Capital Market Integration, Labor Market Distortion, and Labor Misallocation*

Tsenguunjav Byambasuren, Nancy H. Chau, Anh Nguyen, and Vidhya Soundararajan†

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Abstract

This paper examines the labor market effects of foreign capital liberalization in India, with a focus on labor misallocation and employer market power, a key source of misallocation. We estimate the firm-specific wage markdowns, the gap between MRPL and wage, as a proxy for monopsony power. Leveraging a foreign direct investment liberalization episode as a natural experiment, we use difference-in-differences and event study designs to identify the causal impacts of capital market integration on the labor market. For firms with ex ante high MRPL, liberalization increases employment by 17% and labor costs by 14% and decreases MRPL by 13% and wage markdowns by 15%, with insignificant impact on wages relative to low MRPL firms. These effects are driven by female workers. Our findings are robust to a battery of robustness checks.

Keywords: Foreign direct investment, Monopsony power, Markdowns, Gender, India
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†Byambasuren: Department of Economics, Trinity College Dublin (email: byambast@tcd.ie); Chau: Dyson School of Applied Economics and Management, Cornell University and Institute for the Study of Labor (IZA) (email: hyc3@cornell.edu); Nguyen: School of Economics, Finance and Accounting, International University-VNUHCM (email: nptanh@hcmiu.edu.vn); Soundararajan: Centre for Advanced Financial Research and Learning (CAFRAL) and Institute for the Study of Labor (IZA) (email: vidhya.srajan@cafral.org.in).

1 Introduction

An inefficient allocation of production inputs, or factor misallocation, is a significant source of welfare loss in a country and plays a key role in explaining differences in income and productivity across countries (Restuccia and Rogerson, 2008; Hsieh and Klenow, 2009). A large and growing literature on sources of resource misallocation, typically measured by dispersion in marginal productivity of factors, has examined the role of economic factors, such as adjustment costs (Asker et al., 2014) and dispersion in markups (Edmond et al., 2015; Peters, 2020).[1] Globalization via the liberalization of trade and capital changes the allocation of production factors across firms and the efficiency of allocation, which influences a country's aggregate production efficiency. Studies analyzing the role of globalization in misallocation of individual production factors focus on the effect of trade liberalization on misallocation of labor (Xie et al., 2024) and capital (Kilumelume et al., 2025).[2] Except Bau and Matray (2023), who study the impact of foreign capital liberalization on misallocation, primarily focusing on capital misallocation, in India, the effect of foreign capital liberalization on misallocation, particularly labor misallocation, is understudied, and the associated mechanisms and heterogeneity are less clear. No study yet examines the role of foreign capital liberalization in labor misallocation through the lens of imperfect competition in the labor market.

This paper thus provides micro-level evidence on the role of labor market distortions and labor market competition in labor misallocation by estimating the effects of foreign capital inflows on domestic firms' labor market outcomes across heterogeneous firms, with heterogeneous workers. We utilize India's foreign capital liberalization as a natural experiment to study the labor market effects of capital market integration. From the early 1990s until the mid-2000s, India has been deregulating the foreign ownership of domestic firms in several episodes. The policy introduces an automatic approval of foreign equity investments up to 51% of domestic firms' equity shares for given industries in several episodes. This reform is unique as it increases the overall size of financial capital available for firms and changes the allocation of resources across firms, enabling us to investigate the effects on misallocation, particularly labor misallocation, and labor market power. The labor market power and the misallocation of labor will decrease as the within-industry dispersion of MRPL declines if the increased foreign capital allows firms with high MRPL to grow faster than those with low MRPL.

Leveraging industries to which the program was introduced in 2006 as the treatment group and never-treated industries as the control group,[3] we employ a difference-in-differences (DID) and

[1] The causes and effects of misallocation have been extensively reviewed in Restuccia and Rogerson (2017).

[2] Specifically, Xie et al. (2024) suggest that China's trade liberalization via lower input tariffs after its accession into the World Trade Organization (WTO) in 2001 improves labor misallocation by reducing the labor market power (equivalently, employer power or monopsony power). In South Africa, Kilumelume et al. (2025) find that a one standard deviation increase in import tariffs creates 2.8-6.2 percent aggregate productivity losses via increasing the misallocation of capital among import-competing domestic firms.

[3] For robustness check, we also consider including industries treated in 1991 in the control group and find that the results remain the same.

event-study methods to estimate the causal effects of capital market integration on labor market outcomes. Our regression model also controls for rich sets of fixed effects, including firm, year, firm age, and pre-treatment firm size-by-year fixed effects, which capture various unobserved determinants and enable us to isolate the impacts of the reform. Similar to other studies that implemented a similar strategy to identify the effects of the same policy, we find that the identification assumptions, primarily no selection and the common pre-trends assumptions, are plausible in our setting.

We use two firm-level panel datasets to measure the labor market power and MRPL and quantify the labor market effects of the policy reform. The primary source of the firm panel is India's Annual Survey of Industries (ASI), a nationally representative survey of manufacturing establishments. The ASI data has at least three unique features enabling us to introduce innovations to the literature. First, the national representativeness of the data enables us to infer welfare implications, which were restricted in some earlier works, such as Bau and Matray (2023), due to data limitations. Second, it reports the headcount of workers at the firm, enabling us to disentangle the effects on labor cost into employment and wage effects. Third, the data provide information on heterogeneous workers, for example, labor cost and headcounts of male and female workers at the firm, allowing us to analyze the impact of the policy change on different workers. The secondary firm-level data is Prowess, a panel of large and medium-sized publicly traded Indian companies over the same period. This data provides at least two advantages. First, it allows us to check the robustness of estimates of key variables, such as wage markdowns. Second, the data collects information on firms' equity composition, enabling us to calculate the fraction of foreign equity shares in total equity shares and estimate heterogeneous effects by ex ante foreign equity exposures. Both firm-level panel datasets provide detailed information necessary to quantify the wage markdowns as a measure of labor market distortions or a proxy for labor market power using the production function approach (Yeh et al., 2022).[4]

We find that the 2006 reform reduced the MRPL by 13% for high MRPL firms compared to low MRPL firms. Moreover, this paper decomposes the effect of FDI on labor cost, leveraging detailed information on employment and wages. In particular, we show that an increase in labor costs by 14% due to FDI inflows is driven by an increase in employment by 17%, not through an increase in wages, as we fail to find significant wage effect. The positive employment effect indicates a pattern of labor reallocation towards high MRPL firms, a reason for improving misallocation. For firms with initially high MRPL, liberalization reduces wage markdowns by 15% relative to low MRPL firms. The markdown-reducing impact of the policy coincides with the labor reallocation and misallocation reduction, suggesting lower labor market distortions due to the reform as a potential mechanism. Our analysis with heterogeneous workers suggests that liberalization increases employment by 36% and reduces MRPL by 33% and markdowns by 41% for women at high MRPL firms. For men, we find that liberalization has an insignificant effect on their labor market outcomes.

[4]The ASI and Prowess provide establishment- and firm-level data, respectively, and we interchangeably use the terms establishment, firm, and employer throughout the paper.

This study contributes to several strands of literature. Our first contribution is to the literature on misallocation, especially on labor misallocation in developing countries. The literature investigating how capital account liberalization affects domestic firms focuses on its impact on productivity, sectoral and capital misallocation, and welfare (Gopinath et al., 2017; Varela, 2018; Xu, 2022; McCaig et al., 2022). We make two contributions to this literature. First, we quantify the labor wedge, a potential source of misallocation, and measure MRPL consistent with the wedge. Existing studies, such as Bau and Matray (2023), focus on measuring MRPL based on a simple Cobb-Douglas production function estimated without accounting for the endogeneity problem due to unobserved firm-specific productivity. Our paper instead estimates the production function using the "proxy variable" method (Ackerberg et al., 2015) and measures MRPL, allowing for imperfect competition in both output and labor markets. The foreign capital liberalization may affect labor misallocation through several channels, such as (i) relaxing firms' financial constraints (Fonseca and Van Doornik, 2022), (ii) labor reallocation and the associated changes in labor market power, and (iii) general equilibrium effects.[5] We thus, second, quantify how much the change in labor wedge contributes to changes in labor misallocation due to foreign capital integration. Leveraging India's foreign capital liberalization episodes in 2001 and 2006, Bau and Matray (2023) show that the reforms reduced MRPL by 28% and increased labor costs by 24% for firms with high ex ante MRPL relative to low MRPL firms. While our results on the MRPL and labor cost effects of the 2006 foreign capital liberalization episode are consistent with their findings, we provide new results by presenting effects on employment, wages, and wage markdowns.

Second, we contribute to the literature on the firm's monopsony power. We add to this literature in two ways. First, we contribute to the fast-growing literature exploring the determinants of labor market power, particularly the strand that examines foreign capital in terms of trade and investment. Existing studies focus on trade shock to investigate labor reallocation as a source of the firm's wage-setting power (Mertens, 2020; Felix, 2022; Pham, 2023; Kusaka, 2023; Kondo et al., 2024; Xie et al., 2024; Hoang et al., 2024).[6] However, there is no study yet on the link between FDI and monopsony. An exception is Lu et al. (2020) who show that the FDI liberalization increases

[5]Related to the second channel, Lu et al. (2020) examine the effect of China's foreign direct investment (FDI) reform around its WTO accession in the early 2000s on labor market power. But their focus is not on misallocation. They find that capital liberalization increases firms' market power in the labor markets, which is opposite to the effect of trade liberalization suggested by the existing studies in the monopsony literature.

[6]Specifically, Pham (2023) shows that export expansion through bilateral trade agreement (BTA) between the U.S. and Vietnam reduces wage markdown, measured by directly dividing marginal revenue product of labor (MRPL) to wage, by 3.0% in Vietnam at the firm level. The reduction happens mostly for female workers, i.e., markdown on female workers decreases by 13.5% due to the BTA shock. Following Gandhi et al. (2020) approach, the author estimates a mean markdown level of 2.14 and a median of 1.49. Using firm-level data in China, Kondo et al. (2024) suggest that input tariff reductions due to joining WTO are associated with lower wage markdowns and that the wage markdown decline is more for skill-intensive firms than non-skill-intensive firms. On the contrary, Felix (2022) finds that Brazil's 1990s trade liberalization increased labor market concentration and thus the firm's labor market power, using Brazil's matched employer-employee data. Using administrative data from the AFiD database on German manufacturers, Mertens (2020) suggests that an increase in export demand increases a firm's labor market power. Import competition, however, reduces the labor market power.

wage markdowns using the relaxation of FDI regulation during China's WTO accession in 2001. Our results suggest the average effect of India's FDI reform on labor market power is essentially zero, but the effects are highly heterogeneous across firms with different ex ante MRPL and wage markdowns. Second, we relate to studies measuring monopsony power for heterogeneous workers such as male and female workers. In terms of measurement, we estimate markdowns in a developing country using two alternative datasets and introduce worker heterogeneity, focusing on the gender dimension. For example, existing studies measuring monopsony power for male and female workers estimate labor supply elasticities (Hirsch et al., 2010; Caldwell and Oehlsen, 2023; Sharma, 2023). However, we provide the first estimate on wage markdowns for male and female workers to provide a gender-specific measure of monopsony power. We find that women are subject to higher labor market power than men, consistent with other studies.

Third, we complement the literature on the social and economic impacts of foreign firms via multinational firms and foreign capital, especially on the labor market and a vulnerable group like women. Studies on the gender-specific labor market effects of multinational firms provide mixed results. For example, Siegel et al. (2019) suggest that multinationals improve profitability and productivity by intensively hiring from a marginalized social group, women, in the local managerial market in South Korea. Using the same FDI liberalization episode that we study in this paper, Li et al. (2026) show that economic integration reduces rape via empowering women due to an increase in women's relative income and an introduction of gender equality norms. However, González and Kong (2025) find that foreign firms pay women disproportionately higher wages than domestic firms at the cost of disproportionately lower amenities relative to men. Women sort away from foreign firms due to poorer amenities, leading to a smaller share of women at foreign firms than domestic firms. Consistent with the first set of findings, our results suggest that India's foreign capital liberalization benefits women more than men. Specifically, the effects of FDI inflows on manufacturing firms' labor market outcomes are mainly driven by female workers.

The rest of the paper is structured as follows. Section 2 provides background on the Indian context of foreign capital liberalization. Section 3 describes the data and discusses the estimation of markdown and MRPL for manufacturing firms. Section 4 presents the empirical strategy used to estimate the causal effects of capital market integration. Section 5 discusses the empirical results, while Section 6 examines worker heterogeneity by gender. Finally, Section 7 concludes.

2 Study Settings

2.1 Background

In the early 1990s, India implemented a structural adjustment program to overcome the balance of payment crisis in 1991. This reform covers trade liberalization, the relaxation of licensing restrictions, and the opening to foreign investment. Before 1991, most of the manufacturing industries in India were heavily regulated by the Foreign Exchange Regulation Act of 1973, which restricted foreign ownership to below 40% in most sectors and placed restrictions on the use of foreign brand names, the remittances of dividends generated in India by foreign affiliates and the use local content in output (Sivadasan, 2009).

The reform in 1991 relaxed such restrictions by allowing automatic approval of foreign ownership up to 51% in given industries. This new policy was controversial given that India had been accustomed to protectionist policies for decades, and many businesses were concerned about the potential domination of foreign capital (Singh, 2005). Thus, only certain industries were chosen for the reform, and politics played a significant role in the industry selection. Firms in concentrated industries and state-owned firms were more successful in preventing the entry of foreign capital (Chari and Gupta, 2008).

More industries joined the liberalization over time, increasing the cap on foreign equity and/or enabling the automatic approval route. According to the NIC industry classification, there are 400 five-digit industries in India. The policy was introduced into 124, 1, 5, 5, and 19 five-digit industries during the episodes of the reform in 1991, 1998, 2000, 2001, and 2006, respectively (Bau and Matray, 2023). In contrast to the 1991 episode, the 2000s reform was motivated by a political consensus to boost foreign investments to counteract a declining trend in FDI during the 1998-1999 period (India Ministry of Finance, Economic Division, Government, 2002, 2007).

