$

which expresses $\mu^*(\theta)$ in terms of the exogenously given densities, g, h_e, h_r and h_n, and the exogenous mass of potential entrants, N.[25] The variance of productivity across firms is the one associated with the measure in (9) and is denoted by $\sigma_{\text{TFPQ}}^{*2}$. Note also that the variance of TFPR is related to the variance of TFPQ as

$$\sigma_{\text{TFPR}}^{*2} = p^2\sigma_{\text{TFPQ}}^{*2}. \tag{10}$$

The change $\sigma_{\text{TFPQ}}^{*2}$ induced by a change in c_e, c_f or c_r, depends on the nature of the operators $\mathcal{L}_e^*$, $\mathcal{L}_r^*$, and $\mathcal{L}_n^*$. As shown in (8), the operator $\mathcal{L}_e^*$ truncates the density g, and then maps it to a measure of productivity for entrants, $\mathcal{L}_e^* g$, through the density, h_e. The operators $\mathcal{L}_r^*$ and $\mathcal{L}_n^*$ then map $\mathcal{L}_e^* g$ to a new measure, through the densities h_r and h_n.

The equilibrium density of productivity, h^*, associated with μ^* can be expressed as a mixture of the density of productivity for new entrants, and the densities of productivity for restructuring and non-restructuring incumbents. The density of TFPQ for new entrants can be written as

$$h_e^*(\theta) = \frac{1}{1 - G(\phi_e^*)}\int_{\phi_e^*}^1 h_e\,(\theta|z)g(z)dz$$

Similarly, the density for restructuring incumbents is

$$h_r^*(\theta) = \frac{1}{H^*(\theta_r^*) - H(x^*)}\int_{x^*}^{\theta_r^*} h_r\,(\theta|z)h^*(z)dz.$$

Finally, the density for non-restructuring incumbents is

$$h_n^*(\theta) = \frac{1}{1 - H^*(\theta_r^*)}\int_{\theta_r^*}^1 h_n\,(\theta|z)h^*(z)dz.$$

The density of productivity is then a mixture of the three densities defined above

$$h^*(\theta) = \alpha_e^* h_e^*(\theta) + \alpha_r^* h_r^*(\theta) + \alpha_n^* h_n^*(\theta), \tag{11}$$

where α_i^* $(i \in \{e, r, n\})$ is the fraction of firms that are new entrants, restructuring incumbents, and non-restructuring incumbents, respectively, given by

$$\alpha_e^* = H^*(x^*), \alpha_r^* = H^*(\theta_r^*) - H(x^*), \alpha_n^* = 1 - \alpha_e^* - \alpha_r^*.$$

The variance of productivity can then be written as

$$\sigma_{\text{TFPQ}}^{*2} = \textstyle\sum_{i\in\{e,r,n\}} \alpha_i^*\left[\sigma_i^{*2} + (\mu_i^* - \mu_{\text{TFPQ}}^*)^2\right], \tag{12}$$

where $\mu_{\text{TFPQ}}^* = \sum^{i\in\{e,r,n\}} \alpha_i^*\,\mu_i^*$ is the average productivity.[26] Equation (12) makes it clear that the

[25] Note that $\sum_{k=0}^{\infty}(\mathcal{L}_r^* + \mathcal{L}_n^*)^k = (I - \mathcal{L}_r^* - \mathcal{L}_n^*)^{-1}$, where I is the identity operator. The notation $(\mathcal{L}_r^* + \mathcal{L}_n^*)^k$ is equivalent to the repeated application of the operator $\mathcal{L}_r^* + \mathcal{L}_n^*$ for k times. The existence of an invariant measure μ^* hinges on the existence of the inverse operator $(I - \mathcal{L}_r^* - \mathcal{L}_n^*)^{-1}$.

[26] The expression in (12) is a straightforward application of the identity $Var(X) = E[Var(X|Y)] + Var(E[X|Y])$.

change in $\sigma^{*2}_{\text{TFPQ}}$ in response to a change in c_e, c_f, or c_r depends on how $\alpha_i^*, \sigma_i^{*2}$, and μ_i^* change. Because (12) is the variance of a mixture, it incorporates not only the variances, σ_i^{*2}, but also the means, μ_i^*. Thus, the effects of entry and restructuring processes not just on the second moment of productivity, but also on the first, matter for the overall variance. For example, even in the case where restructuring only raises the average productivity of an incumbent without altering its variance, the overall productivity dispersion can change.

Now, let $\mu_e(\phi)$ and $\sigma_e^2(\phi)$ be the mean and variance of productivity for an entrant with prior ϕ. In other words, these are the mean and variance associated with the distribution $H_e(\theta|\phi)$. Analogously, define $\mu_r(\theta)$ and $\sigma_r^2(\theta)$, and $\mu_n(\theta)$ and $\sigma_n^2(\theta)$, as the mean and variance of productivity for a restructuring and a non-restructuring firm with initial productivity θ, respectively. Again, these moments are associated with the distributions $H_r(\theta'|\theta)$ and $H_n(\theta'|\theta)$, respectively. The variances σ_i^{*2} $(i \in \{e, r, n\})$ can then be written as

$$\sigma_e^{*2} = \frac{1}{1-G(\phi_e^*)}\left[\int_{\phi_e^*}^{1}\{\sigma_e^2(\phi) + (\mu_e(\phi) - \mu_e^*)^2\}g(\phi)d\phi\right], \tag{13}$$

$$\sigma_r^{*2} = \frac{1}{H^*(\theta_r^*)-H(x^*)}\left[\int_{x^*}^{\theta_r^*}\{\sigma_r^2(\theta) + (\mu_r(\theta) - \mu_r^*)^2\}h^*(\theta)d\theta\right], \tag{14}$$

$$\sigma_n^{*2} = \frac{1}{1-H^*(\theta_r^*)}\left[\int_{\theta_r^*}^{1}\{\sigma_n^2(\theta) + (\mu_n(\theta) - \mu_n^*)^2\}h^*(\theta)d\theta\right]. \tag{15}$$

To impose some more structure, suppose now that each variance $\sigma_i^2(\cdot)$ is strictly decreasing in its argument. Also, the earlier assumptions of the model imply that the mean $\mu_i(\cdot)$ is strictly increasing in its argument. In other words, restructuring or non-restructuring incumbents with higher productivity achieve a higher productivity on average, and face a lower dispersion of productivity, compared with incumbents that have lower productivity. Similarly, potential entrants with higher priors obtain, on average, a higher productivity and a lower dispersion in productivity, compared with those with lower priors. These features can emerge, for example, in an environment where more productive firms engage in innovative activities that are less risky and that yield better outcomes on average. Clearly, other scenarios are possible. For instance, restructuring incumbents with higher productivity may face riskier outcomes, or entrants with higher priors may have larger dispersion in their initial productivity. We proceed with the understanding that alternative assumptions on the nature of entry and restructuring processes can alter the exact nature of the analysis to follow, but do not make a material difference for the main purpose of demonstrating the potentially ambiguous effects of the entry and restructuring processes on the distribution of productivity.

Now, consider the effect of a decline in the entry cost c_e on $\sigma^{*2}_{\text{TFPQ}}$. In response to this decline, ϕ_e^* decreases, while x^* increases, as discussed above. Assume also that θ_r^* decreases. A decline in ϕ_e^* implies that the variance of initial productivity for the marginal entrant type, $\sigma_e^2(\phi_e^*)$, increases, leading to an increase in the overall variance of productivity for entrants, σ_e^{*2}. To see this, note that the differentiation of (13) yields

$$\frac{d\sigma_e^{*2}}{d\phi_e^*} = \frac{g(\phi_e^*)}{1-G(\phi_e^*)}[\sigma_e^{*2} - \sigma_e^2(\phi_e^*) - (\mu_e(\phi_e^*) - \mu_e^*)^2] < 0,$$

because $\sigma_e^{*2} < \sigma_e^2(\phi_e^*)$ as a result of the assumption that $\sigma_e^2(\phi)$ is strictly decreasing.

Similarly, a lower θ_r^* implies that there is a wider range of firm types that choose not to restructure. Given that $\sigma_r^2(\theta)$ is decreasing, this effect alone can lead to a rise in the variance of productivity for non-restructuring incumbents, as in the case of entry. However, for restructuring incumbents while a decline in θ_r^* can increase the variance of productivity, an increase in x^* can counteract this effect. Overall, if an increase in x^* does not reduce the variance of the productivity for non-restructuring firms substantially, the variance $\sigma_{\mathrm{TFPQ}}^{*2}$ can increase in response to a decline in c_e.[27] In general, if there is a large increase in the heterogeneity of firms induced by entry and restructuring processes, the overall productivity dispersion increases.

The average productivity of firms can also increase or decrease in response to a decline in the entry cost, and hence the threshold, ϕ_e^*. For instance, average productivity rises if the exit threshold x^* increases sufficiently and the decline in θ_r^* does not lead to a large fall in the mass of restructuring firms. In addition, aggregate productivity can also increase or decrease.[28] Overall, then, it is possible to observe a larger variance of productivity, along with a higher average or aggregate productivity, as a result of a decline in entry barriers, represented by c_e. The main message is that a reduction in entry barriers need not result in lower dispersion *and* higher productivity at the same time.

Consider next the effect of an increase in the fixed cost, c_f, on productivity dispersion. In response to such an increase, the exit and entry thresholds, x^* and ϕ_e^*, increase, as discussed in the previous section. In addition, suppose that θ_r^* declines. The variance of productivity for new entrants then declines, but that of non-restructuring incumbents can increase or decrease. A lower θ_r^* also works to increase the variance of restructuring firms, because $\sigma_r^2(\theta)$ is decreasing in its argument. However, the increase in x^* can counter this effect. The overall effect depends on the relative magnitudes of these effects. Both the average and the aggregate productivity can also change in either direction. Similar arguments imply that a decline in the restructuring cost, c_r, can result in either an increase or decrease in the variance of productivity, $\sigma_{\mathrm{TFPQ}}^{*2}$, along with an increase in average or aggregate productivity.

Overall, the analysis suggests that institutional changes such as reductions in entry barriers, increases in regulatory costs, or declines in the costs of business investment or restructuring can result in a variety of outcomes for average productivity and productivity dispersion. We can distinguish separate selection and experimentation mechanisms. Reduced entry costs, for instance, raise the productivity threshold for survival, so that selection is tougher and, ceteris paribus, dispersion is lower. But they also lead to more entry, a form of experimentation that raises productivity dispersion. Changes in these frictions can therefore lead to *both* higher average or aggregate productivity and higher dispersion of productivity at the same time. The main message

[27] The nature of these various effects depend on the productivity distributions involved. In some cases, a definitive statement can be made about the direction of change. For instance, if a productivity distribution is log-concave, an increase in the truncation point on the left (the exit threshold, x^*) leads to a lower variance – see Proposition 1 in Heckman and Honore (1990), and Theorem 9 in Bagnoli and Bergstrom (2005).
[28] See Appendix B for a derivation of aggregate productivity for the special case of Cobb-Douglas production functions. The appendix also highlights how the aggregate productivity depends on the costs $c_e, c_r,$ and c_f.

of the theoretical analysis is that a negative relationship between average productivity and the dispersion of productivity does not necessarily emerge. The correlation between the two can go either way, depending on the relative magnitudes of the forces that determine entry, exit, and restructuring.

3. Empirical Analysis

Our empirical analysis focuses on cases of large-scale deregulation and liberalizing reforms: the telecommunications equipment manufacturing sector in the U.S., and the whole manufacturing sectors in six transition economies. As is true for most cases of significant reforms, the policy changes in each of these cases involved reductions in frictions as well as in idiosyncratic taxes (static distortions), so the effects of all these simultaneous changes cannot be distinguished. Presumably the reforms did serve to reduce misallocation, however, and our interest is in assessing how productivity dispersion changed, and thus whether dispersion is an indicator of misallocation. We consider not only second-order moments, but also compute full distributions to assess the impact of reforms on the tails of the distribution. Our model implies that reforms that reduce entry costs will strengthen the selection mechanism, in the sense of raising the productivity threshold for survival. Greater market pressures that make survival more difficult will tend to truncate the left tail of the productivity distribution. The model also implies that lowering entry costs will increase experimentation which would tend to fatten both tails. Selection and experimentation have opposing effects on the left tail, while only experimentation affects the right tail, so changes in the latter are especially interesting consequences of reforms. We also assess the contributions of three types of firms – entrants, continuers, and exiters – to the changes in productivity dispersion by constructing counter-factual distributions that exclude each type of firms, in turn. Finally, exploiting variation across transition countries and over time in the pace and extent of liberalization, we estimate some simple relationships among reforms, productivity growth, and productivity dispersion.

3.1. Data and Measurement

The paper uses annual census-type data for manufacturing firms in each of the seven countries. Though the data sources and variables are similar, we have taken steps to make them sufficiently comparable to justify cross-country comparisons.

The U.S. data come from the establishment-level Censuses of Manufactures (CM) in 1963, 1967, 1972, 1977, 1982, 1987, 1992, 1997, 2002, and 2007. We use the universe of establishments mailed the Census survey. Very small single-establishment firms (typically fewer than five employees) are excluded from the mail universe, and we omit them here since their output and capital stock are often imputed.

The basic sources for the Hungarian and Romanian data are balance sheets and income statements associated with tax reporting: to the National Tax Authority in Hungary and the Ministry of Finance in Romania. The Romanian data are supplemented by the National Institute for Statistics' enterprise registry. For both countries, all legal entities engaged in double-sided bookkeeping are supposed to report. The Hungarian data are annual from 1986 to 2003, and the Romanian data from 1992 to 2006. The sum of employment across all firms in the database is similar to the statistical yearbook number in both countries.

