DOES FDI REDUCE MISALLOCATION? EVIDENCE FROM INDIA*

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Abstract

A growing literature suggests that capital misallocation is a major contributor to low productivity in low-income countries. Yet little is known about what policies reduce misallocation, and measuring misallocation is challenging. Motivated by the fact that frictions in the domestic capital market may cause misallocation, we test whether foreign direct investment liberalization reduced misallocation in the Indian manufacturing sector. We use a difference-in-differences strategy exploiting industry and time variation in the liberalization of FDI to estimate the effects of the policy. This strategy credibly identifies changes in firms' input wedges. For domestic firms with initially high marginal revenue products of capital (MRPK), FDI liberalization increased assets by 50%, sales by 20%, wage bills by 26%, and reduced the marginal revenue product of capital by 40%. The liberalization had no effect on firms with low MRPK. Aggregating our causal estimates of the effect of the policy change on firms inputs, we find that the policy increased manufacturing sector gross output by 1%.

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1 Introduction

Half of the cross-country differences in income are explained by differences in productivity (Caselli, 2005). These differences in productivity can in turn be driven by the distortion of inputs like labor and capital across heterogenous firms (Restuccia and Rogerson, 2008). For example, Hsieh and Klenow (2009) find that misallocation can explain 40-60% of the differences between Indian and US manufacturing productivity. Yet, despite mounting evidence of the importance of misallocation, important challenges also remain. First, as pointed out by Syverson (2011), the specific sources of distortion cannot be identified from aggregate comparisons by definition, leaving policymakers with limited information about the specific levers they can activate to reduce misallocation and increase prosperity. Second, credible measurement of the implicit frictions that are distorting firms' input choices has been a serious hurdle in quantifying the impact of misallocation. Indeed, it is common to attribute all the cross-sectional heterogeneity in the marginal returns to firms' inputs to misallocation, which may upwardly bias measures of misallocation due to measurement error (Rotemberg and White, 2017), model mis-specification (Haltiwanger et al., 2018), and variation in the volatility of productivity (Asker et al., 2014).

An unusual natural experiment allows us to make progress on both these issues and to better understand the extent to which imperfect capital market frictions lead to misallocation. We study the effects of the staggered introduction of automatic approval of foreign direct investments up to 51% of firms' capital in different industries in India. The structure of the policy allows us to exploit cross-industry and cross-time variation in a difference-in-differences framework to estimate the effects of the policy on misallocation. Thus, we examine whether FDI liberalization policies led capital to be allocated to the firms with the highest marginal returns, improving the allocation of resources. Beyond allowing us to identify the effects of FDI policy, our natural experiment allows us to isolate the change in wedges (and their aggregate effects on output) caused by the policy under relatively weak identifying assumptions, without relying on cross-sectional variation.

The effect of this policy on misallocation is a priori unclear. On the one hand, capital account liberalizations have become increasingly common among developing countries as a way of correcting imperfections in local financial markets (e.g. Chinn and Ito (2006)). Foreign investment may reduce misallocation if foreign investors have better screening technologies or are less likely than domestic capital to invest in firms with low marginal returns due to regulation or for historical, political, or institutional reasons (e.g. Baneerjee and Munshi (2004), Cole (2009a), or Cole (2009b)). On the other hand, foreign investors may be less able to process and monitor soft information, particularly in low-income countries (Detragiache et al., 2008).[1] Thus, empirical work is needed to establish not just the magnitude but the direction of the effect of FDI liberalization on misallocation.

To measure the effects of FDI liberalization policies, we collect data on industry-level FDI liberalization episodes in 2001 and 2006 that allowed automatic approval of foreign investments. Combining this policy variation with a panel of large and medium-sized Indian firms, we find that liberalizing FDI had a moderate positive effect on assets and sales for firms in the treated industries. Next, to test whether FDI liberalization reduces misallocation, we examine whether the policy had different effects depending on firms' marginal revenue products of capital ($MRPK$) prior to the treatment. To estimate these differential effects, we classify firms as having $MRPK$ above or below the median for an industry and allow the effects of the policy change to differ depending on this classification. This empirical strategy now relies on within-industry comparisons. The policy effects for high $MRPK$ firms are given by comparing the change in outcomes for high $MRPK$ firms in industries that did and did not liberalize, and the same is true for low $MRPK$ firms. However, our key effect of interest, the differential effect on high $MRPK$ firms is given by the difference between these two policy effects. Thus, this strategy controls for the average effect of belonging to a deregulated industry, and therefore, requires milder identification assumptions. For instance, the differential effect of liberalization on high $MRPK$ firms is identified even if liberalized and non-liberalized industries have different industry-level time trends.

In this setting, we find that high $MRPK$ firms increase their assets by 50%, sales by 20%, wage bills by 26%, and reduce their marginal revenue returns to capital by 40% in response to the policy. In contrast, low $MRPK$ firms are not affected by the policy. Since high $MRPK$ firms initially have 140% higher marginal revenue returns to capital, misallocation declines. However, firm-level measures of $TFPR$ are unaffected for both types of firms. Event study graphs confirm that these effects are not driven by differential pre-trends between the two types of firms and provide evidence that the reductions in misallocation are not due to mean reversion.

The differential effects of the policy that we identify are remarkably robust. We replicate the

[1] In the context of foreign banks' behavior in poor countries, several studies have found that foreign banks lend essentially to large domestic firms, potentially increasing credit constraints for local firms (e.g. Mian (2006) for Pakistan, Gormley (2010) for India, or Detragiache et al. (2008) for a cross-section of countries). Similarly, Gopinath et al. (2017) shows that foreign capital inflows led to greater misallocation in Southern Europe in the early 2000s.

same patterns using three different measures of $MRPK$. The first of these measures is deflated sales divided by deflated assets, which is proportional to $MRPK$ within 2-digit industries under the assumption that all firms in an industry have Cobb-Douglas production functions with the same capital and labor intensity. The other two measures of $MRPK$ are computed using the parameters from production function estimates produced by the methods of Levinsohn and Petrin (2003) and Ackerberg et al. (2015). Beyond showing that the results are robust to different methods of computing $MRPK$, we find that the results are similar when we include either 2-digit or 5-digit industry by year fixed effects or 4-digit industry by linear time trend controls in our regressions to account for differential time trends. Similarly, the results are robust to restricting the dataset to a balanced panel of firms to account for any differential attrition due to the policy.

Because reductions in distortion should reduce marginal cost for affected firms, we then explore if firms pass these gains on to consumers in the form of lower prices. Exploiting the fact that our panel of firm-level data provides detailed data on each firm's product-mix, as well as information about product-level prices, we find evidence of pass through on unit-prices in the treated industries. FDI deregulation led to lower prices for consumers, with the price effect concentrated in the firms that were initially high $MRPK$.

Motivated by the strong effects of the FDI liberalization on capital misallocation, we next consider whether the liberalization policies affected labor misallocation. We reclassify firms as having high or low marginal revenue products of labor ($MRPL$) and re-estimate the heterogeneous effects of the policies based on this classification. We again find that the reform had greater effects on firms with high $MRPL$, and in particular, that wage bills only increased for firms with above median pre-treatment $MRPL$. While wage bills increased by 30-38% for these firms, $MRPL$ fell by 29-38%. Since high $MRPL$ firms had 96%-145% higher levels of $MRPL$ prior to the treatment, this effect suggests that labor misallocation fell along with capital misallocation as a result of FDI liberalization.

Finally, combining our production function parameter estimates with our reduced-form estimates of the policy effect, we generate estimates of the aggregate effect of the 2001 and 2006 liberalizations on gross output. We find that the liberalizations, which affected 10% of manufacturing firms, raised the manufacturing sector's yearly gross output by 1%. Furthermore, to understand where the effects were the largest, we estimate the policies' effect on gross output at the 2-digit industry-level. We find that the largest effects are in the pharmaceutical industry, whose gross output increases by 24%. Importantly, our approach to identifying these aggregate effects does not rely on cross-sectional variation in the marginal returns to inputs. Instead, we are able to use the difference-in-differences framework to identify the policy's effects on inputs for different types of firms. Thus, observed changes in the dispersion to the returns to inputs or changes in inputs themselves due to changes in the production function, measurement error, or volatility will not affect our calculations as long as they are not correlated with the policy.

This paper contributes to three literatures. First, it contributes to a growing literature quantifying the importance of misallocation for aggregate outcomes (e.g. Restuccia and Rogerson (2008), Hsieh and Klenow (2009), Bartelsman et al. (2013), Restuccia and Rogerson (2013) and Baqaee and Farhi (2018)).[2] Much of this literature has focused on measuring the effect of all sources of misallocation on aggregate output by exploiting cross-sectional dispersion in productivity or firm/establishment size across industries or geographic areas in order to estimate wedges in output or input uses. The principal advantage of this "indirect approach" (Restuccia and Rogerson (2017) is that it allows for the estimation of the cost of misallocation without identifying the underlying source of the distortion and to include in the total cost of misallocation sources that are not necessarily observable by the researchers. However, this approach faces important challenges as it is subject to measurement error (Rotemberg and White, 2017), model misspecification (Haltiwanger et al., 2018), or variation in the volatility of productivity (Asker et al., 2014). In addition, this approach has focused on quantifying misallocation rather than identifying particular sources of distortion, which limits its policy implications (Syverson, 2011). We make two contributions to this literature. First, since we exploit a liberalization episode that affected only certain industries, we can estimate the effect of deregulation on misallocation using weaker identification assumptions. Our difference-in-difference estimation only requires that measurement error or other unobserved attributes are uncorrelated with the policy to identify *changes* in input wedges. Second, our approach isolates the distortions produced by a specific policy, the restriction on FDI, which allows us to isolate the effect of access to the foreign equity market, holding constant access to foreign debt market and other macroeconomic determinants that might affect the cost of capital.[3]

[2] A survey of this literature can be found in Restuccia and Rogerson (2017).

[3] In the context of India, several recent papers have aimed at estimating specific characteristics of the Indian economy that might explain the high degree of misallocation we observe in the country: the role of property rights and contract enforcement (Bloom et al. (2013) and Boehm and Oberfield (2018)); land regulation (Duranton et al.