2.2 Foreign Equity Inflows

The inflow of foreign equity into India has been increasing over time. However, industries that have been liberalized experienced a sudden increase in foreign capital. As shown in Figure 1, foreign equities in industries that liberalized in 2006 increased visibly faster than in industries that never liberalized. For the reform industries, the dynamic pattern shows that transmission of the policy to the observed foreign capital is not immediate and takes some time. Specifically, foreign equities in 2006 reform industries presented a noticeable jump in 2009. Despite the lag, the foreign capital integration strongly increases foreign equities.

3 Data and Measurement

In this section, we first describe the data used in our empirical analysis. Second, we present the methodological approach for measuring the key variables, including markdowns and MRPL, and discuss the estimated measures. Third, we provide descriptive statistics on the main variables in our analysis.

3.1 Firm-Level Data

ASI Data. The primary firm-level data in our empirical analysis is the panel version of the Annual Survey of Industries (ASI), which offers comprehensive data for India's industrial sector. It is a nationally representative survey of all factories registered under The Factories Act of 1948, which are defined as factories employing at least 10 workers and do not use electricity or employing at least 20 workers without electricity. Our panel version spans from 1998-1999 to 2017-2018, where each round covers the financial year from April 1 to March 31.

We restrict the sample to manufacturing firms and extract detailed information on the firm's production inputs and output for firm-level markdown estimation. It includes labor headcount, labor cost (total expenses on wages and salaries), material expenditures, capital stock, and sales revenue. The data also contains information on employment and labor cost for various types of heterogeneous workers, which allows us to estimate markdowns and examine the heterogeneity of FDI effects on labor market outcomes for different types of workers, particularly male and female workers.

Prowess Data. To complement our analysis, we also use the firm panel from Prowess database conducted by the Center for Monitoring the Indian Economy (CMIE). This dataset covers medium and large private firms and all publicly traded firms, which account for more than 70% of the organized industrial activities (Bau and Matray, 2023). The Prowess is thus not representative data, which is why it is our secondary firm-level data. It also contains detailed information on the firm's income and balance sheet, allowing us to estimate the markdowns. However, Prowess only provides consistent information for labor cost and other labor-related expenses, while data for labor headcount is largely missing. For example, there are 52,809 (83%) observations with missing headcount in the raw sample. Thus, we use labor cost as labor input in our markdown estimation on this data to maintain a large sample size.

One advantage of Prowess data is that it contains detailed information on the ownership composition of firms' equity shares, which allows us to calculate the percentage of foreign equity shares in total equity shares at the industry level and estimate the heterogeneous FDI effects by an ex ante exposure to foreign investment.[7,8] To define the fraction of foreign equity shares in total equity shares, we sum the foreign equity shares and total equity shares of all Prowess firms at the 4-digit NIC industry level. We then calculate the percentage of foreign equity shares and take the value of the 2000-2001 financial year to determine the industry's foreign equity exposure prior to the reform. We merge this measure of ex ante exposure to foreign investment with the ASI data by 4-digit NIC industries to conduct the heterogeneity analysis. Since Prowess contains the firm's industry at the

[7] Prowess reports the number of equity shares of the company sourced from the stock exchanges. The number of foreign equity shares is the sum of shares held by (i) foreign promoters, (ii) foreign institutional investors as non-promoters, and (iii) foreign venture capital investors as non-promoters.

[8] A caveat when using Prowess to estimate industry-level foreign equity is that the data only reports equity information for publicly-listed firms, which consists of only 24% of the firms in our sample.

5-digit level, we merge it with FDI data by industry.

3.2 FDI Liberalization Data

We extract the FDI policy change data at the industry level from Bau and Matray (2023). Industries in India are classified based on the National Industry Classification (NIC). The data contains a list of 5-digit NIC-2008 industries liberalized by the FDI reform and the year of reform for each treated industry. An industry is classified as under reform or treated if the policy allows automatic approval of foreign investment and/or increases the cap on foreign ownership to at least 51%. We thus define the treatment as the combination of these two reforms. To match the firm-level data, we use concordance tables to convert the industries to the NIC-2004 and NIC-1998 versions.

Since our ASI data only reports 4-digit NIC industries at the most granular level, we first collapse the 5-digit level FDI data to the 4-digit level and define the year of reform of a 4-digit industry as the last reform year of its 5-digit sub-class. The 4-digit industry is defined as treated if any 5-digit sub-industry had reform. At the 4-digit level, there are 51, 1, 1, 1, and 6 industries liberalized during the 1991, 1998, 2000, 2001, and 2006 reforms, respectively. Given that some industries were randomly missing in our ASI dataset, we have 48, 1, and 5 four-digit industries treated during the 1991, 1998, and 2006 reforms. The 4-digit industry that has been liberalized in 2000 was not present in our ASI data, and the wage markdown was not estimated for firms operating in the 2001 reform industry. We exclude a 4-digit industry that experienced a liberalization in 1998. This amounts to only 1.89% of the sample. Additionally, since our ASI sample begins early 2000s, the industries liberalized in 1991 are always treated, and their treatment status remains the same during our sample period. In our baseline analysis, we exclude these industries from the sample, and we check the robustness of our results by including those industries in the control group like Bau and Matray (2023).

The Prowess contains the firm's industry information at the 5-digit level. We match it with FDI data by industry at the 5-digit level.

3.3 Additional Data

We also control for other changes in trade policy at the industry level that might be correlated with the FDI reform treatment and the firm's outcomes. Specifically, we use trade data from the UN-COMTRADE database to construct the industry's exposure to Chinese import competition. The industries from UN-COMTRADE are classified based on the Standard Industrial Classification (ISIC) Revision 3.1 at 4-digit level, which can be mapped one-to-one onto the NIC-2004. We also use the tariff data from the World Integrated Trade Solution (WITS) database and the input-tariff table to construct a measurement for input and output tariffs at the industry level.

3.4 Measuring Markdowns and MRPL

Estimation Approach. The firm-level wage markdowns have been estimated using production approach under different assumptions. In the misallocation literature, Bau and Matray (2023) employ marginal returns to production inputs, including capital and labor, as a measure of misallocation of the inputs or input wedge for India's manufacturing. They assume a Cobb-Douglas (revenue) production function and thus do not impose the assumptions associated with production function estimation techniques. The wage markdown–wedge between the marginal revenue product of labor (MRPL) and the wage, has been estimated by Pham (2023) for Chinese manufacturing industry based on Gandhi et al.'s (2020) nonparametric approach, relaxing the functional form assumption.

However, neither of these approaches takes the price-cost markups in the output market into account. Accounting for the markups, Brooks et al. (2021) estimate markdowns in the contexts of India and China by estimating production function under various functional forms, with a third-order translog production function specific across 2-digit industries as the primary form. In their markdown estimation, they impose a condition in which small firms with negligible share in the market have no market power in the labor market. The relationship between firm size and markdown has been shown positive and statistically significant for several counties, such as the U.S. (Yeh et al., 2022), Germany (Byambasuren, 2025), and Mexico (Estefan et al., 2024). This assumption is intuitive; however, it might not hold in all contexts. For example, Byambasuren et al. (2025) find that markdown and firm size, measured by employment share similar to the aforementioned studies, is negatively correlated in India's manufacturing sector. Additionally, Byambasuren (2025) show that markdown-firm size relationship is negative in East Germany, which is less developed than West Germany where the relationship is positive. Thus, in this paper, we use a more flexible production approach by Yeh et al. (2022) that does not assume the firm size-markdown link is positive or firms with small market share do not have market power in the labor markets.

From the firm's profit maximization and cost minimization problems, the wage markdown η_{it} is defined as follows

$$\eta_{it} = \frac{\theta^L_{it}}{\alpha^L_{it}} \mu^{-1}_{it}, \tag{1}$$

where θ^L_{it} is the output elasticity of labor for firm i in time t, α^L_{it} is the share of labor cost in revenue, and μ_{it} is the price-cost markup in the output market, which is further defined as

$$\mu_{it} = \frac{\theta^M_{it}}{\alpha^M_{it}}, \tag{2}$$

where θ^M_{it} is the output elasticity of a variable input other than the labor, intermediate materials in our case, and α^M_{it} is the share of cost on materials in revenue.[9] The shares of labor and material costs

[9] Solving for the profit maximization problem shows that the markdown, the ratio of the MRPL to the wage, is the inverse labor supply elasticity plus one. The cost minimization problem with respect to labor input yields that the inverse labor supply elasticity plus one is the ratio of output elasticity of labor to the share of labor cost in revenue,

in revenue can be calculated directly from the data. The output elasticities of labor and materials are derived from the production function estimation. We use the standard method of Ackerberg et al. (2015), as in De Loecker and Warzynski (2012), to estimate the production function. For our baseline analysis, we employ a second-order translog production function at the 2-digit industry level and use an industry-specific Cobb-Douglas functional form as a robustness check. Identifying the consistent estimates of the production parameters and, thus, the output elasticities requires that the firms dynamically optimize their decisions in discrete times and the intermediate materials are fully flexible. The details of the estimation procedure and underlying assumptions are provided in Yeh et al. (2022). We also include the policy reforms we investigate, FDI liberalization defined at the industry-year level, in our markup and markdown estimation procedure, like Brandt et al. (2017).

Consistent with the definition of markdown in Equation (1), essentially the ratio of the marginal revenue product of labor (MRPL) to wage, the MRPL, our measure of labor misallocation, is defined as the wage multiplied by the markdown,

$$\mathrm{MRPL}_{it} = \eta_{it} w_{it}, \tag{3}$$

where w_{it} is the average wage of workers at firm i at time t.

Markdown Estimates in India's Manufacturing. We estimate the markdown in India's manufacturing using two datasets, including the ASI and Prowess. Table 1 presents the estimated wage markdowns using the ASI data from 2000-2018. The results suggest that the labor market in the manufacturing sector is imperfectly competitive as the median and average markdowns are more than unity. In particular, workers in an average firm receive 0.54 rupees for each rupee generated. This estimate is 25% lower than the estimate of 0.72 rupees on the marginal rupee suggested by Byambasuren et al. (2025) using the ASI data from 2000-2008. This could be because markdown has been high in a decade between 2008 and 2018.

To analyze this, Figure 2 depicts the trend of aggregate markdowns between 2000 and 2018, indicating several intriguing results. First, although the markdown trend after 2008 is generally downward, the markdown has been high for most of the post-2008 periods, potentially yielding a higher average markdown from 2000-2018. Second, the upward trend from 2000-2008 is strongly consistent with Byambasuren et al. (2025), who show a similar pattern over the same period. Third, the aggregate wage markdown presents an inverted U-shape pattern over the period 2000-2018 and peaks in 2009 at the dawn of the Great Recession, with a sharp and continuous decline since then. This pattern of markdown after the 2009 Global Financial Crisis (GFC) is strongly similar to the markdown pattern in Germany (Byambasuren, 2025; Dustmann et al., 2024). Figure 3 shows that

multiplied by the inverse of price markup. Equating these two gives Equation (1). In this approach, we assume that intermediate materials are flexible input or that the market for intermediate materials is perfectly competitive. The cost minimization problem with respect to this flexible input of intermediate materials yields Equation (2), showing that the price markup is the ratio of the output elasticity of intermediate materials to the costs of intermediate materials in revenue.

the aggregate markdown presents a similar trend under an alternative Cobb-Douglas production function.

The wage markdowns estimated based on Prowess data from 1995-2019 are presented in Table 2. The results suggest that the publicly-listed manufacturing firms operate in a monopsonistic environment, similar to the findings from the ASI data above. The markdown estimates, however, are persistently higher than those from the ASI data, with a median (average) markdown estimate of 2.692 (3.283). Despite the level differences, as shown in Figure 4, the median wage markdown estimates based on the two datasets are fairly correlated at the two-digit industry level, with a pairwise correlation of 0.45 (SE: 0.22, p-value: 0.05).

We evaluate several potential reasons for the difference between the two estimates. First, fixing the time coverage from 2000 to 2018, similar to the ASI data, we find the median (average) estimate of 2.704 (3.297), suggesting that the difference is not due to the difference in time coverage. Second, the markdown has not been estimated for two 2-digit manufacturing industries, including (i) office, accounting, and computing machinery and (ii) medical, precision, and optical instruments, watches, and clocks. So, we calculate the median and average markdowns estimated using the ASI data by excluding these two sectors, which yield 1.374 and 1.871, respectively. These estimates are not different from the baseline estimates, suggesting that the absence of the two industries also does not explain the difference between markdown estimates from the ASI and Prowess data. Third, the labor inputs in the production function are approximated by labor cost when using Prowess data that does not report the headcount of workers. Thus, we estimate the wage markdowns by using the ASI data and proxying labor inputs with real labor cost as we did in Table 2, and the median and average markdowns are estimated at 1.536 and 1.762, which are not too different from the estimates that leverage headcounts. The measure of labor inputs thus seems to be not the main driver of the difference. However, the distinction between the markdowns from the ASI and Prowess could be due to the underlying differences between the two data sources, i.e., the publicly-listed and non-publicly-listed firms. Although the ASI data does not report whether the establishment is publicly listed, we compute the average and median markdowns for public and private limited companies that can sell shares to the public and trade them on stock exchanges. The median and average markdowns for such firms are 1.490 and 2.031, respectively, higher than national estimates and estimates for other corporations. Hence, it is likely that the manufacturers included in Prowess data inherently have higher markdowns.

Despite the level differences, the evolution of wage markdowns using Prowess data (Figure 5) is highly comparable to the trend of aggregate markdown based on the ASI data. In particular, the time evolution presents an inverted U-shaped trend under both translog and Cobb-Douglas production functions, with a peak around the Great Recession. While we estimate the firm-level markdowns using Prowess data from 1995-2019, the aggregate markdown has been plotted after 2000 because the markdown trend was noisy over the pre-2000 period, for which very few manufacturing firms covered in Prowess data.