The other four transition countries are former Soviet Republics. Their data come from their national statistical offices, the descendants of the former State Statistical Committee (*Goskomstat*), and therefore tend to be quite similar to one another. The Georgian and Lithuanian data cover most firms outside the budgetary and financial sectors in 1995-2005 (Lithuania) or 2000-2004 (Georgia). The Georgian and Lithuanian databases include roughly three-fourths of total manufacturing employment reported in the yearbooks.

The main sources in Russia and Ukraine are industrial enterprise registries from their national statistical offices, supplemented by balance sheet data.[29] The data span 1985-2005 for Russia, and 1989 and 1992-2006 for Ukraine. Prior to 1991, the registries include all firms in the industrial sector, but afterward the Russian registry coverage was revised to include all industrial firms with over 100 employees as well as those that are more than 25 percent owned by the state and/or legal entities that are themselves included in the registry. In practice, it appears that once firms enter the registries, they continue to report even if these conditions no longer hold. The Russian data can therefore be taken as corresponding primarily to the "old" firm sector (and their successors) inherited from the Soviet period. The 1992-1996 Ukrainian registries contain all industrial firms producing at least one unit of output, where a unit is defined differently depending on the product. All legal entities outside the budgetary and financial sectors are included in the 1997-2006 registries. The Ukrainian coverage is fairly complete: the sum of employment across firms in the database is very similar to the corresponding yearbook figure each year. The Russian data cover nearly all activity through 1994; then the coverage declines to about 75 percent in more recent years as the de novo sector has grown.

Some truncation is necessary to make the samples comparable across countries. The data in all countries are limited to manufacturing (NACE 15-36). We exclude the tobacco industry (NACE 16) due to insufficient observations in four of the seven countries and the recycling industry (NACE 37) because of non-comparability with the classification system used until recently in Russia and Ukraine. Following the literature on productivity growth decompositions, we analyze productivity within industries, avoiding problems of comparisons across industries with very different technologies. Ideally one would prefer to use industries disaggregated to the level of product markets, so as to compare firms only to their competitors. On the other hand, since the productivity analyses rely on deviations from the industry average, it is important to have sufficient numbers of firms in each sector to ensure reliable estimates. We have compromised by dividing manufacturing into 19 sectors, which are 2-digit NACE industries (except that 23 and 24 are combined, as are 30 and 32).

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[29] The units of observation in these data are firms, except for multi-plant entities where individual plants are listed as "subsidiaries" (*dochernye predpriyatiya* or "daughter companies") in the Russian registries. Apparently most but not all cases of multiple plants are treated individually in Russia: the 1993 registry contains a variable indicating the number of plants, which equals 1 in 99.91 percent of the 18,121 non-missing cases. To avoid double-counting, we have dropped the consolidated records of entities with subsidiaries from the analysis.

In Russia and Ukraine we exclude firms in regions that are completely missing in the data in one of the two adjacent years, and those in industries with implausibly high entry or exit rates in that year (suggesting a change in sample coverage).[30] Entry and exit associated with firms that were members of Soviet-era production associations or that belong to multi-establishment firms are also excluded in Russia.[31]

Sample sizes are shown in Table 1. We use several variants of the U.S. manufacturing sectors as a comparison, or benchmark, for specific analyses. The large sample sizes reflect the population coverage of these databases.

Variables are defined as follows: Employment in the U.S. data is total employment in the payroll period including March 12; in the transition economies, it is the average annual number of all registered employees, except in Russia, where it excludes personnel working in non-industrial divisions. Output or sales refers to sales in Georgia, Hungary, Lithuania, Romania, and post-2003 Ukraine, and to value of production in Russia, pre-2004 Ukraine, and the U.S. (for the U.S. this is calculated as sales + ending inventories of finished goods – beginning inventories of finished goods). Capital stock is the book value of fixed assets.[32] Output or sales and capital stock are expressed in constant final-year prices (thousands of 2004 GEL for Georgia, millions of 2005 HUF for Hungary, thousands of 2005 LTL for Lithuania, millions of 2006 ROL for Romania, millions of 2004 RUB for Russia, and millions of 2006 UAH for Ukraine), except in the U.S., where they are in thousands of 1987 USD (using output deflators from the National Bureau of Economic Research and book value of capital stock deflators from the Bureau of Economic Analysis).

For the U.S. telecommunications sector and the comparison to all U.S. manufacturing, we compute multifactor productivity (MFP) as follows: $lnMFP_{et} = lnQ_{et} - \alpha_K lnK_{et} - \alpha_L lnL_{et} - \alpha_M lnM_{et}$, where Q_{et} is real gross output, K_{et} is real capital (separate terms are included for structures and equipment), L_{et} is labor input (total hours for production workers plus an imputed value for nonproduction workers' total hours),[33] and M_{et} is real materials (separate terms are included for energy and other materials). We use industry cost shares to measure factor elasticities. The cost shares come from a combination of industry-level data from the NBER Productivity Database and Bureau of Labor Statistics (BLS) capital rental prices. For the manufacturing censuses prior to 1972, when the capital variables are unavailable, we compute labor productivity as value added per worker. Labor productivity (LP) is real output (adjusted for changes in final and unfinished good inventories), minus real material costs (cost of materials and parts, cost of resales, and cost

[30] The size-related exclusions amount to no more than 0.3 percent of the sample in any country. The changes in industry and regional coverage result in the exclusion of about 2 percent of observations in Russia and Ukraine.
[31] The reason for excluding production association entry and exit during the Soviet period and multi-establishment firm entry and exit during the transition period is that many of these firms report inconsistently in the data. In one year a consolidated entity may appear, in the next each of the establishments may report separately, or vice versa. These exclusion rules result in a conservative bias. Of course some production associations may be starting new establishments or closing others down, and there may be some true entry and exit in industries with implausibly high rates and in regions that enter and exit the dataset.
[32] For the U.S. telecommunications sector and its comparison to all U.S. manufacturing, we use capital stock calculated by the perpetual inventory method.
[33] The imputation uses the ratio of total payroll to production worker payroll multiplied by production worker hours.

of contract work), divided by total hours worked, using the same imputation mentioned above for MFP labor input.

For the comparative analysis of the U.S. and East European manufacturing sectors, we compute MFP as the residual from a two-digit-industry-specific and country-specific Cobb-Douglas production function of gross output (or sales) in capital and labor, controlling for year effects. Material costs are unavailable for Russia and in the early years for Ukraine, so this approach is necessary to ensure cross-country comparability, but the results are very similar for years and countries where material costs are taken into account. Moreover, our use of industry-specific production functions implies that the results are identical under both approaches as long as the output-materials ratio is common within two-digit industries, controlling for capital and labor.[34]

While these MFP measures are within country-industry-years, they do not distinguish firm-level quantity and price variation, which are unavailable as in most data sets, and thus they conflate technical efficiency and firm-specific price variation, thus representing revenue productivity.[35] Our productivity dispersion measures include its standard deviation and the 90-10 percentile range, unweighted.[36]

3.2. Deregulation in U.S. Telecommunications Equipment Manufacturing

Prior to deregulation of the telecommunications equipment sector, AT&T was a monopoly provider of telecommunications services, and it extended the monopoly to the equipment manufacturing industry via its requirement that any equipment attached to the Bell system network had to be supplied by AT&T. A series of antitrust decisions and Federal Communications Commission (FCC) policy changes in the late 1960's and 1970's loosened entry into the equipment sector. After being divested by AT&T in January 1984, the regional operating companies became free to purchase equipment from any supplier, while being prohibited from manufacturing equipment themselves. Arguably, this led to a reduction in entry costs and in the implicit tax faced by equipment manufacturers other than AT&T.

Olley and Pakes (1996) provide further details on the deregulation process, and they study the sector's productivity dynamics from 1972 to 1987, finding evidence of major reallocation via

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[34] Foster et al. (2016) consider the implications of two estimation approaches for TFPR and show that a factor share measure corresponds to true TFPR only under CRS (and therefore reflects distortions, under the rest of the Hseih and Klenow (2009) assumptions), while the regression residuals reflect idiosyncratic demand shocks and TFPQ dispersion as well as distortions (again, under the same assumptions). Nonetheless, they find the two measures are highly correlated, with similar magnitudes of dispersion. Our work does not address these measurement issues, although we use both of these measurement approaches (with similar results), but instead we focus on productivity dispersion in a dynamic setting with adjustment frictions.

[35] See Eslava, Haltiwanger, Kugler, and Kugler (2004) and Foster, Haltiwanger, and Syverson (2008) for analyses of firm-specific revenue and physical productivity.

[36] The unweighted calculation follows the procedures of Hsieh and Klenow (2009), but we find similar results if we calculate dispersion measures separately by industry and then weight the industries by their shares in either output or number of firms. Bartelsman and Wolf (2016) emphasize that some productivity measurement approaches are more robust to measurement error and suggest inter-quantile differences to avoid the influence of outliers. We present both the standard deviation and inter-decile ranges for robustness and find little qualitative difference in the results.

entry and incumbent plant size changes, which they report to be productivity enhancing. They do not measure productivity dispersion, however, and their analysis ends in 1987, not long after the January 1984 break-up of AT&T. We extend the data, calculating total factor productivity (MFP) through 1997 and labor productivity (LP) from 1963 to 1997.[37] Extending the data through 1997 allows an assessment of the longer term consequences of deregulation, while extending the data backward to 1963 is especially valuable as the analysis then includes observations prior to deregulation (i.e., for 1963 and 1967) not exploited in Olley and Pakes' (1996) analysis.

Figures 1a and b contains results for the evolution of MFP dispersion among firms in the U.S. telecommunications equipment manufacturing sector from 1972 to 1997. The dispersion measure is alternately the standard deviation and the 90-10 percentile range. The comparable dispersion measures for the U.S. manufacturing sector as a whole are also provided as a baseline. While measured dispersion in U.S. manufacturing overall is declining slightly over time, it increases throughout the deregulation period for telecoms equipment. The telecoms equipment SD(MFP) rises from 0.27 to 0.45, and the 90-10 range increases from 0.68 to 0.93 between 1972 and 1997.

Figures 1c and 1d contain similar measures for LP for the longer time period of 1963 to 1997. Although the LP measures are slightly more volatile compared to MFP, they show a similar upward trend from the pre-deregulation period of the 1960s through the last consistently available observation of 1997. The upward trend is evident for both the SD and 90-10 measures and in both absolute terms and relative to the average for all manufacturing industries. These results are inconsistent with the hypothesis that deregulation reduces productivity dispersion. While, following Olley and Pakes (1996), the reform reduced misallocation, productivity dispersion actually rose during this time period.

To shed more light on the nature of the rise in dispersion, we examine kernel densities of productivity in early and late years. Figure 2a shows the U.S. telecommunications equipment sector MFP distribution in 1972 (the early deregulation period), 1982 (just before the break-up of AT&T) and 1997 (post-deregulation). Figure 2b shows the same years plus 1963 for the LP distribution. Both figures show widening of the distribution over time, but the right tail fattens more.[38] This may be a sign that deregulation facilitated experimentation.

Exiting establishments in the U.S. telecommunications equipment sector in 1972-1997 tend to be less productive than the average in the sector, by 0.195 log points; the difference is 0.072 log

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[37] We do not extend the data past 1997, because the telecom equipment sector's industry classification changed significantly during the conversion from the SIC to NAICS classifications. Olley and Pakes (1996) include not only SIC sector 3661 (telephone and telegraph apparatus), but also selected establishments from the 5-digit product class 36631, including fiber optics communication equipment, microwave communication equipment, facsimile communication equipment, and carrier line equipment not elsewhere classified, while excluding military space satellites, amateur radio communications equipment, and other products. We do not have access to the product data used by Olley and Pakes to distinguish between establishments in 36631 that are relevant for telecommunications and those that are not. We limit the analysis to SIC sector 3661 to be sure that all the establishments are affected by the deregulation.

[38] For LP, the change depends on which moment of the distribution is chosen to represent dispersion: in 1997, the 25-75 percentile ratio is smaller than in 1982, but the tails (especially right tail) are fatter, and both SD and 90-10 are larger in 1997 compared to 1982.

points for exiting establishments in U.S. manufacturing as a whole during that period. Exiting establishments have 0.057 log points lower labor productivity than average both in the 1963-1967 exit cohorts and the 1972-1992 cohorts; the analogous numbers for U.S. manufacturing as a whole are 0.069 log points lower in 1963-1967 and 0.087 log points lower in 1972-1992.

To analyze the effects of entry, exit, and continuers on productivity dispersion change over a five-year period, we produce counterfactual productivity distributions focusing on each effect separately. One distribution includes all establishments in year t except those that have entered since the previous census in year $t-5$. We subtract the SD (90-10 range) of this distribution from the SD (90-10 range) of the actual distribution in year t to get an estimate of the entry effect on productivity dispersion. For the exit effect, we add establishments exiting between $t-5$ and t to the productivity distribution in year t, using exiting establishments' productivity in $t-5$, and we subtract the standard deviation (90-10 range) of this distribution from that of the actual distribution in year t. To estimate the continuer effect, we replace year t productivity of establishments present in both $t-5$ and t with their productivity in $t-5$ and subtract the SD (90-10 range) of this distribution from that of the actual distribution in year t.[39]

Figures 3a and b show these calculations for the standard deviation and 90-10 percentile range of MFP, respectively. The results imply that establishment turnover (both entry and exit) works to raise productivity dispersion in the early deregulation period in the U.S. telecom equipment sector. Post-deregulation, continuers push dispersion upwards, while entry dampens it. Turnover lowers dispersion in the earlier years in U.S. manufacturing as a whole (Figures 3c and d), while continuers keep dispersion from making further declines in the later period. The effects in the telecom equipment sector are generally larger in magnitude than those in manufacturing in the aggregate.