Second, since we are studying the effect of a specific capital reform, the deregulation of FDI on misallocation and aggregate output, we also contribute to the literature on the effect of capital account liberalization on growth and misallocation. A recent strand of this literature has explored how increased foreign financial flows affect domestic firms' productivity and misallocation (Gopinath et al., 2017, Varela, 2017, Larrain and Stumpner, 2017, and Saffie et al., 2018).[4] Our paper contributes to this literature in several ways. First, while much of the previous literature exploits country-level variation in access to foreign investment, this paper exploits variation across industries over time within the same country. This allows us to hold institutional differences constant that may be important in cross-country comparisons. Second, since the Indian deregulation only affected FDI, it allows us to cleanly isolate the effect of foreign investment in *equity* on misallocation holding fixed access to foreign *debt*.[5]

Finally, more broadly, this paper contributes to the literature on the effect of financial frictions on misallocation (Buera et al., 2011, Midrigan and Xu, 2014, Midrigan and Xu, Moll, 2014, and Catherine et al., 2018). We add to this literature by estimating the effect of an exogenous shock to access to the foreign equity market and aggregating our estimates to measure effects at the macro-level.

The remainder of the paper is organized as follows. Section 2 describes the data and the context of the policy change. Section 3 reports our main empirical strategy, while Section 4 reports our estimates of the average effect of the FDI liberalization policy and its heterogeneous effects on high and low $MRPK$ firms. Section 5 replicates the analysis for high and low $MPRL$ firms to test whether the policy also reduced labor misallocation, and Section 6 investigates the drivers of the effects documented in Section 4 by estimating the correlations between firms' pre-treatment characteristics and whether they are capital constrained. Section 7 reports estimates of FDI liberalization policies' aggregate effects on gross output. Section 8 concludes.

(2017)); industrial licensing (Chari (2011) and Alfaro and Chari (2015)); privatisation DINC and GUPTA (DINC and GUPTA); highway infrastructure (Ghani et al. (2016)); and electricity shortage (Collard-wexler and Connell (2016)).

[4] Varela (2017) shows that financial liberalization can increase productivity, while Saffie et al. (2018) find that financial liberalization also accelerates the reallocation of resources across sectors, promoting the development of service/high-income sectors. On the other hand, Gopinath et al. (2017) find that better access to capital markets can amplify misallocation.

[5] By contrast for instance, Varela (2017) studies an episode in Hungary of deregulation of capital control, in a context where FDI were already integrated and did not move around the deregulation. Gopinath et al. (2017) exploit the drop in interest rate for Southern Europe countries following the adoption of the Euro that did not change the equity market.

2 Data & Policy Change

In this section, we first describe the FDI liberalization policy that we study and the process we used to collect data on that policy. We then provide details on Prowess, the firm-level panel data set that we combine with the policy data. Finally, we describe the restrictions imposed on the combined data set used in our analyses.

2.1 Indian FDI Liberalization

Following its independence, India became a closed, socialist economy, and most sectors where heavily regulated.[6] In 1991, India experienced a severe balance of payments crisis, and in June 1991, a new government was elected. Under pressure from the IMF, World Bank, and the Asian Development Bank, which offered funding, this government engaged in a series of structural reforms. These reforms in turn led India to become more open and free-market oriented. In addition to initiating FDI reforms in this period, India also liberalized trade (Topalova and Khandelwal, 2011 and Goldberg et al., 2010) and dismantled extensive licensing requirements (Aghion et al., 2008 and Chari, 2011).

Before 1991, most industries were regulated by the Foreign Exchange Regulation Act (1973), which required every instance of FDI to be individually approved by the government and foreign ownership rates were restricted to below 40% in most industries. With the establishment of the initial liberalization reforms in 1991, several industries experienced a liberalization of FDI whereby FDI was automatically approved up to 51% of equity under the automatic approval route.[7] In the following years, further liberalization of FDI occurred in different industries at different times and the cap for automatically approving foreign investment was increased.

To study the effects of allowing automatic FDI approval, we hand-collected data on the timing of industry-level FDI policy changes from different editions of the *Handbook of Industrial Policy and Statistics*. We match this data to industries at the 5-digit NIC level. We code an industry as having been treated by FDI liberalization if a policy change occurred that allowed automatic approval for investments up to at least 51% of capital (though, in some cases, the maximum is higher). We then merge this data at the industry-level with the firm-level dataset described below.

[6] See Panagariya (2008) for a thorough review of the Indian growth experience and government policies.

[7] This policy is described by Topalova (2007) and Sivadasan (2009).

2.2 Firm Data

Our firm-level data comes from the Prowess database compiled by the Centre for Monitoring the Indian Economy (CMIE). Unlike the Annual Survey of Industries (ASI), which is the other main source of information used to study dynamics in the Indian manufacturing sector, Prowess is a firm-level panel data set.[8] The data are therefore particularly well suited for examination of how firms adjust over time and how their responses may be related to policy changes. The dataset contains information from the income statements and balance sheets of listed companies comprising more than 70% of the economic activity in the organized industrial sector of India and account for 75% of all corporate taxes collected by the Government of India. The database is thus representative of large and medium-sized Indian firms. We retrieve yearly information about sales, capital stock, income from financial and non financial sources, consumption of raw materials and energy, compensation of employees, and ownership group for each firm in this data. In addition to the panel of firms, we also retrieve firm-product-year level data on unit sales and unit quantities sold from the Prowess database. This allows us to also construct a separate panel of product-level prices from 1995-2015.

2.3 Combined Data Sets

To arrive at our final data for analysis, we merge the firm-level panel data set and product-level panel data set with the industry-level policy data. We then make three restrictions to these samples, which we describe below.

First, as is common in the literature estimating production functions, we restrict our analysis to manufacturing firms. Second, we further restrict the sample to firms observed between 1995 and 2015. Restricting the sample to 1995-2015 has two advantages. First, focusing on this later period avoids potential bias from other liberalization reforms during the early-1990s – the main liberalization period. While liberalization occurred for 47% of manufacturing firms in the data, by restricting our sample to observations after 1995, we only exploit policy variation for the 10% of manufacturing firms who experienced FDI liberalization in the 2000s. Second, although the data set we use in our analyses, the Prowess data, technically starts in 1988, its coverage in the first few

[8] ASI is a plant-level repeated cross-section and does not include information on whether plants are owned by the same firm.

years of data collection is much more limited, and it grows substantially over the initial years. In 1988, Prowess only included 1,057 firms total, but it had grown to 7,061 firms by the beginning of our study period in 1995. In contrast, from 1995 onward, our study period, the coverage of the database is more stable, with similar numbers of firms observed across subsequent years (7,526 firms observed in 1996, 7,286 in 1997, and 7,717 in 1998). Appendix Table A1 provides a list of the different industries in the manufacturing sector affected by the deregulation during this restricted period. As the table shows, the only liberalization episodes occurred in 2001 and 2006.

Finally, we restrict the sample to the set of firms for whom we can compute marginal products of capital and labor ($MRPL$ and $MRPK$) prior to the earliest policy change in 2001. These pre-policy change measures are needed to estimate the effects of the policy on misallocation. Thus, we restrict the sample to firms observed before 2001 with non-missing, positive data on both assets and sales.[9] These three restrictions leave us with 62,194 observations.

Table 1 documents summary statistics for the final firm-level sample used in our analysis. As the table shows, classifying firms based on the owner's name, we find that the typical firm in our analysis is a privately-owned domestic firm (58%), while 4% of firms are private, foreign-owned firms, and 3% are state-owned. The table also shows that 10% of firms are in industries that went through FDI deregulation over the course of the sample.

3 Empirical Strategy

3.1 Classifying Firms as High or Low MRPK

We are interested in understanding whether FDI liberalization reduces misallocation by allowing firms with higher marginal revenue products of capital to grow faster. Thus, while we also estimate the effects of FDI liberalization on the average Indian manufacturing firm, we are particularly interested in estimating its heterogeneous effects on firms with high and low $MRPK$ prior to the policy changes. In this subsection, we document the three different methods we use to measure firms' $MRPK$, which we in turn use to assign firms a binary classification as high or low $MRPK$.

[9] This is the minimal requirement to calculate $MRPK$. As we document in the next subsection, we use three methods to estimate marginal revenue returns to capital. The least data intensive method exploits the fact, under Cobb-Douglas production functions, deflated sales divided by deflated capital will be proportional to $MRPK$ within an industry as long as capital intensity is the same for all firms in an industry.

As is standard in the production function estimation literature,[10] we assume that firms have Cobb-Douglas production functions, such that

$$Y_{ijt} = A_{ijt} K_{ijt}^{\alpha_j} L_{ijt}^{\beta_j} M_{ijt}^{\kappa_j}, \tag{1}$$

where i denotes a firm, j denotes a 2-digit industry, and t denotes a year, Y_{ijt}, K_{ijt}, L_{ijt}, and M_{ijt} are measures of output, assets, wage bill, and materials, and A_{ijt} is firm-specific unobserved productivity. In practice, we proxy these measures with deflated ruppee amounts, so that Y_{ijt} is proxied with deflated sales.[11] Our first and primary method for estimating $MRPK$ takes advantage of the fact that, under the Cobb-Douglas production function, $MRPK = \frac{\partial Y_{it}}{\partial K_{it}} = \alpha_j \frac{Y_{it}}{K_{it}}$. Thus, $\frac{Y_{it}}{K_{it}}$ provides a within-industry measure of $MRPK$, under the assumption that all firms in a 2-digit industry share the same α_j. This is our preferred method because it imposes the fewest data requirements, and therefore, allows us to use the largest sample for estimation.

As an alternative, we also use the methods of Levinsohn and Petrin (2003) (LP), using the GMM estimation proposed by Wooldridge (2009), and Ackerberg et al. (2015) (ACF) to estimate the full parameters of firms' production functions. LP assume the same Cobb-Douglas production function, as in equation (1), while ACF propose a value-added production function that only directly depends on K_{it} and L_{it}. Once we estimate the full parameters of these production functions, $MRPK$ is given by the derivative of the production function with respect to K_{it}. Both these methods rely on observing L_{it} and M_{it}.[12] Additionally, the methodology of ACF requires observing the same firm in multiple consecutive years. Thus, while our primary method imposes the least stringent data requirements to calculate $MRPK$, ACF imposes the most stringent. Using LP and ACF methods, we also estimate A_{it}, which we refer to as $TFPR$. By estimating the effect of the reforms on $TFPR$, we will be able to determine if FDI liberalization affected within-firm productivity with the caveat that changes in prices that are not captured by industry-level deflators will also affect $TFPR$.

