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# GLOBALIZATION, GENDER, AND THE FAMILY

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## ABSTRACT

This paper shows that in the presence of labor market shocks, child-bearing and child-rearing have far-reaching implications for gender inequality, household specialization and family structure. Using population register data on all births, marriages, and divorces together with employer-employee linked data for Denmark, we show that reduced labor market opportunities due to Chinese import competition lead to a move towards family, with higher rates of fertility, parental leave, and marriage, as well as lower rates of divorce. This move is driven by women, not men. We document substantial long-run earnings losses concentrated on women, and gender inequality increases. The gender-specific effects are due to a woman's ability to give birth during a fixed period of life–her biological clock. Women have a higher reservation value for staying in the labor market when young, and a negative trade shock induces women to substitute more to family activities than men. High-earning women in their late 30s contribute strongly to the gender difference in fertility because switching to new comparable employment would require high initial commitment which is incompatible with having a newborn in the short time remaining on the biological clock. There is no gender difference (1) for workers past their fertile age, (2) in the size of the negative labor shock, and (3) due to occupational composition since we exploit within-worker variation. Despite lower labor earnings, positive family responses in Denmark are also sustained by insurance payments and government transfers so that workers can afford the shift to family.

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# 1 Introduction

The factors driving differences in the labor market behavior of men and women are not fully understood. This paper makes progress by examining the balance between labor market and family activities in response to a trade shock. Worker skill and the conditions of the local labor market are important determinants of how workers adjust to negative labor demand shocks.[1] We focus on gender differences in the market versus family choice in determining adjustment costs.[2] Studying workers exposed to rising import competition from China in the 2000s, we show that as the trade shock reduces employment opportunities, gender and age play central roles in determining how workers adjust to the shock and how they re-balance market and family activities.

Employing population register and labor market information on workers matched to their firms, we provide a longitudinal picture of individual-level family *and* labor market adjustment to rising import competition in Denmark from 1999-2009. We document a strong gender difference in long-run labor market outcomes and a shift towards family due to reduced labor market opportunities. Workers exposed to increased import competition are disproportionately more likely to have newborn children and to take parental leave; they are also more likely to form new marital unions, as well as to avoid breaking up existing ones. Women, not men, drive this shift caused by trade exposure.

We study two 1999 cohorts of workers, Denmark's entire private-sector labor force and a subset, the textile and clothing workers. The latter are specifically affected by the removal of Multi-fibre Arrangement (MFA) quotas on Chinese exports following the country's entry into the World Trade Organization (2002). This trade liberalization leads to an increase in fertility and parental leave among single women by about one quarter, and their marriage probability increases strongly, while exposed married women respond by reducing divorce rates. These family responses go hand in hand with long-run labor earnings losses for women, almost 120 percent of one year's salary, while men do not significantly lose earnings over an eight-year period (2002 - 2009). The findings from this trade policy change are broadly confirmed for Denmark's labor force as a whole using an instrumental-variables approach exploiting industry variation in the change in import penetration. Gender differentials along the labor market-family margin are a source of inequality on which generally little is known.

[1] Jacobson, Lalonde, and Sullivan (1993), Artuç, Chaudhuri, and McLaren (2010), Autor, Dorn, Hanson, and Song (2014), Dix-Carneiro (2014), Utar (2018), and Traiberman (2019).

[2] Synonymous to family in our paper is the term household. An early contribution is Becker (1973)

Investigating this gender difference with worker, firm, and partner information, the primary reason why women disproportionately shift towards family is not that their pre-shock employment is concentrated in highly exposed firms or in more vulnerable occupations compared to men. The respective earnings losses at the initial firm, and the likelihood of displacement from it, are similar for men and women. Rather, men and women follow different paths of adjustment to the shock, with women moving consistently towards family following displacement.

Our explanation for this gender difference is the biological clock of a woman. Because women are unlikely to conceive beyond their early forties, fertile-age women have a higher reservation value to stay in the labor market than men. Consequently, a given negative labor demand shock due to trade exposure will raise a woman's incentive of moving towards family by more than it does for a man. Furthermore, because having and caring for a young child is female-intensive, making investments to succeed in a new career before the biological clock runs out may be impossible, and lead women to focus on family. Evidence for this comes from the finding that it is mostly women in their fertile age who account for the gender difference, especially those women who are closer to the end of their fertility period. In contrast, the adjustment of women past their fertile age is similar to that of men. Also, it is those women who would have to make the highest investments into new careers that contribute most to the higher female birth rates.

Much progress has been made in understanding adjustment costs to workers' re-establishing promising career paths after a trade shock, with a growing literature studying heterogeneity in adjustment experiences. On the one hand, young workers tend to perform better than older workers (Artuç, Chaudhuri, and McLaren 2010, Dix-Carneiro 2014), in part because they lose comparatively little industry- or occupation-specific human capital (Utar 2018). On the other, young workers may face higher adjustment costs because of weak attachment to the labor market (Autor, Dorn, Hanson, and Song 2014). By showing that women and men subject to the same-sized shock respond differently in their labor and non-labor market choices, our analysis shifts the focus to gender differences, with child bearing and rearing playing a key role. In a non-trade context, Becker's (1960, 1965) model of fertility lays out two effects in response to lower market earnings. It could decrease fertility through an income effect, or increase fertility through a substitution effect because the opportunity costs of time have fallen. Furthermore, recent research has highlighted dynamic effects that provide incentives for women to postpone fertility, such as the adverse impact of early-stage career interruptions, and that a young child negatively affects the probability of finding a new job (Del Bono, Weber, and Winter-Ebmer 2012, 2015, Huttunen and Kellokumpu 2016). We study how these static and dynamic effects of lower labor market opportunities interact with a woman's

biological clock. Women in manufacturing occupations adversely affected by import competition, such as machine operators, need to learn new skills, while women in high-level professional occupations need to show special commitment early in their new jobs. These requirements are less compatible with having a newborn the closer a woman is to the end of her fertile period. This explains why the trade-induced fertility impact increases with age before women are in their 40s, and is stronger for women with occupations that require high levels of commitment and investment to adjust to the shock.