3.3. Reforms in Transition Economies

Our second illustration of how market liberalization affects productivity dispersion examines East European manufacturing firms that we follow from the 1980s, when they were governed by central planning, through the reforms of the 1990s, and into the late transition period up to 2005. We use U.S. manufacturing, constructing productivity and dispersion measures comparably, as a benchmark. Studying the East European transition is fruitful not only because of the drastic nature of the reforms and the long annual time series data available for all registered firms (in most countries), but also because of the large variation across countries and over time in the pace and depth of the reform process. To provide a measure of this variation, we draw upon data from the European Bank of Reconstruction and Development (EBRD), which has tracked annual market reform progress in Eastern European economies. The covered reforms include small- and large-scale privatization, governance and enterprise restructuring; liberalization of

[39] For the U.S. telecom equipment and all U.S. manufacturing analysis in Figure 3, we use the residual from a regression of MFP on year dummies as the MFP measure, so that productivity is relative to mean productivity in the particular year. Thus, the inclusion of t-5 productivity for exiting establishments or continuers in the year t productivity distribution abstracts from aggregate productivity shocks occurring between the two periods. All the MFP measures for the comparative analysis of Eastern Europe and the U.S. control for year effects.

prices, foreign exchange, interest rates, and trade; and reform of banking and infrastructure ranging from 1 (unreformed, centrally planned economy) to 4.3 (developed market economy). The composite (average) index is displayed in Figure 4 for the countries we analyze. The six countries' reform paths are quite heterogeneous. Hungary liberalized most quickly and maintained a lead throughout the period. Though there are some changes in rankings over time, by the end of the period the other two European Union accession countries have implemented the next most reform, while the Commonwealth of Independent States (CIS) countries (Georgia, Russia, and Ukraine) have done less.

The evolution of manufacturing sector productivity dispersion before and during market liberalization in Eastern Europe is shown in Figures 5 (standard deviation), Figure 5b (90-10 percentile range), Figure 6a (pre-liberalization MFP distribution), and Figure 6b (MFP distribution during market liberalization period), using the U.S. manufacturing five-year census numbers as a benchmark. In all three countries for which 1980s (pre-liberalization) data are available, Hungary, Russia, and Ukraine, we find that manufacturing productivity dispersion is very similar to that in the U.S., despite their very different economic systems. For Hungary and Ukraine, both the standard deviation and 90-10 percentile range are essentially identical to the U.S. in the 1980s, while in Soviet Russia, the measured dispersion is only marginally higher. Our interpretation of this finding is that while static distortions were rampant and selection processes worked poorly under central planning, as enterprises faced soft budget constraints, experimentation was also strongly discouraged.[40] The negative experimentation effect seems to be strong enough to offset the dispersion-raising forces, resulting in similar productivity dispersion in the U.S. and the Soviet Union – the most liberalized and the least liberalized economies.

Figures 5a and 5b also show that productivity dispersion in the transition countries rises sharply post-liberalization. Hungary liberalized faster than the other Eastern European countries, as indicated in Figure 4, and its dispersion rises much quicker than that of the other economies during the early liberalization period (the early 1990's).[41] There is some evidence of dispersion plateauing and in some cases declining, after different lengths of time and at different levels. The peak and decline occurs earlier in countries that liberalized faster - first in Hungary, followed by Romania. Both tails of the distribution fatten, as shown in Figures 6a and 6b.[42]

Snapshots of the relative productivity dispersion across countries at different points in time therefore show varying correlations between frictions and dispersion. Prior to liberalization, the

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[40] On soft budget constraints and incentives for innovation in the socialist system, see Kornai (1992, especially pp. 140 and 297).

[41] Price liberalization was largely accomplished very early in the transition process (for instance, in a "big bang" liberalization of almost all prices in Hungary on January 1, 1990), so while price reforms likely raised price and TFPR dispersion, they cannot account for the time pattern of later and continuing rises in dispersion. All these countries experienced very high inflation in the early 1990s (hyper-inflation in some cases), but productivity dispersion rose significantly later.

[42] Hsieh and Klenow (2009) consider the possibility that dispersion is driven by measurement error; against this hypothesis, they show that productivity differs systematically by state versus private ownership in China. Our transition economy data also show strong productivity differences associated with ownership, as documented in Brown, Earle, and Telegdy (2006, 2016) and, for Russia, Brown, Earle, and Gehlbach (2013).

distribution is similar to that in the U.S. Dispersion is positively associated with liberalization early in the reform process. At the end of the period, the association reverses, as the slower reformers partially catch up in the extent of reforms and their dispersion overtakes that of the faster reformers.[43]

Next, we investigate the contribution of entry and exit to dispersion separately. Productivity dispersion of age 1 firms is much greater in Eastern European economies during the market liberalization period than that of U.S. age 1 firms, as shown in Figure 7; the difference is especially noticeable in the right tail of the distribution, which is much fatter in the Eastern European economies.[44] One component of entry in these economies is foreign direct investment, which was non-existent under central planning but rose rapidly when entry was liberalized. As with overall dispersion, entrant dispersion rises more quickly in countries that liberalized faster.[45]

Figure 8 examines the correlation between exit and one-year lagged productivity, shedding light on the strength of the selection mechanism. Lithuania and Georgia actually have higher exit rates among low-productivity firms than does the U.S. The productivity-exit relationship in Russia is almost exactly the same as in the U.S., suggesting similar selection of firms for survival or exit based on their relative productivity.

The one country where exit is large enough in the 1980s to permit analysis of the productivity distribution is Hungary, and Figure 8 includes the productivity-exit relationship for three years of the late 1980s to compare with the transition years starting in 1993. Exit rates are quite low under central planning, and except for a bump in the second lowest productivity category, the slope of the relationship is essentially flat for most of the range. The exit rate rises substantially at the highest productivity intervals, implying negative selection (for survival) under central planning. However, after the transition reforms, the profile takes on a distinctly negative slope. To take another approach, we have also analyzed 3-year exit rates for 1986-1989 versus 2001-2004, finding again that the profile slope changes from flat to negative: under planning, exit rates from the bottom two-fifths and top two-fifths of the productivity distribution are 4.6 and 4.3 percent, respectively, and after reforms, the corresponding rates are 16.4 and 11.6 percent.

Figures 9a-9f show five-year entry, exit, and continuer effects analogous to those in Figure 3 for the transition economies with long time series (Hungary, Romania, Lithuania, Russia, and Ukraine), as well as for U.S. manufacturing as a baseline. We show only the calculations for the 90-10 MFP percentile range, as those for the standard deviation are very similar. The results suggest that experimentation by entrants is primarily responsible for the jump in productivity dispersion, though the degree to which it comes from incumbent firms and new entry varies considerably across countries. Entry is the dominant effect in Hungary and Romania, continuers

[43] The industrial compositions of these economies differ from each other and change over time, but we obtain qualitatively similar results when we fix the industrial structure.

[44] We measure productivity at age 1, because productivity in the first year is poorly measured due to partial-year operation for many entrants.

[45] Both entry rates and productivity dispersion among entrants rose during the transition, and there is some evidence that entry shifted from the right towards the left tail of the productivity distribution. In Hungary, for instance, the 3-year entry rate in 1986-1989 was 4.6 percent for the bottom two quintiles and 7.1 percent in the top two quintiles; these rates became 30.5 and 24.3 percent in the period 2001-2004.

are responsible for the increase in Russia, and both entry and continuers contribute roughly equally in Lithuania and Ukraine. The magnitudes dwarf those in U.S. manufacturing. Reductions in all three effects contribute to the leveling off and decline of productivity dispersion in Eastern European countries. The exit effect is sometimes positive in the early transition years, but it turns negative later in the transition everywhere, likely due to hardening budget constraints. The continuer effect also becomes negative in Hungary, Lithuania, and Ukraine, while the entry effect remains positive everywhere. The entry effect has very different patterns across countries, with a clear peak and decline in Hungary (and also in Ukraine six years later), two peaks and declines in Romania, and little trend in Lithuania and Russia.

Although the time plots suggest a positive relationship between market liberalization and productivity dispersion, particularly in the early years of transition, we may quantify the strength of the relationship with regressions of productivity dispersion in country-year cells on the EBRD reform index (as well as its square in some specifications), along with country and year fixed effects. The results shown in Table 2 suggest that reforms raise productivity dispersion but the relationship is concave with market-oriented reform. Using the late 1980s starting values of 0.6 for the standard deviation and 1.5 for the 90-10 range in the transition economies (which are the same as the U.S. throughout the period), the quadratic specifications imply that the dispersion measures more than double for the first one-point increase in the EBRD index. Thereafter, dispersion increases at a decreasing rate as the EBRD index rises. Only when the EBRD index passes 3, does the positive impact of market liberalization on productivity dispersion peak, after which it falls back somewhat for the final one-and one-third points of the index to reach the "market economy" standard. The concave shape, increasing up to EBRD=3, holds whether dispersion is measured as standard deviation or 90-10 percentile range in MFP.

The analysis so far has pooled firms across industries to calculate dispersion, but industries may differ systematically in frictions such as entry and restructuring costs, even in the absence of differences in the policy regimes that affect them. If we take the U.S. as a benchmark economy with dispersion mostly the result of such non-policy factors, then an interesting question is whether reforms lead the relative productivity dispersion patterns across industries in Eastern European economies to resemble those in the U.S. For this purpose, we estimate regressions of productivity dispersion in the East European industry-country-year on an interaction of the U.S. productivity dispersion value (calculated as the mean for the industry with the EBRD reform index (and its square in some specifications)). Controls include three sets of fixed effects: country-industry interactions, country-year interactions, and industry-year interactions. The results in Table 3 suggest that reforms tend to bring greater alignment, but at low levels of liberalization there is some suggestion of divergence that disappears with further reform.

A final question relevant to the interpretation of productivity dispersion is its correlation with the level and growth of aggregate productivity. If dispersion primarily reflects static distortions, then higher dispersion should be associated with lower aggregate productivity. Or if dispersion primarily reflects the strength of selection mechanisms that weed out poor performers, such as harder budget constraints, this again implies a negative relationship between dispersion and aggregate productivity and growth. But if dispersion predominantly reflects experimentation – innovation with uncertain outcomes – then it may be positively related with subsequent productivity growth. The results from simple regressions of aggregate MFP level and growth as alternative dependent variables on productivity dispersion, measured alternatively as standard deviation and 90-10 percentile range, are shown in Table 4. The regressions also include country and year fixed effects. The estimated coefficients for MFP level are small and statistically insignificant, but the coefficients for MFP growth are positive and highly significant. The magnitudes are not small, suggesting as much as an additional half-point to a point in percentage aggregate productivity growth associated with the range of variation of dispersion present in the data.

4. Conclusion

Persistent dispersion of productivity among firms within narrow industries may reflect a variety of factors. Recent research has focused on the possibility that firm-specific taxes on output or inputs create idiosyncratic differences in the effective prices faced by firms, leading to variation in marginal costs or products. If these factors were the only source of dispersion, then policies to eliminate static distortions, for instance by unifying tax rates, would equalize effective prices and increase aggregate output. On this basis, for example, Hsieh and Klenow (2009) find that equalizing productivity within industries would raise GDP in China by 87 percent, in India by 128 percent, and in the U.S. by 43 percent. However, a different set of factors points to a more ambiguous conclusion: in a dynamic setting with uncertainty, where firms experience idiosyncratic shocks and make decisions about entry, exit, investment, and restructuring, productivity dispersion arises naturally. Frictions affect the amount of reallocation taking place along each margin, but policies to reduce frictions will generally not reduce productivity dispersion.

Our theoretical model demonstrates these contentions. We show in particular that lowering the cost of entry will raise productivity dispersion among entrants and possibly in the overall distribution. Lowering the cost of investment or restructuring may also raise dispersion as some investments are successful, leading to productivity in the right tail, while others may fatten the left tail. Lowering the cost of exit (raising fixed operating costs) would raise the exit threshold, and dispersion may increase among continuing firms. The model shows that in each of these cases, reducing the friction raises aggregate output but may raise productivity dispersion at the same time. In the dynamic setting we focus on, the productivity distribution is influenced by forces of selection that tend to reduce dispersion and by changing opportunities for experimentation, which tend to raise dispersion.

Our empirical analysis considers the case of a major deregulation of a U.S. industry, telecommunications equipment manufacturing, and the drastic liberalization of the 1990s in six East European transition economies. In both cases, we find that the policy reforms raise productivity dispersion, however measured. Dispersion rises both absolutely and relative to a benchmark of overall U.S. manufacturing productivity dispersion. The analysis of East European economies during the socialist period shows levels of productivity dispersion very similar to those in the U.S. Productivity dispersion rises with reforms in all six countries. Evidence also suggests that productivity dispersion is associated with future aggregate productivity growth rather than decline. The results are consistent with a large role of experimentation in driving the heterogeneity of productivity outcomes.