[10] Duranton et al. (2017) describe the variety of methods used to estimate production functions and the revenue returns to capital and labor.

[11] We use deflators for India made available by Allcott et al. (2016). Revenue is deflated using three-digit commodity price deflators. The materials deflators are measures of the average output deflator of a given industry's suppliers using the 1993-4 input-output table. Finally, Allcott et al. (2016) calculate the capital deflator using an implied national deflator.

[12] Though M_{it} does not appear directly in ACF's value-added production function, it is key for identifying the parameters of the production function.

To determine whether firms had a high or low pre-reform $MRPK$, we average each firm's measures of $MRPK$ over time from 1995 to 2000 (prior to the first policy change). We then classify a firm as capital constrained (high $MRPK$) if it is above the industry median for the averaged measure. Since we have three measures of $MRPK$, this produces three measures of whether a firm is capital constrained or not.

Before turning to our main econometric specifications, we report the baseline levels of misallocation in the Indian manufacturing sector based on the dispersion of $MRPK$. However, we caution that dispersion in the cross-sectional distribution of $MRPK$ may be upwardly biased by measurement error or misspecified production functions. Figure 1 reports the distribution of $\log(MRPK)$ as measured using the LP methodology during 2000,[13] the last pre-policy year in the sample. In line with Hsieh and Klenow (2009), there appears to be substantial misallocation. A firm at the 90th percentile has a $\log(MRPK)$ 22 times greater than the MRPK of a firm at the 10th percentile.

3.2 Econometric Specification

In this section, we first describe the regressions we use to measure the effects of FDI liberalization on firm-level outcomes. We view these regressions, which allow us to directly measure whether the FDI liberalization policy reduced misallocation, as our main set of results. In addition to these key specifications, we also outline the analogous regressions used to measure the policies' effects on product-level prices. Since reducing capital wedges for high $MPRK$ firms would reduce those firms' marginal costs, we expect that prices may also be affected if misallocation falls. Thus, this secondary set of results provides us with an additional test of whether FDI liberalization affected misallocation.

Firm-level Outcomes

We first estimate the effect of FDI liberalization on the outcomes of the average treated firm i, in industry j at year t using a difference-in-differences strategy with the following regression:

$$y_{ijt} = \beta_1 \ Reform_{jt} + \mathbf{\Gamma X_{it}} + \alpha_i + \delta_t + \epsilon_{ijt}, \tag{2}$$

[13] Our primary measure only allows us to compare $MRPK$ within-industries, as opposed to across industries.

where i denotes a firm, j denotes an industry, t denotes a year, and y_{ijt} is the outcome variable of interest, consisting of the logs of capital, assets, income, the wage bill, sales, $TFPR$, and $MRPK$. $Reform_{jt}$ is an indicator variable equal to one if FDI has been liberalized in industry j, α_i are firm fixed effects, δ_t are year fixed effects, and $\mathbf{X_{it}}$ are firm-level controls. In our main specifications, $\mathbf{X_{it}}$ consists of firm age fixed effects. In this expression, our key variable of interest is β_1, which measures the effect of belonging to a liberalized industry. The fixed effects account for several important sources of variation in firms' outcomes that would otherwise bias the estimates. The firm fixed effects, α_i, absorb all time-invariant determinants of the outcome variable at the firm-level. Thus, if for example, more productive industries or firms are more likely to be liberalized, this will not bias the estimated effects of the policy. Time fixed effects, δ_t, ensure that we absorb any macro-economic fluctuations over time. So, for example, if the liberalization coincides with a boom period or a recession, this also will not bias the estimated effects. Thus, β_1 is only identified by comparing changes in outcomes for the liberalized firms between the pre- and post-periods to the non-liberalized firms. Finally, because our treatment of interest occurred at the industry level, we cluster our standard errors at the 4-digit industry-level to account for any serial correlation that might bias our standard errors downward.

While equation (2) estimates the average effect of the reform on firms' outcomes in treated industries, this may mask substantial heterogeneity within a given industry. In particular, in a country like India, where the dispersion in $MRPK$ is high, liberalizing FDI may differentially affect firms depending on their initial level of frictions in accessing capital. If firms who appear to be more capital constrained grow faster, this will ultimately reduce misallocation. Therefore, to estimate the effect of FDI on the reallocation of resources within industries, we estimate:

$$y_{ijt} = \beta_1 \; Reform_{jt} + \beta_2 Reform_{jt} \times I_i^{High\,MRPK} + \mathbf{\Gamma X_{it}} + \alpha_i + \delta_t + \epsilon_{ijt}, \tag{3}$$

where $I_i^{High\,MRPK}$ is an indicator variable equal to 1 if a firm had a high pre-treatment $MRPK$ according to our measures in Section 3.1. In this specification, β_2 captures the differential effect of FDI liberalization on ex-ante constrained firms relative to unconstrained firms in the same industry. If FDI reduces misallocation, we should expect $\beta_2 > 0$ for outcomes like assets and sales, since capital should flow disproportionably more to constrained firms, allowing them to expand faster.

If this is the case, we should expect $\beta_2 < 0$ when $log(MRPK)$ is the outcome, since dispersion in $MRPK$ should decline as constrained firms (i.e. firms with low ex-ante $MRPK$) receive more capital.

Just like equation (2), this specification controls for firm-level variation and macro-economic fluctuations over time. However, an additional strength of this specification is that it exploits within-industry variation in firms' pre-treatment $MRPK$'s to estimate the differential effect of $Reform_{jt}$ on high $MRPK$ firms. Thus, we can control for any industry-level time trends or shocks as long as they do not vary with pre-treatment $MRPK$. By controlling for $Reform_{jt}$ in our main specification, we control for both the overall impact of the FDI liberalization on firms' outcomes, irrespective of whether firms are constrained or not, and for potential sources of bias that are correlated at the industry-level with liberalization. Indeed, in our most conservative specifications, we control non-parametrically for industry-level shocks/time trends by including 5-digit industry by year fixed effects in the estimating equation. This empirical strategy accounts of any correlation between the policy and 5-digit level industry growth rates. Even if the Indian government liberalized industries that are growing more quickly earlier, this would not bias β_2 as long as high $MPRK$ firms were not growing differentially more quickly within these industries. The identifying assumptions for equation (3) are therefore milder than in the classic difference-in-differences framework (e.g. equation (2)), which would require that the liberalization policy was uncorrelated with industry-level time trends.

Our estimates could still be biased if high $MPRK$ firms in treated industries would have grown at a different rate than high $MPRK$ firms in untreated industries in the absence of the policy. This might occur if the Indian government targeted the policy toward industries where misallocation was already decreasing, although it is not clear why this would be the case. We can test for this source of bias directly by estimating and plotting the year-by-year relative treatment effect for high $MRPK$ firms in event study graphs. If the outcomes of high $MPRK$ firms were indeed changing faster in treated industries relative to in untreated industries prior to the policy change, we should see an effect of belonging to an industry that would be deregulated in the future on high $MPRK$ firms prior to the policy change. The dynamics of the FDI deregulation are obtained by estimating the following equation:

$$y_{ijt} = \sum_g \beta_{1,g}\ Reform_{jt} \times I^g_{it} + \sum_g \beta_{2,g} Reform_{jt} \times I_i^{High\,MRPK} \times I^g_{it} + \mathbf{\Gamma X_{it}} + \alpha_i + \delta_t + \epsilon_{ijt}, \tag{4}$$

where an industry's policy change is normalized to take place in period 0, and $\sum_g$ is a summation over the years that firms were observed before and after the policy event. I^g_{it} is an indicator variable equal to 1 if in year t a firm was observed g years after the policy event. Then, our event study graphs plot the set of coefficients $\beta_{2,g}$, which estimate the relative effect of being in a treated industry on a high $MPRK$ firm for each year from 7 years before the policy change to 7 years afterward. In the absence of differential pre-trends between capital constrained and unconstrained firms, we expect $\beta_{2,g}$ to be indistinguishable from 0 if $g < 0$ since belonging to a treated industry should have no differential effect on the high $MPRK$ firms prior to the treatment occurring.

Product-level Outcomes

We next adapt equations (2) and (3) to measure the effect of FDI liberalization on prices and its differential effect on the prices of high $MRPK$ firms. To measure the average effect on prices, we now estimate

$$log(price_{ipjt}) = \beta_1\ Reform_{jt} + \mathbf{\Gamma X_{it}} + \alpha_{ip} + \delta_t + \epsilon_{ipjt}, \tag{5}$$

with the additional subscript p denoting a product, and the fixed effect α_{ip} denoting a firm-by-product fixed effect. The remaining notation and terms are unchanged, with β_1 now capturing the effect of the policy on log(prices).

Similarly, the regression

$$log(price_{ipjt}) = \beta_1\ Reform_{jt} + \beta_2 Reform_{jt} \times I_i^{High\,MRPK} + \mathbf{\Gamma X_{it}} + \alpha_{ip} + \delta_t + \epsilon_{ipjt}, \tag{6}$$

captures the differential effect of the policies on prices for high $MRPK$ firms. As with the firm-level outcomes, δ_t accounts for any macroeconomic variation that could affect prices. α_{ip} accounts for any firm-level or product-level variation, such industries with cheaper products being more likely to be liberalized.

4 Results

Having outlined our empirical strategy, in this section, we estimate the effects of the FDI liberalization policies. We first document the average effects of the policy (estimates from equation (2)) on firm-level outcomes. Then, to estimate the effect on capital misallocation, we estimate the differential effects of the policy for high $MPRK$ firms on these outcomes. Next, we show that our estimates of the effect of the policy on misallocation are highly stable across a series of robustness test. Finally, we turn to our product-level data to estimate the effect of the policy on prices.