The relationships of the firm's idiosyncratic characteristics, including size, age, and productivity with the wage markdown based on the ASI data from 2000-2018 (Figure 6) are strongly consistent with the results from Byambasuren et al. (2025) who used the ASI data from 2000-2008. In particular, markdown is negatively correlated with firm size and age and is positively correlated with productivity. These relationships are similar when using Prowess data, except for the markdown-size relationship, which is positive and statistically significant at the top part of the distribution (Figure 7). This difference between the markdown-size relationship using the ASI and Prowess data could be another indication of the difference between firms included in the two datasets. The relationship for publicly-listed firms in Prowess is similar to that in developed countries like the U.S. and Germany.[10]

MRPL Estimates. The MRPL is positively associated with markdowns by construction as in Equation (3). Specifically, the correlation coefficient between log MRPL and log markdown is 0.69 (SE: 0.003, p-value: 0.000) after controlling for firm and year fixed effects among ASI firms. This indicates that the two measures are closely related and that firm-specific wages are relatively stable over time. Changes in markdown is therefore a strong predictor of changes in MRPL.

Assuming that all firms in an industry share the same output elasticity of labor in the Cobb-Douglas revenue function, Bau and Matray (2023) measure as $MRPL_{it} = Revenue_{it}/L_{it}$. The OLS coefficient from estimating log revenue per labor on our MRPL measure based on a Cobb-Douglas and Translog production function is 0.73 (SE: 0.001) and 0.63 (0.001), respectively, conditional on firm and year fixed effects.[11]

3.5 Descriptive Statistics

Table 3 shows the summary statistics for ASI firms' characteristics, such as markdown and MRPL, by the treatment (FDI liberalization) status. An observation is at the firm-year level. The two groups are relatively similar in capital, sales revenue, employment, and wages, except that the treated (liberalized) firms are slightly older and have higher markdowns on average. As a supplement, we also show the summary statistics for Prowess sample in Table A.1. Compared to the ASI firms, Prowess firms are much older and higher in all characteristics.[12] Compared to the control group, the treated firms in Prowess data are also relatively similar in capital and sales, while their employment is smaller on average.

[10] Figure A.1 shows that the labor wedge (τ_i^L) as in Bau and Matray (2023) and our estimate of wage markdowns (η_i) are highly correlated both in pre- and post-treatment periods. The OLS coefficient from estimating markdowns on labor wedge, with firm and year fixed effects controlled, is 0.92 (SE: 0.000). This suggests that the results are not too much different across the two approaches.

[11] Figure A.2 shows that the MRPL based on revenue per labor is comparable to our MRPL measures based on Cobb-Douglas and Translog production functions in the pre- and post-treatment periods.

[12] Since Prowess records the monetary value in US dollars, we convert it to Indian Rupees to be comparable with the ASI data. The wages in Prowess are measured by labor cost instead of wages per worker due to limited information on employment headcount.

As the ASI contains employment information for different types of workers, we also explore how gender-specific (male and female) employment evolved for the control and treated industries. As shown in Figure 8, the employment share of female workers has been declining in treated industries, while that in control industries moderately grew after the treatment.

4 Empirical Strategy

In this section, we first describe the empirical strategy we employ to estimate the causal impact of FDI liberalization on labor misallocation, monopsony, and other labor market outcomes. We then discuss the identification assumptions.

4.1 Empirical Specification

We estimate the following difference-in-differences (DID) specification to investigate the labor market effects of capital liberalization:

$$Y_{ijt} = \alpha + \beta \text{Reform}_j \times \text{Post}_t + \mathbf{X}'_{it}\delta + \mu_i + \gamma_t + \varepsilon_{ijt}, \tag{4}$$

where Y_{ijt} is one of the outcomes for firm i in industry j at time t, including the logs of employment, labor cost, average wage per worker, MRPL, and markdowns. Reform$_j$ is a dummy variable equal to 1 if the industry is liberalized during the 2006 FDI reform and to zero if the industry is never received FDI liberalization. Post$_t$ is an indicator variable equal to 1 if the period is 2006-2007 financial year or after, zero otherwise.[13] Our treatment variable is coded at the 4-digit level. The vector $\mathbf{X}'_{it}$ contains firm-level covariates, including firm age and pre-treatment firm size-by-year fixed effects. The firm fixed effects μ_i control all time-invariant unobserved factors, and the year fixed effects γ_t account for all time-varying factors common across firms. The standard errors are two-way clustered at the 4-digit industry and year level to account for any serial correlation.

To analyze the heterogeneous effects of liberalization across firms with different ex ante MRPL, we estimate the following specification:

$$\begin{aligned} Y_{ijt} = \alpha &+ \beta_1 \text{Reform}_j \times \text{Post}_t \times I_i^{\textit{High MRPL}} + \beta_2 \text{Reform}_j \times \text{Post}_t \\ &+ \beta_3 \text{Post}_t \times I_i^{\textit{High MRPL}} + \mathbf{X}'_{it}\delta + \mu_i + \gamma_t + \epsilon_{ijt}, \end{aligned} \tag{5}$$

where $I_i^{\textit{High MRPL}}$ is an indicator if the firm i's ex ante marginal revenue product of labor (MRPL) is high (or above the within industry median), and other variables are the same as those in Equation (4). The average MRPL over the 1999-2000 and 2000-2001 financial years has been used as an

[13] To check the robustness of our results and assess the anticipation effect, we also consider the periods 2005-2006 financial year and after as post-treatment period.

ex ante MRPL in the baseline analysis to avoid the potential spillover effect of the FDI reform introduced in mid-2001 across industries.[14] The time-invariant and firm-specific terms, including $\text{Reform}_j \times I_i^{\text{High MRPL}}$, Reform_j, and $I_i^{\text{High MRPL}}$, are captured by the firm fixed effects, and the terms changing over time but common across firms like Post_t are captured by the year fixed effects.

4.2 Identification and Assumptions

In our empirical analysis, we focus on the set of industries liberalized in 2006 because our ASI firm panel starts in 1999 (or the 1999-2000 financial year), and thus it provides more pre-reform periods in the event-study analysis. We also restricted our regression sample to 2002-2018 to avoid the confounding effects of other episodes of liberalization in the early 1990s and 2001. The use of later years is also advantageous because the number of firms grew over time and became more stable. Since the FDI shock that we study in this paper occurs once and involves a binary treatment, the ordinary least squares (OLS) two-way fixed effects estimator is unbiased for the average treatment effect on the treated (ATT) (De Chaisemartin and d'Haultfoeuille, 2020).

Our main empirical specification is the triple difference regression in Equation (5) which compares firms with high MRPL and high markdowns with low MRPL and low markdowns between treated and untreated industries before and after 2006. The key identification assumption is the parallel trends assumption. In the absence of the 2006 reform, firms with high and low MRPL in treated and untreated industries evolve in parallel. Consistent with Olden and Møen (2022), we use the following event-study specification to test the parallel pre-trends for our triple difference design:

$$\begin{aligned} Y_{ijt} = \alpha + & \sum_{\tau \neq -1; \tau=-5}^{\tau=11} \beta_{1\tau} \times I_\tau \times \text{Reform}_j \times I_i^{\text{High MRPL}} + \sum_{\tau \neq -1; \tau=-5}^{\tau=11} \beta_{2\tau} \times I_\tau \times \text{Reform}_j \\ & + \sum_{\tau \neq -1; \tau=-5}^{\tau=11} \beta_{3\tau} \times I_\tau \times I_i^{\text{High MRPL}} + \mathbf{X}'_{it}\delta + \mu_i + \gamma_t + \epsilon_{ijt}, \end{aligned} \tag{6}$$

where I_τ are lags and leads in event time, with $\tau = -1$ as the base period. The remaining variables are similar to those in Equation (5). Below, we discuss potential identification threats.

Selection of Treated Firms. The rationales for liberalized industries are not provided in official documents such as the Economic Surveys. However, using our data, we calculated industry-level characteristics before the 2006 reform and computed the relationship between the reform status and those characteristics to examine whether the initial conditions can predict the likelihood of receiving the reform. All regressions are weighted by the industry's total capital stock in 2005. We report the regression results in Table 4. There is little evidence that an industry's reform status is correlated with its ex ante dispersion in labor or capital misallocation (columns1–2). Columns 3 and 4 show that industries with a higher number of firms or a larger capital stock are not more likely to

[14] The 2000-2001 financial year starts on April 1, 2000, and ends on March 31, 2001.

receive a reform. Lastly, columns 5 and 6 show that political-economy characteristics such as the Herfindahl Index and sales share of state-owned firms are also not predictors for the 2006 reform. Thus, the potential selection bias is implausible in our setting.

We also examine the distribution of employment before and after the 2006 reform by different initial characteristics. We classify the firms according to whether they were initially exposed to (i) high foreign investment, (ii) high MRPL, and (iii) high TFPR. Figure 9 shows that before 2006, on average, there was no difference in prereform employment between firms with high or low labor wedge and productivity. However, firms in industries with higher initial exposure to foreign investment tend to have higher employment. After 2006, the employment gap shrinks among high and low foreign equity groups, while it is now higher for firms with high initial MRPL and TFPR. It could indicate that firms with higher initial labor distortion or higher productivity could expand their employment after the FDI reform, which is in line with the trade literature on the heterogeneous effects of trade along firm productivity, which reallocate production resources to more productive firms (Melitz, 2003; Rodriguez-Lopez and Yu, 2024).

Endogeneity of Foreign Equity Flows. Foreign equities can flow into specific firms within an industry. However, our empirical strategy does not exploit an exogenous variation in the amount of foreign capital received, but leverages an exogenous shifter to the potential amount of foreign capital. Thus, we do not require foreign capital inflows to be randomly allocated among firms within treated industries. Instead, our estimates are valid as long as the parallel trends are satisfied.

Measurement Error in MRPL and Markdowns. The MRPL and markdowns are our main outcomes in equations (4) and (5), but the ex ante values of those variables are also on the right-hand side of Equation (5). Thus, we discuss potential biases from measurement error in both the outcome and explanatory variables. In general, measurement errors either firm-specific and time-invariant or time-varying but common across firms should not create bias, since we control for firm and year fixed effects. However, measurement error can bias the point estimates if it is idiosyncratic or firm-specific and time varying.

Classical measurement errors in MPRL and markdowns as outcome variables would not bias our point estimates. However, the idiosyncratic measurement error in MRPL and markdowns as independent variables may introduce bias as we might misclassify firms with high and low ex ante MRPL and markdowns. To avoid this potential attenuation bias, we remove outliers by removing firms with extreme values of MRPL and markdowns. Additionally, we conduct a battery of robustness checks using various values of ex ante MRPL to estimate the heterogeneous effects.

5 Baseline Results

This section presents the baseline results. First, we discuss the average effects of the policy. Second, we present the heterogeneous effects across firms with different levels of ex ante MRPL. Third, we evaluate several potential mechanisms. We then perform a battery of robustness checks of our baseline results.

5.1 Average Effects

To examine the labor market consequences of foreign capital integration, we focus on the effects on employment, labor costs, average wages, MRPL, and wage markdowns. Table 5 presents the baseline results on the average effects of liberalization on these outcomes. We do not find strong average effects, except for the positive wage effects (column 3). The estimate of the coefficient on employment is surprisingly negative, but this average employment effect is essentially zero. The average impact on the labor cost is positive, while the coefficients on MRPL and wage markdowns are negative. However, these estimates are also not statistically different from zero. Using our firm-level data from the ASI, which reports the number of workers, we calculate the firm's average wage and estimate the wage effect, one of the key innovations of this paper. We find that India's FDI liberalization increases the average wage by 5.2 percent in the manufacturing industry.

5.2 Heterogeneous Effects by Ex ante MRPL

Table 6 presents the estimates of the differential effects of the policy by the pre-reform MRPL using our main specification in Equation (5). In response to liberalization, employment increases by 17% at high MRPL firms compared to low MRPL firms (column 1). Consistent with Bau and Matray (2023), high MRPL firms also experience a relative increase in their labor costs by 14% (column 2). Our finding on the wage effect reveals that the effect of foreign capital liberalization on labor costs is explained by the employment effect, as the wage effect for high MPRL firms is positive but not statistically significant (column 3). The null effect on average wages could be due to opposing forces. For example, intensifying labor market competition would increase the average wage, while an increase in the share of low-paying workers at high MRPL firms could dampen the average wage through the composition effect. For firms with low MRPL firms, the liberalization increases average wages relative to high MRPL firms, but the coefficient estimate is weakly significant at the 10% level.

For firms with a high ex ante MRPL, the MRPL decreases by 13% relative to firms with a low initial MRPL (column 4). This misallocation-correcting effect appears to be driven by a 15% decline in labor market power, measured by wage markdowns, for high MRPL firms compared to low MRPL firms (column 5). This is consistent with the wage effect of the policy, that is, from Equation (3)

$$\underbrace{\%\Delta \mathrm{MRPL}_{it}}_{\text{labor misallocation effect}} = \underbrace{\%\Delta \eta_{it}}_{\text{markdown effect}} + \underbrace{\%\Delta w_{it}}_{\text{wage effect}},$$

where $\%\Delta$ indicates proportional changes. The underlying cause of declines in markdowns and labor misallocation is labor reallocation from low MRPL firms to high MRPL firms. For domestic firms with initially low MRPL, liberalization reduces the number of workers by 12% relative to high MRPL firms. More financial resources allow high-MRPL firms to expand on hiring from low-MRPL firms. It leads to reductions in labor market power and labor misallocation within treated industries.

We estimate the heterogeneous effects by ex ante MRPL using the event study specification in Equation (6) to show the dynamic effects and formally test the parallel pre-trends assumption. Figure 10 illustrates the impacts for high-MRPL firms relative to low-MRPL firms. The reference year denoted by "-1" is the year before the 2006 reform, the 2005-2006 financial year in our baseline specification. We find no pre-reform effect on all labor market outcomes of our interest in the treated industries except for a few pre-treatment years for some outcomes. Therefore, we consider that our estimated effects are valid and causal because a common pre-trends assumption is plausible for most pre-treatment periods. The positive effects on employment and labor cost at high ex ante MRPL are different from zero and persistent over time. Although slightly less precisely estimated, the dynamic impacts on MRPL and markdowns for high MRPL firms are negative, particularly in the medium and long term. For high MRPL firms, liberalization generally presents a null effect on wages, with some negative impacts in the medium and long term.

