Family Leave Law and the Demand for Female Labor: Evidence from a Trade Shock[1]

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Abstract

This paper estimates a labor demand effect of mandated, job-protected family leave on female workers. We use confidential microdata of matched employers and employees for the universe of U.S. non-farm, private sector firms and separately identify firms required to provide leave under the 1993 Family and Medical Leave Act (FMLA) and firms that are exempt (non-FMLA). We study the difference in the demand for female workers between FMLA and non-FMLA firms, following a trade-induced, exogenous labor demand shock. We find that the demand for female relative to male workers is lower at FMLA firms compared to non-FMLA firms in response to the trade shock. This difference is most pronounced for less than college-educated female workers and women in their childbearing ages. The difference is mitigated at firms with female managers.

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Keywords: job-protected leave, gender gap, regression discontinuity, trade shock, gender norms

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1. Introduction

Job-protected family leave programs offer individual workers the ability to balance competing family and career responsibilities. Women tend to shoulder a larger share of family and child care tasks than men, and are thus more likely to avail these leave. On the one hand, job-protected leave provisions may result in an increase in the supply of female workers. On the other hand, employers may reduce the demand for female workers if the costs associated with hiring temporary replacement workers or coordinating work schedules of existing employees exceed the benefits from higher retention of workers who would otherwise have quit. This paper uses confidential, microdata on firms and employees in the United States to isolate a labor demand impact of mandated, job-protected family leave on female workers exploiting a large scale, trade-induced labor demand shock.

We study the U.S. Family and Medical Leave Act (FMLA) enacted in 1993 at the federal level that mandates employers to provide twelve weeks of unpaid, job-protected family leave to qualifying workers. An employer is mandated to provide leave if they employ at least 50 employees within a 75 mile radius of the employment location.[2] The firm size threshold provides a sharp regression discontinuity setting to study the labor demand impact of FMLA. We implement a regression discontinuity difference-in-differences (RD-DD) design to identify a labor demand impact of FMLA by evaluating responses at FMLA and non-FMLA firms before and after the surge in Chinese imports following China's accession to the World Trade Organization in 2001 (Autor, Dorn, Hanson, 2013).

We examine how the share of female workers and the share of female earnings at FMLA and non-FMLA firms change between 2000 and 2003 across industries that were exposed to varying levels of the Chinese import competition shock during this period. We attribute the differential change in response between FMLA and non-FMLA firms to a labor demand effect of FMLA. We find that, when faced with the same Chinese import competition shock within an industry, FMLA firms exhibit

[2] The employment threshold is the minimum of the federal and any state level family and medical leave provisions.

differentially lower shares of female workers and lower share of female earnings than non-FMLA firms.

Our analyses utilize confidential, administrative microdata on the universe of all non-farm, private sector firms and their employees sourced from the Longitudinal Employer-Household Dynamics (LEHD) at the U.S. Census Bureau. The data contain information on workers' quarterly earnings by employer and demographic characteristics. We aggregate worker level outcomes - employment and earnings - differentiated by gender at the firm-state level combined with measures of Chinese import competition at the firm's six-digit NAICS. We identify firm FMLA status using the total number of workers employed at a firm in a given state. A firm is categorized as FMLA if it employs 50 or more workers in a state and non-FMLA if it employs less than 50 workers in a state. In our analyses, we only consider firms with a single establishment in a given state such that the final sample contains firm-states that can be assigned FMLA status based on the employment criteria only.

We find that, when faced with a trade-induced labor demand shock resulting from a rise in Chinese import competition between 2000 and 2003, firms required to provide job-protected leave under FMLA exhibit lower shares of female workers and shares of female earnings compared to non-FMLA firms. A one percentage point larger increase in Chinese import penetration is associated with a 2.16 (2.83) percentage points smaller increase or larger decrease in the growth rate of the percentage of female employment (earnings) at FMLA compared to non-FMLA firms.

Further, we find that the lower differential demand for female labor is concentrated among female workers between ages 25 and 45. Women in this age group are in their childbearing years and are also more likely to have young children. Separating female workers into college-educated and less than college-educated, we find that the negative differential labor demand effect is concentrated on the less than college-educated group. Finally, we study firms with female managers (defined as top three earners) separately from firms with no female managers. We find that that the negative differential labor demand effect is concentrated at firms with no female managers, indicating a possible role of women in decision-making positions in ensuring gender equity at firms.

Our work contributes to three strands of literature. First, we contribute to the large literature that has examined the impact of family leave policies on women's U.S. labor market outcomes.[3] Family leave policies have been found to be associated with increased leave-taking and higher incidence of return to work after childbirth (Berger and Waldfogel, 2004; Rossin-Slater, Ruhm, Waldfogel, 2013; Ruhm, 1997; Waldfogel, 1998). However, leave policies have been associated with a small or statistically insignificant increase in employment and decrease in wages (Baum, 2003a; Blau, Klerman, Leibowitz, 1997). Studies analyzing the impact of FMLA in particular find that it had small and statistically insignificant effects on employment and wages (Baum, 2003b; Waldfogel, 1999) but a negative impact on promotions (Thomas, 2016).

A potential concern in these studies is that labor demand effects cannot be causally distinguished from the labor supply effects of leave policy. Our paper focuses on isolating and identifying a labor demand impact of job-protected leave provisions by exploiting the shock to labor demand from import competition following China's accession to the WTO. Moreover, earlier work relied on self-reported survey data on wages and employer size that suffered from small sample sizes. We utilize administrative data on the universe of private-sector employers in the United States and their workers, thereby providing a more comprehensive picture of the impact of FMLA.

Second, we contribute to the nascent literature examining firm responses to family-friendly workplace policies. Rossin-Slater (2018) notes that there are very few studies focusing on firm responses to maternity and family leave policies. The few papers that have begun to examine the impacts of leave taking on U.S. firms have found no significant impacts on firm's total wage costs or

[3] See Olivetti and Petrongolo (2017) for a survey of the impacts of family policies on gender outcomes in high-income countries.

turnover rates, but they face identification and sample size challenges (Bedard and Rossin-Slater, 2016; Bartel, Rossin-Slater, Ruhm, Waldfogel, 2016). The lack of evidence of leave take-up affecting firm's total wage costs and employee turnover rates as well as output, gross profits and survival are corroborated in Danish data (Brenøe, Canaan, Harmon, Royer, 2018). However, these studies do not separately examine the impact of leave provision on firm's relative demand for female to male workers as an outcome. We are able to causally identify a demand effect of job-protected leave on female relative to male workers in response to a trade shock. We provide, what we believe to be, new evidence on the impact of job-protected family leave on labor demand for female workers by employers.