4.1 Average Effects for Firm-Level Outcomes

Table 2 reports the results from estimating (2). The estimates indicate that the liberalization policy had moderate, positive effects on the average firm's development. For the average firm, assets increased by 24%, sales increased by 22%, and income increased by 24%. We also observe a marginally significant average decline in $MRPK$ by 17%. On the other hand, there is no meaningful effect on the average firms' $TFPR$.[14] However, we caution that this identification strategy could underestimate gains in firm-level productivity, since $TFPR = TFPQ \times P$, where P is firm-level prices. If prices fell in response to the policy, $TFPR$ could fall or remain unchanged even if $TFPQ$ increased.

4.2 Differential Effects by MRPK for Firm-Level Outcomes

We next turn to our key results of interest – the heterogeneous effect of the policy on capital constrained and unconstrained firms. Table 3 reports the estimates from equation (3). Panel A uses our primary method for classifying whether firms are capital constrained, while Panel B reports the results using LP to estimate $MRPK$, and Panel C uses ACF.

Regardless of the method used to identify capital constrained firms, we find that the positive effect of the deregulation on firm expansion is mainly driven by its effects on capital constrained firms. Following the FDI deregulation, capital constrained firms become bigger relative to capital

[14] This result differs from the findings of Varela (2017), who shows that allowing domestic firms access to foreign capital in Hungary increased R&D and $TFPR$.

unconstrained firms, as measured by their total assets (+50-55% depending on how $I_i^{High\,MRPK}$ is defined). Consistent with the deregulation relaxing credit constraints, this expansion in firm size is partially made possible by an increase in capital (+18-30%). This higher investment does not crowd-out labor, as constrained firms also experience a relative increase in their wage bills by 25-33%.

The expansion of ex-ante constrained firms also appears to allow them to get closer to their optimal size. We observe a decrease in $MRPK$ by 40-56% for this group of firms following the deregulation. This effect indicates that the inflow of foreign capital reduces misallocation, since prior to the treatment, high $MRPK$ firms have between 136% (ACF classification) and 173% (LP classification) greater $MRPK$ than low $MRPK$ firms. Thus, the decline in $MRPK$ for the initially high $MRPK$ firms reduces the dispersion of $MPKR$, resulting in lower misallocation. While leading to substantial reallocation, the reform did not affect within-firm productivity as proxied by $TFPR$ (Columns (7) and (8)). This finding implies that during this more recent period, the FDI liberalization led to a substantial efficient reallocation within industries rather than an acceleration of productivity growth within firms (Sivadasan (2009), Bollard et al. (2013)).

Finally, to assess whether these results are driven by pre-trends, using our primary classification for high $MRPK$ firms, we plot the event study graphs described by equation (4) for our key outcomes of interest. Figure 2 reports these results for the logs of assets, sales, wage bill, and $MRPK$. Three facts are noteworthy. First, for all of these outcomes, being treated by the policy had no differential effect on high $MRPK$ firms before the policy was adopted, providing evidence that the results are not driven by pre-trends.

Second, the figure provides suggestive evidence that our results are not driven by mean reversion, which could be a source of bias, since high $MRPK$ firms could have high $MRPK$ in the pre-treatment for idiosyncratic reasons. If the results were driven by mean reversion, we would expect a decline in $MRPK$ to occur even before the liberalization policy went into effect. However, the figure shows that there is no relative decline in $MRPK$ prior to the initiation of the FDI liberalization.

Third, the effect of FDI on the different firm outcomes grows over time. This is reasonable, since the reallocation of resources (e.g. adjustment of worker flows, adaptation of production) is likely to take time. In addition, some of the reallocation we observe might also come from competitive effects where the relaxation of credit constraints allows firms with higher returns to capital to expand at the expense of the less efficient and ex-ante less constrained firms. We also expect this phenomenon to be progressive and to only be fully observable after some time has passed.

4.3 Robustness of Firm-level Results

In this subsection, we consider two remaining potential sources of bias for our estimates of the effects of the FDI liberalization policies on firm-level outcomes. First, we test whether differential time trends between treated and untreated industries could be driving our results. Second, we account for differential attrition by restricting our sample to a balanced panel.

Differential Time Trends

Differential time trends pose a threat to our estimation strategy. The key identifying assumption of our reduced-form misallocation results is that, in the absence of the FDI deregulation, the gap between constrained and unconstrained firms would evolve in the same way in deregulated and non-deregulated industries. Figure 2 already suggests that this is likely to be the case since the parallel trend assumption appears to hold in our setting. Furthermore, controlling for $Reform_{jt}$ already partially accounts for industry-level time trends that are correlated with the liberalization policy. Nonetheless, in this section, we further show that the results are robust to accounting for industry-specific time trends more flexibly.

Beyond showing that our reduced-form misallocation estimates are valid, we are also interested in estimating the total effects of the policy, including on low $MRPK$ firms. Doing so is necessary for estimating the aggregate effects of the policy on output in Section 7. The key identifying assumption for measuring the policy's total effects is that industries that were affected by the policy have the same time trends as industries that were not liberalized. Thus, in this subsection, we also provide further evidence that our results on the total effects of the policy are unlikely to be driven by differences in industry trends.

First, we include 2-digit industry by year fixed effects in equation (3). Including these fixed effects ensures that our differential effect is identified from comparing firms with different levels of $MRPK$ *within* the same 2-digit industry-year. Similarly, the total policy effect is identified by comparing firms that did and did not experience liberalization in the same 2-digit industry. This strategy effectively accounts for any sector-level shocks that may vary over time and also controls for any sector-level time trends. Appendix Table A2 reports the results including these controls and show that the magnitudes of the point estimates are the same as in Table 3. In Table A3 we also report results from an even more conservative specification, which includes 5-digit industry by year fixed effects to control for any time-varying unobserved characteristics at a very narrow industry level. This specification non-parametrically controls for any differential time trends between treated and non-treated industries. With these more stringent controls, the baseline effect of the program (captured by $Reform_{jt}$) is no longer identified since $Reform_{jt}$ is collinear with the controls, but the differential effect on the high $MRPK$ firms is. The estimates show that the differential effect of the policy on high $MRPK$ firms remains qualitatively and quantitatively similar despite these stringent controls.

Additionally, we can control for differential time trends by including linear time trends whose coefficients are allowed to vary by 4-digit industry in equation (3). Appendix Table A4 reports the new estimates including these additional controls. The qualitative pattern of the results is again very similar. The only departure from the results in Table 3 and Appendix Table A2 is that there are now negative effects on the low $MRPK$ firms' assets, leading to increases in $MRPK$ for these firms and further reductions in the dispersion of the marginal revenue returns to capital. One possible explanation for this result is that there are competitive spillovers from the high $MRPK$ firms to the low $MRPK$ firms. When high $MRPK$ firms grow and compete with low $MRPK$ firms, the low $MRPK$ firms shrink.

Attrition

Another concern for identification is that our results are affected by differential attrition between treated and untreated industries or between the two classifications of firms. To examine if this is the case, we instead re-estimate equation (3) using a balanced panel of firms who appear in both 1995 and 2015. Appendix Table A5 reports the results from this exercise. While the balanced samples are substantially smaller for all three classifications, the same pattern as before is evident. In all three cases, assets for high $MRPK$ firms increase by approximately 50% relative to low $MRPK$ firms, and for the Y/K and ACF classifications, the results are also similar for the remaining outcomes. The only deviation from the pattern in the non-balanced panel is that there is no longer a negative differential effect for $MRPK$ as an outcome in the ACF classification. However, we view the balanced ACF results as the least reliable, since imposing a balanced panel and using the ACF classification reduces the sample size to only 4,227 firm-year observations (relative to 21,665 for the Y/K classification and 12,678 for the LP classification).

4.4 Product-level Outcomes: Prices

We next consider whether FDI reforms reduced the prices faced by consumers. Most directly, if FDI liberalization reduced capital distortions for high $MRPK$ firms, these firms' marginal costs would fall, and lower marginal costs may be passed on to consumers in the form of lower prices. FDI liberalization will also lead to an overall increase in productivity at the industry-level by allowing more constrained firms to invest more, become bigger, and produce more output. Thus, the reform could also affect prices by increasing competitive pressure on firms and reducing mark-ups.

To estimate the effects of the reform on prices, we take advantage of one rare feature in firm-level datasets that is available in Prowess: the dataset reports both total product sales and total quantity sold at the firm-product level. Using this information, we can deduce the unit-quantity price of products that we define as total sales over total quantity. We then test if FDI deregulation affects the pricing strategy of treated firms by re-estimating equations (5) and (6) with log unit price as the outcome variable.

Table 4 reports the results of these regressions. On average, the reform reduces prices by a marginally significant 10% (column 1). We find that this reduction is concentrated among the high $MRPK$ firms, who reduced their prices by 15% (column 2) according to the Y/K classification of high $MRPK$. The alternative classifications of $MRPK$ yield similar patterns, although the ACF classification greatly reduces both the sample size and the precision of the estimates. This finding is consistent with high $MRPK$ firms reducing their prices when a reduction in capital wedges reduced their marginal costs.

5 Extension to Labor Misallocation

Our results so far point toward the existence of complementarity between capital and labor and show that FDI liberalization allowed firms not only to invest more (as seen by their higher amount of capital) but also to expand their wage bill.[15] This raises the question as to whether the FDI liberalization also reduced labor misallocation. To investigate if this is the case, we use the same estimation strategy as before but now compare the effects of the policy on firms with higher or lower marginal revenue products of labor ($MRPL$) prior to the policy change. We classify high and low $MRPL$ analogously to how we classify high and low $MRPK$.

Table 5 reports the results. We find some evidence that following the FDI deregulation, labor constrained firms grew faster as measured by their total assets (+25-30%). This expansion is particularly strong and statistically significant when we focus on the firm wage bill, with a relative increase of 30-40%. By allowing labor-constrained firms to grow faster, the FDI deregulation led to a reduction in misallocation, with a decrease in $MRPL$ by approximately 30% for ex-ante labor constrained firms.

6 Determinants of Capital Constraints

Our results so far suggest that FDI allows capital to be reallocated to firms with higher marginal revenue products of capital, reducing misallocation. This may be because domestic capital is trapped in firms with low $MRPK$ for institutional or political reasons. However, this does not necessarily imply that capital increased for high $MRPK$ firms specifically because they had high marginal returns to capital.