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Appendix A: Proofs

Existence and uniqueness of stationary equilibrium. Let $R = (1 - G(\phi_e))N$ be the total mass of entrants corresponding to a given entry threshold ϕ_e. Note that for any given ϕ_e there exists a unique corresponding R by the assumption that N is exogenously given and the fact that G is monotonic. Therefore, R and ϕ_e can be used interchangeably to denote the extent of entry. Let $\mu \equiv \mu(R, x, \theta_r)$ be an invariant measure that corresponds to a given triplet $\{R, x, \theta_r\}$. Consider now the pair $\{R(\theta_r), x(\theta_r)\}$ such that given $\theta_r \in U \equiv [0,1]$, $\{R(\theta_r), x(\theta_r)\}$ satisfies the free entry condition (3) and the exit condition (2) with equality for the associated invariant measure μ. Denote by $\mathcal{T}_1: U \to [0, N] \times U$ the mapping that yields a pair $\{R(\theta_r), x(\theta_r)\}$ for any given $\theta_r \in U$. Next, let $\theta_r(R, x)$ be the value that satisfies the restructuring condition (4) with equality for a given pair $\{R, x\}$ and the associated invariant measure μ. Denote by $\mathcal{T}_2$: $[0, N] \times U \to U$ the mapping that associates a given pair $\{R, x\}$ with some $\theta_r \in U$ that satisfies the restructuring condition with equality. The proof of existence and uniqueness then amounts to showing that the composite mapping $\mathcal{T} = \mathcal{T}_1 \circ \mathcal{T}_2$ possesses a unique fixed point that lies in the interior of U. Some of the arguments in the proofs below follow closely the related arguments in Hopenhayn (1992). Note that the model satisfies all the basic assumptions A1-A5 in Hopenhayn (1992) and, in addition, the conditions U1 and U2 therein. In particular, the model reduces to Hopenhayn's (1992) framework when the restructuring cost, c_r, is prohibitively high, the mass of potential entrants, N, is infinite, and the distribution of entrants' priors, G, is degenerate at some value ϕ.

Existence. First, note that the invariant measure μ is defined by

$$\mu(\theta) = (\mathcal{L}\mu)(\theta) + N \int_{\phi_e}^{1} \left(\int_0^{\theta} h_e(z|\phi) dz \right) g(\phi) d\phi,$$

where $\mathcal{L}$ is the operator such that, for any set $S \subset U$,

$$\mathcal{L}(S) = \begin{cases} \displaystyle\int_{y \in S} h\,(y|z)\mu(dz), & \text{for } z \geq x. \\ 0, & \text{otherwise.} \end{cases}$$

The steps similar to Lemma 4 in Hopenhayn (1992) guarantee the existence of μ. Also, following Lemma 5 in Hopenhayn (1992), μ is jointly continuous in its arguments, strictly decreasing in x, and strictly increasing in R (strictly decreasing in ϕ_e). Now, let $R_1(\theta_r)$ be the entry mass that satisfies (3) with equality for a given exit rule $x(\theta_r)$. Similarly, let $R_2(\theta_r)$ be the entry mass such that the exit rule $x(\theta_r)$ satisfies (2) with equality. The properties of R_1 and R_2 follow from Lemmas 6 and 7 in Hopenhayn (1992). Theorems 2 and 3 in Hopenhayn (1992) then imply the existence of a pair $\{R(\theta_r), x(\theta_r)\}$ such that $R(\theta_r) > 0$ and $x(\theta_r) \in (0,1)$, for any given θ_r, as long as c_e is not too high. Therefore, $\mathcal{T}_1$ is a well-defined, continuous operator that maps θ_r into a pair $\{R(\theta_r), x(\theta_r)\}$ that satisfies (2) and (3) with equality. Next consider the mapping $\mathcal{T}_2$. Given a pair $\{R, x\}$, the left hand side of (4) is continuous and strictly decreasing in θ_r by the assumptions of the model. Therefore, there exists a unique value θ_r that satisfies (4) with equality, as long as $c_r < E_r[V(\theta')|0]$, i.e. the benefit from restructuring exceeds the cost of doing so for the least productive firm. Thus, $\mathcal{T}_2$ is a well-defined, continuous function that maps any $\{R, x\}$ into a θ_r that satisfies (4) with equality. Given the continuity of $\mathcal{T}_1$ and $\mathcal{T}_2$, the composition $\mathcal{T} = \mathcal{T}_1 \circ \mathcal{T}_2$ is then a continuous function that maps U onto itself. The existence of a fixed point θ_r^* then follows from the Brouwer fixed point theorem. This fixed point is in the interior of U and satisfies $\theta_r^* > x^*$, as long as $c_r < E_r[V(\theta')|0]$. Consequently, there exists a triplet $\{\theta_r^*, x^*, R^*\}$ and the associated invariant measure μ^*, that constitute a stationary equilibrium with positive entry, exit, and restructuring.

Uniqueness. Suppose that the stationary equilibrium is not unique, i.e., the mapping $\mathcal{T}$ has more than one fixed point. Let $x_1^* < x_2^*$ denote the two exit thresholds for two different stationary equilibria with the corresponding measures μ_1^* and μ_2^*. By the definitions of x_1^* and x_2^*, $V(x_1^*; \mu_2^*) < V(x_2^*; \mu_2^*) = 0$, and $V(x_1^*; \mu_1^*) = 0$. Thus, there must be some firm type θ such that $\tilde{\pi}(\theta; \mu_2^*) < \tilde{\pi}(\theta; \mu_1^*)$. However, the free entry condition implies that $V^e(\phi_1^*; \mu_1^*) = V^e(\phi_2^*; \mu_2^*) = c_e$. Therefore, while profits $\tilde{\pi}$ are lower for some firm type θ under μ_2^*, they cannot be lower for all θ, for otherwise $V^e(\phi_2^*; \mu_2^*) < c_e$. This argument implies that if profits move in the same direction for all θ going from one equilibrium to another, then the free entry condition cannot be satisfied for both equilibria – a contradiction. Assumptions U1 and U2 in Hopenhayn (1992), both of which are also assumed here, ensure that profits for all firm types move in the same direction and hence, the equilibrium is unique.

Proposition 1. *As a result of a decline in c_r, x^* and ϕ_e^* either both increase or both decrease.* ■

Proof. Let

$$\hat{V}(\theta; \mu) = \frac{V(\theta; \mu)}{c_r},$$

be the rescaled value function for an establishment, where the dependence on the measure μ is made explicit. One can then rewrite (1) as

$$\hat{V}(\theta; \mu) = \frac{u(\theta)m(p(\mu), w, r)}{c_r} - \frac{c_f}{c_r} + \beta \max\{0, E_r[\hat{V}(\theta')|\theta] - 1, E_n[\hat{V}(\theta')|\theta]\},$$

where we used the assumption that the profit function is separable in productivity and prices, i.e. $\tilde{\pi}(\theta; \mu) = u(\theta)m(p(\mu), w, r)$, for some functions u and m. Now consider two industries such that c_r is lower in the second industry: $c_r^2 < c_r^1$. The aim is to show that if $x_2 \geq x_1$ ($x_2 < x_1$) then it must be the case that $\phi_e^2 \geq \phi_e^1$ ($\phi_e^2 < \phi_e^1$). Towards that end, let the measures of firms be μ^1 and μ^2. If $x_2 \geq x_1$

$$\int \hat{V}(\theta'; \mu_1) h_n(\theta'|x_2) \geq \int \hat{V}(\theta'; \mu_1) h_n(\theta'|x_1) = 0 = \int \hat{V}(\theta'; \mu_2) h_n(\theta'|x_2),$$

where the first inequality follows from the fact that $H_n(\theta'|\theta)$ is strictly decreasing in θ, and the equalities from the fact that x_1 and x_2 are the marginal firm types so they must have zero expected value from continuing. But the inequality $\int \hat{V}(\theta'; \mu_1) h_n(\theta'|x_2) \geq \int \hat{V}(\theta'; \mu_2) h_n(\theta'|x_2)$ can hold can only when

$$\frac{m(p(\mu_1), w, r) - c_f}{c_r^1} \geq \frac{m(p(\mu_2), w, r) - c_f}{c_r^2},$$

which implies

$$m(p(\mu_1), w, r) - c_f \geq \frac{c_r^1}{c_r^2}\left(m(p(\mu_2), w, r) - c_f\right) > m(p(\mu_2), w, r) - c_f,$$

where the last inequality follows because $c_r^1 > c_r^2$. Therefore, period profit for each firm type, $\tilde{\pi}(\theta; \mu)$, is higher in industry 1, and so is the value of each firm type
$V(\theta, \mu_1) \geq V(\theta, \mu_2)$.

The expected profit for any potential entrant type ϕ must then also be higher in industry 1, implying a lower entry threshold in industry 1, i.e. $\phi_e^2 \geq \phi_e^1$. The steps of the proof so far can be repeated to show the other combination, $x_2 < x_1$ and $\phi_e^2 < \phi_e^1$. ■

Appendix B: The derivation of aggregate productivity

For general production functions, it is not possible to represent the aggregate production function in the economy using the exact same form of the firm-level production function.[46] To derive an explicit expression for aggregate productivity, TFP, assume, as done commonly in the literature, that a firm's production function is of Cobb-Douglas type, $f(k,l) = k^{\lambda} l^{\gamma}$, $\lambda + \gamma < 1$. Suppose also that the fixed cost entails both overhead labor and capital: $c_f = rk_f + wl_f$, where k_f and l_f are the amount of fixed capital and labor per firm, respectively. Let Q^* be the aggregate output, and let K^* and L^* be the total capital and labor used. The industry's production function can then be written as $Q^* = \text{TFP} \times K^{*\lambda} L^{*\gamma}$. Define $o_l^* = \frac{M^* l_f}{M^* l_f + \int_0^1 l^*(\theta)\mu^*(d\theta)}$ and $o_k^* = \frac{M^* k_f}{M^* k_f + \int_0^1 k^*(\theta)\mu^*(d\theta)}$ as the fraction of labor and capital used as overhead, respectively. TFP can then be expressed as

$$\text{TFP} = \frac{Q^*}{K^{*\lambda} L^{*\gamma}} = \left(\frac{(1 - G(\phi_e^*))N}{H^*(x^*)}\right)^{1-\lambda-\gamma} (1 - o_l^*)^{\gamma} (1 - o_k^*)^{\lambda} \left(\int_0^1 \theta^{1/(1-\lambda-\gamma)} h^*(\theta) d\theta\right)^{1-\lambda-\gamma}. \quad (1)$$

Note that TFP is a geometric average of firm-level TFPQ, θ. TFP is higher when there is a larger mass of potential entrants (N), lower entry threshold (ϕ_e^*), lower fixed costs (l_f and k_f), lower exit threshold (x^*), and a higher productivity distribution H^*, in a first-order stochastic dominance sense. Note that θ_r^* also affects TFP through its effect on H^*, implicit in the expression (1). Because ϕ_e^*, x^*, θ_r^*, o_l^*, o_k^* and H^* are all functions of the costs c_e, c_f, and c_r, TFP is also a function of the costs c_e, c_f, and c_r.

[46] See, e.g., Osotimehin (2013).

Figure 1a. Evolution of Productivity Dispersion in the U.S. Telecommunications Equipment Sector: Standard Deviation

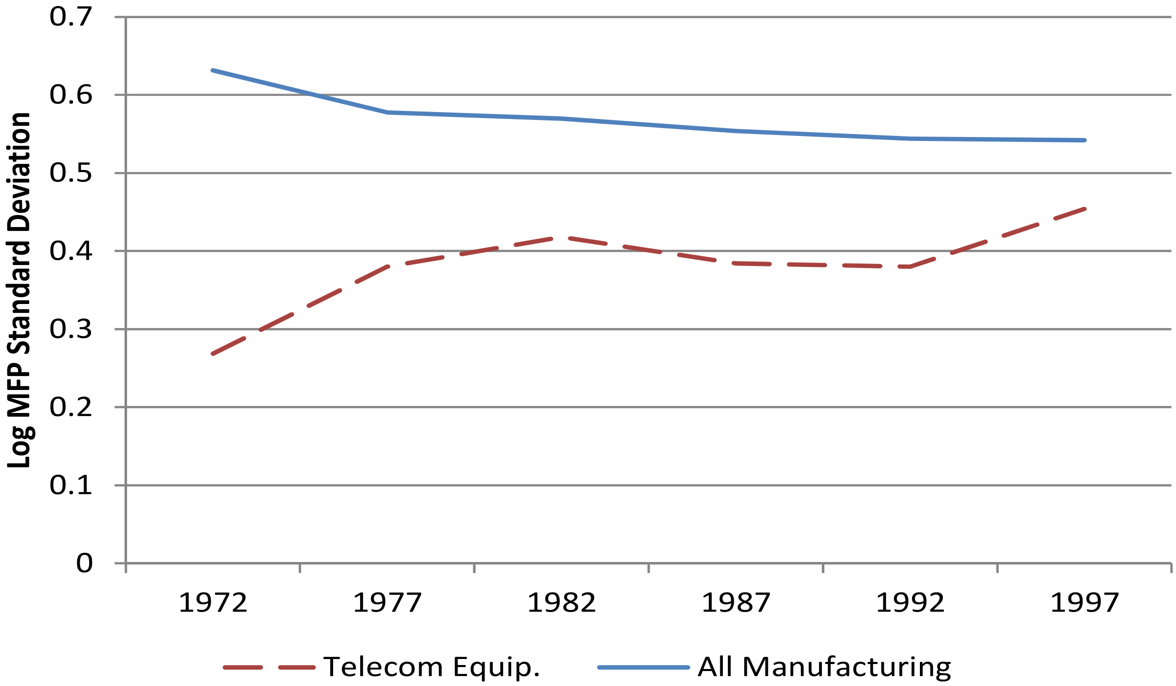


Figure 1b. Evolution of Productivity Dispersion in the U.S. Telecommunications Equipment Sector: 90-10 Percentile Range