Finally, our paper contributes to the literature on the impact of trade liberalization on gender gaps. The empirical evidence point to trade liberalization as an attenuating force, however, the mechanisms vary by country. For example, Gaddis and Pieters (2014) document that the comprehensive tariff cuts across Brazil led to a reduction in the gender gap through a larger relative decrease in male labor force participation and employment. Juhn, Ujhelyi, and Villegas-Sanchez (2014) find that tariff reductions associated with the North American Free Trade Agreement (NAFTA) induced Mexican firms to enter export markets, adopt new technologies that lower the demand for physically demanding tasks, and replace blue-collar male with blue-collar female workers. Hakobyan and McLaren (2018) find the opposite where NAFTA tariff reductions reduced the relative wage growth for married blue-collar women in areas of the U.S. that faced stiffest competition from Mexican imports. Our results underscore the role of labor market policies in mediating the impact of trade shocks on domestic labor market outcomes. Specifically, we highlight that the impact of trade on differential outcomes for women relative to men interact with provisions for family-friendly leave policies.

The rest of the paper is organized as follows. Section 2 provides a brief history and description of the provisions under FMLA and presents a simple conceptual framework to interpret our results.

Section 3 describes our identification strategy and empirical specification. Section 4 describes the data and construction of key variables, and provides summary statistics of our analysis sample. Section 5 discusses validity of the regression discontinuity design, main results, and robustness checks. Section 6 explores the potential mechanisms proposed in Section 2.2. Section 7 offers an extension to our main results. The final section concludes.

2. Background and Conceptual Framework

2.1 The 1993 Family and Medical Leave Act

The 1993 FMLA entitles eligible employees of covered employers to take up to twelve weeks of unpaid, job-protected leave for family and medical reasons. Covered employers refer to all private-sector firms with 50 or more workers located within 75 miles.[4] An eligible employee must have worked for a covered employer for at least twelve months with a minimum of 1,250 hours of service preceding the leave. Leave under FMLA can be taken for care of a new-born, sick family member (spouse, child, or parent), employee's own sickness, or a family member with active military status.[5]

While both male and female employees can avail of leave under FMLA, survey evidence suggests that female compared to male workers are more likely to utilize FMLA for caregiving purposes. In particular, women with young children are considerably more likely to be leave-takers relative to men with young children.[6] FMLA does not require paid leave, however, the employer must continue providing any health insurance coverage. The employers must also bear the cost of replacing

[4] Covered employers also include all public agencies and all public or private elementary or secondary schools. In this study, we focus only on the universe of non-farm, private-sector employers.
[5] In addition to childbirth, adoption and foster care are also eligible events. Eligible employees may take up to twenty-six weeks of leave to care for a family member on active military duty.
[6] The Survey of Employees in 2000, conducted by the U.S. Department of Labor, finds that 76 (45) percent of women (men) with children 18 months or younger take leave under FMLA. Accessed on October 22, 2018 at https://www.dol.gov/whd/fmla/chapter4.htm.

the employee for the period of leave, either by hiring a temporary employee or by spreading the employee's workload across other existing employees.

2.2 Conceptual Framework

In this section, we sketch a theoretical framework to highlight the trade-offs faced by an employer when required to provide job-protected family leave to its employees. Consider a firm operating in perfectly competitive product and labor markets. We posit a constant elasticity of substitution production function where the firm employs female and male labor, L_F and L_M, respectively, as follows:

$$Y = [L_F^{\rho} + L_M^{\rho}]^{\frac{1}{\rho}} \qquad (2.1),$$

where ρ is the elasticity of substitution between male and female workers and $0 < \rho < 1$. Female and male workers earn a market wage, w_F and w_M, respectively. Suppose that the firm believes that with exogenous positive probabilities, g_F and g_M, female and male workers will be called upon to meet caregiving responsibilities. Moreover, $g_M = \tau g_F$ where $0 < \tau < 1$. The term, τ, captures the firm's beliefs that men are strictly less likely to take leave to fulfil caregiving responsibilities.

The firm provides a leave of duration l to its employees for caregiving purposes. If an employee is on leave, the firm must replace the employee at a replacement cost of $r(l)w_j$ where $j = F, M$ and $r'(l) > 0$. Using a panel survey of California businesses in 2003 and 2008, Dube, Freeman, and Reich (2010) document that employee replacement costs can be substantial relative to annual wages at between approximately USD 3,000 and 4,500 (in 2003 dollars) per recruit. Assume further that with probability $h(l)$, the employee may not return to work from leave, where $h'(l) < 0$. In the event that the workers quits, the firm loses the worker's firm-specific human capital vw_j for its lifetime and discounted at the rate δ. The firm then maximizes profits,

$$\max_{L_F, L_M, l} \left(p[L_F^{\rho} + L_M^{\rho}]^{\frac{1}{\rho}} - w_F L_F - w_M L_M \right) - \lambda(l)(w_F g_F L_F + w_M \tau g_F L_M) \qquad (2.2).$$

The product price is p and $\lambda(l)$ is a loss function as follows,

$$\lambda(l) = [r(l) + \frac{\delta}{1-\delta}h(l)v] \quad (2.3).$$

The loss function (2.3) captures the trade-offs the firm must weigh when choosing optimal leave duration l^*. Choosing higher l^* means higher replacement costs but also lower probability of losing the worker and the worker's firm-specific human capital for the future. The first-order conditions are given by,

$$\lambda'(l^*) = \left[r'(l^*) + \frac{\delta}{1-\delta}h'(l^*)v\right](w_F g_F L_F^* + w_M \tau g_F L_M^*) = 0 \quad (2.4),$$

$$w_F(1 + \lambda(l^*)g_F) = p[L_F^\rho + L_M^\rho]^{\frac{1}{\rho}-1} L^{*\rho-1}_F \quad (2.5),$$

$$w_M(1 + \lambda(l^*)\tau g_F) = p[L_F^\rho + L_M^\rho]^{\frac{1}{\rho}-1} L^{*\rho-1}_M \quad (2.6).$$

Hence,

$$\left(\frac{L_F^*}{L_M^*}\right)^{\rho-1} = \frac{w_F(1+\lambda(l^*)g_F)}{w_M(1+\lambda(l^*)\tau g_F)} \quad (2.7).$$