The unprecedented growth of imports from China and rising competition has generated a large literature that documents significant decline in manufacturing employment in many advanced countries.[3] Among the smaller set of analyses of non-labor outcomes, Autor, Dorn, and Hanson (2019) find that import competition lowers marriage rates, while we find that import competition increases marriage rates. Our analysis highlights how the empirical approach and the structure of welfare policies can lead to different results in response to the China trade shock. First, we focus on gender differences in worker adjustment, showing that the market-family margin is a major factor leading to gender inequality. This is difficult to study without individual-level measures of exposure and outcomes, as in Autor, Dorn Hanson (2019), and their region-level cross-sectional analysis is better suited to measuring aggregate effects including spillovers. Furthermore, the institutional setting matters. Our results suggest that effective insurance and transfer payments can dampen the negative income effect of trade exposure and shape the response of an economy.[4]

We also contribute to work on the reasons behind gender differences in various settings (Bertrand 2010, Blau and Kahn 2017). We document that an important source of gender wage gaps, namely 'child penalties' for women, are due to differential gender responses because of the woman's biological clock (see also Goldin and Katz 2002, Kleven, Landais, and Sogaard 2019).[5] By placing child-bearing and child-rearing at the center of the argument, our biological clock explanation relates to the temporal flexibility argument (Goldin 2014, Goldin and Katz 2016). One difference is that to the extent that the gender gap is driven in part by biology, greater family friendliness and social policies cannot fully eliminate the gender gap. Furthermore, temporal flexibility is typ-

[3] See Autor, Dorn, and Hanson (2013), Bloom, Draca, and van Reenen (2016), Ebenstein, Harrison, McMillan, and Phillips (2014), Keller and Utar (2016), Pierce and Schott (2016), Utar and Torres-Ruiz (2013), and Utar (2014). None of these studies focus on gender. See also Che, Lu, Pierce, Schott and Tao (2016) and Autor, Dorn, Hanson, and Majlesi 2017) on political outcomes.

[4] Consistent with that, cross-country analysis indicates that generous benefits tied to previous employment in Scandinavian countries are associated with relatively high female fertility (Adsera 2004).

[5] We also show that globalization, here in form of rising import competition, need not reduce gender inequality. See Pieters (2015) for an overview of the relationship between trade liberalization and gender inequality, as well as Boler, Javorcik, and Ullveit-Moe (2018) and Hakobyan and McLaren (2018).

ically studied in the context of long-term career planning of highly paid workers (e.g., Bertrand, Goldin, and Katz 2010), whereas we study the career choices of manufacturing workers subject to a sudden, plausibly exogenous shock with data that allows us to rule out potentially important composition effects. Fertility responses have been studied in the context of job displacement following recessions or plant closures (e.g., Del Bono, Weber, and Winter-Ebmer 2012, Huttunen and Kellokumpu 2016, Huttunen, Møen, and Salvanes, 2018). Our analysis suggests that compared to our approach, the focus on job displacement and unemployment leaves little room for estimating positive family responses.

The remainder of the paper is as follows. The next section reviews the recent evolution of imports in Denmark and discusses identification of the impact of rising import competition. We also introduce the most important recent developments regarding family formation and fertility in Denmark. Section 3 lays out the econometric framework for the quasi-experiment. Section 4 shows that the labor shock due to increased import competition has caused a significant gender gap in labor market outcomes with far higher earnings losses for women than for men. Next, we show that it has also caused increased childbirths and parental leave, increased marriage probabilities, and reduced divorce rates. Further, we document the key gender difference by demonstrating that all family impacts are largely due to women and establish that reduced market work is the flip side of increased family activities. Section 5 shows that our findings on textile workers generalize to Denmark's entire private-sector labor force. Section 6 discusses the main mechanism for our finding, a woman's biological clock, and explains why we see a more positive family response for women than others before us. Section 7 contains a concluding discussion. The Appendix presents supplemental results and analysis.

# 2 Import Shocks and Integrated Data on Individual-Level Market and Family Outcomes

The following provides background on recent trends in import competition and family structure in Denmark and how they are captured in our data.

## 2.1 Rising Import Competition for Denmark's Workers

Since the late 1990s, Denmark, like many other advanced countries, has experienced increased import competition from lower-wage countries. To estimate gender differences in trade adjustment, we first employ a quasi-experiment that uses a concrete policy change, the lifting of quotas on China's textile exports that were part of the Multi-Fibre Agreement (MFA) due to China's entry into the WTO (December 2001). Due to its compelling nature, this policy change has been extensively employed in the literature to address various questions (Bloom, Draca, and van Reenen 2016, Khandelwal, Wei, and Schott 2013, and Harrigan and Barrows 2009, Utar 2014). We also generalize the analysis to the entire private-sector labor force by employing an instrumental variables approach that exploits changes in import penetration due to a sudden expansion of production capabilities in China around the time of its WTO accession. The following summarizes our two approaches, with details given in the Appendix (sections A and C).

### 2.1.1 Textile and Clothing Liberalization

The MFA was established in 1974 as the cornerstone of a system of quantitative trade restrictions on developing countries' textile and clothing exports with the intention to protect this relatively labor-intensive sector in advanced countries. With the conclusion of multilateral trade negotiations in the year 1994, it was agreed to bring trade in textiles in line with the rules of other world trade at the time, and thus import quotas were to be removed.[6] Specifically, it was agreed that MFA quotas were to be abolished in four phases: 1995, 1998, 2002, and 2005.

An advantage of this policy change is that neither Denmark nor China was directly involved in negotiating the creation or removal of the textile quotas (as well as which goods would be covered in which of the four phases). This is because negotiations took place at the level of the EU, where Denmark's influence as a relatively small country is limited. Also, China did not influence the process because it was not a member of the WTO in 1995. Similarly, because it was not a member of the WTO, China did not benefit from the first two trade liberalization phases of 1995 and 1998. At the same time, China stood out in comparison to other countries subject to the MFA quotas in terms of the number of binding quotas. While there was considerable uncertainty about the if, when, and how of China's entry into the WTO, China did become a member of the WTO in

[6] These 1986-1994 multilateral trade negotiations took place under the auspices of the General Agreement on Tariffs and Trade (GATT) and are commonly referred to as the Uruguay Round. The Round led to the creation of the WTO as the successor of the GATT agreements.