If, for instance, foreign capital systematically flows into the largest firms because of large asymmetric information (e.g. Detragiache et al. (2008)), and these firms have also higher value of $MRPK$, then FDI liberalization could reduce misallocation without implying that foreign capital specifically targets high $MRPK$ firms. Similarly, Gopinath et al. (2017) shows that foreign investment flows to firms with higher asset values. If these firms also have higher $MRPK$ in India, then foreign capital could reduce misallocation without being "smart" and identifying high $MRPK$ firms. Thus, to understand why FDI liberalization leads capital to flow to firms with greater $MRPK$, we next examine what pre-treatment characteristics of firms correlate with being high $MRPK$.

To do so, using the sample years prior to the first liberalization policy in 2001, we regress each

[15] Unfortunately, we only observe the total wage bill and cannot distinguish between existing workers being better paid and an increase in the number of workers.

firm characteristic on firm age fixed effects, 4-digit industry fixed effects, and an indicator variable for whether a firm was coded as high $MRPK$. Table 6 reports the results of this exercise. We find that, prior to the first reform, capital constrained firms had greater sales and were more productive (greater $TFPR$), but their assets measures are 17-50% lower. Thus, it is not the case that foreign capital merely flowed to firms who already had higher asset values. Instead, the results in Table 6 suggest that investors were able to identify the firms with the greater $MRPK$.

The results in Table 6 also provide evidence that reduction in misallocation from FDI liberalizatoin are not driven by pre-treatment misallocation due to the existence of a large, state-owned sector with especially low $MRPK$. There is no correlation (negative or positive) between being high $MRPK$ and being a state-owned enterprise. This is not entirely surprising since only 3% of the firms in the sample are state-owned.

7 Aggregate Effects

So far, our results provide robust evidence that FDI liberalization leads to a reduction in misallocation for treated industries relative to untreated industries. However, our difference-in-differences results only generate firm-level estimates of the policy's effect. To estimate the effect of the policy on total output, we need to aggregate these firm-level effects.

In this section, we show how to use our reduced-form estimates of the policies' firm-level effects to estimate FDI liberalization's effects on gross output. To do so, we first define gross output of industry I, $P_{G_I}G_I$ as

$$P_{G_I}G_I = \sum_i p_i y_i,$$

where i denotes a firm, P_{G_I} is the unit price-level, p_i is the deflated price of firm i's output and y_i is the quantity that firm i produces. Then, change in real output due to the policy, $\mathrm{d}\log G_I$, is defined as

$$\mathrm{d}\log G_I = \sum_{i \in I} \lambda_{Ii}^{G} \mathrm{d}\log y_i, \tag{7}$$

where the market share of firm i in its industry λ_{Ii}^G is given by:

$$\lambda_{Ii}^G = \frac{p_i y_i}{\sum_{j \in I} p_j y_j}. \tag{8}$$

A first order approximation of $\mathrm{d}\log y_i$ yields

$$\mathrm{d}\log y_i = \alpha_{im}\mathrm{d}\log m_i + \alpha_{il}\mathrm{d}\log l_i + \alpha_{ik}\mathrm{d}\log k_i \tag{9}$$

Note first that λ_{Ii}^G, each firm's market share, is a function of firms' sales and is observed in the Prowess data. Thus, if equation (9) is estimable, we can approximate the effect of the policy to the first order with equation (7).

To estimate equation (9), we first recognize that, if we assume a Cobb-Douglas production function, α_{im}, α_{il}, and α_{ik} are its parameters. Thus, we can estimate each parameter using the methods of Levinsohn and Petrin (2003) and substitutes these estimated parameters into equation (9). Second, the changes in capital, labor, and materials due to the policy, $\mathrm{d}\log k_i$, $\mathrm{d}\log l_i$, and $\mathrm{d}\log m_i$ can be estimated using our difference-in-differences estimation strategy.

To estimate $\mathrm{d}\log k_i$, we regress

$$\begin{aligned}\log(k_{ijt}) =& \beta_1 Reform_{jt} + \beta_2 Reform_{jt} \times I_i^{High\,MRPK} + \beta_3 Reform_{jt} \times I_i^{High\,MRPL} \\ &+ \beta_4 Reform_{jt} \times I_i^{High\,MRPM} + \mathbf{\Gamma X_{it}} + \alpha_i + \delta_t + \epsilon_{ijt},\end{aligned}$$

where $I_i^{High\,MRPM}$ is an indicator variable equal to 1 if the marginal revenue product of materials according to the LP production function estimation is above the industry-level median in the pre-treatment period. All the remaining covariates are defined in the same way as in Section 3.2. Then, we estimate the change in $\log k_i$ due to the policy, $\widehat{\mathrm{d}\log k_i}$, with

$$\begin{aligned}\widehat{\mathrm{d}\log k_i} =& \widehat{\beta}_1 Reform_j + \widehat{\beta}_2 Reform_j \times I_i^{High\,MRPK} + \widehat{\beta}_3 Reform_j \times I_i^{High\,MRPL} \\ &+ \widehat{\beta}_4 Reform_j \times I_i^{High\,MRPM},\end{aligned}$$

where $Reform_j$ is an indicator variable equal to 1 if an industry is liberalized between 1995 and 2015. We use an analogous strategy to estimate $\mathrm{d}\log l_i$ and $\mathrm{d}\log m_i$. Appendix Table A6 reports the results of these regressions.

Using these estimates and using sales data from 2000 (the last pre-treatment year) to predict λ^G_{Ii}, we define I as the manufacturing sector and approximate $\mathrm{d}\log G_I$. We find that the two liberalization episodes that we observe (in 2001 and 2006) increased manufacturing's yearly gross output by 1%. These liberalization episodes affected about 10% of the manufacturing industry, while 39% of the industry was liberalized in 1991 and 50% of the industries has not yet been liberalized. Thus, we expect that both the initial liberalization and liberalizing the remaining industries could have larger effects.

To see which sectors drive the increase in gross output, we can also re-define I to denote a 2-digit industry. We then estimate the effects of the policy on total gross output at the 2-digit industry-level using the same strategy as before. Table 7 reports the effects of FDI liberalization on gross output by industry for the 2-digit industries that experienced a deregulation during our sample period. Our estimates suggest that the policy had the largest effects in the pharmaceutical industry, with gross output increasing by 24%. In contrast, the effects were substantially smaller in tobacco and food products, which increased by only 0.2 and 1% respectively.

8 Conclusion

This paper provides new evidence, exploiting within-country, cross-industry and cross-time variation, that access to FDI can reduce the misallocation of capital and labor in a low-income country. We show that reforms that automatically approved FDI investments in the 2000's increased capital investment in the affected industries. However, the effects of the reforms on the average firm mask important heterogeneity in the policies' effects. The entirety of the FDI liberalization's effect on firms' outcomes is driven by increased investment in firms that previously had high marginal revenue products of capital. Thus, the inflow of foreign investment reduces the marginal revenue returns to capital for these firms, reducing the dispersion of $MRPK$. Aggregating our reduced-form estimates, we find that the policy increased manufacturing's total gross output by 1%. Thus, we make two important contributions to the misallocation literature. First, we identify a policy lever that can be used to reduce misallocation. Second, we exploit a natural experiment to arrive at measures of the aggregate effect of reducing misallocation that do not rely on cross-sectional variation.

Furthermore, we provide evidence that the reduction is misallocation is not merely because capital flows to the firms with the largest pre-existing capital stock. Instead, the firms most affected by the reform have higher productivity and marginal revenue returns to capital but smaller pre-treatment capital stocks. Thus, investors seem to be able to identify firms with high marginal revenue products of capital. Taken together, these results suggest that FDI can play an important role in equalizing the revenue returns to capital across firms in low-income countries, where the domestic financial sector may be unable to reallocate capital for political or institutional reasons.