We assume that a unique solution for l^*, L_F^*, L_M^* exists and that the second-order conditions (SOC) are satisfied. This implies that $\lambda''(l^*) > 0$. Taking the natural log of (2.7), we get

$$(\rho - 1)ln\frac{L_F^*}{L_M^*} = ln\frac{w_F}{w_M} + ln\frac{(1+\lambda(l^*)g_F)}{(1+\lambda(l^*)\tau g_F)} \quad (2.8).$$

Totally differentiating (2.8), we get

$$(\rho - 1)dln\frac{L_F^*}{L_M^*} = dln\frac{w_F}{w_M} + dln\frac{(1+\lambda(l^*)g_F)}{(1+\lambda(l^*)\tau g_F)} \quad (2.9).$$

Let the relative market labor supply response of female to male labor be given by[7],

$$\eta_{s,F/M} = \frac{dln\frac{L_F}{L_M}}{dln\frac{w_F}{w_M}} > 0 \quad (2.10).$$

[7] We abstract from modelling the impact of leave provision on female and male labor supply given our focus on isolating the labor demand effects of FMLA. The implicit assumption is that any effect of leave provision on labor supply does not vary differentially by worker gender.

Labor market equilibrium requires that relative female labor demand equals relative female labor supply. Substituting for the relative equilibrium market wage from (2.10), we get

$$(\rho-1)dln\frac{L_F^*}{L_M^*}=\frac{1}{\eta_{s,F/M}^*}dln\frac{L_F^*}{L_M^*}+dln\frac{(1+\lambda(l^*)g_F)}{(1+\lambda(l^*)\tau g_F)} \quad (2.11).$$

Also, we can re-write[8],

$$dln\frac{(1+\lambda(l^*)g_F)}{(1+\lambda(l^*)\tau g_F)}=\frac{\mathrm{d}\lambda(l^*)g_F(1-\tau)}{(1+\lambda(l^*)g_F)(1+\lambda(l^*)\tau g_F)} \quad (2.12).$$

Hence, (2.11) can be written as,

$$\left(\rho-1-\frac{1}{\eta_{s,F/M}^*}\right)dln\frac{L_F^*}{L_M^*}=\frac{\mathrm{d}\lambda(l^*)g_F(1-\tau)}{(1+\lambda(l^*)g_F)(1+\lambda(l^*)\tau g_F)} \quad (2.13).$$

Finally, we can express the change in relative demand for female labor as follows:

$$dln\frac{L_F^*}{L_M^*}=\frac{\eta_{s,F/M}^*}{\left(\eta_{s,F/M}^*\right)((\rho-1)-1)}\frac{\mathrm{d}\lambda(l^*)g_F(1-\tau)}{(1+\lambda(l^*)g_F)(1+\lambda(l^*)\tau g_F)} \quad (2.14).$$

Proposition 1: A mandate of job-protected leave induces employers to decrease the relative demand for female workers.

Suppose that the government mandates a leave duration of $\bar{l}$. If $l^*\geq \bar{l}$ then the mandate has no effect on leave provision or employment. If the mandate is binding and $\bar{l}>l^*$, then the SOC implies that $\lambda(\bar{l})>\lambda(l^*)$ for $\bar{l}>l^*$, or that $d\lambda(l^*)>0$ for $dl^*>0$. Together with (2.10), $0<\tau<1$, and $0<\rho<1$, (2.14) implies that $dln\frac{L_F^*}{L_M^*}<0$. Also, from (2.10), $dln\frac{w_F^*}{w_M^*}=\frac{1}{\eta_{s,F/M}^*}(dln\frac{L_F^*}{L_M^*})<0.$

Proposition 2: The impact of a mandated job-protected leave on the relative demand for female workers varies by worker skill.

[8] $dln\frac{(1+\lambda(l^*)g_F)}{(1+\lambda(l^*)\tau g_F)}=\frac{(1+\lambda(l^*)\tau g_F)}{(1+\lambda(l^*)g_F)}\frac{(1+\lambda(l^*)\tau g_F)\mathrm{d}\lambda(l^*)g_F-(1+\lambda(l^*)g_F)\mathrm{d}\lambda(l^*)\tau g_F}{(1+\lambda(l^*)\tau g_F)^2}$

$=\frac{(1+\lambda(l^*)\tau g_F)\mathrm{d}\lambda(l^*)g_F-(1+\lambda(l^*)g_F)\mathrm{d}\lambda(l^*)\tau g_F}{(1+\lambda(l^*)g_F)(1+\lambda(l^*)\tau g_F)}=\frac{(1+\lambda(l^*)\tau g_F)\mathrm{d}\lambda(l^*)g_F-(1+\lambda(l^*)g_F)\mathrm{d}\lambda(l^*)\tau g_F}{(1+\lambda(l^*)g_F)(1+\lambda(l^*)\tau g_F)}.$

From (2.4) we get $\frac{d\lambda'(l^*)}{dv} = \frac{\delta}{1-\delta} h'(l^*)(w_F g_F L_F^* + w_M \tau g_F L_M^*) < 0$. Hence, higher values of v correspond to lower values of $\mathrm{d}\lambda(l^*)$. We hypothesize that high- compared to low-skilled workers are more likely to accrue firm-specific human capital v, for instance, by receiving on-the-job training (Altonji and Spletzer, 1991). Therefore, within this framework, any negative impact on the relative demand for female workers is mitigated (magnified) for high- (low-) skilled workers as long as relatively higher (lower) replacement costs do not offset the respective firm-specific skill advantages.

Proposition 3: A larger gender differential between the perceived likelihoods of taking leave to fulfil caregiving responsibilities (smaller τ) is associated with a more negative effect of mandated, job-protected leave on the relative demand for female workers.

This follows directly from how τ enters equation (2.14). An implication of Proposition 3 is that the effect of a mandated, job-protected leave on relative labor demand for women is more likely to be exacerbated during childbearing ages, when likelihood of leave taking is higher.