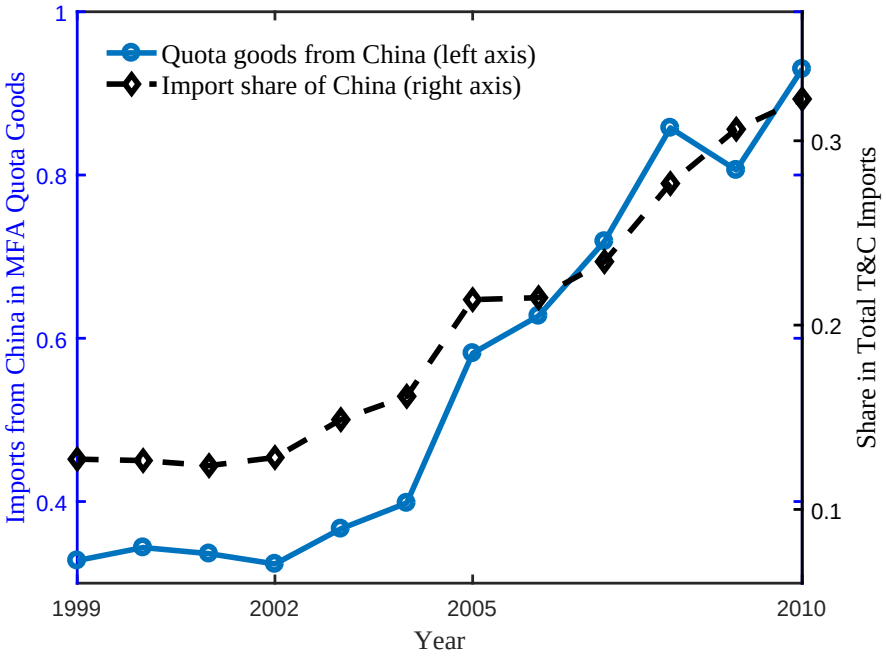


Figure 1: **Evolution of Chinese Imports in Response to Quota Removal**

**Notes**: Solid line shows imports from China of MFA quota goods in Denmark, measured in the total value added in the textile and clothing industry. Dashed line shows China's share in all imports of textiles and clothing (T&C) goods in Denmark.

December 2001. Only once that had happened, China benefited from the first three liberalization phases (1995, 1998, and 2002), and subsequently, from the fourth liberalization phase of 2005.[7] The lifting of quotas led to a surge of Chinese textile imports starting in the year 2002, as Figure 1 shows. Between 2002 and 2010, imports of quota goods from China relative to the 1999 domestic value added of the Danish textile and clothing industries (about 1.3 billion Euros) tripled. These trade liberalizations generate a plausibly exogenous increase in import competition.

This trade liberalization episode generates a worker-level measure of exposure to import competition. A worker is defined to be exposed if, as of 1999, her or his firm has domestic production in any 8-digit Common Nomenclature (CN) goods for which China sees her quota lifted due to her 2002 WTO entry. Workers employed in 1999 in textile and clothing manufacturing firms that

[7]The EU had some influence on which goods to liberalize first, and it chose mostly those with non-binding quotas vis-à-vis other WTO countries in the first two phases. Additionally, as Phase I and II removals did not cover China which had the highest number of binding quotas, the first two removals did not trigger more competition in the industry (Utar 2014). This is consistent with our identification strategy that focuses on the WTO accession of China because the large majority of Phase III quotas (2002) were binding, and the firms that manufacture Phase III quota products are largely the same as those that produce Phase IV quota products. See Appendix A for additional information on the quota liberalizations including a temporary reversal by the EU in 2005 and the relationship between the first three and the fourth phase of liberalization.

produce none of the affected 8-digit CN products constitute our control group of workers. We also define an alternative exposure variable using the firms' revenue share in quota products in 1999, finding similar results (see Appendix B).

It is useful to compare our approach with the literature on the effects of job displacement, which contrasts workers who lose their employment (due to a recession or mass layoffs, for example) with workers who do not. In the job displacement literature, displaced workers are compared with employed workers. In our approach, treatment is defined based on an ex-ante criterion, workers' exposure to competition based on the firm's product portfolio several years before the quota lifting. The hazard of being displaced is relatively high for exposed workers. Still, we do not have the one-to-one mapping that is common in job displacement papers (our control group also includes displaced workers). We will examine the role of this for our empirical results in Section 6.

### 2.1.2 Import Competition in the Economy-wide Labor Force

In our second approach, exploiting the rise of imports from China around the time of China's WTO accession, we study the impact of changes in import penetration from China across industries in the entire private sector economy. Figure X1 shows the change in Chinese imports between 1999 and 2009 over absorption at the six-digit (NACE) industry-level, against the industry's employment share in 1999 among manufacturing industries.

The change in import penetration varies widely, ranging from 0.3 in some to 0 in other industries. Even within two-digit industries there is substantial variation in import penetration (six-digit industries belonging to the same two-digit industry have the same marker). To eliminate the influence of broader technology shocks, we will rely below on variation within two-digit industries. Furthermore, possible correlation between domestic industry demand shocks and changes in import penetration will be addressed with an instrumental variables approach based on geography-based trade costs, the strength of the pre-existing distribution channels, and imports of other advanced countries as instruments. Additional information on this approach is given in section 5 and in section C of the Appendix.

## 2.2 Workers and their Firms

We employ the Integrated Database for Labor Market Research (IDA database) of Statistics Denmark, which contains administrative records on virtually all individuals and firms in Denmark.[8] To construct our sample, we apply some sample restrictions. First, to ensure that workers would typically not retire during the sample period, we focus on workers who were between 18 and 56 years old in 1999. Second, we concentrate on employees who worked as a full-time employee in 1999 so that both female and male workers have comparably strong labor market attachment.[9] This yields a 1999 cohort of about 1.6 million workers, essentially Denmark's entire private-sector labor force. Of these workers, about 10,000 were employed in the textile and clothing industry that experienced the quota liberalization. We will refer to this for short as the quasi-experiment sample.

We follow the 1999 cohort for a decade as they change jobs, switch firms, industries, or occupations, and move out of the labor force to take parental leave, give birth, or become unemployed. In this way the analysis yields a worker-level picture of labor market and family adjustments throughout the entire economy. Table 1 shows a number of key sample characteristics for the quasi-experiment.