To further examine the identification strategy, Figure 11 shows the relative effects for both high and low MRPL firms. The liberalization has no or negligible, if not null, effects on the five outcomes for most of the post-treatment periods, except for employment. Unless new workers enter the domestic labor market from unemployment or abroad, workers would reallocate due to liberalization because the workers pool would be the same.

Although our focus is on the labor market effects of the foreign capital liberalization, we also estimate the impacts on firms' revenue and capital. Table 7 presents the results. The results are generally consistent with Bau and Matray (2023), who suggest that the reform hardly affected the revenue and capital of firms with high MRPL relative to those with low MRPL.

Importance of the Industrial Relations Climate. Our results so far show that opening up to foreign capital allows high MRPL firms to employ more and grow faster. If foreign capital liberalization is acting as a substitute for a more pro-worker labor market, a natural implication is that the liberalization affects less among ex ante high MRPL firms in areas with labor markets with regulations more favorable for workers prior to the reform. We directly test this hypothesis by using the measure of industrial relations climate, which is originally proposed by Besley and Burgess (2004), defined at the 2001 level in the state s. The data are obtained from Dasgupta (2023), who refined the measure. We focus on the cumulative number of Industrial Disputes Acts in a pro-worker direction net of those in a pro-employer direction and then interact this measure with all the single and cross-terms in equation (5). The coefficient of interest is the coefficient for the four-way interaction $\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}} \times \text{Labor Regulation}_s$, which captures the differential effect of the policy on high MRPL firms located in states with more pro-worker labor regulations.

Table 8 reports the results.[15] For employment, labor cost, and wages, the interaction $\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}} \times \text{Labor Regulation}_s$ is negative and significant at least at the 5% level. For MRPL and markdowns, the four-way interaction is positive and significant at the 5% level. Taken together, these results suggest that the reform had more impact for high MRPL firms located in states with less pro-worker labor regulations. These results also imply that pro-employer regulations are an important source of labor misallocation in India and that foreign capital liberalization can act as an alternative for amending regulations in a pro-worker direction.

5.3 Mechanisms

Direct or Indirect Effects. We first analyze whether the effect of liberalization is direct–misallocation reduced at firms that directly received the foreign equities–or if the effect is indirect–misallocation reduced at firms that did not directly receive the foreign equities, e.g., through general equilibrium effects. In Table 9, we analyze which firms received foreign equity inflows by estimating foreign equity regressions across industries with different initial exposure to foreign equities. Foreign equity flows into industries with high ex ante foreign equity exposure (column 1). However, treated industries with low exposure received foreign equity, while treated industries with high exposure experienced foreign equity outflows (column 2).

Then, we estimate the labor market effects of the reform, heterogeneous by foreign equity exposure. Table 10 reports the results. The number of observations was reduced by more than 50% due to a lack of information on the amount of foreign equities in Prowess data. This could lead to less precisely estimated effects, and indeed, the estimates are noisier for employment and labor costs. The effects on employment and labor costs become essentially zero (columns 1–2). The impact on wages is also not statistically significant, similar to the result under heterogeneity by MRPL (column 3). FDI inflows relatively reduce MRPL and markdown by 49% and 50%, respectively, for firms with low foreign capital exposure compared to those with high exposure (columns 4–5). These findings generally indicate that the effects of liberalization in treated industries are more direct, rather than indirect or general equilibrium effects.

Dispersion in Markdowns. Given that labor market power is a potential source of misallocation, we analyze how the reform affected the dispersion in labor market power or markdown distortions. In doing so, we estimate a specification similar to Equation (5) but with $I_i^{High\ Markdown}$ instead of $I_i^{High\ MRPL}$. Table 11 reports the results. The foreign capital liberalization reduced dispersion in markdowns. For firms with high ex ante markdowns, liberalization reduces markdowns by 12% relative to low markdown firms (column 5). The heterogeneous effects on the other four outcomes are generally the same. Among employers with high initial markdowns, employment

[15] The sample sizes are somewhat reduced from Table 6 since the Industrial Dispute Act information is not available for all states.

and labor costs increase by 15% and 19%, respectively, and MRPL decreases by 11% relative to employers with low initial markdowns. The wage effect is positive but not statistically significant, similar to the heterogeneity by MRPL.

Dispersion in Markups. Existing studies in the misallocation literature, such as Edmond et al. (2015), suggest that opening to trade can reduce misallocation by reducing markup distortions. Foreign capital liberalization could also expose producers to greater competition in the product market. We thus examine the potential role of dispersion in markups in the impact of India's foreign capital liberalization on misallocation by estimating the effect of the reform on markups across firms with different levels of ex ante markup. Table 12 presents the results. We find that liberalization reduces markup distortions by 1-3%, depending on the functional form of the production function. The effect is essentially zero for the translog production function (column 1). The impact of liberalization on the markup for high-markup firms relative to low-markup firms is 3% at the 1% significance level under the Cobb-Douglas production function (column 2). However, this result should be taken cautiously and considered as suggestive evidence, given that our firm-level markup measure is biased because we estimate the revenue function instead of the output production function (Klette and Griliches, 1996; Bond et al., 2021).

5.4 Robustness

We focus on assessing the robustness of our main findings regarding the heterogeneous effects of the reform by ex ante MRPL. To achieve this, we conduct a comprehensive battery of robustness checks, including: (i) using an alternative control group, (ii) employing an alternative reference period in the event study analysis, (iii) adopting a Cobb-Douglas production function, and (iv) utilizing alternative measures of ex ante MRPL.

Alternative Control Group. In the baseline analysis, our treatment variable is based on a treatment group with industries liberalized during the 2006 reform and a control group with never-treated industries. Here, we check the robustness of our main results by including the 1991 reform industries in the control group like Bau and Matray (2023) and keeping the treatment group the same as the baseline. Table 13 reports the estimates of the heterogeneous effects of the policy from Equation (5) using the alternative treatment variable. As we add firms operating in 1991 reform industries in the control group, the observations increased by 46% from 18,784 to 27,362. The qualitative results remarkably remain the same.

Alternative Reference Period. To examine the robustness of our event study results and test the no anticipation effect assumption, we estimate the specification in Equation (6) by replacing the reference year of 2005-2006 with the 2004-2005 financial year. Figure 12 presents the results. Moving the base period to a year earlier does not change the qualitative findings. The validation of common pre-trends assumption even improves when using this alternative reference period. The immediate impact in the first period after the base year (or period "0") is statistically insignificant at the 5% level for the five outcomes, except for employment and labor costs. The estimated coefficients on employment and labor costs are marginally significant at period "0" in the event study graphs. However, overall, we consider that there is no substantial anticipation effect.

Cobb-Douglas Production Function. In our baseline analysis, we use markdown and MRPL based on the Translog production function, which is a more flexible functional form, both for the regressions and the measure of ex ante MRPL and markdowns. As we estimate the markdown using the Cobb-Douglas production function to check the robustness of our markdown estimate in Section 3.4, we also check the robustness of our regression results using markdown and MRPL based on Cobb-Douglas production function. The heterogeneous impacts among firms with high and low ex ante MRPL are shown in Table 14. The baseline results qualitatively stay the same when using this alternative production function form to measure wage markdowns and MRPL.

Alternative MRPL Measures. We further check the robustness of our results on the heterogeneous effects by ex ante MRPL using three different measures of ex ante MRPL.

MRPL as Revenue per Worker. Under the revenue Cobb-Douglas production function, MRPL $= \frac{\partial \text{Revenue}_{it}}{\partial L_{it}} = \alpha_j^l \frac{\text{Revenue}_{it}}{L_{it}}$, where α_j^l is the production parameter for the labor input. Assuming that all firms in an industry j share the same α_j^l, one can measure within-industry MRPL with $\frac{\text{Revenue}_{it}}{L_{it}}$ (Bau and Matray, 2023). First, we check the robustness of our results using this alternative measure of MRPL. As shown in Table 15, the heterogeneous effects of the FDI reform on employment, labor costs, average wages, and MRPL by ex ante MRPL are qualitatively the same when using this simple measure. Both the MRPL used for classifying firms into high- and low-MRPL firms and the MRPL used as an outcome are measured as the sales revenue per worker. Thus, the markdown and MRPL outcomes are not linked by construction in this particular analysis.

MRPL in 2001 as ex ante MRPL. The baseline analysis uses the average MRPL between 2000 and 2001 as ex ante MRPL to estimate heterogeneous effects by MRPL. The baseline regressions estimating heterogeneous effects by industry's ex ante foreign equity exposure are based on the share of foreign equities in 2001, which is the first period available in Prowess that reports firms' equity composition. To make these sets of regressions consistent, we use the 2001 value of MRPL to estimate the heterogeneous effects by ex ante MPRL. As shown in Table 16, the results are substantially robust.

Residualized Ex ante MRPL. Finally, we check the robustness of heterogeneous effects by ex ante MRPL using an alternative measure of ex ante MRPL, which is the baseline measure residualized with selected controls. Since the treatment is at the 4-digit level and ex ante MRPL is defined as the average MRPL over 2000-2001, we residualize the baseline measure using 2-digit industry fixed effects. Table 17 presents the results, and the baseline findings remain unchanged.

6 Worker Heterogeneity

This section extends the analysis by allowing worker heterogeneity, focusing on gender, and estimates the differential effects of foreign capital liberalization on male and female workers. We first estimate average effects using Equation (4), and Table 18 presents the results. Regressions are estimated on a reduced sample in which gender-specific markdowns have been measured. The sample is approximately one-fourth of our baseline sample. As shown in Panel A, the effects on employment, labor costs, and average wages are positive, and the effects on MRPL and markdowns are negative for male workers. However, these effects for men are not statistically significant. As reported in Panel B, the mean effects for female workers are quite different. The liberalization reduces employment and labor costs (columns 1 and 2) but increases average wages (column 3), with no average impact on MRPL and markdowns for female workers.

Next, we estimate heterogeneous effects by ex ante MRPL. Given that our measure of markdown and thus MRPL are gender-specific or computed for male and female workers separately from the production function with two types of labor. To estimate the heterogeneous effects, we compute the average firm-level ex ante MRPL as a weighted average of ex ante MRPLs for male and female workers,

$$\overline{\text{MRPL}}_i = s_i^f \text{MRPL}_i^f + (1 - s_i^f)\text{MRPL}_i^m, \tag{7}$$

where s_i^f is the share of female workers in the workforce at firm i, and MRPL_i^f and MRPL_i^m are MRPL for female and male workers, respectively. There is no time subscript because the variables are expressed at the ex ante level. Since the ex ante period is defined as 2000 and 2001 in our baseline analysis, the variables in Equation (7) are the average between these two years before the 2001 and 2006 episodes of the liberalization.[16]

The heterogeneous effects by firm-level ex ante average MRPL ($\overline{\text{MRPL}}_i$) are also gender-specific and are mainly concentrated among female workers (Table 19). The firms growing in size due to the inflows of foreign equities demand women more than men. The effects on the five outcomes for male workers (Panel A) are not statistically different from zero. The labor market effects of liberalization for female workers at firms with initially high MRPL (Panel B) are all in the same direction as the heterogeneous effects in Table 6. However, the estimated effect on labor costs is noisy and is not statistically significant. These results reveal that the reductions in labor misallocation and labor market distortions are driven by female workers, rather than male workers.

We then estimate heterogeneous effects by ex ante exposure to foreign capital for male and female workers. Table 20 presents the estimation results. First, for firms operating in industries with initially low exposure to foreign equity, liberalization affects men and women similarly. Foreign capital inflows increase employment and labor cost and reduce MRPL and markdowns for both

[16] Both gender-specific and gender-neutral markdowns and MRPL have been estimated for some firms. We calculate the relationship between the gender-neutral MRPL and weighted average MRPL. The pairwise correlation between the two firm-level MRPL measures is 0.41 (SE: 0.039, p-value: 0.000).

types of workers. In terms of magnitude, the coefficient of interest ($\widehat{\beta}_1$) in the employment and labor cost regressions are relatively the same for men and women (columns 1-2), and the estimated coefficients in the MRPL and markdown regressions for women are about twice as large as those for men (columns 4-5). The men's wage weakly increases at these firms, while the positive effect on women's wage is statistically insignificant (column 3). Second, the effects of liberalization among firms in industries with initially high exposure to foreign capital are quite different for men and women. For women, the effects among firms in industries with high and low exposure are highly heterogeneous. Female workers are reallocated from high-exposure to low-exposure firms, which drives the effect on labor costs.

Robustness. We check the robustness of heterogeneous effects by ex ante MRPL for male and female workers. First, Table A.2 presents the results when using an alternative control group and shows that our baseline results for heterogeneous workers are substantially robust. In this specification, the positive effect on labor costs for female workers became weakly significant at the 10% level. Second, as shown in Table A.3, the baseline results generally remain the same when using an alternative ex ante period, except that labor costs and markdown effects for women became statistically insignificant. Third, Table A.4 shows that our baseline results are remarkably robust to using a residualized ex ante MRPL, except that the effect on labor costs over female workers becomes statistically insignificant. Overall, the gender-specific effects are a bit noisy, potentially due to a smaller sample size; however, our baseline results are quite robust.

7 Conclusion

This paper examines the impacts of foreign capital integration on labor markets, focusing on labor misallocation among firms and imperfect competition in the labor markets. In doing so, we leverage the industry-time variation from India's 2006 foreign capital liberalization episode. The policy allowed some industries to receive automatic approval for foreign equity investments and raised caps on foreign equity. We introduce three innovations to the literature. First, we quantify wage and employment effects, which enables us to disentangle the effect on labor costs. Second, we investigate worker heterogeneity and focus on gender differences. Third, we examine the potential mechanisms through which foreign capital liberalization affects the misallocation of labor and provide direct evidence on the impacts on wage markdowns.