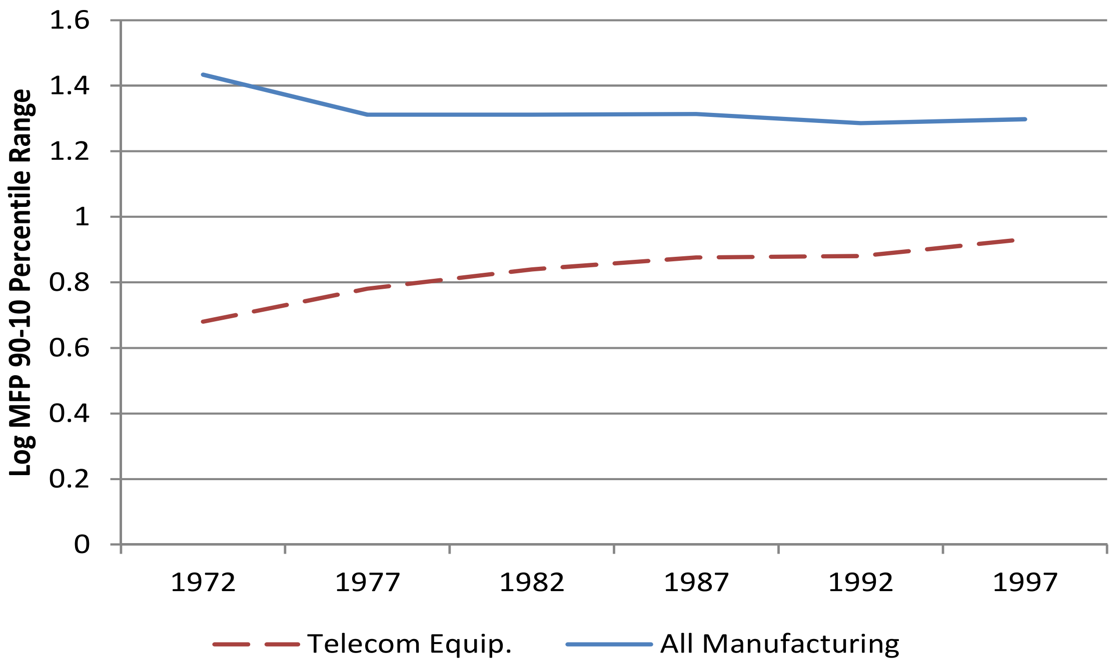


Figure 1c. Evolution of LP Dispersion in the U.S. Telecommunications Equipment Sector: Standard Deviation

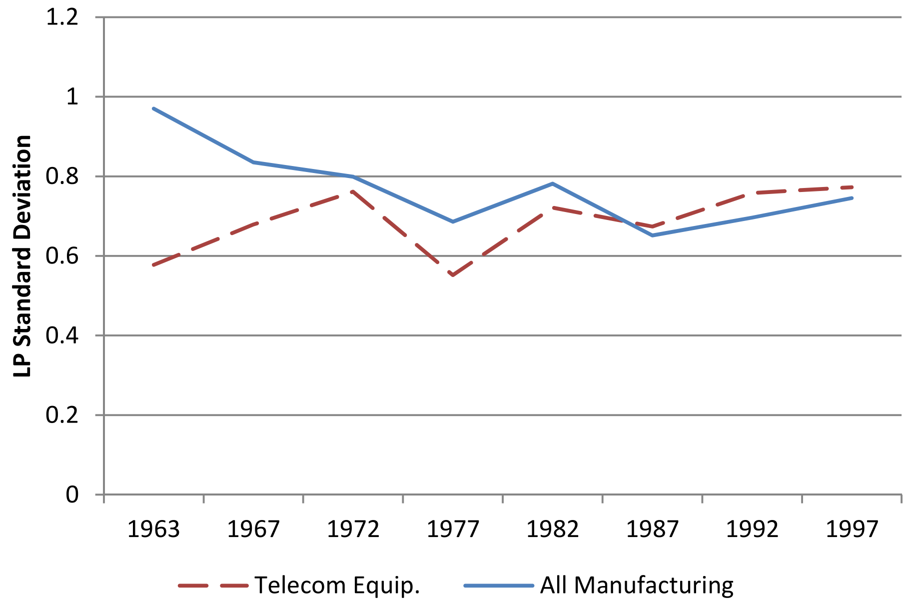


Figure 1d. Evolution of LP Dispersion in the U.S. Telecommunications Equipment Sector: 90-10 Percentile Range

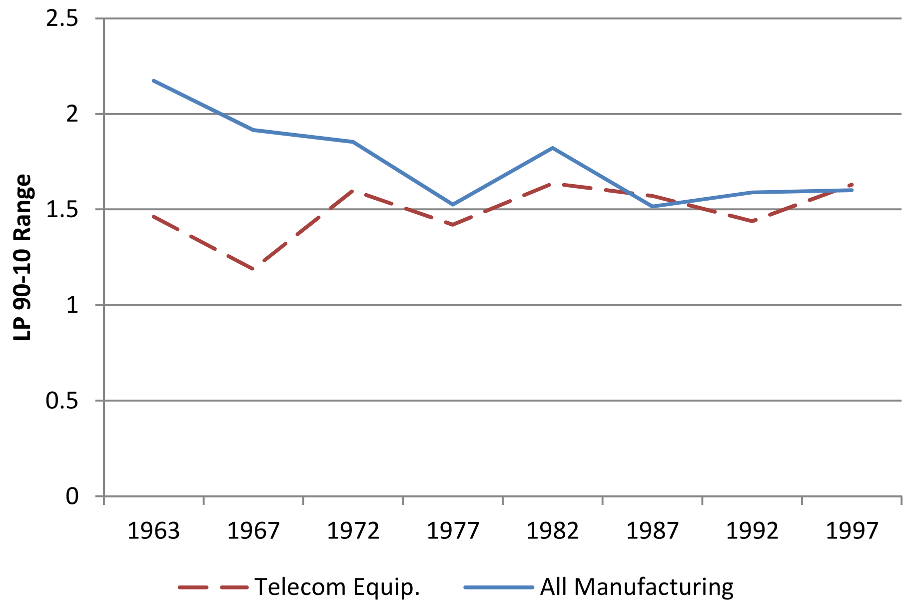


Figure 2a. MFP Distribution in the U.S. Telecommunications Equipment Sector

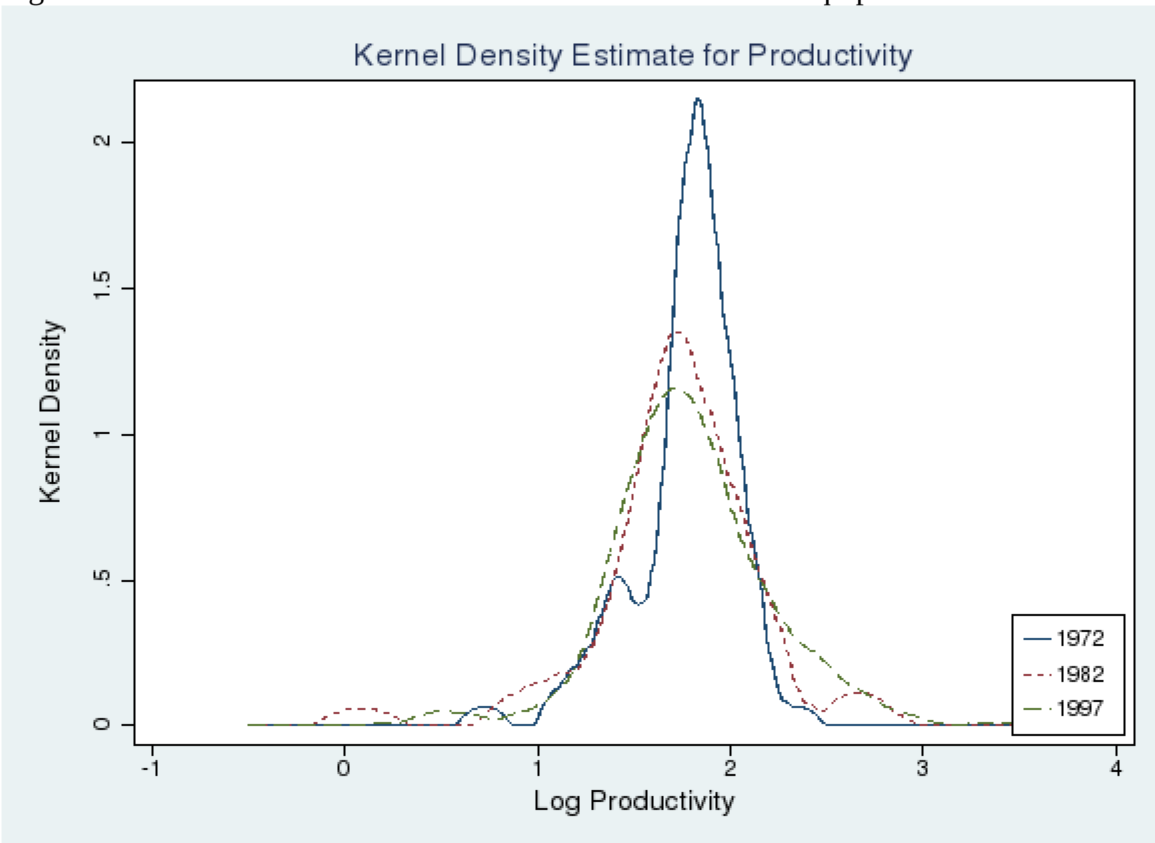


Figure 2b. LP Distribution in the U.S. Telecommunications Equipment Sector

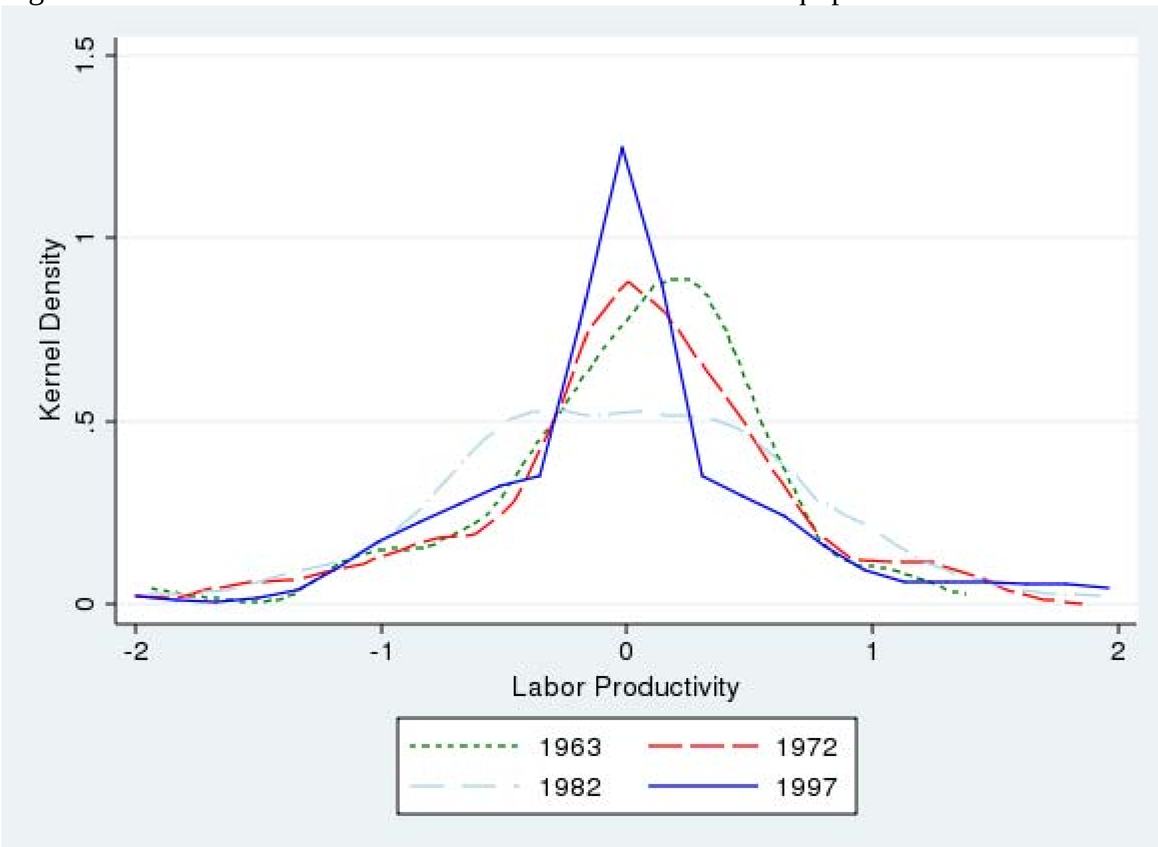


Figure 3a. Components of 5-Year MFP Standard Deviation Change
U.S. Telecommunications Equipment Sector