The employer-employee link allows us to control for a number of firm-level variables that may be important for the workers' labor market and family choices. They include firm size (measured by employment) and firm quality (proxied by the average firm wage). Specific industry, firm, and job characteristics are important for assessing the importance of selection at different margins for our results. We match the firm IDs with the domestic production database (VARES) that reports firms' sales of each CN-10 digit product that they produce domestically to identify quota producing firms before the removals. Furthermore, for those workers who are not single, our analysis accounts for partner (married spouse as well as cohabitant) characteristics, including earnings, income, and the degree to which the partner is exposed to rising import competition.

According to Table 1, roughly half of the workers were employed at firms that manufactured products subject to quota removals for China (exposed in Table 1). The average age of both treated and untreated workers is the same at 39.2, which is important because adjustment costs vary between older and younger workers not least because older workers typically have a harder time to learn the skills needed in new jobs than younger workers. We also see that both sets of workers have

[8]The database provides information on all persons of age 15 to 70 residing in Denmark with a social security number, on all establishments with at least one employee in the last week of November of each year, and on all jobs that are active in that same week. See Bunzel (2008) for a description of this data.

[9]Nevertheless, results that include part-time workers are similar (available upon request).

between 14 and 15 years of labor market experience. Average annual earnings are quite similar in 1999 for exposed and control workers. Also notice that around 60 percent of exposed workers are married, compared to about 58 percent for the control group.[10] The average number of children as of 1999 in our sample is 1.46, with no significant difference between the sets of exposed and control workers. About one in 20 workers has a newborn child or takes parental leave in 1999, which is similar for exposed and control workers. Table 1 also shows summary statistics for education and occupation.[11]

Table 1: **Worker Characteristics by Exposure: Quota Removal Experiment**

| | **Exposed** N = 4,743 **Average** | **Control** N = 5,255 **Average** | **Diff.** | **t-stat** |
|---|---|---|---|---|
| Age | 39.206 | 39.228 | -0.022 | -0.111 |
| Labor Market Experience | 14.912 | 14.491 | 0.421 | 3.694 |
| Log Annual Earnings | 12.165 | 12.154 | 0.011 | 0.843 |
| Married | 0.604 | 0.576 | 0.028 | 2.802 |
| No. of Children | 1.448 | 1.480 | -0.032 | -1.387 |
| Birth Event | 0.040 | 0.045 | -0.004 | -1.099 |
| Parental Leave Take | 0.053 | 0.050 | 0.003 | 0.687 |
| College Educated | 0.130 | 0.107 | 0.023 | 3.580 |
| Vocational Educated | 0.361 | 0.360 | 0.001 | 0.127 |
| Machine Operator | 0.353 | 0.359 | -0.007 | -0.685 |
| Manager | 0.059 | 0.052 | 0.008 | 1.680 |

**Notes**: Shown are averages of the 1999 characteristics of textile workers by exposure to rising import competition due to the quota removals for China. Exposed workers are those whose firm manufactured in Denmark in 1999 a product protected by a quota that would be removed with China's entry into the WTO; correspondingly, Not exposed workers are those who did not do so. Labor market experience measured in years. Married, Birth Event, Parental Leave Take, College, Vocational, Machine Operator, and Manager are indicator variables. Log earnings is measured in 2000 Danish Kroner; the mean is about 40,000 current US Dollar.

Given our interest in gender, Table X1 shows sample information separately for men and women, as well as by family status (married versus unmarried). There are some differences, for example

[10] The share of single workers is about 28 percent for both sets of workers.

[11] We employ the Danish version of the International Standard Classification of Occupation (D-ISCO) at the four-digit level; it has about 400 different job types. See https://www.dst.dk/en/Statistik/dokumentation/nomenklaturer/.

women overall account for a larger share of textile workers than men (about 5,600 versus 4,400 workers, respectively). Textile and clothing production tends to be female intensive; however, women perform a wide range of occupations in the textile industry. Figure A3 in the Appendix shows the distribution of female workers across occupations by exposure to import competition. The majority of women, treated or control, work in production as machine operators, similar to men, as this is a typical manufacturing industry. We also observe the position of a worker in the job hierarchy of the firm, which will be employed in section 6 below.

Table X1 shows that there is essentially no age difference between exposed and control workers, whether among married women, unmarried women or married men. We also see that married workers are generally older than unmarried workers. The average difference in age between married and unmarried is about seven years for women and nine years for men. In line with this age difference, hourly wages are relatively high for married workers. However, note that the gender wage difference is larger for married than for unmarried workers. This could be an indication that family activities require more time away from the labor market for women in comparison to men. Hourly wages are quite similar across different demographic samples by exposure to import competition, indicating treated and control workers have very similar qualifications even within demographic groups. Overall, gender and other differences in our sample by treatment status are limited (Table X1).[12]

We turn to summarizing our economy-wide sample now. This cohort of 1999 workers consists of N = 1,651,774 individuals, of which are about 45 percent women. In 1999 these workers are 38 years old on average and 26 percent of them have college education, in contrast to textile workers, for whom these figures are 39 years and 12 percent respectively. Furthermore, workers in the economy-wide sample overall have similar rates of fertility as textile workers (between four and five percent).

The fraction of workers with college education is higher among women than men (thirty two versus twenty five percent respectively), while men are relatively more frequently vocationally trained.[13] The sample also confirms some well-known occupational differences, for example men are more likely to be managers than women (five compared to two percent, respectively), while women are

[12] Analogous to Table 1 for the year 1999, three years before the onset of rising import competition, we show descriptive statistics for the year 1995. The results for up to seven years before the shock are similar in that they show no clear evidence for pre-trends (see Appendix, Table A2). But we also conduct an extensive set of pre-trends analysis, see Section 3 and Appendix Section A.5.

[13] Vocational education is an important institution in Denmark, combining on the job training at firms with formal education at schools.

more likely to be office workers than men (twenty-two percent versus six percent, respectively). See Table X2 for additional information on our private sector sample.

## 2.3 Family Trends in Denmark

We begin this section by providing an overview of marriage trends in Denmark. The age at first marriage has increased for both men and women in Denmark since the 1960s, as it did in many other countries. Education goals and increased life expectancy are contributing factors. An important aspect of family life by now is co-habitation (living together without being married), which for many couples is the stage of life before marrying. In 2003, 22 percent of all couples in Denmark were cohabitating. In our sample, the figure is the same.

Divorce rates for Danes have fallen from the mid-1980s to the mid-2000s.[14] Marriage and divorce information for all Danish residents comes from Denmark's Central Population Register. We match this data to worker data with a unique person identifier.