Our analysis shows that the resulting labor reallocation across firms reduced labor misallocation in treated industries by weakening the firm's labor market power—measured by wage markdowns using the production function approach. For firms with high ex ante MRPL, the inflow of foreign equity increased employment and labor costs relative to low MRPL firms. However, the policy has an insignificant impact on wages, indicating that the effect on labor costs is completely explained by the employment effect. MRPL and markdowns relatively declined for firms with high ex ante

MRPL. Analysis with heterogeneous workers suggests that the labor market and misallocation effects are driven by female workers.

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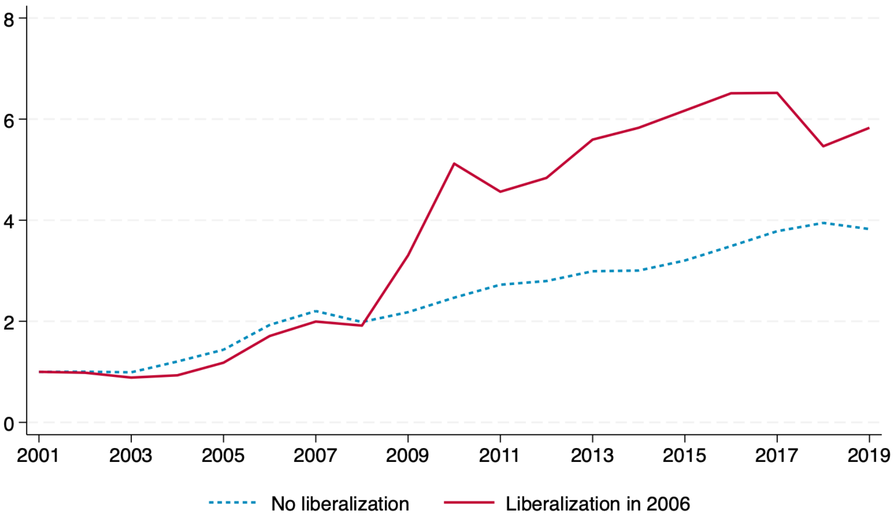
Figures

Figure 1: Foreign Equities

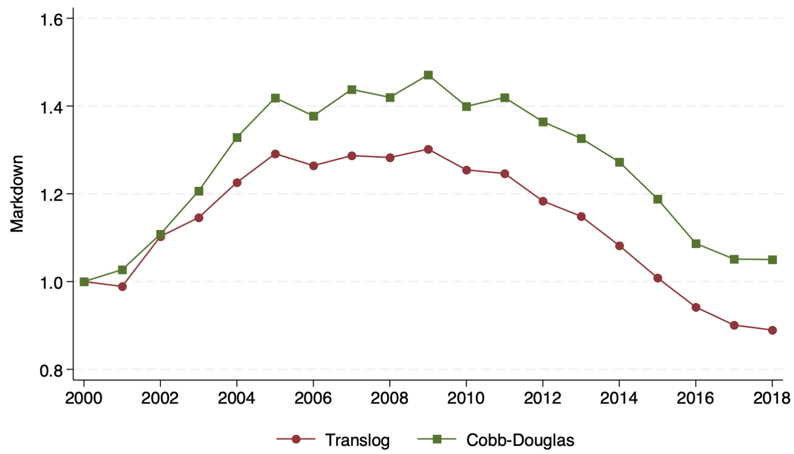


Notes: The figure plots the total amount of foreign equities in Prowess for industries that have been liberalized in 2006 (red solid line) and for those that have not been liberalized during the period 1995-2019 (blue short dashed line). The amounts for these two groups of industries are normalized to their 2001 level.

Figure 2: Trend of the Aggregate Markdown

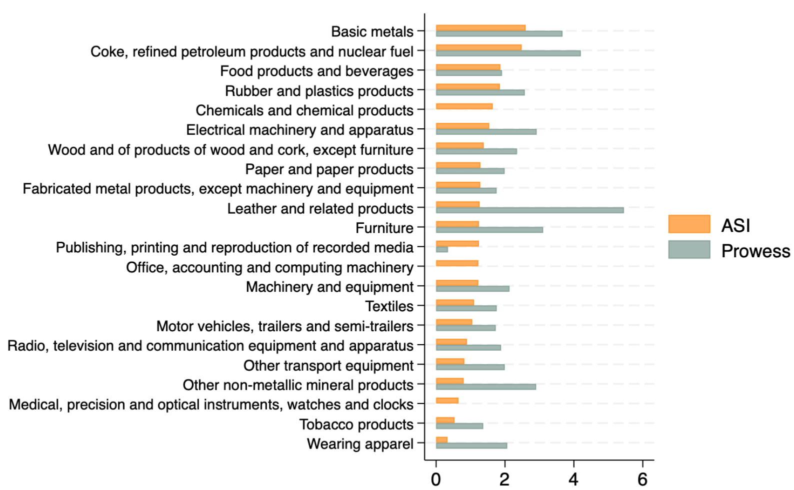
Notes: The establishment-level markdowns are constructed using the ASI data from 2000-2018 under the assumption of translog production where labor inputs are measured by headcount. The establishment-level markdowns are aggregated at the year level using employment shares of the labor market (combination of 4-digit NIC-1998 industry and states).

Figure 3: Trend of the Aggregate Markdown using ASI Data under Translog and Cobb-Douglas Specifications



Notes: The establishment-level markdowns are constructed using the ASI data from 2000-2018 under the assumption of translog and Cobb-Douglas production where labor inputs are measured by headcount. The establishment-level markdowns are aggregated at the year level using employment shares of the labor market (combination of 4-digit NIC-1998 industry and states).

Figure 4: Correlation between Median Wage Markdowns Estimated using ASI and Prowess Data



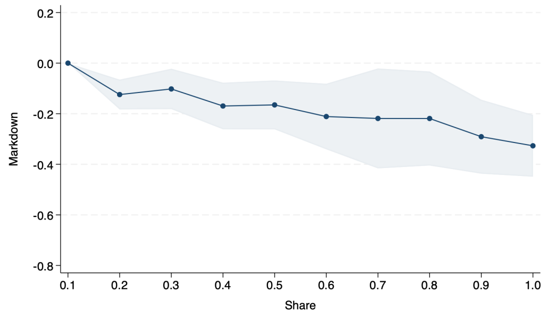
Notes: The figure plots the median wage markdowns at the two-digit NIC-1998 industry level where the firm-level markdowns are constructed using the ASI data from 2000-2018 (orange) and Prowess data from 2000-2019 (teal) under the assumption of translog production function. The median estimate is calculated using sampling weights provided in the data for the ASI data, while no weights are provided in Prowess data.

Figure 5: Trend of the Aggregate Markdown using Prowess Data under Translog and Cobb-Douglas Specifications

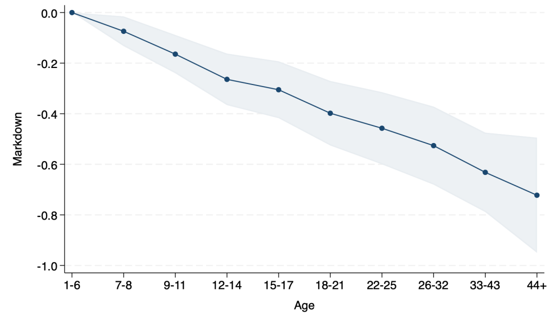
Notes: The figure depicts the aggregate markdowns (normalized to their value by 2000) from 2000-2019. The firm-level markdowns are constructed using Prowess data from 1995-2019 under the assumption of translog and Cobb-Douglas production, where labor inputs are measured by real labor cost. The firm-level markdowns are aggregated at the year level using employment shares of the labor market (combination of 2-digit NIC-1998 industry and districts).

Figure 6: Relationship between Markdown and Firm Characteristics (ASI Data)

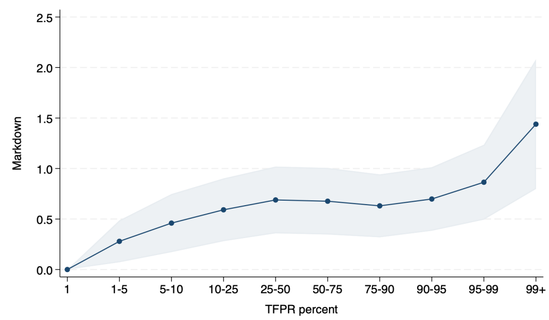
(a) Markdown-Size Relationship



(b) Markdown-Age Relationship



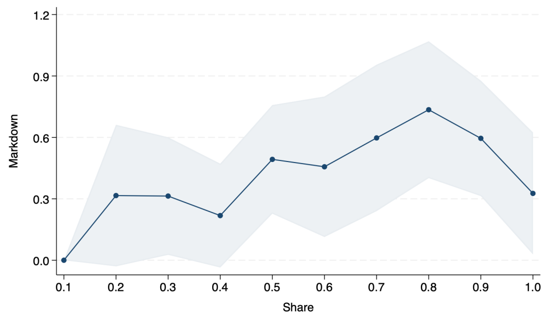
(c) Markdown-TFPR Relationship



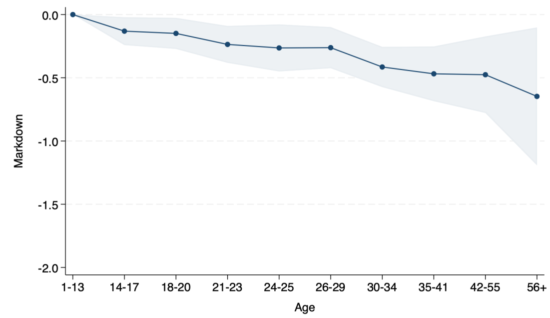
Notes: Based on the ASI data from 2000-2018, and 2000 is the financial year between 1 April 1999 and 31 March 2000. Panel (a) illustrates the point estimates and 95% confidence intervals from estimating establishment-level markdowns on size (measured by employment share) indicators. In the production function estimated separately for each two-digit industry group, labor inputs are measured by headcount. The regression controls for indicators for establishment age and 4-digit industry, state, and year fixed effects. The smallest size indicator is omitted, and thus coefficients reflect deviations relative to this reference group. The reference group labeled "0.1" includes establishments with employment shares $s \in (0, 0.1]$. Other indicator variables are similarly defined. Panel (b) shows the point estimates and 95% confidence intervals from estimating establishment-level markdowns on indicators of age deciles. The regression controls for indicators for establishment size and 4-digit industry, state, and year fixed effects. The first age decile is omitted; thus, coefficients reflect deviations relative to this reference group. Firm ages included in each decile are shown on a horizontal axis. Panel (c) shows the point estimates and 95% confidence intervals from estimating establishment-level markdowns on productivity. The regression controls for establishment size and 4-digit industry, state, and year fixed effects. The first percentile of productivity is omitted; thus, coefficients reflect deviations relative to this reference group. Standard errors (SEs) are clustered by 4-digit NIC-1998 industries. The qualitative results remain the same when the SEs are clustered at the state level (34 clusters).

Figure 7: Relationship between Markdown and Firm Characteristics (Prowess Data)

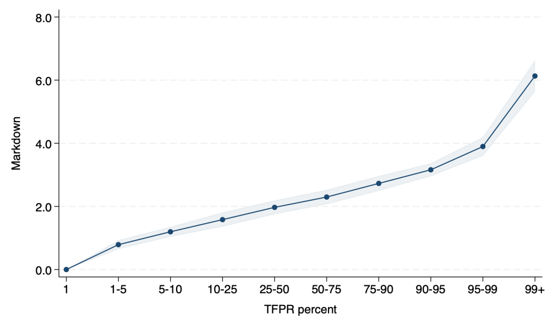
(a) Markdown-Size Relationship



(b) Markdown-Age Relationship

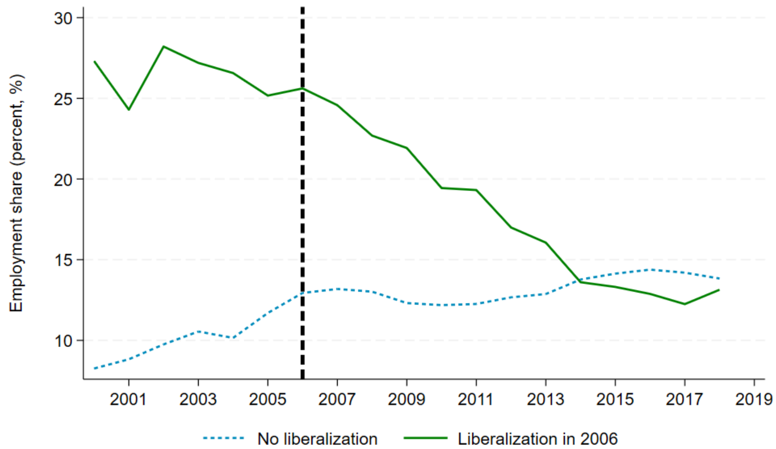


(c) Markdown-TFPR Relationship



Notes: Based on Prowess data from 1995-2019. Panel (a) illustrates the point estimates and 95% confidence intervals from estimating firm-level markdowns on size (measured by employment share) indicators. In the production function estimated separately for each two-digit industry group, labor inputs are measured by real labor cost. The regression controls for indicators for firm age and 2-digit industry, district, and year fixed effects. The smallest size indicator is omitted, and thus coefficients reflect deviations relative to this reference group. The reference group labeled "0.1" includes firms with employment shares $s \in (0, 0.1]$. Other indicator variables are similarly defined. Panel (b) shows the point estimates and 95% confidence intervals from estimating firm-level markdowns on indicators of age deciles. The regression controls for indicators for firm size and 2-digit industry, district, and year fixed effects. The first age decile is omitted; thus, coefficients reflect deviations relative to this reference group. Firm ages included in each decile are shown on a horizontal axis. Panel (c) shows the point estimates and 95% confidence intervals from estimating firm-level markdowns on productivity. The regression controls for firm size and 2-digit industry, district, and year fixed effects. The first percentile of productivity is omitted; thus, coefficients reflect deviations relative to this reference group. Standard errors (SEs) are clustered by districts. The qualitative results remain the same when the SEs are clustered at the 2-digit industry level (20 clusters), except for the markdown-size relationship, which becomes generally statistically insignificant, especially in the bottom part of the distribution.