3. Identification Strategy and Empirical Specification

Our main identification strategy is to exploit the employment cut-off requirement for employers to offer FMLA in a sharp RD design. Further, a key contribution of our study is to isolate a labor demand impact of FMLA. Our argument is that several previous studies that found no effect of family leave on outcomes for women may have conflated a labor supply and labor demand effect.[9] To achieve this goal, we augment the basic RD set-up with a difference-in-differences approach. Specifically, we compare FMLA firms (have 50 or more workers) to non-FMLA firms (have less than 50 workers), before and after a large-scale, exogenous trade-induced labor demand shock.[10]

[9] This is confirmed in the regressions presented in Appendix Table A1, where we estimate the direct effect of firm FMLA status on total male and total female employment at the firm-level. The coefficient on firm FMLA status is statistically insignificant and we cannot reject the null hypothesis that FMLA is not related to employment.

[10] Results in Appendix Table A1 show that the change in Chinese import penetration was associated with significantly lower male and female employment at firms in our sample.

The surge in Chinese exports to the U.S. in the early 2000s was exogenous in that it was predominantly a result of internal Chinese supply shocks and falling global trade barriers rather than U.S. demand shocks (Autor, Dorn, Hanson, 2013). Nonetheless, to capture the supply-driven component of Chinese import penetration in U.S. industries, we check the robustness of our results by instrumenting for trade exposure using Chinese exports to other non-U.S., high-income countries (Acemoglu, Autor, Dorn, Hanson, Price, 2016).

The baseline empirical specification is as follows:

$$\Delta Y_i = \beta_0 + \beta_1 FMLA_{i,2000} + \beta_2 \Delta IP_j + \beta_3 \left(FMLA_{i,2000} * \Delta IP_j \right) + \beta_4 f(r_i) + \beta_5 \left(f(r_i) * FMLA_{i,2000} \right) + \beta_6 \left(f(r_i) * \Delta IP_j \right) + \beta_7 \left(f(r_i) * FMLA_{i,2000} * \Delta IP_j \right) + \beta_8 X_{i,2000} + \beta_9 X_{k,2000} + \delta_s + \Delta \varepsilon_i \quad (3.2.1).$$

The outcome variable of interest at firm i is denoted by Y_i. The firm is located in state s and operates in six-digit industry j within a four-digit industry k. Δ denotes change between 2000 and 2003. We focus on the change between 2000 and 2003 to minimize potential labor supply effects of the trade shock that are more likely to manifest over a longer time horizon (Autor, Dorn, and Hanson, forthcoming). Our main outcomes of interest are female share of employment and female share of earnings, respectively, as follows:

$$\Delta Y_i \equiv ln\left[\left(\frac{y_f}{y_f + y_m}\right) * 100\right]_{i,2003} - ln\left[\left(\frac{y_f}{y_f + y_m}\right) * 100\right]_{i,2000} \quad (3.2.2)$$

where y denotes employment (earnings) of female workers at firm i; y_f and y_m indicate the total number (earnings) of female and male employees at firm i, respectively. We also consider the female share of employment and female share of earnings differentiated by age and skill categories, c, as follows,

$$\Delta Y_i \equiv ln\left[\left(\frac{y_{fc}}{y_f + y_m}\right) * 100\right]_{i,2003} - ln\left[\left(\frac{y_{fc}}{y_f + y_m}\right) * 100\right]_{i,2000} \quad (3.2.3).$$

We group workers into the following age categories: 24 and less, 25-45, and 46+.[11] We group workers into the following skill categories: high-skilled (with bachelor's or higher degrees) and low-skilled (with less than bachelor's degree).

[*Note: In ongoing work, we are examining two additional outcome variables - changes in female share of hires and changes in female share of promotions.*]

The measure of Chinese import penetration within a six-digit NAICS industry j is given by IP_j. Firm i's FMLA status denoted by $FMLA_{i,2000}$ is determined as of the initial year, 2000, of the analysis. $X_{i,2000}$ is a set of firm-level controls as of 2000 that includes firm age and multi-unit status. All our regressions also include the share of female workers $X_{k,2000}$ in a four-digit NAICS industry, k, and a set of state fixed effects, δ_s. The rating variable, r_i, is centered around the employment cut-off such that $r_i = (employment_i - 50)$, and ε_i is an idiosyncratic error term.

The control function, $f(r_i)$, is a continuous *n*-order polynomial function of the rating variable on each side of the cut-off point. Our preferred specification uses a second-order polynomial. The control function captures the relationship between firm size and the outcome variables. Including an interaction between the rating variable and FMLA status accounts for the fact that FMLA status may impact both the intercept and the slope of the regression line (Jacob and Zhu, 2012). We allow the slope to vary on either side of the cut-off. In our baseline analyses, we focus on the sample of firms closest to the employment cut-off - firms with 45-55 workers. Standard errors are clustered at the four-digit NAICS level.

The coefficient of interest is β_3 which isolates the differential effect of an increase in Chinese import competition on FMLA compared to non-FMLA firms. We attribute the differential effect to

[11] Age bins were selected based on 1999 fertility rates in the United States (Ventura, Martin, Curtin, Menacker, and Hamilton, 2001; Table 3).

a labor demand effect of the job-protected leave mandate (FMLA) on female workers. We expect $\beta_3 < 0$ following Proposition 1 in Section 2.2.

4. Data Description

We utilize the Longitudinal Employer Household Dynamics (LEHD) data on the universe of U.S. non-farm, private sector firms and their employees (McKinney and Vilhuber, 2014). Employers are identified at the state-level and for purposes of this study we use the term firm interchangeably with firm-state. Worker level information includes quarterly earnings and demographic characteristics including gender, age, and education. We restrict attention to workers with a strong labor force attachment. We begin by restricting the sample to workers that have three or more quarters of reported earnings at an employer. We then construct a measure of annual earnings for each worker. First, we compute average quarterly earnings using earnings in all quarters worked. Then, we only retain quarters where earnings are at least equal to half of the average quarterly earnings. We re-compute average quarterly earnings by summing the earnings in quarters where earnings are at least equal to or greater than half of the average earnings. Finally, we annualize the adjusted quarterly earnings to construct annual earnings for a worker employed at a given firm. This restriction further ensures that lower earnings for a worker does not reflect periods of leave-taking.

Measures of Chinese import penetration at the six-digit NAICS level are constructed using information on imports and exports from Schott (2008) and domestic output from the CES-NBER manufacturing productivity database.[12]

4.1 FMLA status

Firms are required to provide unpaid, job-protected leave to eligible employees under FMLA if they employ 50 or more workers within 75 miles. The employment threshold may vary by state.

[12] Accessed at http://faculty.som.yale.edu/peterschott/sub_international.htm and http://www.nber.org/nberces/.