One goal of household formation is to raise children. While the total fertility rate in Denmark since the year 1990 has been broadly stable, there have been fluctuations, such as during the period 2002 to 2008, when Denmark's total fertility rate increased by almost 10%.[15] Childbirth information is from Statistics Denmark's Fertility Database, which provides parental information with personal IDs on every child born in Denmark. Our results below indicate that import competition contributed to the decline in divorce at the same time when it increased fertility.

Another indicator of family activities is parental leave, where parents reduce time spent in the labor market to care for children. This can be newborn or older children. Although there have been some fluctuations in parental leave provision over time, by international standards parental childcare leave is generous in Denmark.[16] From the year 2002 on, there is a maximum of 112 weeks of job-protected parental leave per child. Of this, the mother can take up to 64 weeks–18 weeks of maternity leave plus 46 weeks of parental leave– while the father can take a maximum of

[14] In 1986, the chance that a marriage would last for five years was about 86%, rising to above 89% by 1998 and above 91% by the year 2007.

[15] The total fertility rate is defined as the number of children that would be born alive per 1,000 women during the reproductive period of their lives (ages 15 through 49), if all 1,000 women lived to be 50 years old, and if at each age they experienced the given year's age-specific fertility rate. The rate for Denmark is estimated at around 1,871 in 2010, compared to 1,925 for the United States. Source: Human Fertility Database, Max Planck Institute for Demographic Research (Germany) and Vienna Institute of Demography (Austria). Available at www.humanfertility.org.

[16] Specifically, during the 1990s there was a step-by-step decrease of parental leave support, which was reversed in the early 2000s.

48 weeks, composed of 2 weeks of paternity leave and 46 weeks of parental leave (OECD Family Database).[17] Our information on childcare leave spells comes from the Social Benefits Database of Statistics Denmark.[18]

We now turn to discussing a number of descriptive findings about women's birth rates both during and before the sample period, see Figure 2. The figure shows annual birth rates for two subsamples of women, those who are unmarried as of 1999 versus those women who are married in 1999.[19] For the two sets of unmarried female workers on the left of Figure 2, we see that while the birth rates of exposed and control workers are generally similar during the 1990s, by the year 2002 exposed women have typical birth rates of six and a half percent, in contrast to women in the control group who have birth rates about one percentage point lower. Furthermore, this difference is by and large present for the entire treatment period of 2002 to 2009. This is consistent with fertile age women responding with family activities to a negative labor shock, as summarized by our biological clock argument. In contrast, there is no comparable difference for the two sets of older women who where married as of the year 1999. Note that by plotting data on birth rates, Figure 2 does not show the causal effect of rising import competition.[20] As we will see below, the effect of rising import competition on fertility is larger than the difference shown in Figure 2.

# 3 Estimation Approach

This section describes the quasi-experiment and explains how we address challenges to identification. Our difference-in-differences framework exploits the drastic change in import competition as China entered the WTO (first full year: 2002) together with the longitudinal structure of the data that allows us to employ worker fixed effects. We aggregate the annual data into pre- and post-shock periods to address the concerns noted in Bertrand, Duflo, and Mullainathan (2004). The impact of import competition on outcome $X_{is}$ of worker $i$ in period $s$ is estimated as follows:

$$X_{is} = \alpha_0 + \alpha_1 Post_s \, x \, Exposure_{i,99} + \alpha_2 Post_s + \delta_i + \varphi_{is}, \tag{1}$$

[17] See https://www.oecd.org/els/family/.

[18] Source: Social benefits statistics database is a part of the income statistics registers (SHSS - *Sammenhængende socialstatistik* and OF *Offentligt forsørgede*).

[19] In 1990, these married women are typically 33 years old, and have correspondingly higher birth rates than in 1999, when they are typically 42 years old.

[20] See Figure A2 in the Appendix for a difference-in-differences analysis based on pre-sample data.

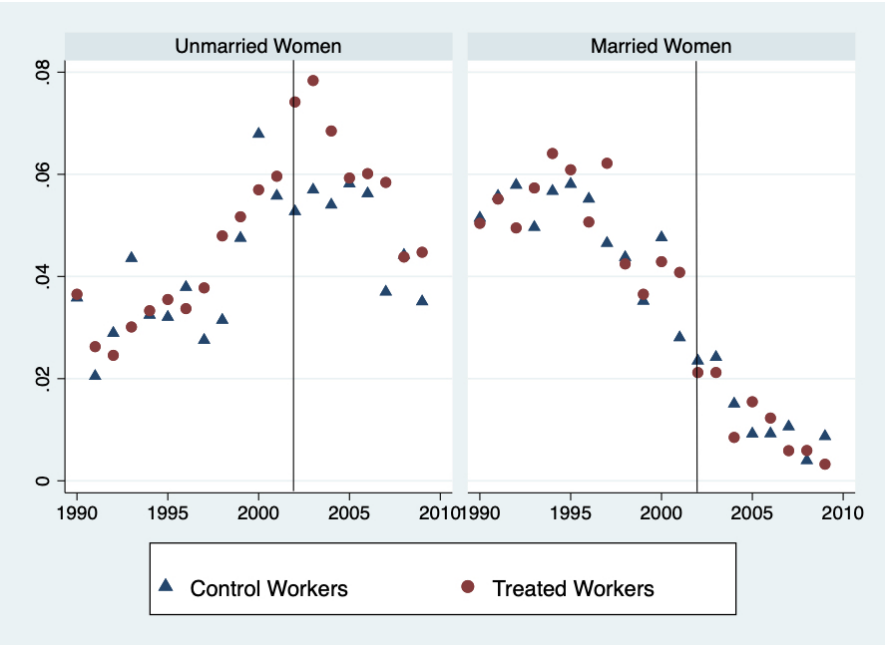


Figure 2: **Birth Rates by Exposure and Marriage Status**

**Notes**: Figure shows fraction of women having a newborn child from 1990 to 2009, by exposure to rising import competition and depending on whether they were married in the year 1999 or not.

where $Exposure_{i,99}$ is an indicator for exposure to rising import competition that takes one if the worker was employed in 1999 in a firm domestically producing a quota-protected good from China, $Post_s$ is an indicator variable for the post-liberalization period (years 2002-2009) that captures the influence of aggregate trends affecting all workers, and $\delta_i$ is a fixed effect for each worker $i$. The subscript $s$ identifies the pre- and post-liberalization periods (years 1999-2001 and 2002-2009, respectively).[21] The error term $\varphi_{is}$ is assumed to be mean zero, and we allow for correlation within groups of workers employed by the same firm by clustering standard errors by 1999 firm.[22] For ease of exposition, we denote the difference-in-differences term $Post_s \, x \, Exposure_{i,99}$ by $ImpComp_{is}$, mnemonic for import competition.