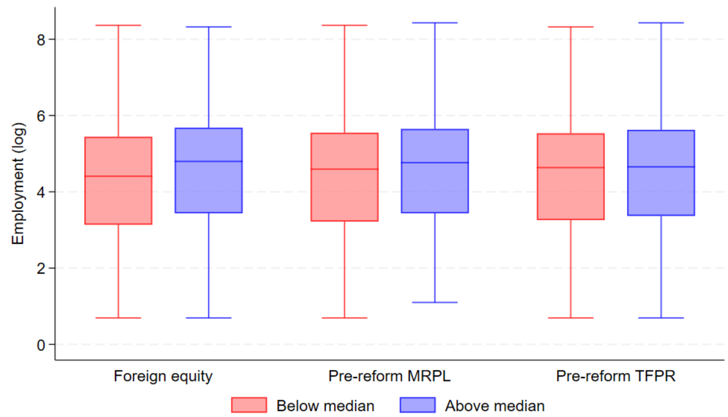
Figure 8: Trend in Share of Female Workers by Treatment Status



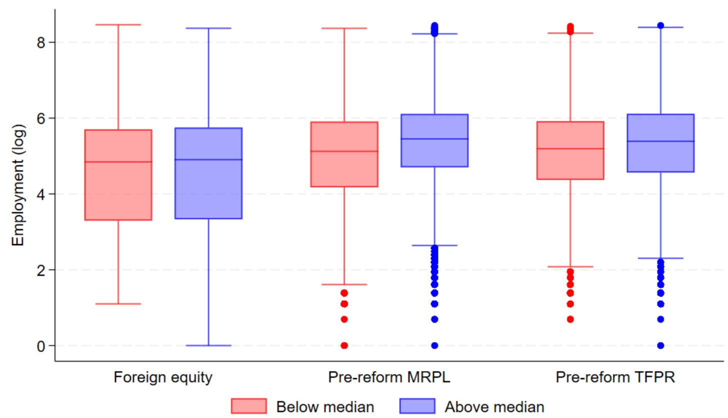
Notes: Based on ASI data. The figures plots the share of female workers at the establishment over time for industries that were liberalized in 2006 (green solid line) and for industries without liberalization (blue short dashed line).

Figure 9: Distribution of Firm-Level Employment by Initial Characteristic

(a) Before 2006 reform

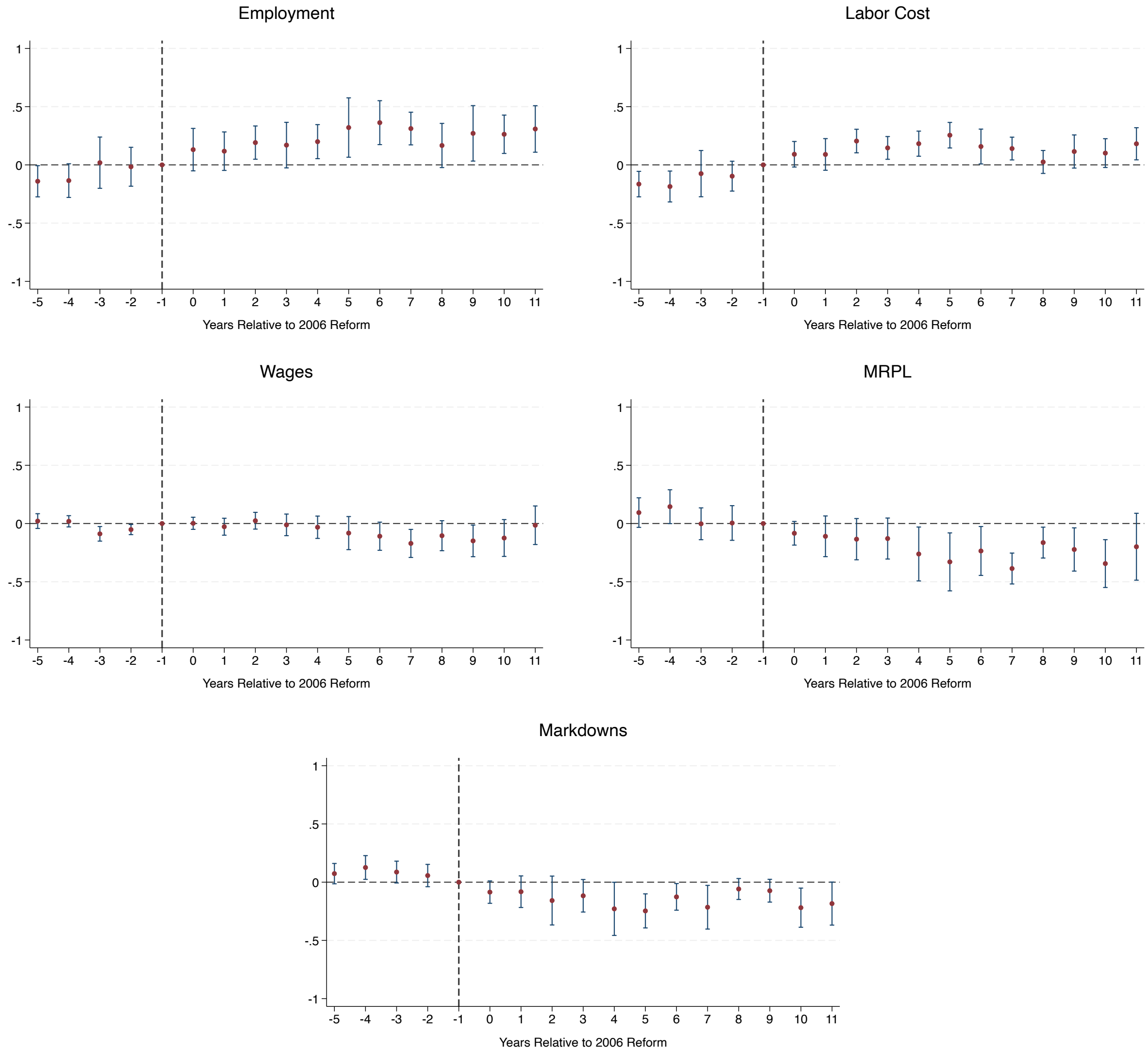


(b) After 2006 reform



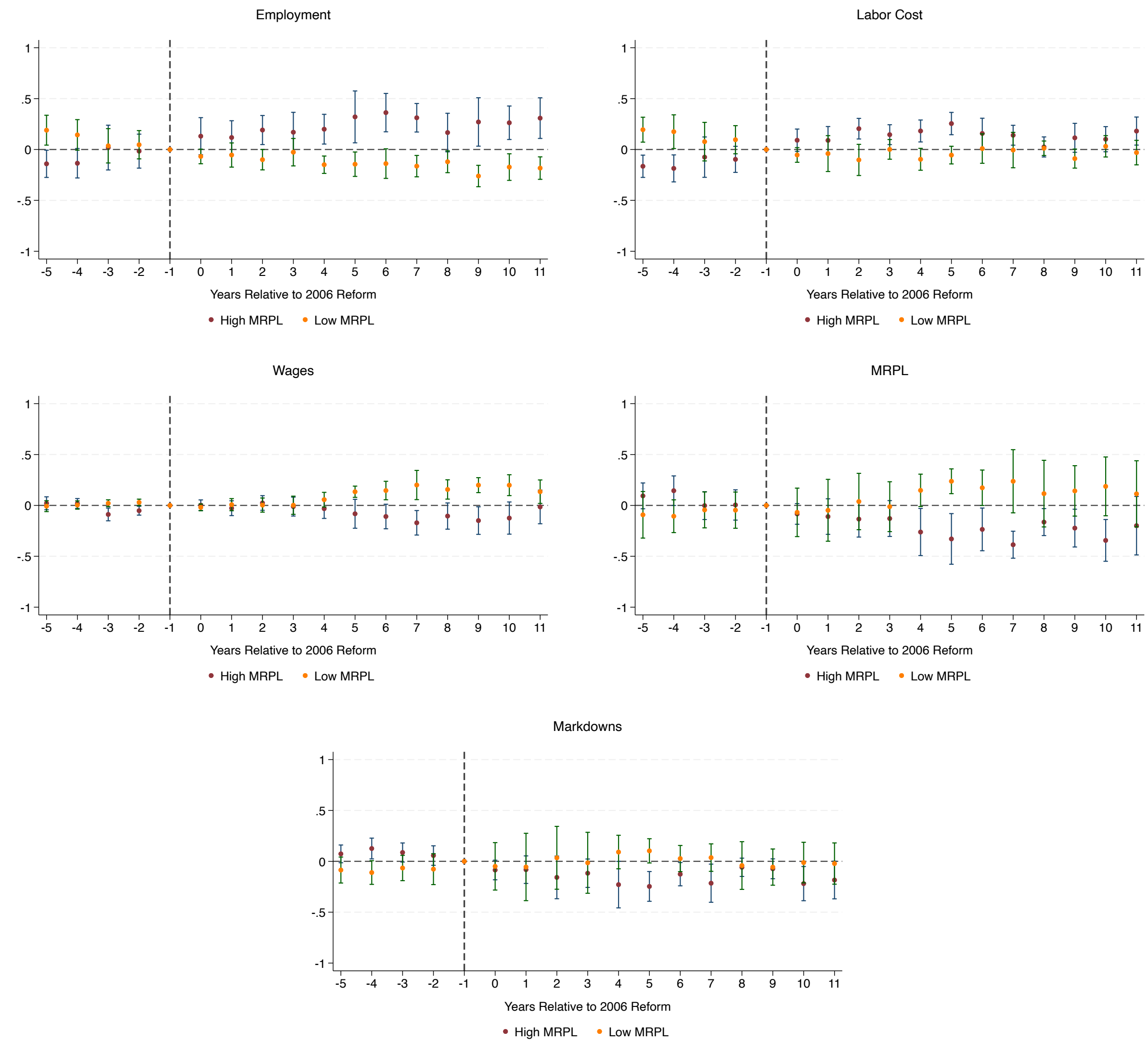
Notes: The figure plots the distribution of firm-level employment before and after the 2006 reform by initial characteristics. Based on the ASI data from 2000-2018 on which markdown has been estimated. The firms are classified based on whether they are below or above the median of initial share of foreign equity, marginal revenue product of labor (MRPL) and revenue productivity (TFPR).

Figure 10: Event Study: Relative Effects of Foreign Capital Liberalization on High MRPL Firms



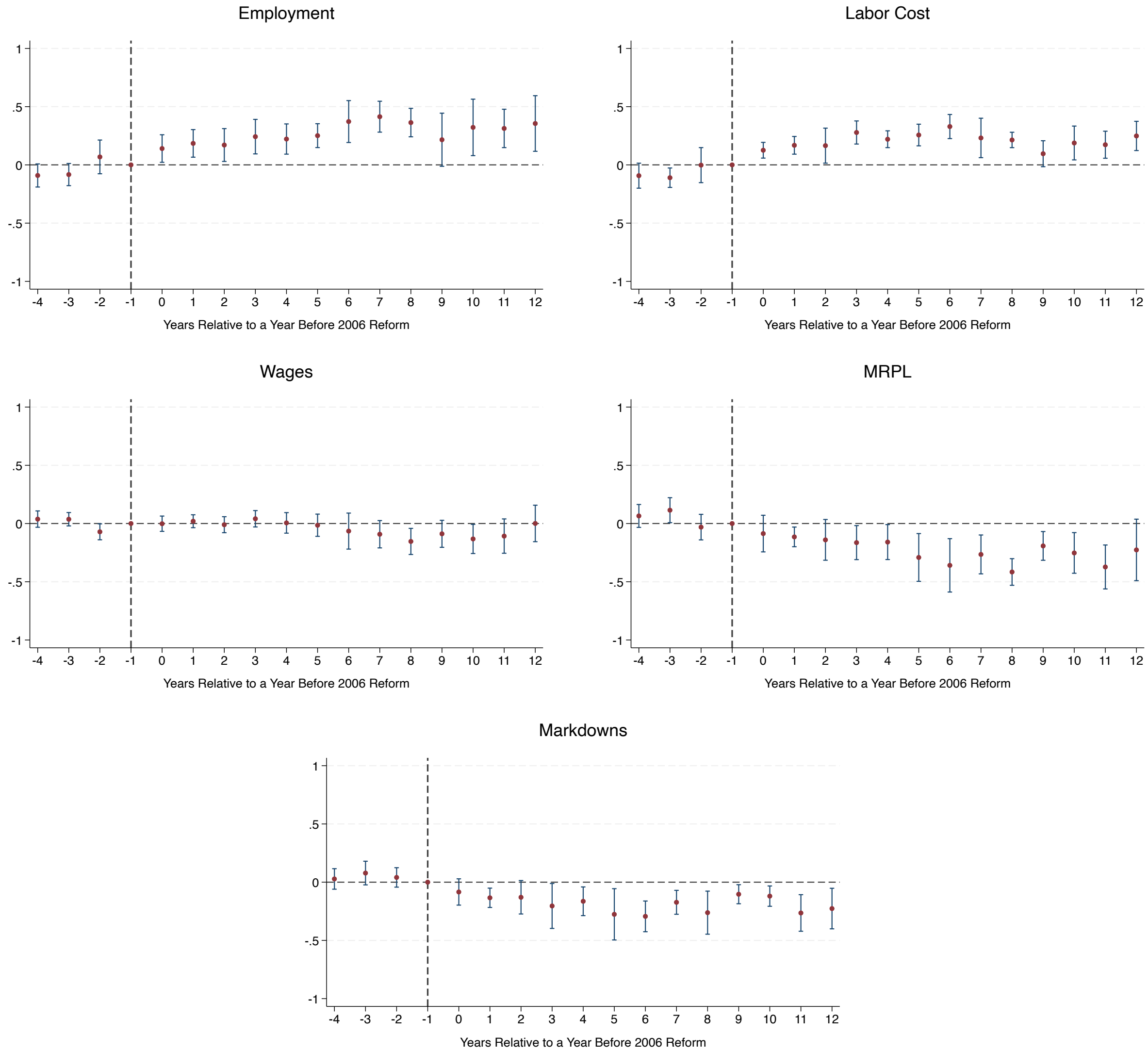
Notes: The figure plots results from event study regressions estimating the effects of India's 2006 FDI reform for high ex ante MRPL firms relative to low MRPL firms. Each point is the estimated coefficient $\beta_{1\tau}$ in regression (6) of employment, labor cost, average wage, MRPL, and markdown regressions, where τ represents the year relative to the reform period. All outcomes are in logs, and the reform is normalized to take place in year 0, which is the 2006-2007 financial year. Standard errors are two-way clustered at the 4-digit industry and year level, and 95% confidence intervals are shown.