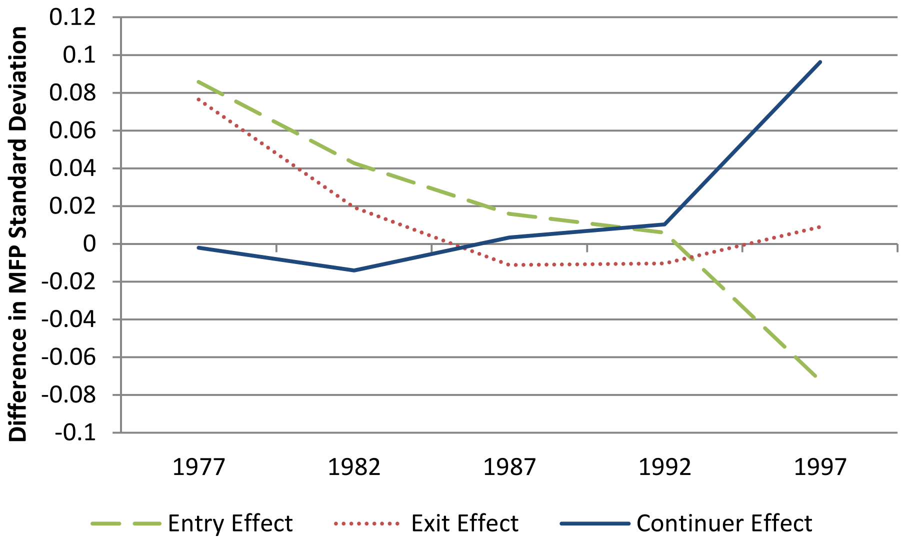


Figure 3b. Components of 5-Year MFP 90-10 Range Change
U.S. Telecommunications Equipment Sector

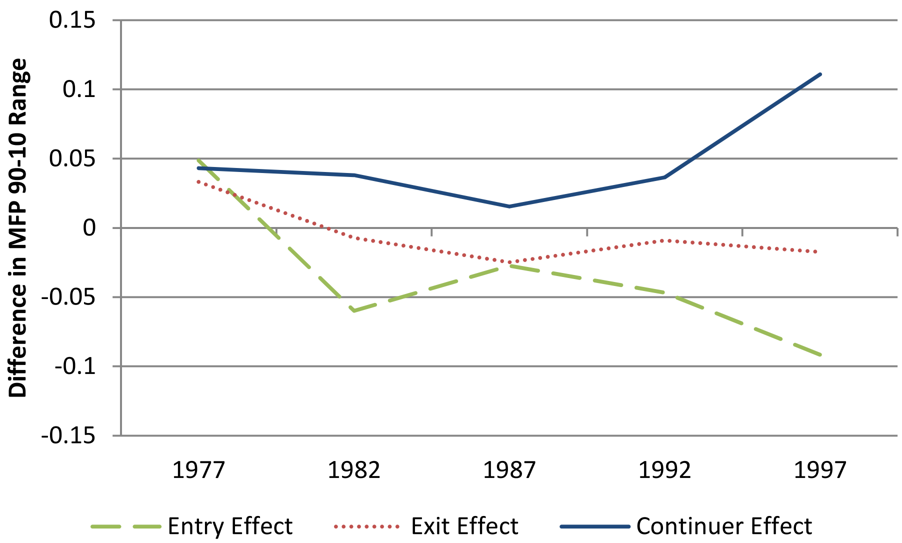


Figure 3c. Components of 5-Year MFP Standard Deviation Change
U.S. Manufacturing Sector

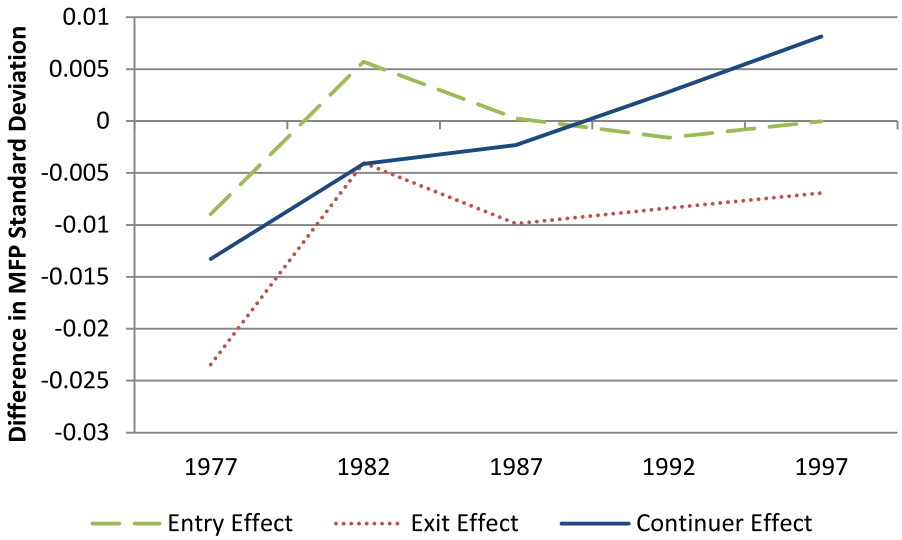


Figure 3d. Components of 5-Year MFP 90-10 Range Change
U.S. Manufacturing Sector

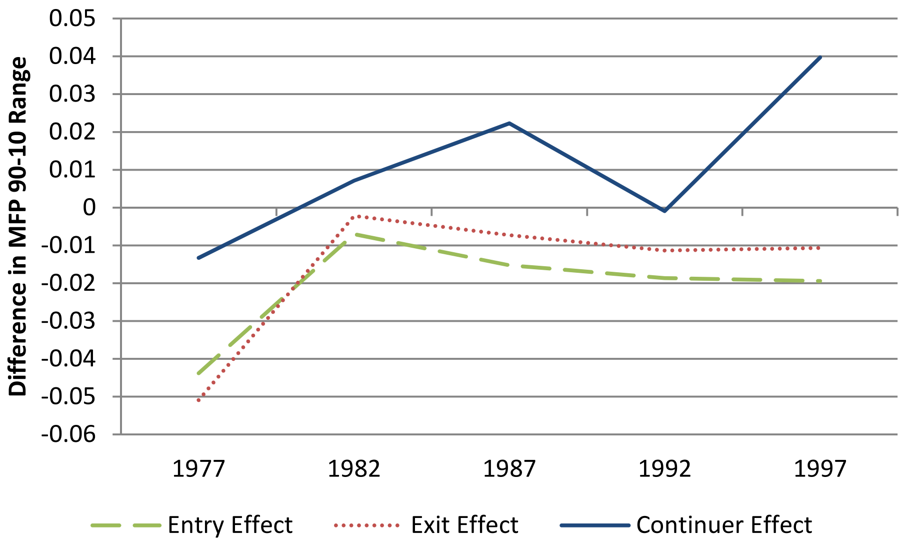
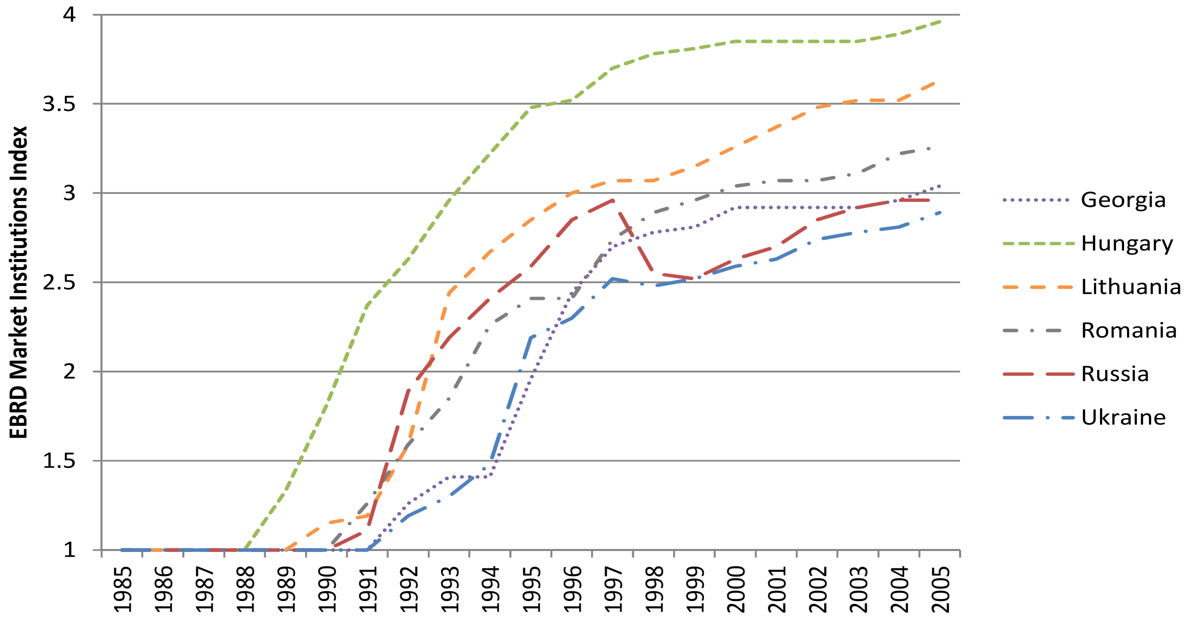


Figure 4. EBRD Market Institution Index



The index ranges from 1 (central planning) to 4.3 (developed market economy). The indices can be downloaded from http://www.ebrd.com/what-we-do/economic-research-and-data/data/forecasts-macro-data-transition-indicators.html.

Figure 5a. Evolution of Productivity Dispersion with Market Liberalization, MFP Standard Deviation

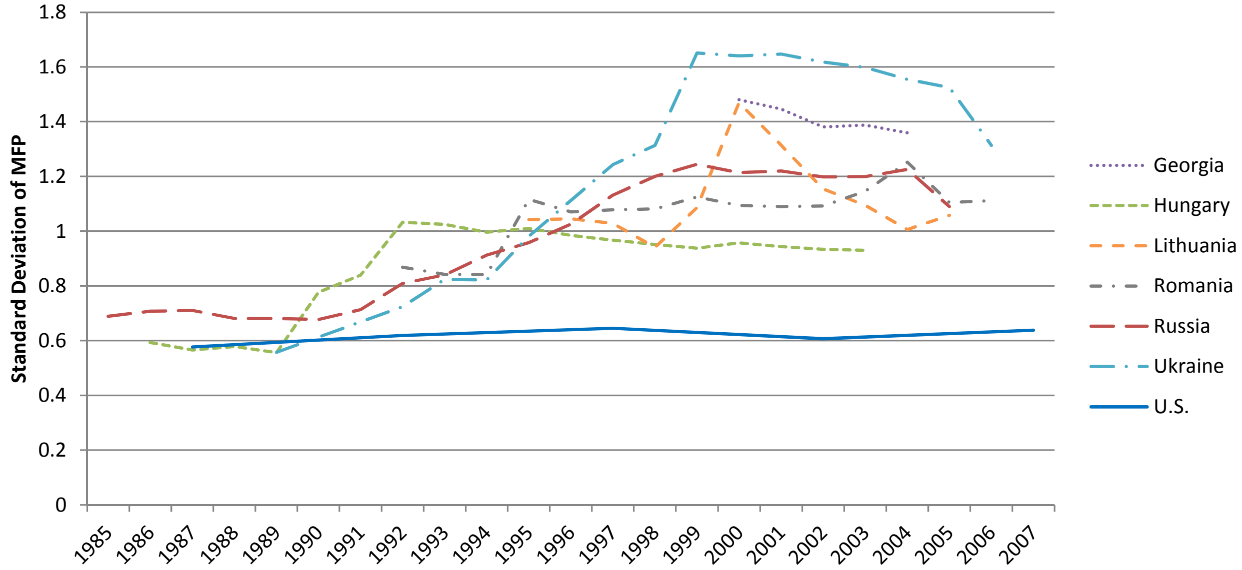


Figure 5b. Evolution of Productivity Dispersion with Market Liberalization, MFP 90-10 Percentile Range