Given our focus on gender differences in labor and family responses, we form a triple difference-in-differences estimation equation to distinguish the possible differential effect of import competition on either gender. The estimation equation is then:

[21] Since changes in family status and the number of children are relatively rare, discrete events, in addition to OLS regressions that allow us to control for worker fixed effects, we have also performed probit analyses (see Appendix section D).

[22] To examine the evolution of the effect over time, we also vary the endpoint of the analysis yearly from 2002 until 2009 and estimate equation (1) with different endpoints.

$$X_{is} = \alpha_0 + \alpha_1 ImpComp_{is} + \alpha_2 ImpComp_{is} x Female_i + \alpha_3 Post_s + \alpha_4 Post_s x Female_i + \delta_i + \nu_{is}, \quad (2)$$

where $Female_i$ is equal to one if worker $i$ is a woman. In this specification, $\alpha_2$ measures the differential effect of rising import competition on women.

**Identification** The inclusion of worker fixed effects implies that the coefficient $\alpha_1$ is estimated from within-worker variation over time. This has the advantage that the influence of any observed or unobserved worker characteristic as of year 1999 that may be correlated with workers' future exposure to competition, such as occupation, education, unobserved abilities, or differential technological trends across firms, is eliminated.

The coefficient $\alpha_1$ in equation (1) is the well-known linear difference-in-differences estimator, which gives the treatment effect under the standard identification assumption that in the absence of treatment the workers would have followed parallel trends. This assumption would not hold, for example, if removal of quotas for other developing countries in 1995 and 1998 (MFA quota removal Phase I and II, respectively) had led to increased competition and differential trends between exposed and control workers. Furthermore, the second half of the 1990s is also a period where several countries were added to the European Union and hence there was increased trade integration with Eastern European countries.

To address this we conduct a falsification exercise for the period 1990-1999, during which rising import competition due to the removal of import quotas on China associated with China's entry into the WTO was absent (placebo test). Collecting data for our workers back to the year 1990, we run specifications analogous to equation (2) without changing the definition of treatment (a worker's firm produces a MFA quota product as of 1999). In this placebo analysis for the period 1990-1999, the years 1990-94 are assumed to be the pre- and the years 1995-99 is assumed to be the post-shock period. The results, presented in Tables A3 and A4 of the Appendix, show no differential trends in earnings, income, hours worked and unemployment as well as in marriage, divorce and birth outcomes between treated and other textile workers. For example, the point estimates for marriage are positive but close to zero, and there is no significant difference between the coefficients for women and men, respectively. Employing annual data instead of aggregating the data into pre- and post-1995 periods yield similar estimates (See Table A5 and Figure A1 in

the Appendix). All placebo results are shown in the Appendix, section A.5.[23] [24]

We conclude from this placebo analysis that there is no evidence that the MFA removal phases I and II, the enlargement of the European Union with the Eastern European Countries, or any other factor generated differential pre-trends that would preclude estimating consistent effects for the period 1999-2009.

What about other threats to identification? Broad differences between men and women, for example the lower propensity of women to work as a manager, are captured by worker fixed effects. We also control for aggregate time trends that are specific to female workers (the term Post x Female in equation (2)).[25] Finally, the trade shock might be correlated with other shocks during this time period, in particular other aspects of globalization, such as technical change. This is important especially for studies exploiting broad cross-industry variation in imports because technology developments tend to vary across industries. It is a less important concern in the quota removal context, first because all workers are employed in 1999 in the same industry and thus are subject to the similar technology shocks and secular trends. Second, our treatment definition is based on the lifting of quotas at a fine (8-digit) product level, and the possibility that other shocks could mimic this variation is small.[26] At any rate, to address this concern we have re-estimated equation (2) with differential time trends across workers' occupations as of 1999 (occupation x time). Since the technology impact increases with routine-task intensity (RTI) of occupations, we also extend equation (2) to include time trends for occupations with different routine task contents (RTI x time). The analysis shows that eliminating such trends does not lead to changes in our main findings.[27]

[23] Extending the placebo analysis, we have also estimated separate annual effects for every year from 1995 to 1999 (Table A5, Figure A1). In the case of marriage, divorce, and new births, 28 out of 30 (93%) of these coefficients are statistical zeros, and those that are not do not generally confirm the gender difference that we find for our sample period of 1999-2009. In the end, these placebo results are not surprising, both because China stood out in terms of the number of binding quotas and because the EU sought to remove in the first two MFA removal phases, in 1995 and 1998, those quotas that were not effective in protecting EU markets, as also earlier studies have noted.

[24] In addition, our placebo results are similar whether we employ the indicator or continuous treatment variable for the quota removal (results available upon request).

[25] We have also analyzed different groups of workers separately, generally finding that the evidence for gender differences in the impact of rising import competition does not strongly vary across occupations and education.

[26] The treatment definition is based on which exact 8-digit products firms produce within Denmark and no single six-digit industry consists entirely of treated or control firms.

[27] See Table X3 and Appendix Table F17 for these results.

# 4 Family and Labor Market Responses to Import Competition: Gender Matters

This section shows that the labor market consequences of import competition are far from gender-neutral, with women experiencing larger labor market adjustment costs than men. Are the greater labor market losses of women the flip side of the shift towards more family activities? We continue in this section by showing that women, not men, respond to rising import competition by increasing family activities. This is in part a substitution effect from market work to family activities driven by the worker's lower opportunity costs.

## 4.1 Gender Inequality Through Trade Adjustment

We show that import competition from China has a substantial impact on workers' long-run labor market outcomes by estimating equation (1) for a number of key outcomes; Table 2 shows the results. Recall that ImpComp denotes the difference-in-differences term *Post x Exposure*. In column (1) the dependent variable is the cumulative earnings of workers over the pre- (1999-2001) and post-periods (2002-2009), measured relative to the worker's initial earnings.[28] Quantitatively, the coefficient of -0.618 means that the import shock leads to a long-term earnings reduction of about 62% of the initial earnings over 2002-2009, or about 8% of the initial earnings per year of exposure.