Figure 11: Event Study: Relative Effects of Foreign Capital Liberalization on High and Low MRPL Firms



Notes: The figure plots results from event study regressions estimating the effects of India's 2006 FDI reform for high and low MRPL firms. Each point is the estimated coefficient $\beta_{1\tau}$ in regression (6) of employment, labor cost, average wage, MRPL, and markdown regressions, where τ represents the year relative to the reform period. All outcomes are in logs, and the reform is normalized to take place in year 0, which is the 2006-2007 financial year. Standard errors are two-way clustered at the 4-digit industry and year level, and 95% confidence intervals are shown.

Figure 12: Event Study: Relative Effects of Foreign Capital Liberalization on High MRPL Firms



Notes: The figure plots results from event study regressions estimating the effects of India's 2006 FDI reform for high ex ante MRPL firms relative to low MRPL firms. Each point is the estimated coefficient $\beta_{1\tau}$ in regression (6) of employment, labor cost, average wage, MRPL, and markdown regressions, where τ represents the year relative to the reform period. All outcomes are in logs, and the reform is normalized to take place in year 0, which is the 2005-2006 financial year. Standard errors are two-way clustered at the 4-digit industry and year level, and 95% confidence intervals are shown.

Tables

Table 1: Estimated Establishment-Level Markdowns in India's Manufacturing using ASI Data

Industry Group	Median	Mean	$IQR_{75\text{-}25}$	SD	N
Basic metals	2.598	3.124	2.800	2.021	15724
Coke, refined petroleum products and nuclear fuel	2.483	2.914	2.569	1.956	2978
Food products and beverages	1.863	2.283	1.864	1.622	44340
Rubber and plastics products	1.850	2.240	1.685	1.473	11642
Chemicals and chemical products	1.643	2.071	1.601	1.490	21023
Electrical machinery and apparatus	1.542	1.963	1.513	1.478	13021
Wood and of products of wood and cork, except furniture	1.383	1.630	1.118	1.091	5108
Paper and paper products	1.285	1.507	0.980	0.913	7833
Fabricated metal products, except machinery and equipment	1.281	1.582	1.139	1.120	13001
Leather and related products	1.262	1.701	1.414	1.419	6314
Furniture	1.245	1.634	1.232	1.340	8667
Publishing, printing and reproduction of recorded media	1.242	1.638	1.281	1.334	3491
Office, accounting and computing machinery	1.224	1.532	1.149	1.273	629
Machinery and equipment	1.223	1.580	1.269	1.233	18646
Textiles	1.103	1.450	1.150	1.224	26118
Motor vehicles, trailers and semi-trailers	1.047	1.216	0.739	0.757	11873
Radio, television and communication equipment and apparatus	0.893	1.294	1.122	1.234	3882
Other transport equipment	0.817	1.241	1.138	1.246	5894
Other non-metallic mineral products	0.798	1.105	0.897	0.973	27051
Medical, precision and optical instruments, watches and clocks	0.649	1.083	0.976	1.256	3549
Tobacco products	0.533	1.100	1.355	1.362	5071
Wearing apparel	0.327	0.582	0.475	0.757	11917
Whole sample	**1.314**	**1.750**	**1.502**	**1.474**	**267772**

Notes: Markdowns are estimated for 71,264 unique manufacturing establishments using the ASI data from 2000-2018 under the assumption of a translog specification for gross output, where 2000 is the financial year between 1 April 1999 and 31 March 2000. The labor inputs are measured by headcount in the production function, estimated separately for each two-digit industry group. Each manufacturing industry group corresponds to the manufacturing categorization of the National Industry Classification (NIC-1998) at the two-digit level. The distributional statistics are calculated using sampling weights provided in the data.

Table 2: Estimated Firm-Level Markdowns in India's Manufacturing using Prowess Data

Industry Group	Median	Mean	$IQR_{75\text{-}25}$	SD	N
Leather and related products	5.451	5.969	5.956	3.735	328
Coke, refined petroleum products and nuclear fuel	4.200	4.417	2.372	1.926	7100
Basic metals	3.668	4.160	2.967	2.471	4248
Furniture	3.107	3.372	2.408	1.942	796
Electrical machinery and apparatus	2.920	3.235	2.117	1.718	2755
Other non-metallic mineral products	2.905	3.817	3.550	3.112	1790
Rubber and plastics products	2.574	2.575	1.100	1.000	3646
Wood and of products of wood and cork, except furniture	2.348	2.658	1.775	1.420	262
Machinery and equipment	2.128	2.188	1.042	0.840	3818
Wearing apparel	2.063	2.900	2.500	2.426	904
Other transport equipment	1.991	3.057	3.399	2.871	472
Paper and paper products	1.987	2.285	1.864	1.454	1416
Food products and beverages	1.910	2.204	1.657	1.784	4326
Radio, television and communication equipment and apparatus	1.884	2.600	1.989	2.109	1803
Textiles	1.759	2.459	2.051	2.176	4714
Fabricated metal products, except machinery and equipment	1.754	1.964	1.496	1.206	1638
Motor vehicles, trailers and semi-trailers	1.731	1.969	1.223	1.275	3087
Tobacco products	1.369	1.441	1.776	1.417	116
Publishing, printing and reproduction of recorded media	0.342	0.836	0.168	1.278	33
Whole sample	**2.522**	**3.016**	**2.312**	**2.119**	**43252**

Notes: Markdowns are estimated for 4,778 unique manufacturing establishments using Prowess data from 1995-2019 under the assumption of a translog specification for gross output. The labor inputs are measured by the real labor cost in the production function, estimated separately for each 2-digit industry group. Each manufacturing industry group corresponds to the categorization of the National Industry Classification (NIC-1998) at the 2-digit level. The markdown was not estimated for two 2-digit manufacturing industries, including (i) office, accounting, and computing machinery, and (ii) medical, precision, and optical instruments, watches, and clocks, because no firms from these industries are included in Prowess data. The markdown was not estimated for the manufacturing of pharmaceuticals, medicinal chemical and botanical products due to data limitation.

Table 3: Summary Statistics

	No liberalized			Liberalized		
	Mean	SD	Median	Mean	SD	Median
Firm Age	22.1	19.3	17.0	25.2	18.6	21.0
Capital stock (log)	16.2	2.3	16.4	16.0	2.3	16.0
Sales revenue (log)	18.7	2.1	18.9	18.7	2.0	18.8
Employment (log)	4.6	1.5	4.8	4.7	1.3	4.9
Wage (log)	5.4	0.7	5.4	5.3	0.9	5.4
Markdown	1.6	1.4	1.2	1.8	1.5	1.4
MRPL (log)	5.6	0.9	5.6	5.6	1.1	5.8
N	234,653			12,151		

Notes: The table presents the summary statistics for firms' characteristics by treatment (FDI liberalization) status. The sample is based on the ASI data from 2000-2018 on which wage markdown has been estimated. An observation is at the firm-year level. Capital stock is the net fixed assets. Employment is the number of workers. Wage is the compensation per worker. The establishment-level wage markdowns are estimated under the assumption of translog production function in the manufacturing industry. The marginal revenue product of labor (MRPL) is computed by multiplying the average wage by establishment-level markdowns.

Table 4: Correlation between Industry Pre-Reform Characteristics and Reform Status

	Liberalized = 1					
	(1)	(2)	(3)	(4)	(5)	(6)
Log (Variance in MRPL)	0.008					
	(0.011)					
Log (Variance in MRPK)		0.008				
		(0.015)				
Log (Num. Firms)			0.001			
			(0.015)			
Log (Avg. Capital Stock)				-0.042		
				(0.031)		
Herfindahl Index					-0.099	
					(0.148)	
State-Owned Firms Share of Total Sales						-0.091
						(0.061)
N	130	130	130	130	130	129

Notes: The table presents the correlation between industry pre-reform characteristics and the liberalization status. The sample is based on the ASI data from 2000 to 2018, on which wage markdowns have been estimated. The marginal revenue product of labor (MRPL) is computed by multiplying the average wage by establishment-level markdowns. The marginal revenue product of capital (MRPK) is the ratio between sales revenue and capital. Capital stock is the real net fixed assets.

Table 5: Average Effects of the Foreign Capital Liberalization

	$\ln L_{it}$ (1)	$\ln W_{it}$ (2)	$\ln w_{it}$ (3)	$\ln MRPL_{it}$ (4)	$\ln \eta_{it}$ (5)
$\text{Post}_t \times \text{Reform}_j$	-0.037	0.009	0.052**	-0.016	-0.068
	(0.031)	(0.054)	(0.020)	(0.138)	(0.138)
N	18784	18784	18784	18784	18784
R^2	0.95	0.97	0.90	0.87	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the average effects of FDI liberalization on employment (headcount), labor cost, average wage, marginal revenue product of labor (MRPL), and markdown in columns 1–5, respectively. All dependent variables are in logs. The treatment is a dummy variable indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function in the manufacturing industry. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 6: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante MRPL

	$\ln L_{it}$ (1)	$\ln W_{it}$ (2)	$\ln w_{it}$ (3)	$\ln MRPL_{it}$ (4)	$\ln \eta_{it}$ (5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.166***	0.143**	0.012	-0.134***	-0.146***
	(0.048)	(0.066)	(0.025)	(0.035)	(0.033)
$\text{Post}_t \times \text{Reform}_j$	-0.118**	-0.061	0.047*	0.050	0.003
	(0.043)	(0.077)	(0.026)	(0.142)	(0.136)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.131***	0.113***	-0.057***	-0.143***	-0.086**
	(0.042)	(0.035)	(0.013)	(0.042)	(0.032)
N	18784	18784	18784	18784	18784
R^2	0.95	0.97	0.90	0.87	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions that estimate the heterogeneous effects of FDI liberalization by firms' ex ante marginal revenue product of labor (MRPL). The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function. The MRPL is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 7: Effect of the Foreign Capital Liberalization by Firms' Ex ante MRPL on Revenue and Capital

	$\ln Revenue_{it}$ (1)	$\ln K_{it}$ (2)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	-0.046	-0.142
	(0.105)	(0.123)
$\text{Post}_t \times \text{Reform}_j$	0.108	0.141
	(0.094)	(0.134)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.052	0.032
	(0.032)	(0.043)
N	18784	18784
R^2	0.96	0.95

Notes: The table presents the results from OLS regressions that estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL on revenue (column 1) and capital (column 2). All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 8: Heterogeneous Effects of the Foreign Capital Liberalization by Labor Regulation

	$\ln L_{it}$ (1)	$\ln W_{it}$ (2)	$\ln w_{it}$ (3)	$\ln MRPL_{it}$ (4)	$\ln \eta_{it}$ (5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}} \times \text{Labor Regulation}_s$	-0.057**	-0.081***	-0.016***	0.071**	0.087**
	(0.020)	(0.018)	(0.004)	(0.031)	(0.037)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.008	-0.019	0.036	0.017	-0.019
	(0.043)	(0.049)	(0.027)	(0.092)	(0.102)
N	16395	16395	16395	16395	16395
R^2	0.94	0.96	0.90	0.87	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions that estimate the heterogeneous effects of FDI liberalization among high ex ante MRPL firms by labor regulation. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. Post_t is an indicator variable equal to 1 if the year is after 2006. Reform_j is an indicator variable equal to 1 if the industry has liberalized access to the international capital market in the 2006 liberalization episode and 0 if the industry is never reformed. Firms are classified as high MRPL if their average MRPL in the pre-treatment period from 2000 to 2001 is above the 4-digit industry median. MRPL is calculated by multiplying the average wage by the markdown. The markdown is estimated under the assumption of a translog production function. $\text{Labor Regulation}_s$ is the pre-treatment cumulative net score of Besley-Burgess index that measures the friendliness of the labor market to workers in 2001. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 9: Heterogeneous Effects on Foreign Equity Inflows by Firms' Ex ante Exposure to Foreign Equity

	Dependent variable: Foreign equities	
	(1)	(2)
$\text{Post}_t \times I_j^{\text{High Exposure}}$	156.635**	83.434**
	(64.771)	(37.182)
$\text{Post}_t \times \text{Reform}_j \times I_j^{\text{High Exposure}}$		-177.661***
		(46.066)
$\text{Post}_t \times \text{Reform}_j$		126.280***
		(20.561)
Year FE	✓	✓
4-digit Industry FE	✓	✓
N	26683	14144
R^2	0.79	0.67

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization on foreign equities by industry-level ex ante exposure to foreign equities. The ex ante foreign equity exposure is based on the fraction of foreign equities shares in total equity shares, defined at the 4-digit industry (NIC-2004) in 2001. The foreign equity exposure is high (low) if the fraction is above (below) the within-industry median. The outcome variable is the total foreign equities at the 4-digit level in million Rupees. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 10: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante Exposure to Foreign Equity

	$\ln L_{it}$ (1)	$\ln W_{it}$ (2)	$\ln w_{it}$ (3)	$\ln MRPL_{it}$ (4)	$\ln \eta_{it}$ (5)
$\text{Post}_t \times \text{Reform}_j \times I_j^{\text{Low Exposure}}$	-0.058	0.044	0.005	-0.490***	-0.495***
	(0.078)	(0.096)	(0.036)	(0.081)	(0.084)
$\text{Post}_t \times \text{Reform}_j$	-0.007	-0.025	0.021	0.120**	0.100
	(0.057)	(0.069)	(0.023)	(0.057)	(0.060)
$\text{Post}_t \times I_j^{\text{Low Exposure}}$	0.073	0.053	0.014	-0.030	-0.044
	(0.058)	(0.066)	(0.040)	(0.074)	(0.064)
N	9019	9019	9019	9019	9019
R^2	0.94	0.97	0.91	0.87	0.81