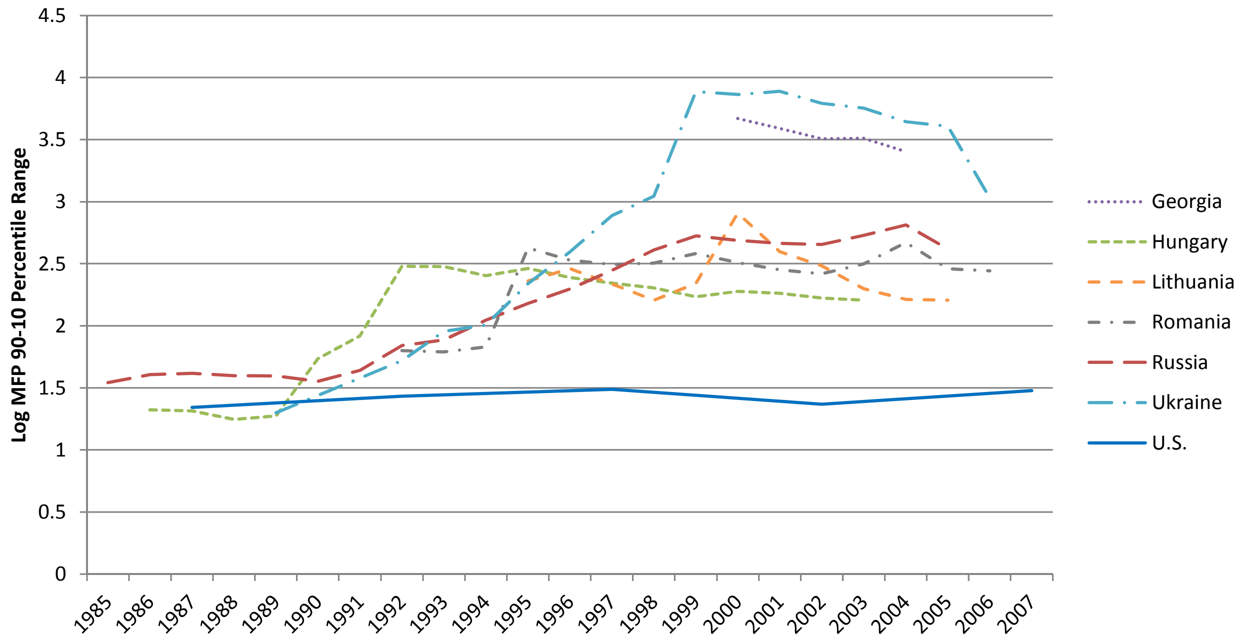


Figure 6a. Productivity Distribution Prior to Transition

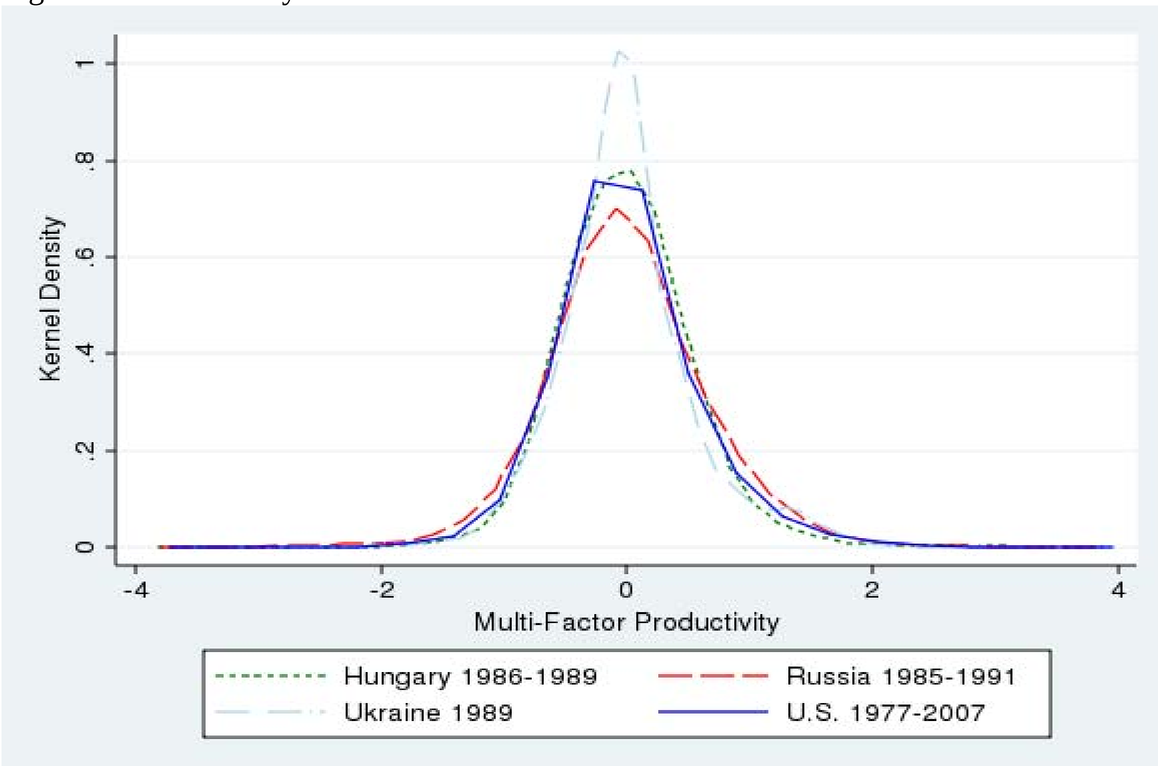


Figure 6b. Productivity Distribution during Transition

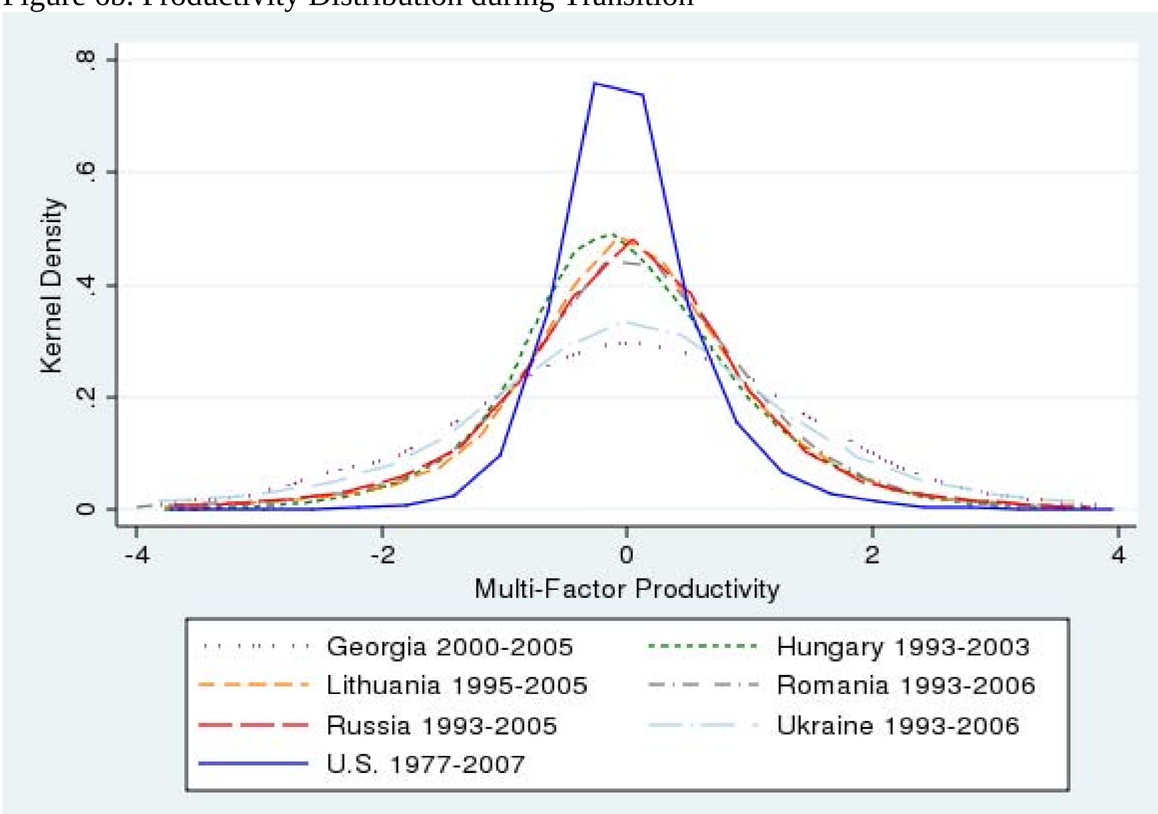
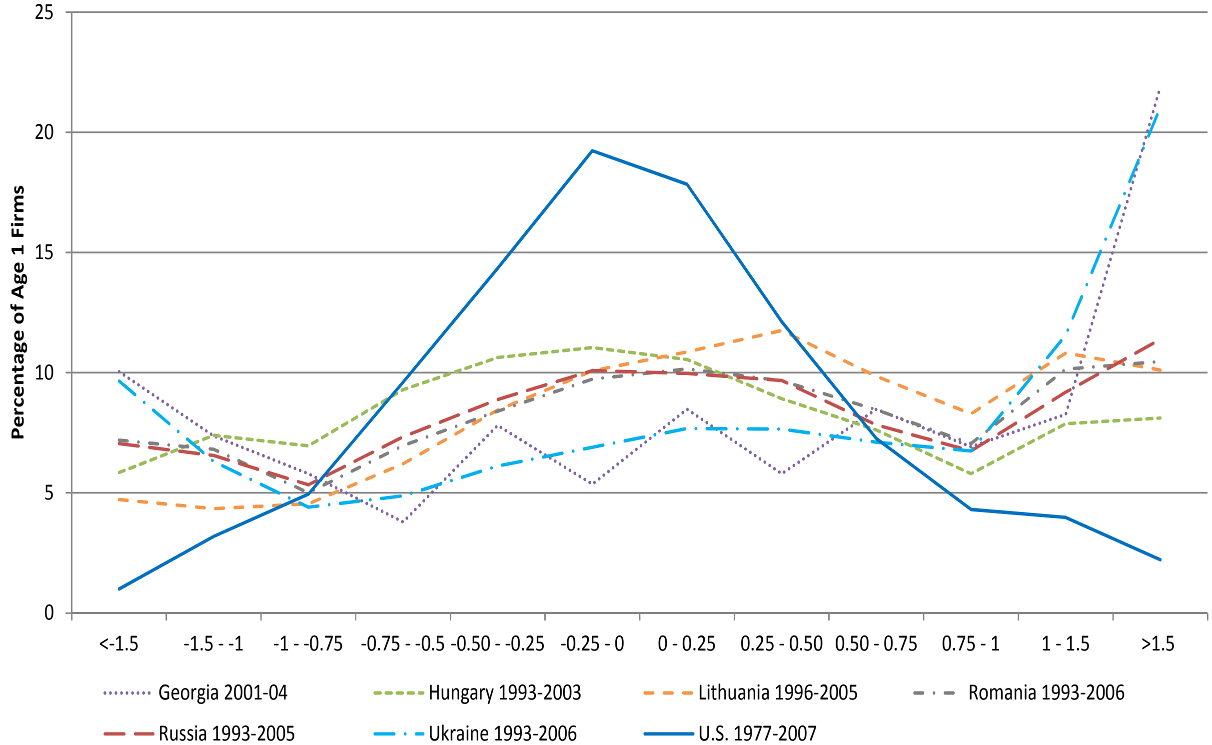
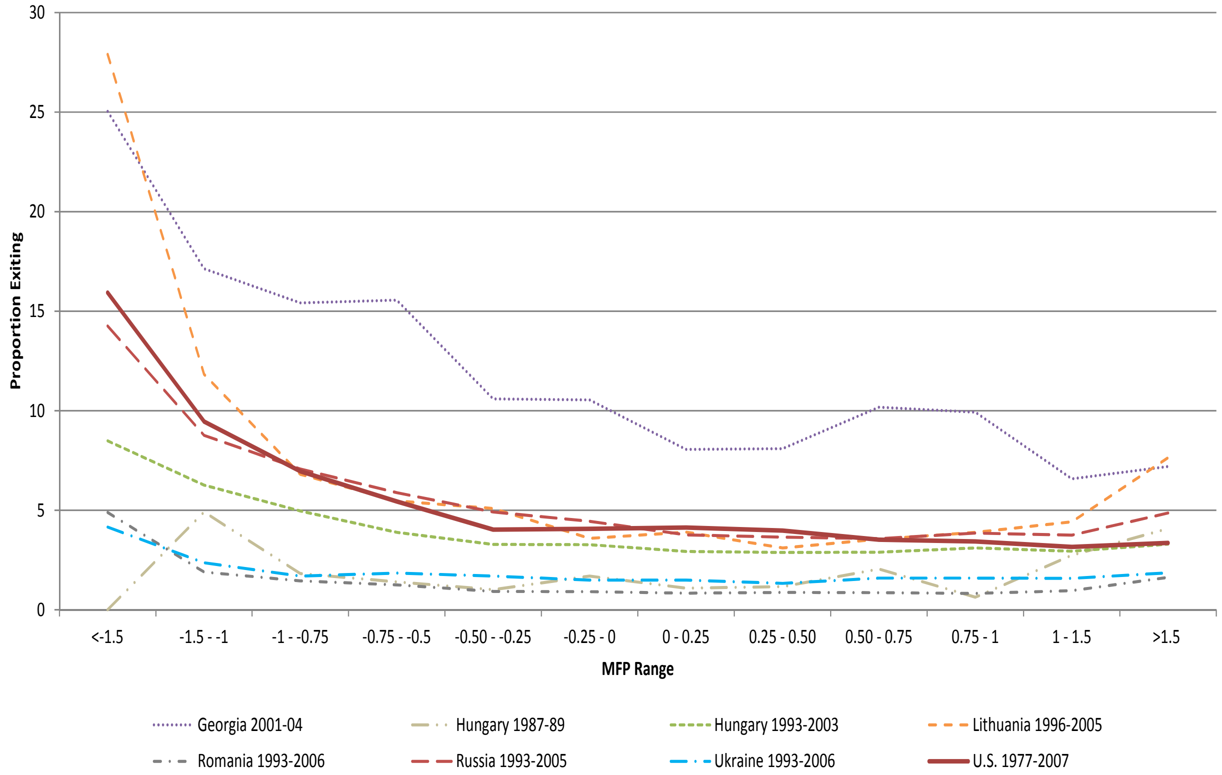


Figure 7. New Entrant Productivity Distribution



Note: Sample is restricted to firms entering the data for the first time in the previous year. Productivity is measured as MFP at age one to avoid mismeasurement associated with partial year operation in the first year.

Figure 8. Productivity and Exit Rates



Note: Productivity is measured as MFP in the last year before the firm permanently exits the data. For the transition economies, the observations are annual, while for the U.S. observations are every five years (but the productivity lag is still one-year).