Some states mandate more generous employment thresholds and firms operating in these states must use the state-specific threshold to determine FMLA provision.[13] We exclude these states in our current analyses. We construct a binary FMLA status indicator at the firm-state level using the Longitudinal Business Database (LBD). The LBD consists of data on all private, non-farm U.S. establishments in existence that have at least one paid employee (Jarmin and Miranda, 2002).

In order to facilitate comparison of firms of similar sizes, we focus on firms with a single establishment in a state. Thus, FMLA status is determined solely based on firm-state employment. A single-unit firm is assigned a non-zero FMLA status if it employs 50 or more workers and zero otherwise. A multi-unit firm is assigned FMLA status in a similar manner using its firm-state employment. The restriction of only considering firms with single establishments in a state ensures that the distance rule does not determine FMLA status. Using the distance rule could potentially result in FMLA firms where employment is less than 50 and hence invalidate the RD setup. Thus, the baseline analysis is conducted on a sample of single-unit firms and only those multi-unit firms that have a single establishment in a state and employ between 45 and 55 workers. [*Note: In ongoing work, we are constructing a propensity score matched estimate on the sample of all multi-unit firms to exploit the fact that with the same number of total employees, multi-unit firms may be FMLA (treatment) or non-FMLA (control) based on the distance between their establishments*.]

4.2 Chinese import penetration

We construct a measure of trade exposure, following Acemoglu, Autor, Dorn, Hanson, and Price (2016), as change in Chinese import penetration for a U.S. manufacturing industry over the period 2000 and 2003. The measure is given as follows (expressed in percentage changes):

$$\Delta IP_j = \frac{\Delta M_j^{U.S.-CHINA}}{Y_{j,1997}+M_{j,1997}-E_{j,1997}}, \qquad (4.2.1)$$

[13] The following states have a more generous employment threshold than the federal mandate: Maine (15+), Massachusetts (6+), Minnesota (21+), Oregon (25+), Vermont (10+), and Washington DC (20+).

where $M_j^{U.S.-CHINA}$ is Chinese imports to the U.S. by six-digit NAICS industry j, $M_{j,1997}$, $E_{j,1997}$, and $Y_{j,1997}$ are total U.S. imports, total U.S. exports, and total U.S. domestic production in 1997 by six-digit NAICS industry j, respectively. Δ denotes change between 2000 and 2003.

There may be concerns that (4.2.1) reflects domestic shocks to U.S. industries that affect U.S. import demand rather than being driven purely by supply shocks within China. To isolate the supply-driven component of Chinese exports to the U.S., in robustness checks, we instrument for trade exposure with the following:

$$\Delta IPO_j = \frac{\Delta M_j^{OTH.-CHINA}}{Y_{j,1997}+M_{j,1997}-E_{j,1997}}. \tag{4.2.2}$$

$\Delta M_j^{OTH.-CHINA}$ is the growth in imports from China in industry j between 2000 and 2003 in eight other high-income countries excluding the United States.[14] High-income countries are similarly exposed to growth in imports from China that are driven by domestic supply shocks in China.

4.3 Summary statistics

Table 1 provides descriptive statistics of our main outcome variables for FMLA and non-FMLA firms that are close to the employment cut-off of 50: firms that employ between 45 and 55 workers. We compare the following outcomes: total number of female workers as a percentage of total workers at the firm, total earnings of female workers as a percentage of total earnings of all workers at the firm, total number of female workers, total number of male workers, firm age, and firm multi-unit status. The first two columns display the mean values and associated standard errors in parentheses. The third column reports the difference in the mean between FMLA and non-FMLA firms and the associated t-statistic from testing whether the difference is statistically significant.

We can see that FMLA and non-FMLA firms are very similar in their share of female workers at roughly 41 percent. This is also true for the share of female earnings at roughly 34 percent. The

[14] The eight countries are: Australia, Denmark, Finland, Germany, Japan, New Zealand, Spain, and Switzerland.

average FMLA (non-FMLA) firm in our sample employs about 20 (16) women and 29 (25) men. The difference in total female and male workers at FMLA and non-FMLA firms is statistically significant which is not surprising given that FMLA status is determined by firm size. We also see that non-FMLA firms tend to be a year younger, on average, compared to FMLA firms. FMLA firms are also more likely be part of a multi-unit firm. In all our regressions we control for firm age and multi-unit status as of 2000 and also confirm the robustness of our results using the sample of single-unit firms only.

5. Results

This section presents checks of the internal validity of the RD design, baseline results, and robustness checks.

*5.1 Preliminary check*s

In this section we present two validity checks for the RD design to ensure that there are no significant jumps in the rating variable at the cut-off. First, we examine whether the density of the rating variable is continuous at the discontinuity. We perform a McCrary density test of the density of the rating variable. [*Note: we will disclose this result in the next draft.*] We also provide a count of the number of firms within each employment bin, from 45 to 55, in Figure 1. We observe no significant bunching at the cut-off point. In fact, there is a slight increase in the number of firms that employ 50 workers, the employment threshold at which firms are mandated to offer job-protected leave under FMLA.

To select the appropriate functional form of the rating variable and its interactions, $f(r_i)$, in our parametric regression, we compute the Akaike Information Criterion (AIC) (Lee and Lemieux, 2010). The AIC captures the bias-precision trade-off of using a more complex model specification and thus measures the relative goodness of fit of a statistical model. Table 2 reports the difference in the AIC between the linear and other three specifications where linear is coded as a baseline of 0. We can see that when the dependent variable is the change in share of female employment, the linear has the lowest AIC, although the AIC associated with the quadratic specification is only 0.54 larger than the AIC associated with the linear specification. When the dependent variable is the change in share of female earnings, the quadratic specification displays the lowest AIC and by a large margin. Taken together, we select the second-order polynomial as our preferred specification.

5.2 Baseline regression discontinuity and difference-in-differences

Table 3 presents our baseline results from estimating equation (3.2.1). The top panel presents results for the change in the log female share of employment and the bottom panel for the change in the log female share of earnings. The coefficient on FMLA x ΔIP is the RD-DD coefficient, which isolates the response of firms mandated to provide FMLA relative to non-FMLA firms to a change in Chinese import competition. In the first column we focus on all firms. The second column restricts the sample to manufacturing firms only and the third column to single-unit firms only. All our analyses focus on firms employing between 45 and 55 workers. Across all three columns in the top panel, we find a negative and statistically significant RD-DD coefficient. Focusing on the first column, a one percentage point larger increase in Chinese import penetration is associated with a 2.16 percentage points smaller increase (or larger decrease) in the growth rate of the percentage of female employment at FMLA compared to non-FMLA firms.