Next, we are interested in the role of employment changes in this earnings loss. The dependent variables are the cumulative years with primary employment and the cumulative hours worked in columns (2) and (3), respectively. Exposure to import competition does not reduce years of employment; however, it leads to significantly less hours worked (columns (2) and (3) respectively).[29] In particular, column (3) shows a reduction in cumulative hours worked amounting to 45 % of the initial annual hours worked. This indicates that the majority of the earnings effect is driven by the loss of hours worked because exposed workers spend disproportionately less time at work after the import shock.

Further, we see that the reduction in earnings and hours per year of employment amounts to 7.3% and 6.3% of an initial annual earnings and hours worked respectively over 2002-2009 (columns

[28] We use the average of annual earnings over 1996-1999 to minimize possible measurement error.

[29] Given that the years of employment variable is based on employment status in a particular week in late November, the result indicates that import competition does not cause long-term unemployment.

(4) and (5)). The fact that these figures are similar confirms that the reduction in earnings is mostly driven by fewer hours worked. The difference of one percentage point reflects a shift to lower-paying jobs as the result of rising import competition.

Table 2: **Effects of Import Competition on Gender Inequality in Labor Market Outcomes**

| | (1) Earnings | (2) Employment | (3) Hours | (4) Earnings per year | (5) Hours per year | (6) Unemployment | (7) Personal Income |
|---|---|---|---|---|---|---|---|
| **Panel A.** | Long-run Consequences of Trade Adjustment | | | | | | |
| ImpComp | -0.618** | -0.045 | -0.447** | -0.073* | -0.063** | 1.057*** | 0.157 |
| | (0.314) | (0.086) | (0.212) | (0.038) | (0.025) | (0.360) | (0.120) |
| **Panel B.** | Long-run Consequences of Trade Adjustment by Gender | | | | | | |
| ImpComp | 0.051 | -0.033 | -0.155 | 0.019 | -0.013 | 0.777* | 0.205 |
| | (0.425) | (0.097) | (0.295) | (0.045) | (0.028) | (0.445) | (0.208) |
| ImpComp x Female | -1.274** | 0.086 | -0.531* | -0.191*** | -0.103*** | 0.034 | -0.074 |
| | (0.515) | (0.118) | (0.32) | (0.06) | (0.038) | (0.449) | (0.218) |
| **For both panels.** | | | | | | | |
| Observations | 19,526 | 19,526 | 19,300 | 19,744 | 18,438 | 19,526 | 19,526 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |

**Notes:** The dependent variables, given on top of columns, are defined as follows: Earnings is the cumulative labor earnings, periods 1999-2001 and 2002-2009; Employment is the total number of years of employment over 1999-2001 and 2002-2009; Hours is the cumulative hours worked over 1999-2001 and 2002-2009. Unemployment is the cumulative time spent in unemployment over 1999-2001 and 2002-2009. Personal Income is the cumulative personal income over 1999-2001 and 2002-2009. Personal income includes labor earnings, unemployment insurance and government transfers. All earnings, hours worked, and income variables are normalized by worker $i$'s average annual outcome over 1996-1999. The sample is all full-time 1999 textile and apparel workers. Panel A: Estimation of equation (1) by least squares. Panel B: Estimation of equation (2) by least squares. Robust standard errors clustered at the level of workers' initial (1999) firm are in parentheses. ***, ** and * indicate significance at the 10 %, 5% and 1% levels respectively.

Does the loss of hours worked represent voluntary leaves or involve spells of unemployment? Results in column (6) indicate that rising import competition increases unemployment; quantitatively, the coefficient implies more unemployment by about 1 month.

The final specification in Panel A of Table 2 shows the impact of rising import competition on workers' personal income, which includes insurance and government transfer payments in addi-

tion to labor earnings. We see that trade exposure does not significantly lower personal income (column (7)). This is in part due to the relatively high level of person- (not job-) oriented support through Denmark's active labor market policy. It is important because it means that responses to rising import competition will relatively strongly reflect a positive substitution effect, rather than a negative income effect. We will return to this issue in section 6 below.

Overall, Panel A of Table 2 documents that rising import competition due to the quota removal for China leads to lower labor market opportunities of workers. Now we will examine the gender dimension in this. Panel B of Table 2 shows results from estimating equation (2) for the same labor market variables.

Our results indicate that the labor market impact of rising import competition varies strikingly by gender. In particular, the earnings point estimate for men is close to zero and not significant, whereas the triple difference-in-difference coefficient (ImpComp x Female) is significantly negative, with women losing on average about 120 percent (0.051 + (-1.274)) of the initial earnings over 2002-2009.

Substantially lower cumulative earnings of women are not the result of fewer years with employment (column (2)). Rather, hours worked fall disproportionately for women, though less than earnings. Thus, women disproportionately lose employment hours that pay relatively well. Results in columns (4)-(5) confirm that women experience significantly lower earnings and hours worked per year of employment, but they also strongly move to less well-paid employment.

The import shock causes significant unemployment for exposed workers (column (6)). However, women are not disproportionately prone to trade-induced unemployment, so it is not a driver for the gender differential in earnings.[30]

Finally, it is important to see that in contrast to labor earnings, there is no significant difference between the income impact of trade exposure for women and for men (column (7)). In a standard model of the family, the income effect, which is expected to have stronger impact on men as opposed to women under household specialization, should decrease family activities (as long as they are normal goods). But in Denmark, the income effect is muted via transfers.[31] On the other hand, the opportunity cost effect of the negative labor demand shock and the resulting substitution

[30] In addition to employment disruptions or work in lower-pay jobs, earnings changes can be due to moving outside of the labor force (e.g. early retirement, sickness, or maternity leaves) or unemployment. It turns out that movements outside of the labor force for early retirement or for sickness leave are not important adjustment dimensions. Results are available upon request.

[31] Utar (2018) shows that trade exposure significantly increases income transfers, especially transfers paid out of the Unemployment Insurance Fund.

from market to family activities, which is expected to be stronger for women, will not be affected by income support.