Figure 9a. Components of 5-Year MFP 90-10 Range Change in Hungary

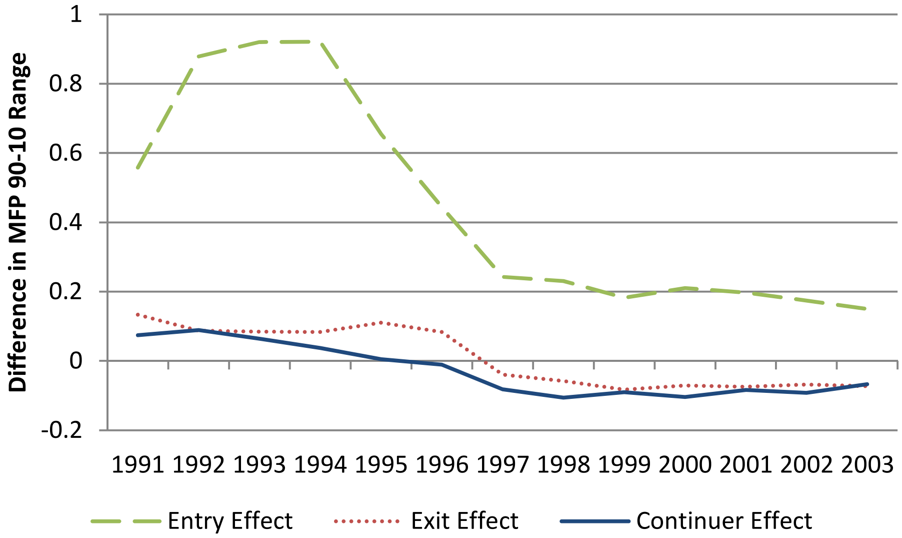


Figure 9b. Components of 5-Year MFP 90-10 Range Change in Lithuania

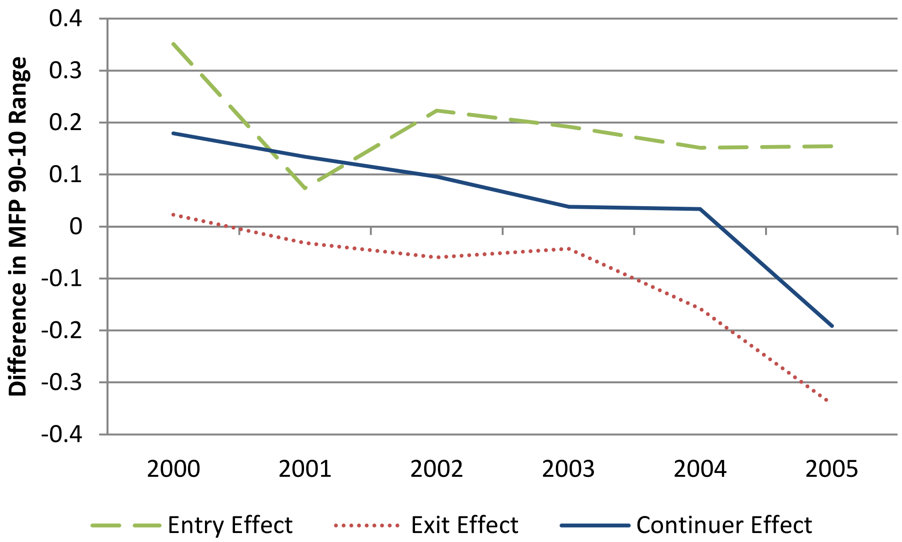


Figure 9c. Components of 5-Year MFP 90-10 Range Change in Romania

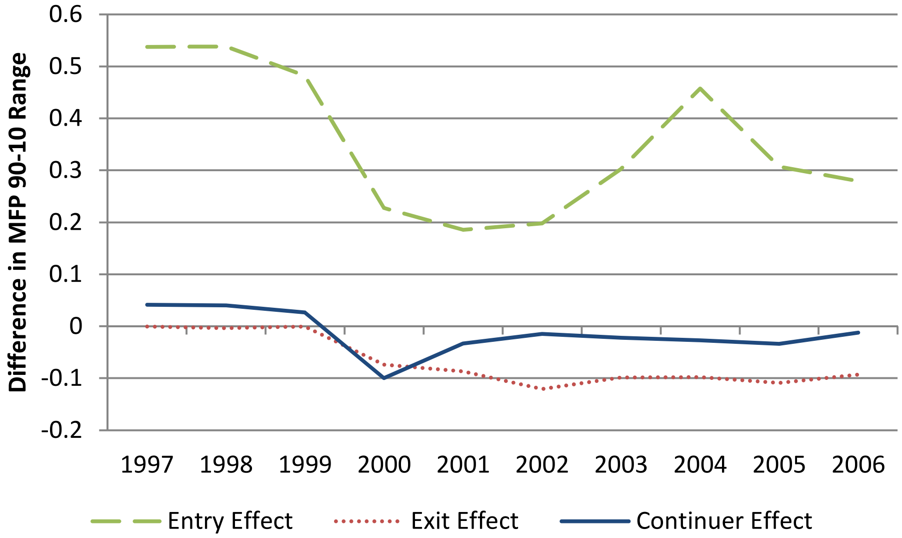


Figure 9d. Components of 5-Year MFP 90-10 Range Change in Russia

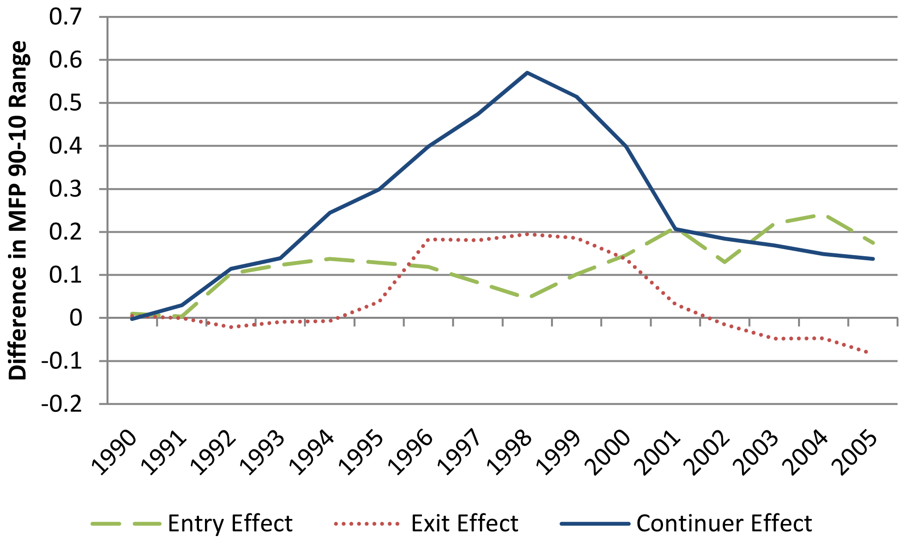


Figure 9e. Components of 5-Year MFP 90-10 Range Change in Ukraine

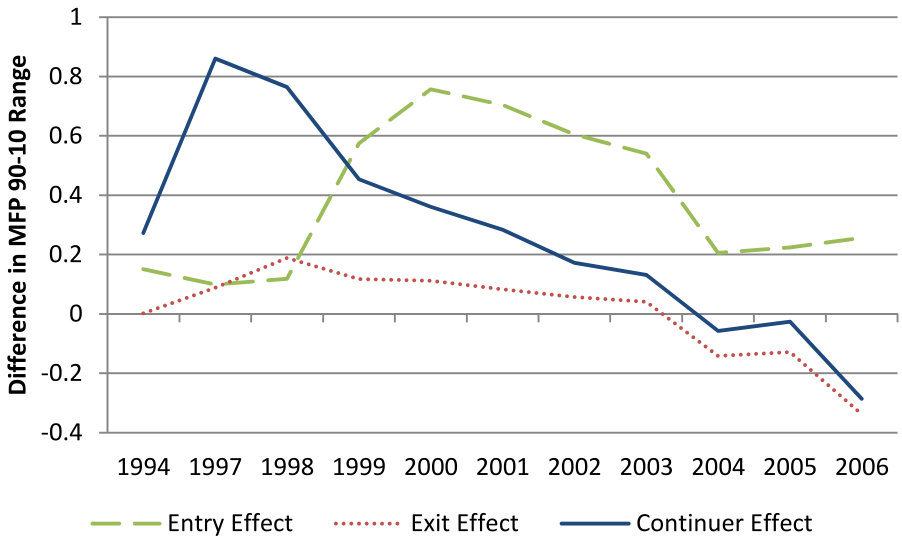


Figure 9f. Components of 5-Year MFP 90-10 Range Change in U.S. Manufacturing

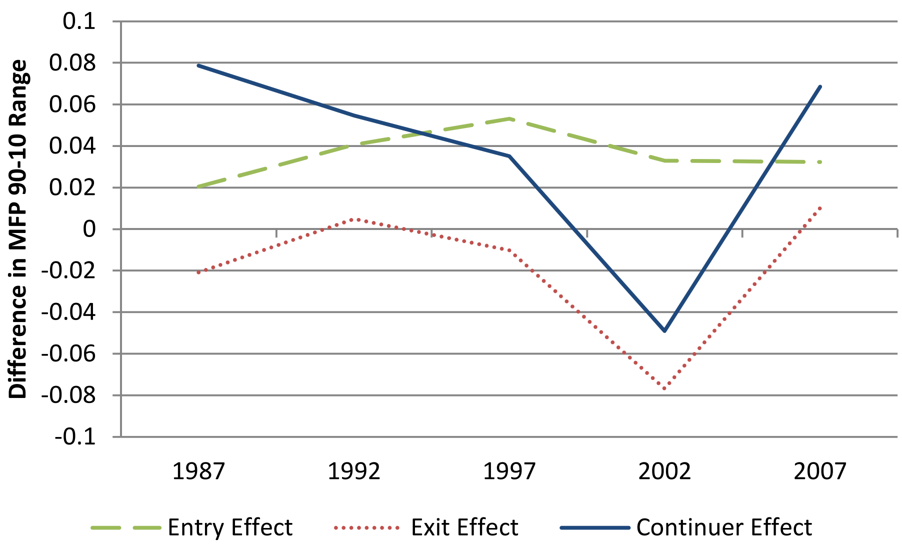


Table 1. Number of Businesses and Business-Year Observations

Country	Years	Number of Businesses	Number of Business-Year Observations
Georgia	2000-2004	2,463	7,566
Hungary	1986-2003	32,482	170,495
Lithuania	1995-2005	7,731	40,596
Romania	1992-2006	69,323	356,838
Russia	1985-2005	35,405	318,535
Ukraine	1989, 1992-2006	43,084	222,473
U.S. Manufacturing MFP (benchmark for Eastern Europe)	1977, 1982, 1987, 1992, 1997, 2002, 2007	551,144	1,310,913
U.S. Manufacturing MFP (benchmark for Telecom Equipment)	1972, 1977, 1982, 1987, 1992, 1997	496,444	1,070,582
U.S. Manufacturing LP (benchmark for Telecom Equipment)	1963, 1967, 1972, 1977, 1982, 1987, 1992, 1997	1,113,427	2,513,087
U.S. Telecom Equipment MFP	1972, 1977, 1982, 1987, 1992, 1997	324	524
U.S. Telecom Equipment LP	1963, 1967, 1972, 1977, 1982, 1987, 1992, 1997	1,416	2,265

Note: MFP = multi-factor productivity. LP = labor productivity. A business is defined as a firm in the transition economy datasets, and it is an establishment in the U.S. data. The transition economy data are annual, while productivity measurement is possible only every five years in the U.S.

Table 2. Productivity Dispersion After Market Liberalization

	MFP Standard Deviation		MFP 90-10 Percentile Range	
EBRD	0.041	1.000	0.177	2.334
	(0.086)	(0.121)	(0.136)	(0.495)
$EBRD^2$		-0.186		-0.419
		(0.020)		(0.113)

Notes: Each column shows the results from a separate country-year regression with a productivity dispersion measure (SD or 90-10 range) as the dependent variable and the EBRD index (and in columns 2 and 4, the index squared) and country and year fixed effects as independent variables. The EBRD reform index is lagged one year. Standard errors, cluster-adjusted by country, are in parentheses. N = 86 country-year observations.

Table 3. Market Liberalization and Correlation of U.S. and Eastern European MFP Dispersion

	MFP Standard Deviation		MFP 90-10 Percentile Range	
U.S. Value* EBRD	0.565	-0.312	0.542	-0.596
	(0.144)	(0.279)	(0.150)	(0.287)
U.S. Value* $EBRD^2$		0.170		0.221
		(0.046)		(0.048)

Notes: Each column shows the results from separate industry-country-year regressions with a productivity dispersion measure (SD or 90-10 range) as the dependent variable, country-industry, and the corresponding U.S. value interacted with the EBRD index (and in columns 2 and 4, the index squared) and country-year and industry-year fixed effects as independent variables. The EBRD reform index is lagged one year. Industry is defined at the 2-digit SIC or NAICS level. The U.S. value is the mean for the dependent variable in the U.S. industry across 1977-2007. Robust standard errors adjusted for clustering by country-year are in parentheses. N = 1,634 industry-country-year observations.

Table 4. Productivity Dispersion and Aggregate Productivity

	MFP Level	MFP Growth
MFP Standard Deviation	-0.124	0.519
	(0.168)	(0.175)
MFP 90-10 Percentile Range	0.116	0.221
	(0.142)	(0.057)

Notes: Each cell shows the results from a separate country-year regression with either aggregate MFP level or growth as the dependent variable and a productivity dispersion measure (SD or 90-10 range) and country and year fixed effects as independent variables. Standard errors, cluster-adjusted by country, are in parentheses. N = 86 country-year observations.