Similarly, from the bottom panel, we find a negative and statistically significant RD-DD coefficient. A one percentage point larger increase in Chinese import penetration is associated with a 2.83 percentage points smaller increase (or larger decrease) in the growth rate of the percentage of female earnings at FMLA compared to non-FMLA firms. Our findings suggest that an increase in import competition is associated with lower female share of employment and earnings at firms mandated to provide FMLA compared to non-FMLA firms. Together, our results are consistent with a negative effect of FMLA on relative demand for female labor in line with Proposition 1.

5.3 Robustness checks

In this section, we report results from robustness checks of our baseline results. We perform a set of placebo tests. We test whether there exists a differential female labor demand response by FMLA and non-FMLA firms if we were to apply an artificial employment cut-off to determine FMLA status. We examine a cut-off of 40 using a sample of single-unit firms employing between 35 and 45 workers where a firm is defined as FMLA if it employs 40 or more workers and non-FMLA otherwise. We also examine a cut-off of 60 using a sample of single-unit firms employing between 55 and 65 workers where a firm is defined as FMLA if it employs 60 or more workers and non-FMLA otherwise. The idea is that if our identification strategy correctly isolates a labor demand effect of FMLA, we should observe no effects in the two placebo regressions. Indeed, from columns 1 and 2 in Table 4, we find that the coefficient on FMLA x ΔIP is statistically insignificant for both employment and earnings outcomes.

Next, we address potential endogeneity concerns about the measure of Chinese import penetration. We believe that any concerns about Chinese import competition being endogenous to relative outcomes for women is mitigated in our case, given our focus on small firms. Nonetheless, we employ an instrumental variables strategy where we instrument for the change in U.S. import competition from China using Chinese exports to other high-income countries as described in equation (4.2.2). We present results in column 3 of Table 4. [*Note: we will disclose this result in the next draft.*]

6. Mechanisms

Our baseline result, presented in Section 5.2, suggests that FMLA firms display lower demand for female labor compared to non-FMLA firms consistent with Proposition 1 in Section 2.2. Under Proposition 1, we argue that when the leave mandate is binding, firms reduce their demand for female labor relative to male, given that they trade-off the replacement cost of a worker on leave for caregiving to the benefit from retaining the worker's firm-specific human capital if they do not quit. In this section, we examine the implications from Propositions 2 and 3.

First, since high-skilled workers are more likely to accrue firm-specific human capital compared to low-skilled workers, we explore the female labor demand response to FMLA separately for college-educated and less than college-educated workers. The underlying hypothesis is that the replacement cost for low-skilled workers is likely to dominate gains from retaining the worker's firm-specific human capital. We decompose female workers into two skill groups: bachelor's degree or higher (high-skilled) and less than bachelor's degree (low-skilled) and estimate our baseline specification. Results are presented in Table 5. We find that the negative labor demand effect of FMLA is driven by reduced demand for low-skilled female workers.

Next, since a worker's likelihood of fulfilling caregiving responsibilities, particularly child rearing, is higher for women during childbearing years, we would expect the negative labor demand effect to arise primarily within the group of female workers between ages 25 and 45. Women in this age group are in their prime, childbearing years and are also more likely to have young children. We decompose our results by age categories as follows: 24 or less, 25-45, and 46+, where fertility is highest within the 25-45 age group. Results are presented in Table 6. Across columns, the dependent variable is the number (earnings) of female workers in each age group as a share of total workers (earnings) at the firm in the first (second) panel. We find that, as hypothesized, the negative labor demand effect is concentrated among women in the 25-45 age group.

7. Extensions

Beliefs about gender roles can affect women's labor market outcomes (Charles, Guryan, Pan, 2018). In this section, we seek evidence of how the presence of norms that prescribe roles based on gender, such as the belief that women are more likely to take leave to fulfil caregiving responsibilities relative to men, might impact the female labor demand response of FMLA compared to non-FMLA firms. We extend our analysis by asking if responses differ across firms that have women in decision-making roles versus not. If women are in decision-making roles and if women are less disposed to beliefs founded on gender norms, the negative labor demand effect of FMLA may be mitigated for firms with female managers.

We examine this hypothesis by identifying firms that have female managers. A firm is defined as female-managed if at least one of the top three earners at the firm is female. The underlying idea is that workers in decision-making roles are more likely to be highly paid. This definition has been used in prior studies using LEHD to identify firm owners (Howell and Babina, 2018; Azoulay, Jones, Kim, Miranda, 2018; Kerr and Kerr, 2017). The results are presented in Table 7. We find that the negative labor demand effect of FMLA is mitigated for firms with female managers. This result contributes to the literature examining the rationale for and impacts of gender quotas on corporate boards, where one of the arguments is that women in positions of power might facilitate more family-friendly workspaces (Bertrand, Black, Jensen, Lleras-Muney, 2017).

8. Conclusion

This paper causally estimates a female labor demand impact of job-protected family leave. We offer evidence for a negative impact of the Family and Medical Leave Act (FMLA) on female shares of employment and earnings among the universe of U.S. private sector, non-farm employers. We propose a framework where employers trade off the replacement cost of workers taking leave with the benefit of retaining firm-specific human capital should the worker quit in the event that they need to fulfil caregiving responsibilities.

Consistent with this framework, we find that the identified negative labor demand effect of mandated leave under FMLA is concentrated among women in their childbearing years and low-skilled women for whom firm-specific human capital might be less relevant. We find that the negative labor demand effects of FMLA are mitigated for firms with female managers, consistent with the idea that women in decision-making roles might facilitate more family-friendly work environments. Our results highlight the importance of disentangling labor demand separately from labor supply effects of job-protected family leave legislation.