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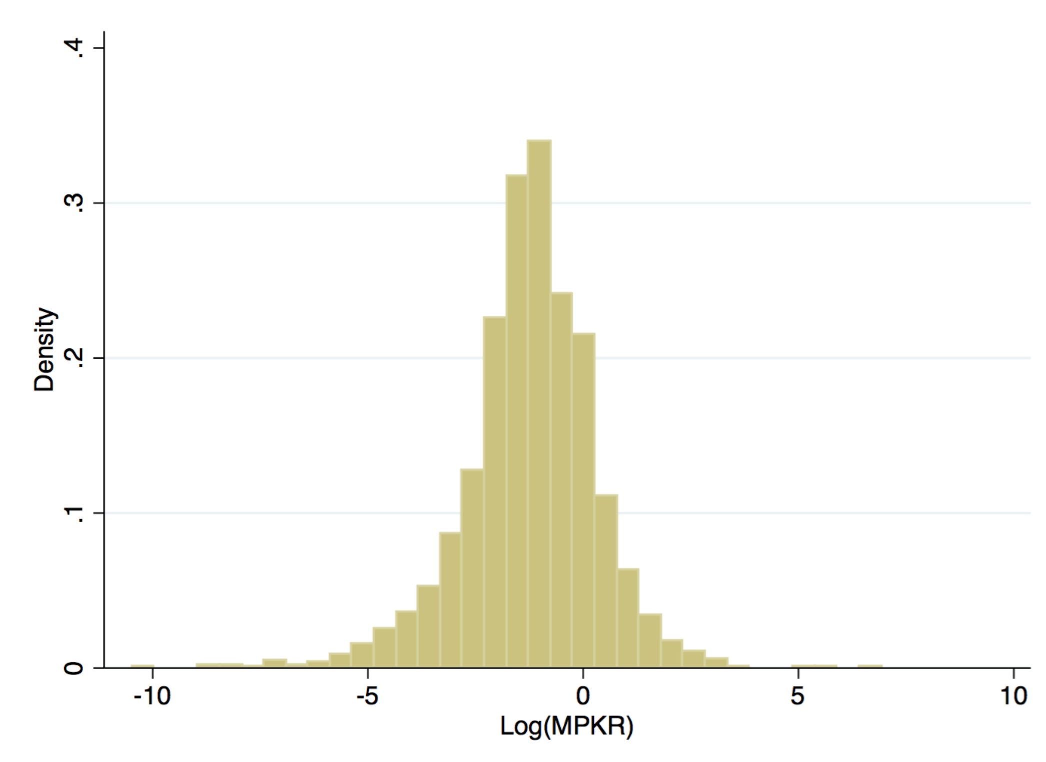
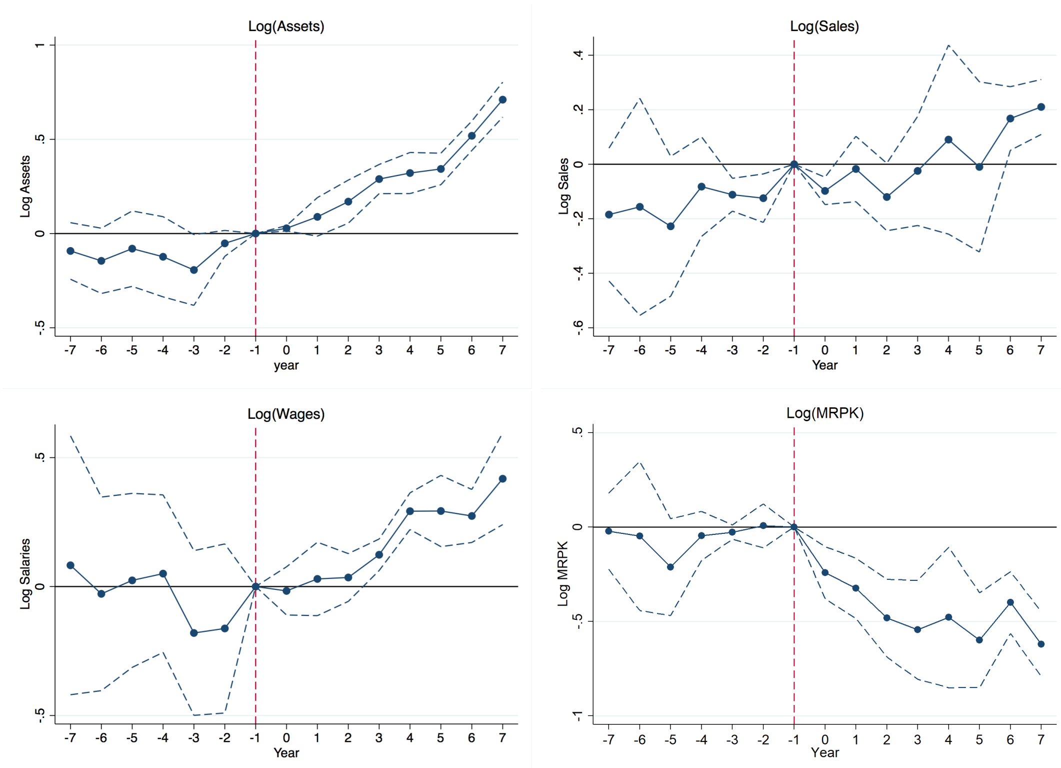
Figures

Figure 1: Distribution of Log(MRPK) in 2000



This figure displays the distribution of log(MRPK) for manufacturing firms in the Prowess data in 2000, the year before the first FDI liberalization episode in 2001. MRPK is computed using the methodology of Levinsohn and Petrin (2003).

Figure 2: Event Study Graphs for the Relative Effect of FDI Liberalization on High $MRPK$ Firms



This figure reports event study graphs for the relative effects of FDI liberalization on firms with high pre-treatment $MRPK$. $MRPK$ is calculated using Y/K as a within-industry proxy for $MRPK$.

Tables

Table 1: Summary Statistics

	Obs.	Mean	p10	p50	p90
Treated during Study Period (%)	62,194	10	0	0	0
Private, Domestic (%)	62,194	58	0	100	100
Private, Foreign (%)	62,194	4	0	0	0
State Owned (%)	62,194	3	0	0	0
Firm Age	62,194	26	8	21	51
Capital	60,907	6	0	1	8
Assets	60,939	79	1	9	114
Sales	58,478	68	1	10	119
Wages	45,821	4	0	1	7
Income	59,790	67	1	10	113

This table reports summary statistics for the manufacturing firms appearing in the CMIE Prowess dataset from 1995 to 2015. An observation is at the firm-year level. Firms' capital, assets, sales, salaries, and income are measured in millions of USD.

Table 2: Effect of FDI Liberalization on the Average Firm in the Prowess Data

	(1) **Log Assets**	(2) **Log Accounting Capital**	(3) **Log Sales**	(4) **Log Income**	(5) **Log Wages**	(6) **Log MRPK YK**	(7) **Log TFPR LP**	(8) **Log TFPR ACF**
$Reform_{jt}$	0.066	0.237**	0.223***	0.238***	0.144	-0.174*	-0.047	-0.181
	(0.051)	(0.108)	(0.079)	(0.077)	(0.101)	(0.106)	(0.057)	(0.143)
Firm Age FE	Y	Y	Y	Y	Y	Y	Y	Y
Firm FE	Y	Y	Y	Y	Y	Y	Y	Y
Year FE	Y	Y	Y	Y	Y	Y	Y	Y
Number of observations	60,790	60,823	58,335	59,662	45,716	49,016	28,666	26,658
Clusters	96	96	96	96	96	96	70	96
Adjusted R^2	0.896	0.850	0.803	0.784	0.887	0.653	0.717	0.587

This table reports difference-in-differences estimates of the effect of the FDI liberalization reforms in the Prowess dataset (equation (2)). Firms are observed between 1995 and 2015. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table 3: Heterogeneous Effects of FDI Liberalization on Firms by Pre-Treatment $MRPK$

	(1) **Log Assets**	(2) **Log Accounting Capital**	(3) **Log Sales**	(4) **Log Income**	(5) **Log Wages**	(6) **Log MRPK**	(7) **Log TFPR LP**	(8) **Log TFPR ACF**
				Panel A: Y/K Classification				
$Reform_{jt}$	-0.032	-0.058	0.118	0.145*	0.003	0.047	-0.033	-0.095
	(0.074)	(0.047)	(0.084)	(0.087)	(0.082)	(0.119)	(0.087)	(0.188)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.498***	0.228***	0.190***	0.170***	0.260**	-0.398***	-0.019	-0.147
	(0.105)	(0.041)	(0.048)	(0.045)	(0.107)	(0.075)	(0.061)	(0.124)
Number of observations	60,833	60,800	58,313	59,632	45,727	48,999	28,674	26,937
Clusters	97	97	97	97	97	97	71	97
Adjusted R^2	0.851	0.896	0.803	0.784	0.888	0.654	0.717	0.587
				Panel B: LP Classification				
$Reform_{jt}$	0.042	0.007	0.033	0.093	0.017	0.133	0.001	-0.017
	(0.068)	(0.046)	(0.106)	(0.092)	(0.084)	(0.121)	(0.097)	(0.187)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.492***	0.173***	0.307***	0.273***	0.329***	-0.474***	-0.086	-0.121
	(0.139)	(0.038)	(0.102)	(0.077)	(0.124)	(0.104)	(0.080)	(0.142)
Number of observations 43,697	43,657	42,698	43,340	34,961	27,095	27,095	21,721	
Clusters	80	80	80	80	80	80	80	80
Adjusted R^2	0.852	0.895	0.805	0.788	0.884	0.995	0.708	0.574
				Panel B: ACF Classification				
$Reform_{jt}$	0.039	-0.011	0.274**	0.268**	0.148	-0.042	-0.005	-0.160
	(0.129)	(0.054)	(0.135)	(0.123)	(0.119)	(0.160)	(0.104)	(0.175)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.525***	0.282***	0.156	0.210	0.346***	-0.356**	-0.142	-0.119
	(0.164)	(0.066)	(0.156)	(0.137)	(0.084)	(0.139)	(0.098)	(0.123)
Number of observations	17,778	17,794	17,177	17,504	15,225	9,234	10,172	9,234
Clusters	82	82	82	82	82	81	60	81
Adjusted R^2	0.838	0.883	0.793	0.779	0.867	0.746	0.674	0.551

This table reports estimates of the heterogeneous effects of FDI liberalization reforms on high and low pre-treatment $MRPK$ firms in the Prowess dataset (equation (3)). Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, and firm age fixed effects. Firms are classified as high $MRPK$ if their average $MRPK$ in the pre-treatment period from 1995-2000 is above the industry median. In Panel A, $MRPK$ is approximated as Y/K. In Panel B, it is calculated by estimating the production function using LP. In Panel C, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table 4: Price Effects of the FDI Reform

	(1)	(2)	(3)	(4)
	Dep Var: Log Output Price			
$Reform_{jt}$	-0.098*	0.007	-0.029	-0.191
	(0.055)	(0.032)	(0.038)	(0.147)
$Reform_{jt} \times I_i^{High\,MRPK}$ (Y/L)		-0.157**		
		(0.072)		
$Reform_{jt} \times I_i^{High\,MRPK}$ (LP)			-0.107**	
			(0.047)	
$Reform_{jt} \times I_i^{High\,MRPK}$ (ACF)				-0.074
				(0.074)
Number of observations	182,674	182,674	127,777	50,539
Clusters	112	112	80	92
Adjusted R^2	0.956	0.956	0.958	0.943

This table reports estimates of the effect of FDI liberalization reforms on unit prices in the Prowess dataset (equation (3)). Each observation is at the firm-product-year level. Firms are observed between 1995 and 2015. All regressions include firm by product fixed effects, survey year fixed effects, and firm age fixed effects. Firms are classified as high $MRPK$ if their average $MRPK$ in the pre-treatment period from 1995-2000 is above the industry median. In column 2, $MRPK$ is approximated as Y/K. In column 3, it is calculated by estimating the production function using LP. In column 4, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table 5: Heterogeneous Effects of FDI Liberalization on High and Low $MRPL$ Firms