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by industry-level ex ante exposure to foreign equities. The ex ante exposure to foreign equity is the percentage of foreign equities in total equities, defined at the 4-digit industry (NIC-2004) in 2001. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 11: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante Markdowns

	$\ln L_{it}$ (1)	$\ln W_{it}$ (2)	$\ln w_{it}$ (3)	$\ln MRPL_{it}$ (4)	$\ln \eta_{it}$ (5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High Markdown}}$	0.146**	0.187***	0.016	-0.106*	-0.122**
	(0.058)	(0.064)	(0.033)	(0.059)	(0.048)
$\text{Post}_t \times \text{Reform}_j$	-0.103**	-0.075	0.045	0.032	-0.013
	(0.045)	(0.066)	(0.029)	(0.154)	(0.149)
$\text{Post}_t \times I_i^{\text{High Markdown}}$	0.091**	0.122**	0.027	-0.069*	-0.095**
	(0.039)	(0.047)	(0.023)	(0.036)	(0.040)
N	18767	18767	18767	18767	18767
R^2	0.95	0.97	0.90	0.87	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante markdowns. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 12: Heterogeneous Effects of the Foreign Capital Liberalization on Markups by Firms' Ex ante Markup

	Translog (1)	Cobb Douglas (2)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High Markup}}$	0.009	-0.031***
	(0.014)	(0.010)
$\text{Post}_t \times \text{Reform}_j$	0.140	0.162
	(0.115)	(0.115)
$\text{Post}_t \times I_i^{\text{High Markup}}$	-0.019**	-0.029***
	(0.007)	(0.010)
N	18764	18753
R^2	0.80	0.80

Notes: Based on ASI data from 2002 to 2018, on which markups have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante markup. The outcomes in columns 1 and 2 are log markups estimated under the assumption of a translog and a Cobb Douglas production function, respectively. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: $*p < 0.10$, $**p < 0.05$, and $***p < 0.01$.

Table 13: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante MRPL (Alternative Control Group)

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.149***	0.154**	0.032	-0.067*	-0.099***
	(0.045)	(0.061)	(0.025)	(0.034)	(0.025)
$\text{Post}_t \times \text{Reform}_j$	-0.145***	-0.109	0.036*	0.039	0.003
	(0.042)	(0.074)	(0.019)	(0.137)	(0.131)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.135***	0.095***	-0.069***	-0.184***	-0.115***
	(0.034)	(0.027)	(0.014)	(0.038)	(0.027)
N	27362	27362	27360	27360	27362
R^2	0.95	0.97	0.90	0.85	0.82

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is the alternative treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated and 1991 reform industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: $*p < 0.10$, $**p < 0.05$, and $***p < 0.01$.

Table 14: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante MRPL (Cobb-Douglas Production Function)

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.229***	0.197***	-0.015	-0.225***	-0.210***
	(0.037)	(0.051)	(0.028)	(0.040)	(0.056)
$\text{Post}_t \times \text{Reform}_j$	-0.158***	-0.095	0.061**	0.109	0.047
	(0.033)	(0.064)	(0.026)	(0.099)	(0.092)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.137***	0.123***	-0.049**	-0.168***	-0.119***
	(0.039)	(0.031)	(0.019)	(0.050)	(0.040)
N	18775	18775	18775	18775	18775
R^2	0.95	0.97	0.90	0.89	0.85

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a Cobb-Douglas production function. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: $*p < 0.10$, $**p < 0.05$, and $***p < 0.01$.

Table 15: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante MRPL ($MRPL = Revenues/L$)

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.173***	0.138*	-0.011	-0.215**	-0.180***
	(0.049)	(0.069)	(0.024)	(0.085)	(0.038)
$\text{Post}_t \times \text{Reform}_j$	-0.126**	-0.062	0.060**	0.234***	0.027
	(0.051)	(0.083)	(0.027)	(0.076)	(0.122)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.141***	0.123***	-0.046***	-0.116***	-0.094***
	(0.033)	(0.030)	(0.007)	(0.035)	(0.030)
N	18764	18764	18764	18764	18764
R^2	0.95	0.97	0.90	0.91	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL. The marginal revenue product of labor (MRPL) is measured by sales revenue per worker. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: $*p < 0.10$, $**p < 0.05$, and $***p < 0.01$.

Table 16: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante MRPL (Ex ante Period = 2001)

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.212***	0.153**	-0.022	-0.205***	-0.183***
	(0.064)	(0.066)	(0.020)	(0.043)	(0.037)
$\text{Post}_t \times \text{Reform}_j$	-0.145***	-0.066	0.067**	0.077	0.010
	(0.049)	(0.075)	(0.024)	(0.142)	(0.133)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.119***	0.111***	-0.050***	-0.104**	-0.054*
	(0.037)	(0.031)	(0.017)	(0.043)	(0.031)
N	16829	16829	16829	16829	16829
R^2	0.94	0.97	0.90	0.87	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL. The ex ante MRPL is defined as the 2001 value of MRPL. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: $*p < 0.10$, $**p < 0.05$, and $***p < 0.01$.

Table 17: Heterogeneous Effects of the Foreign Capital Liberalization by Firms' Ex ante MRPL (Residualized Ex ante MRPL)

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.163***	0.134**	0.013	-0.125**	-0.138***
	(0.042)	(0.056)	(0.025)	(0.044)	(0.040)
$\text{Post}_t \times \text{Reform}_j$	-0.116***	-0.056	0.046	0.045	-0.001
	(0.038)	(0.071)	(0.027)	(0.142)	(0.135)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.114**	0.104***	-0.055***	-0.135***	-0.080**
	(0.041)	(0.034)	(0.014)	(0.042)	(0.031)
N	18784	18784	18784	18784	18784
R^2	0.95	0.97	0.90	0.87	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL. The ex ante MRPL is residualized by 2-digit industry fixed effects. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdown, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns are estimated under the assumption of a translog production function in the manufacturing industry. The marginal revenue product of labor (MRPL) is calculated by multiplying the average wage by the establishment-level markdowns. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: $*p < 0.10$, $**p < 0.05$, and $***p < 0.01$.

Table 18: Average Effects of the Foreign Capital Liberalization on Male and Female Workers

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
	Panel A. Male workers				
$\text{Post}_t \times \text{Reform}_j$	0.157	0.141	0.053	-0.146	-0.199
	(0.113)	(0.098)	(0.040)	(0.098)	(0.121)
N	4126	4126	4126	4126	4126
R^2	0.94	0.96	0.88	0.82	0.82
	Panel B. Female workers				
$\text{Post}_t \times \text{Reform}_j$	-0.241	-0.236**	0.084**	0.152	0.068
	(0.152)	(0.100)	(0.033)	(0.217)	(0.221)
N	4126	4126	4126	4126	4126
R^2	0.89	0.92	0.88	0.87	0.83

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns over male and female workers have been estimated. The table presents the results from OLS regressions, which estimate the average effects of FDI liberalization on employment (headcount), labor cost, average wage, MRPL, and markdowns for male (Panel A, columns 1–5) and female (Panel B, columns 1–5) workers. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns over male and female workers are estimated under the assumption of a translog production function with heterogeneous workers. The marginal revenue product of labor (MRPL) for male and female workers is calculated by multiplying the average wage for each type of worker by the establishment-level markdowns for that type of worker. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 19: Heterogeneous Effects of the Foreign Capital Liberalization on Male and Female Workers by Firms' Ex ante MRPL

	$\ln L_{it}$ (1)	$\ln W_{it}$ (2)	$\ln w_{it}$ (3)	$\ln MRPL_{it}$ (4)	$\ln \eta_{it}$ (5)
	Panel A. Male workers				
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	-0.009	0.075	0.186	-0.018	-0.204
	(0.111)	(0.190)	(0.121)	(0.097)	(0.216)
$\text{Post}_t \times \text{Reform}_j$	0.147	0.081	-0.062	-0.117	-0.055
	(0.112)	(0.187)	(0.096)	(0.136)	(0.226)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.137**	0.084	-0.104***	-0.148*	-0.044
	(0.062)	(0.057)	(0.034)	(0.083)	(0.087)
N	4126	4126	4126	4126	4126
R^2	0.94	0.96	0.88	0.82	0.82
	Panel B. Female workers				
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$	0.363***	0.355	0.086	-0.325***	-0.410***
	(0.107)	(0.229)	(0.136)	(0.003)	(0.127)
$\text{Post}_t \times \text{Reform}_j$	-0.512***	-0.499***	0.034	0.380*	0.346
	(0.130)	(0.171)	(0.090)	(0.205)	(0.273)
$\text{Post}_t \times I_i^{\text{High MRPL}}$	0.207**	0.181**	-0.075*	-0.060	0.014
	(0.074)	(0.070)	(0.038)	(0.062)	(0.066)
N	4126	4126	4126	4126	4126
R^2	0.89	0.92	0.88	0.87	0.83

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns over male and female workers have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by firms' ex ante MRPL. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdowns for male (Panel A) and female (Panel B) workers, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns over male and female workers are estimated under the assumption of a translog production function with heterogeneous workers. The marginal revenue product of labor (MRPL) for male and female workers is calculated by multiplying the average wage for each type of worker by the establishment-level markdowns for that type of worker. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 20: Heterogeneous Effects of the Foreign Capital Liberalization on Male and Female Workers by Firms' Ex ante Exposure to Foreign Equity

	$\ln L_{it}$	$\ln W_{it}$	$\ln w_{it}$	$\ln MRPL_{it}$	$\ln \eta_{it}$
	(1)	(2)	(3)	(4)	(5)
	Panel A. Male workers				
$\text{Post}_t \times \text{Reform}_j \times I_j^{\text{Low Exposure}}$	0.499***	0.528***	0.098*	-0.321**	-0.419***
	(0.165)	(0.144)	(0.048)	(0.114)	(0.115)
$\text{Post}_t \times \text{Reform}_j$	-0.209	-0.293**	-0.030*	0.037	0.067
	(0.120)	(0.119)	(0.016)	(0.120)	(0.115)
$\text{Post}_t \times I_j^{\text{Low Exposure}}$	-0.154	-0.271**	-0.047	-0.058	-0.011
	(0.143)	(0.115)	(0.050)	(0.124)	(0.130)
N	1453	1453	1453	1453	1453
R^2	0.93	0.96	0.93	0.80	0.79
	Panel B. Female workers				
$\text{Post}_t \times \text{Reform}_j \times I_j^{\text{Low Exposure}}$	0.541***	0.526***	0.059	-0.555**	-0.614**
	(0.150)	(0.142)	(0.062)	(0.243)	(0.215)
$\text{Post}_t \times \text{Reform}_j$	-0.347**	-0.374***	0.052	0.368*	0.316
	(0.133)	(0.095)	(0.059)	(0.206)	(0.201)
$\text{Post}_t \times I_j^{\text{Low Exposure}}$	-0.092	-0.224**	-0.049	-0.124	-0.075
	(0.131)	(0.082)	(0.063)	(0.216)	(0.178)
N	1453	1453	1453	1453	1453
R^2	0.92	0.94	0.92	0.90	0.84

Notes: Based on ASI data from 2002 to 2018, on which wage markdowns over male and female workers have been estimated. The table presents the results from OLS regressions, which estimate the heterogeneous effects of FDI liberalization by industry-level ex ante exposure to foreign equities. The ex ante exposure to foreign equity is the percentage of foreign equities in total equities, defined at the 4-digit industry (NIC-2004) in 2001. The outcomes in columns 1–5 are employment (headcount), labor cost, average wage, MRPL, and markdowns for male (Panel A) and female (Panel B) workers, respectively. All dependent variables are in logs. The treatment is our baseline treatment variable, a dummy indicating the 2006 FDI reforms, with never-treated industries in the control group. The establishment-level wage markdowns over male and female workers are estimated under the assumption of a translog production function with heterogeneous workers. The marginal revenue product of labor (MRPL) for male and female workers is calculated by multiplying the average wage for each type of worker by the establishment-level markdowns for that type of worker. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.

Table 21: Effect of the Foreign Capital Liberalization on TFPR

	$\ln TFPR_{it}$ (1)	$\ln TFPR_{it}$ (2)	$\ln TFPR_{it}$ (3)
$\text{Post}_t \times \text{Reform}_j$	0.003	0.020	0.041
	(0.014)	(0.026)	(0.040)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPL}}$		-0.036	-0.046
		(0.037)	(0.039)
$\text{Post}_t \times \text{Reform}_j \times I_i^{\text{High MRPK}}$			-0.036
			(0.028)
N	18784	18784	18784
R^2	0.57	0.57	0.57

Notes: The table presents the results from OLS regressions that estimate the average effect and the heterogeneous effect of FDI liberalization on log firm-level TFPR. TFPR is measured by estimating revenue production function using the methodology of Ackerberg et al. (2015) under the assumption of translog functional form. Post_t is an indicator variable equal to 1 if the year is after 2006. Reform_j is an indicator variable equal to 1 if the industry has liberalized access to the international capital market in the 2006 liberalization episode and 0 if the industry is never reformed. Firms are classified as high MRPL (MRPK) if their average MRPL (MRPK) in the pre-treatment period from 2000 to 2001 is above the 4-digit industry median. MRPL is calculated by multiplying the average wage by the markdown. The markdown is estimated under the assumption of a translog production function. MRPK is estimated with the $Revenue/K$ method. All specifications control for firm, firm age, and firm pre-treatment size-by-year fixed effects. Standard errors are two-way clustered at the 4-digit industry and year level and are presented in parentheses. Significance: *$p < 0.10$, **$p < 0.05$, and ***$p < 0.01$.