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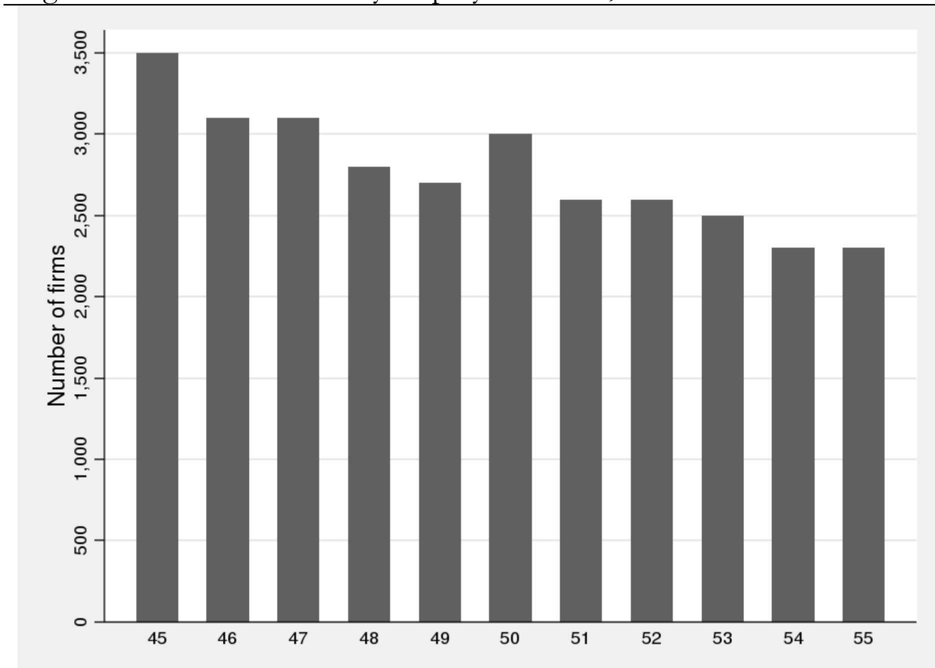
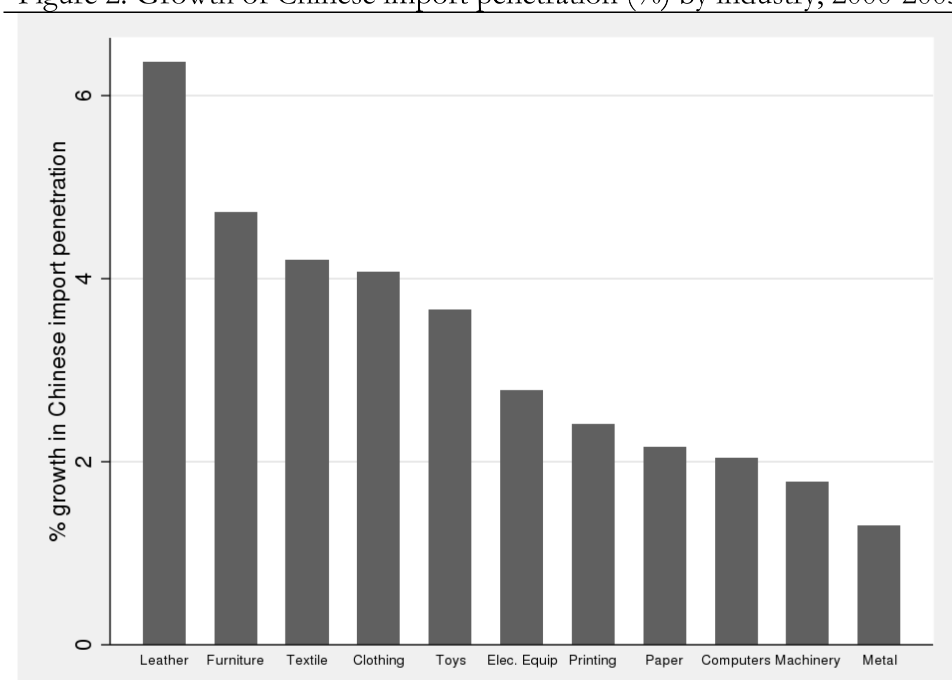
Figures

Figure 1. Number of firms by employment bins, 2000.



Notes: This figure displays the number of firms in the analysis sample by employment bins. Firm counts are rounded for disclosure avoidance.

Figure 2. Growth of Chinese import penetration (%) by industry, 2000-2003.



Notes: This figure displays the top 11 industries with highest growth in Chinese import penetration between 2000 and 2003 at the 3-digit NAICS.

Tables

Table 1. Descriptive statistics, 2000.

	Firms with 45 – 55 Workers		
	FMLA	Non-FMLA	Difference (t-statistic)
Share of female employment	40.70	40.62	0.0738
	(26.25)	(26.10)	(0.25)
Share of female earnings	34.13	34.14	-0.0011
	(25.47)	(25.41)	(0.00)
Female workers	16.45	19.72	-3.277***
	(12.25)	(113.3)	(3.55)
Male workers	24.59	28.73	-4.139***
	(13.90)	(116.2)	(4.37)
Firm age	13.98	14.73	-0.758***
	(8.429)	(8.434)	(7.84)
Firm multi-unit status	0.0135	0.0858	-0.0723***
	(0.115)	(0.280)	(29.45)

Notes: This table displays the means with standard errors in parentheses of outcomes variables at FMLA and non-FMLA firms near the cut-off. The fourth column shows the difference in means between FMLA and non-FMLA firms with t-statistics in parentheses.

Table 2. Specification tests.

Dependent variable	Linear	Quadratic	Cubic	Quartic
Δ Log Share of Female Employment	0	0.54	2.39	6.74
Δ Log Share of Female Earnings	0	-2.25	0.44	5.59

Notes: This table displays the difference in AIC statistic between the linear and all other specifications from estimating equation (3.2.1) where linear is coded as 0.

Table 3. Labor demand impact of FMLA on female workers.

	All	**Manufacturing**	**Single-Unit**
Δ Log Share of Female Employment			
FMLA	0.0201	0.0840*	0.0186
	(0.0199)	(0.0495)	(0.0202)
ΔIP	0.0171***	0.0243***	0.0155***
	(0.0057)	(0.0058)	(0.0057)
FMLA x ΔIP	-0.0216***	-0.0257***	-0.0242***
	(0.0072)	(0.0077)	(0.0070)
Δ Log Share of Female Earnings			
FMLA	0.0179	0.0953*	0.0169
	(0.0212)	(0.0543)	(0.0215)
ΔIP	0.0231**	0.0334***	0.0236**
	(0.0111)	(0.0112)	(0.0112)
FMLA x ΔIP	-0.0283**	-0.0333***	-0.0321***
	(0.0116)	(0.0122)	(0.0114)
Observations	30,500	5,500	29,000

Notes: This table displays results from estimating equation (3.2.1) on the sample of firms that employ 45-55 workers. The first column presents results on the full sample ("All"), and separately for manufacturing only ("Manufacturing") and single-unit only ("Single-Unit") firms. The dependent variable is defined as in (3.2.2). Standard errors are clustered at the four-digit NAICS level. All regressions include firm-level controls for log age and multi-unit status; share of female workers at the four-digit NAICS level, and state fixed effects. Observations are rounded for disclosure avoidance.