	(1) **Log Assets**	(2) **Log Accounting Capital**	(3) **Log Sales**	(4) **Log Income**	(5) **Log Wages**	(6) **Log MRPL**	(7) **Log TFPR LP**	(8) **Log TFPR ACF**
				Panel A: Y/L Classification				
$Reform_{jt}$	0.187***	0.051	0.134*	0.195***	0.023	0.138	0.061	-0.104
	(0.054)	(0.058)	(0.069)	(0.062)	(0.109)	(0.102)	(0.069)	(0.194)
$Reform_{jt} \times I_i^{High\,MRPL}$	0.225	0.091***	0.167	0.145	0.299***	-0.317***	-0.205***	-0.146
	(0.147)	(0.033)	(0.106)	(0.099)	(0.072)	(0.081)	(0.063)	(0.105)
Number of observations	50,663	50,633	49,129	50,069	40,413	34,528	25,667	23,906
Clusters	93	93	93	93	93	93	67	92
Adjusted R^2	0.850	0.893	0.803	0.788	0.883	0.752	0.719	0.583
				Panel B: LP Classification				
$Reform_{jt}$	0.180***	0.049	0.099	0.148**	0.018	0.080	0.043	-0.014
	(0.060)	(0.054)	(0.078)	(0.065)	(0.118)	(0.124)	(0.070)	(0.163)
$Reform_{jt} \times I_i^{High\,MRPL}$	0.264***	0.104***	0.216***	0.199**	0.358***	-0.289***	-0.183***	-0.131
	(0.101)	(0.026)	(0.073)	(0.079)	(0.077)	(0.111)	(0.064)	(0.104)
Number of observations	43,735	43,694	42,748	43,387	35,000	27,129	27,129	21,760
Clusters	79	79	79	79	79	79	79	79
Adjusted R^2	0.851	0.895	0.805	0.788	0.884	0.812	0.708	0.574
				Panel C: ACF Classification				
$Reform_{jt}$	0.167	0.127**	0.370***	0.382***	0.154	0.034	0.102*	-0.038
	(0.140)	(0.055)	(0.089)	(0.086)	(0.160)	(0.161)	(0.061)	(0.152)
$Reform_{jt} \times I_i^{High\,MRPL}$	0.309*	0.014	-0.001	0.022	0.381***	-0.377***	-0.379***	-0.355***
	(0.160)	(0.066)	(0.083)	(0.123)	(0.142)	(0.097)	(0.033)	(0.080)
Number of observations	17,688	17,706	17,083	17,411	15,140	9,183	10,123	9,183
Clusters	81	81	81	81	81	80	59	80
Adjusted R^2	0.837	0.880	0.791	0.778	0.866	0.735	0.673	0.553

This table reports estimates of the heterogeneous effects of FDI liberalization reforms on high and low $MRPL$ firms in the Prowess dataset. Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, and firm age fixed effects. Firms are classified as high $MRPL$ if their average $MRPL$ in the pre-treatment period from 1995-2000 is above the industry median. In Panel A, $MRPL$ is approximated as Y/L. In Panel B, it is calculated by estimating the production function using LP. In Panel C, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table 6: Characteristics of High $MRPK$ Firms

	(1) Log Sales	(2) Log MRPK	(3) Log Assets	(4) TFPR LP	(5) TFPR ACF	(6) Age	(7) State Firm	(8) Private Indian	(9) Private Foreign
				Panel A: Y/K Classification					
$I_i^{High\ MRPK}$	0.488***	1.416***	-0.355***	0.251***	0.456***	7.452***	0.002	0.034**	0.022***
	(0.050)	(0.025)	(0.046)	(0.026)	(0.039)	(0.579)	(0.006)	(0.016)	(0.007)
Number of observations	29,161	27,435	28,778	4,893	1,816	30,331	30,331	30,331	30,331
Clusters	4,628	4,628	4,628	2,770	1,212	4,628	4,628	4,628	4,628
Adjusted R^2	0.280	0.432	0.231	0.287	0.356	0.156	0.103	0.152	0.076
				Panel B: LP Classification					
$I_i^{High\ MRPK}$	0.210***	1.731***	-0.516***	0.295***	0.471***	5.176***	-0.000	0.080***	0.017**
	(0.059)	(0.038)	(0.058)	(0.027)	(0.037)	(0.760)	(0.008)	(0.020)	(0.009)
Number of observations	18,050	4,848	17,732	4,848	1,554	18,583	18,583	18,583	18,583
Clusters	2,737	2,737	2,737	2,737	1,036	2,737	2,737	2,737	2,737
Adjusted R^2	0.287	0.651	0.238	0.298	0.349	0.118	0.118	0.152	0.082
				Panel C: ACF Classification					
$I_i^{High\ MRPK}$	0.480***	1.317***	-0.171**	0.274***	0.522***	7.404***	0.012	-0.023	0.006
	(0.087)	(0.042)	(0.084)	(0.037)	(0.038)	(1.007)	(0.012)	(0.030)	(0.012)
Number of observations	7,910	1,759	7,751	3,062	1,759	8,117	8,117	8,117	8,117
Clusters	1,284	1,161	1,284	1,087	1,161	1,284	1,284	1,284	1,284
Adjusted R^2	0.326	0.649	0.292	0.303	0.375	0.180	0.140	0.215	0.139

This table reports correlations between whether a firm is classified as high $MRPK$ and the firm's pre-treatment characteristics. The sample consists of firm-years between 1995 and 2000. All regressions include firm age fixed effects (unless the outcome is firm age), 4-digit industry fixed effects, and year fixed effects. Standard errors are clustered at the firm-level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table 7: Disaggregate Effects on Gross Output by 2-Digit Industry

Industry	Change in Gross Output
Manufacture of Pharmaceuticals	24%
Manufacture of Beverages	11%
Manufacture of Rubber and Plastics	8%
Manufacture of Tobacco	1%
Manufacture of Food Products	0.2%

This table reports the estimates of the effect of the FDI liberalizations in 2001 and 2006 on gross output at the 2-digit industry-level. The estimates are generated using the Prowess data set.

Appendix Tables

Table A1: List of Industries that Changed FDI Polices Between 1995 and 2015

(1) **NIC 5-Digit Industry Classification**	(2) **Reform Year**
Manufacture of 'ayurvedic' or 'unani' pharmaceutical preparation	2001
Manufacture of allopathic pharmaceutical preparations	2001
Manufacture of medical impregnated wadding, gauze, bandages, dressings, surgical gut string etc.	2001
Manufacture of homoeopathic or biochemic pharmaceutical preparations	2001
Manufacture of other pharmaceutical and botanical products n.e.c. like hina powder etc.	2001
Manufacture of rubber tyres and tubes n.e.c.	2006
Manufacture of essential oils; modification by chemical processes of oils and fats (e.g. by oxidation, polymerization etc.)	2006
Manufacture of various other chemical products	2006
Manufacture of rubber tyres and tubes for cycles and cycle-rickshaws	2006
Manufacture of distilled, potable, alcoholic beverages such as whisky, brandy, gin, 'mixed drinks' etc.	2006
Coffee curing, roasting, grinding blending etc. and manufacturing of coffee products	2006
Retreading of tyres; replacing or rebuilding of tread on used pneumatic tyres	2006
Manufacture of chemical elements and compounds doped for use in electronics	2006
Manufacture of country liquor	2006
Manufacture of matches	2006
Manufacture of rubber plates, sheets, strips, rods, tubes, pipes, hoses and profile -shapes etc.	2006
Distilling, rectifying and blending of spirits	2006
Manufacture of bidi	2006
Manufacture of catechu(katha) and chewing lime	2006
Stemming and redrying of tobacco	2006
Manufacture of other rubber products n.e.c.	2006
Manufacture of rubber contraceptives	2006
Manufacture of other tobacco products including chewing tobacco n.e.c.	2006
Manufacture of pan masala and related products.	2006

This table lists 5-digit NIC industries that changed to automatic FDI approval for investments up to (at least) 51% of a firm's capital and the year that the policy reform took place.

Table A2: Robustness of Heterogeneous Effects of FDI Liberalization to Inclusion of 2-Digit Industry by Year FE

	(1) Log Assets	(2) Log Accounting Capital	(3) Log Sales	(4) Log Income	(5) Log Wages	(6) Log MRPK	(7) Log TFPR LP	(8) Log TFPR ACF
				Panel A: Y/K Classification				
$Reform_{jt}$	-0.181	-0.089***	-0.005	0.016	-0.114	0.173	0.013	-0.076
	(0.133)	(0.033)	(0.093)	(0.111)	(0.123)	(0.112)	(0.088)	(0.181)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.515***	0.237***	0.203***	0.187***	0.287**	-0.369***	-0.034	-0.138
	(0.104)	(0.049)	(0.049)	(0.046)	(0.114)	(0.082)	(0.054)	(0.114)
Number of observations	60,935	60,907	58,447	59,750	45,824	49,153	28,803	27,305
Clusters	97	97	97	97	97	97	71	97
Adjusted R^2	0.855	0.897	0.807	0.788	0.890	0.661	0.728	0.621
				Panel B: LP Classification				
$Reform_{jt}$	-0.029	-0.039	-0.122	-0.064	-0.122	0.129	0.029	0.052
	(0.143)	(0.071)	(0.180)	(0.158)	(0.105)	(0.095)	(0.095)	(0.138)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.519***	0.185***	0.343***	0.313***	0.345***	-0.462***	-0.079	-0.088
	(0.139)	(0.033)	(0.108)	(0.087)	(0.127)	(0.098)	(0.070)	(0.123)
Number of observations	43,718	43,679	42,722	43,361	35,012	27,190	27,190	21,996
Clusters	80	80	80	80	80	80	80	80
Adjusted R^2	0.858	0.896	0.811	0.794	0.888	0.995	0.719	0.613
				Panel C: ACF Classification				
$Reform_{jt}$	0.147	0.117	0.273	0.254	0.121	-0.158	0.077	-0.123
	(0.178)	(0.082)	(0.202)	(0.206)	(0.123)	(0.251)	(0.111)	(0.206)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.563***	0.304***	0.159	0.204	0.349***	-0.306**	-0.152	-0.101
	(0.155)	(0.065)	(0.165)	(0.155)	(0.086)	(0.134)	(0.094)	(0.103)
Number of observations	17,790	17,806	17,193	17,517	15,260	9,409	10,207	9,409
Clusters	82	82	82	82	82	81	60	81
Adjusted R^2	0.843	0.886	0.796	0.782	0.869	0.761	0.684	0.586

This table reports estimates of the heterogeneous effects of FDI liberalization reforms on high and low $MRPK$ firms in the Prowess data set (equation (3)). Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, firm age fixed effects, and 2-digit industry by year fixed effects. Firms are classified as constrained if their average $MRPK$ in the pre-treatment period from 1995-2000 is above the industry median. In Panel A, $MRPK$ is approximated as Y/K. In Panel B, it is calculated by estimating the production function using LP. In Panel C, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table A3: Robustness of Heterogeneous Effects of FDI Liberalization to Inclusion of 5-Digit Industry by Year FE