Table 4. Labor demand impact of FMLA on female workers, robustness checks.

	Placebo: 40	**Placebo: 60**	**IV**
Δ Log Share of Female Employment			
FMLA	0.0374	-0.0900	
	(0.0791)	(0.0955)	
ΔIP	-0.0075	-0.0158	
	(0.0495)	(0.0398)	
FMLA x ΔIP	0.0058	0.0068	
	(0.0476)	(0.0400)	
Δ Log Share of Female Earnings			
FMLA	0.0501	-0.1356	
	(0.0973)	(0.1274)	
ΔIP	0.0846	0.0120	
	(0.0836)	(0.0734)	
FMLA x ΔIP	-0.0935	-0.0193	
	(0.0846)	(0.0718)	
Observations	47,000	18,500	

Notes: The column titled "Male" displays results from estimating equation (3.2.1) on the sample of firms that employ 45-55 workers. Columns titled "Placebo: 40" ("Placebo: 60") display results from estimating equation (3.2.1) on the sample of firms that employ 35-45 (55-65) workers and uses 40 (60) as the employment cut-off. The column titled "IV" presents results where ΔIP is instrumented using (4.2.2). The dependent variable is defined as in (3.2.2). Standard errors are clustered at the four-digit NAICS level. All regressions include firm-level controls for log age and multi-unit status; share of female workers at the four-digit NAICS level, and state fixed effects. Observations are rounded for disclosure avoidance.

Table 5. Labor demand impact of FMLA on female workers, by skill.

	High-skilled	**Low-skilled**
Δ Log Share of Female Employment		
FMLA	0.0210	0.0183
	(0.0356)	(0.0200)
ΔIP	-0.0009	0.0152
	(0.0204)	(0.0109)
FMLA x ΔIP	-0.0023	-0.0254**
	(0.0213)	(0.0126)
Δ Log Share of Female Earnings		
FMLA	0.0397	0.0095
	(0.0303)	(0.0205)
ΔIP	0.0063	0.0171*
	(0.0257)	(0.0096)
FMLA x ΔIP	-0.0108	-0.0271**
	(0.0282)	(0.0117)
Observations	30,500	

Notes: This table displays results from estimating equation (3.2.1) on the sample of firms that employ 45-55 workers. The dependent variable is defined as in (3.2.3). "High-skilled" is defined as workers with a bachelor's degree or higher and "Low-skilled" otherwise. Standard errors are clustered at the four-digit NAICS level. All regressions include firm-level controls for log age and multi-unit status; share of female workers at the four-digit NAICS level, and state fixed effects. Observations are rounded for disclosure avoidance.

Table 6. Labor demand impact of FMLA on female workers, by age.

	24 or less	25-45	46+
Δ Log Share of Female Employment			
FMLA	-0.0089	0.0205	0.0317
	(0.0403)	(0.0297)	(0.0303)
ΔIP	-0.0110	0.0339***	0.0054
	(0.0316)	(0.0107)	(0.0118)
FMLA x ΔIP	-0.0252	-0.0416***	-0.0068
	(0.0312)	(0.0095)	(0.0139)
Δ Log Share of Female Earnings			
FMLA	0.0006	0.0193	0.0162
	(0.0319)	(0.0308)	(0.0284)
ΔIP	-0.0070	0.0216	0.0207
	(0.0215)	(0.0133)	(0.0179)
FMLA x ΔIP	-0.0206	-0.0314**	-0.0220
	(0.0216)	(0.0131)	(0.0193)
Observations		30,500	

Notes: This table displays results from estimating equation (3.2.1) on the sample of firms that employ 45-55 workers. The dependent variable is defined as in (3.2.3). Standard errors are clustered at the four-digit NAICS level. All regressions include firm-level controls for log age and multi-unit status; share of female workers at the four-digit NAICS level, and state fixed effects. Observations are rounded for disclosure avoidance.

Table 7. Labor demand impact of FMLA on female workers, role of female managers.

	Top 3 Earners Female	**Top 3 Earners Not Female**
Δ Log Share of Female Employment		
FMLA	-0.0083	0.0473
	(0.0215)	(0.0288)
ΔIP	0.0141	0.0266**
	(0.0090)	(0.0115)
FMLA x ΔIP	-0.0148	-0.0362***
	(0.0098)	(0.0118)
Δ Log Share of Female Earnings		
FMLA	-0.0071	0.0411
	(0.0238)	(0.0304)
ΔIP	0.0281*	0.0281**
	(0.0144)	(0.0116)
FMLA x ΔIP	-0.0268*	-0.0408***
	(0.0151)	(0.0138)
Observations	13,500	16,500

Notes: This table displays results from estimating equation (3.2.1) on the sample of firms that employ 45-55 workers separately for firms where "Top 3 Earners Female" and "Top 3 Earners Not Female", respectively. The dependent variable is defined as in (3.2.2). Standard errors are clustered at the four-digit NAICS level. All regressions include firm-level controls for log age and multi-unit status; share of female workers at the four-digit NAICS level, and state fixed effects. Observations are rounded for disclosure avoidance.

Table A1. Level effects of FMLA and trade shock.

	Female	Male
Δ Log Employment		
FMLA	0.0161	0.0204
	(0.0212)	(0.0229)
ΔIP	-0.0144***	-0.0117***
	(0.0025)	(0.0024)
Δ Log Earnings		
FMLA	0.0847	0.0911*
	(0.0686)	(0.0526)
ΔIP	-0.0161**	-0.0225**
	(0.0074)	(0.0096)
Observations	30,500	

Notes: This table displays results from regressing the level change in employment and earnings by gender on an indicator of firm's FMLA status and the change in Chinese import penetration using the sample of firms that employ 45-55 workers separately for firms. The dependent variable is defined as $\Delta ln\left[y_g\right]_i$ where Δ denotes change between 2000 and 2003 and g is either female or male; y denotes total employment (earnings) of workers employed at firm i. Standard errors are clustered at the four-digit NAICS level. All regressions include firm-level controls for log age and multi-unit status; share of female workers at the four-digit NAICS level, and state fixed effects. Observations are rounded for disclosure avoidance.