	(1) **Log Assets**	(2) **Log Accounting Capital**	(3) **Log Sales**	(4) **Log Income**	(5) **Log Wages**	(6) **Log MRPK**	(7) **Log TFPR LP**	(8) **Log TFPR ACF**
				Panel A: Y/K Classification				
$Reform_{jt} \times I_i^{High\,MRPK}$	0.645***	0.308***	0.338***	0.307***	0.433***	-0.351***	-0.050	-0.028
	(0.084)	(0.048)	(0.023)	(0.045)	(0.099)	(0.083)	(0.044)	(0.127)
Number of observations	59,527	59,494	57,020	58,367	44,335	47,992	27,851	25,624
Clusters	94	94	94	94	92	94	66	90
Adjusted R^2	0.857	0.897	0.810	0.791	0.893	0.665	0.727	0.648
				Panel B: LP Classification				
$Reform_{jt} \times I_i^{High\,MRPK}$	0.627***	0.259***	0.516***	0.452***	0.461***	-0.416***	-0.087	-0.083
	(0.120)	(0.026)	(0.068)	(0.045)	(0.099)	(0.088)	(0.054)	(0.115)
Number of observations	42,511	42,466	41,508	42,144	33,751	26,123	26,123	20,800
Clusters	78	78	78	78	78	78	78	73
Adjusted R^2	0.860	0.896	0.813	0.795	0.891	0.995	0.725	0.626
				Panel C: ACF Classification				
$Reform_{jt} \times I_i^{High\,MRPK}$	0.722***	0.383***	0.381***	0.367***	0.501***	-0.392***	-0.226***	-0.100
	(0.097)	(0.043)	(0.102)	(0.103)	(0.053)	(0.142)	(0.084)	(0.113)
Number of observations	16,182	16,178	15,568	15,896	13,660	8,052	9,247	8,052
Clusters	75	75	75	75	73	65	55	65
Adjusted R^2	0.844	0.885	0.798	0.779	0.871	0.767	0.696	0.612

This table reports estimates of the heterogeneous effects of FDI liberalization reforms on high $MRPK$ firms in the Prowess data set (equation (3)). Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, firm age fixed effects, and 5-digit industry by year fixed effects. Firms are classified as constrained if their average $MRPK$ in the pre-treatment period from 1995-2000 is above the industry median. In Panel A, $MRPK$ is approximated as Y/K. In Panel B, it is calculated by estimating the production function using LP. In Panel C, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table A4: Robustness of Heterogeneous Effects of FDI Liberalization to Inclusion of 4-Digit Industry-Specific Linear Trends

	(1) Log Assets	(2) Log Accounting Capital	(3) Log Sales	(4) Log Income	(5) Log Wages	(6) Log MRPK	(7) Log TFPR LP	(8) Log TFPR ACF
				Panel A: Y/K Classification				
$Reform_{jt}$	-0.356***	-0.156***	-0.112	-0.089	-0.150*	0.134***	-0.035	-0.190
	(0.061)	(0.029)	(0.071)	(0.074)	(0.075)	(0.046)	(0.071)	(0.142)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.536***	0.246***	0.220***	0.214***	0.277**	-0.377***	-0.024	-0.057
	(0.098)	(0.053)	(0.042)	(0.042)	(0.113)	(0.080)	(0.057)	(0.135)
Number of observations	60,935	60,907	58,478	59,789	45,818	49,179	28,803	27,036
Clusters	96	96	96	96	96	96	70	96
Adjusted R^2	0.858	0.898	0.809	0.790	0.892	0.664	0.728	0.626
				Panel B: LP Classification				
$Reform_{jt}$	-0.249***	-0.141***	-0.128*	-0.115	-0.116***	0.167***	-0.012	-0.148
	(0.048)	(0.019)	(0.074)	(0.079)	(0.043)	(0.064)	(0.076)	(0.118)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.527***	0.207***	0.363***	0.326***	0.328**	-0.451***	-0.062	-0.069
	(0.139)	(0.029)	(0.092)	(0.075)	(0.132)	(0.101)	(0.070)	(0.126)
Number of observations	43,718	43,679	42,722	43,361	35,012	27,190	27,190	21,996
Clusters	80	80	80	80	80	80	80	80
Adjusted R^2	0.860	0.898	0.811	0.795	0.889	0.995	0.720	0.616
				Panel C: ACF Classification				
$Reform_{jt}$	-0.166	-0.115*	0.049	0.024	0.011	-0.165	0.001	-0.243*
	(0.111)	(0.061)	(0.119)	(0.130)	(0.125)	(0.142)	(0.102)	(0.131)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.527***	0.321***	0.149	0.195	0.293***	-0.343**	-0.131	-0.091
	(0.182)	(0.067)	(0.170)	(0.156)	(0.105)	(0.139)	(0.099)	(0.101)
Number of observations	17,790	17,806	17,193	17,517	15,260	9,409	10,207	9,409
Clusters	82	82	82	82	82	81	60	81
Adjusted R^2	0.852	0.891	0.805	0.789	0.875	0.768	0.688	0.596

This table reports estimates of the heterogeneous effects of FDI liberalization reforms on high and low $MRPK$ firms in the Prowess data set (equation (3)). Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, firm age fixed effects, and linear time trends whose coefficients are allowed to vary at the 4-digit industry-level. Firms are classified as constrained if their average $MPKR$ in the pre-treatment period from 1995-2000 is above the industry median. In Panel A, $MRPK$ is approximated as Y/K. In Panel B, it is calculated by estimating the production function using LP. In Panel C, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table A5: Robustness of Heterogeneous Effects of FDI Liberalization to Using a Balanced Panel of Firms

	(1) **Log Assets**	(2) **Log Accounting Capital**	(3) **Log Sales**	(4) **Log Income**	(5) **Log Wages**	(6) **Log MRPK**	(7) **Log TFPR LP**	(8) **Log TFPR ACF**
				Panel A: Y/K Classification				
$Reform_{jt}$	-0.078	-0.005	0.054	0.057	0.088	-0.125	-0.068	-0.079
	(0.175)	(0.060)	(0.116)	(0.152)	(0.087)	(0.131)	(0.068)	(0.153)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.526***	0.054	0.207*	0.222	0.038	-0.205**	0.098**	-0.147
	(0.130)	(0.033)	(0.109)	(0.144)	(0.092)	(0.095)	(0.050)	(0.169)
Number of observations	27,557	27,558	27,276	27,582	21,757	21,665	13,004	13,181
Clusters	90	90	90	90	90	90	65	89
Adjusted R^2	0.875	0.884	0.832	0.826	0.910	0.613	0.724	0.577
				Panel B: LP Classification				
$Reform_{jt}$	-0.008	0.051	-0.046	-0.005	-0.041	-0.016	-0.024	-0.055
	(0.089)	(0.061)	(0.141)	(0.135)	(0.089)	(0.141)	(0.080)	(0.140)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.525***	0.001	0.366***	0.354***	0.290**	-0.213**	0.029	-0.063
	(0.095)	(0.038)	(0.124)	(0.130)	(0.119)	(0.092)	(0.082)	(0.126)
Number of observations	21,572	21,563	21,502	21,654	17,257	12,678	12,678	10,941
Clusters	73	73	73	73	73	73	73	72
Adjusted R^2	0.879	0.892	0.839	0.837	0.915	0.995	0.723	0.543
				Panel C: ACF Classification				
$Reform_{jt}$	-0.092	0.047	-0.127	-0.057	0.057	-0.331**	-0.035	-0.357***
	(0.138)	(0.085)	(0.154)	(0.185)	(0.115)	(0.136)	(0.066)	(0.095)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.562***	0.019	0.471***	0.434***	0.282***	0.053	0.040	0.113
	(0.139)	(0.061)	(0.137)	(0.156)	(0.091)	(0.245)	(0.040)	(0.109)
Number of observations	7,951	7,953	7,857	7,957	6,862	4,227	4,178	4,227
Clusters	67	67	67	67	67	66	51	66
Adjusted R^2	0.870	0.880	0.829	0.826	0.901	0.712	0.661	0.510

This table reports estimates of the heterogeneous effects of FDI liberalization reforms on capital constrained and unconstrained firms in a balanced panel of firms that appear in both 1995 and 2015 from the Prowess data set (equation (3)). Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, and firm age fixed effects. Firms are classified as constrained if their average $MRPK$ in the pre-treatment period from 1995-2000 is above the industry median. In Panel A, $MRPK$ is approximated as Y/K. In Panel B, it is calculated by estimating the production function using LP. In Panel C, it is calculated by estimating the production function using ACF. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.

Table A6: Regression Estimates Used to Estimate the Effect of the Policy on Gross Output

	(1) **Log Assets**	(2) **Log Salaries**	(3) **Log Materials**
$Reform_{jt}$	-0.119	-0.176	-0.172
	(0.083)	(0.114)	(0.104)
$Reform_{jt} \times I_i^{High\,MRPK}$	0.498***	0.354***	0.065
	(0.111)	(0.088)	(0.062)
$Reform_{jt} \times I_i^{High\,MRPL}$	0.278***	0.374***	0.274***
	(0.092)	(0.091)	(0.101)
$Reform_{jt} \times I_i^{High\,MRPM}$	0.054	0.004	0.089
	(0.082)	(0.080)	(0.112)
Number of observations	43,265	34,646	37,108
Clusters	80	80	80
Adjusted R^2	0.852	0.883	0.798

This table reports the difference-in-differences estimates used to estimate the policy's effects on gross output. Firms are observed between 1995 and 2015. All regressions include firm fixed effects, survey year fixed effects, and firm age fixed effects. Firms are classified as $High\,MRPK$ if their average $MRPK$ in the pre-treatment period from 1995-2000 is above the industry median, where $MRPK$ is calculation using the LP production function estimation method. $High\,MRPL$ and $High\,MRPM$ are defined analogously for materials. Standard errors are clustered at the 4-digit industry level. *, **, and *** denote 10, 5, and 1% statistical significance respectively.