Given that we show women that incur substantially larger costs over the long-run, an immediate question is whether this is driven by women being disproportionately more exposed to the shock than men. Workers will respond to the shock by moving to different jobs, industries or occupations. The impact on cumulative earnings will be the long-run impact of the shock inclusive of workers' adjustment to their displacement (if any). Our results could potentially be due to the initial shock, rather than differences in adjustments to the shock. Exposure may trigger lower earnings in the initial firm because workers are laid-off, or have now reduced hours of work or hourly wages.[32] Figure X2 shows how the gender differential in cumulative earnings from all jobs and the cumulative earnings at the initial firm evolved over the period 2002 to 2009. The results show no gender gap in earnings obtained at the initial firm, suggesting that the difference between men and women in labor market outcomes is driven by the differential response of men and women after the initial displacement.[33]

Our results demonstrate that the labor market consequences of import competition are far from gender-neutral, with women experiencing substantially greater labor market losses than men. Are the greater labor market losses of women the flip side of the shift towards more family activities? This is addressed in the next section.

## 4.2 The Fertility Response to Rising Import Competition

In this section we study the relationship between rising import competition and fertility decisions. Our outcome variable is one if the worker has become a parent to a newborn child during a particular period, and zero otherwise.[34] The sample is the set of fertile-age women and men, defined to be between 18 and 39 years old as of the year 1999.[35] Table 3 shows the results from estimating equations (1) and (2).

[32] Among the reasons for this may be voluntary separations out of fear of job loss or reduced earnings in the future.

[33] We confirm this finding also with employment and hours worked at the initial firm; see also Figure G7 in the Appendix, section G.

[34] We only consider biological parents.

[35] Our results are robust to using alternative age limits for fertile age period, see Appendix, section E.

Table 3: **Import Competition and Newborn Children**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Gender | All | All | All | Men | Women | All | Men | Women |
| Sample | | Not married | | | | Single | | |
| ImpComp | 0.022 | 0.061** | 0.053 | 0.053 | 0.077** | -0.019 | -0.019 | 0.109*** |
| | (0.029) | (0.026) | (0.035) | (0.035) | (0.037) | (0.037) | (0.037) | (0.041) |
| ImpComp x Female | 0.008 | | 0.024 | | | 0.128** | | |
| | (0.034) | | (0.05) | | | (0.055) | | |
| Observations | 10,418 | 5,784 | 5,784 | 2,808 | 2,976 | 3,160 | 1,708 | 1,452 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time FE | ✓ | | ✓ | | | ✓ | | |

**Notes:** Dependent variable is one if worker $i$ has a newborn child during period $s$, and zero otherwise. The sample in all columns is textile workers of fertile age (between 18 and 39 years old as of 1999). In columns (2) to (5) the sample is limited to workers not married as of 1999, and in columns (6) to (8) it is limited to workers who are single as of 1999 (neither married nor co-habiting). Estimation of equation (1) in columns (2), (4), (5), (7), (8) and of equation (2) in columns (1), (3) and (6) by least squares. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

A first result is that there is no evidence that import competition leads to lower fertility. On the contrary, the estimates for men and women are positive though insignificant (column (1)). Thus, even though the import shock significantly reduces labor earnings of workers as shown in Table 2, it does not lead to fewer newborn children.

The focus on married couples might influence our findings if the fertility effect of rising import competition for men were negative but for women positive – couple decisions are to some extent joint. Also, in twelve percent of all married couples, both spouses are textile workers. While this does not appear to be a particularly high number, the textile sector in Denmark is concentrated in a few areas (an important cluster is around the city of Herning). Local spillovers may reinforce the potential bias due to couples, and thus prevent us from estimating gender-specific responses to rising import competition.

To examine this, we limit the sample to workers who were not married as of year 1999. They can

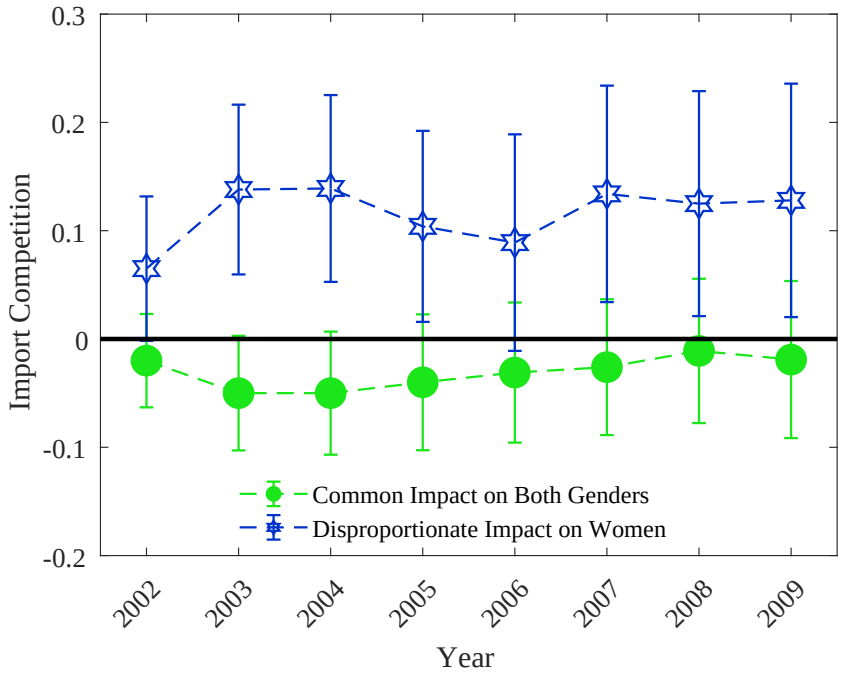


Figure 3: **Import Competition and Fertility**

**Notes**: The figure draws the coefficient estimates obtained from estimating equation (2) on a rolling end-year basis when the dependent is child birth and the sample consists of single fertile age textile workers. Stars indicate the coefficient estimates of ImpComp (Post x Exposure), circles indicate the coefficient estimates of ImpComp x Female (Post x Exposure x Female). Shown is the 95% confidence interval.

be co-habitating with a partner, or they can be single. Column (2) shows that increased import competition increases birth rates for these workers. To understand the size of the fertility impact of import exposure, note that the average of the dependent variable in column (2) is 0.31, so that three in ten workers in the sample have one or more newborn children during the years 1999 to 2009. The coefficient of 0.061 in column (2) means that trade exposure raises the probability of birth by one fifth. Thus, the trade-induced increase in fertility is substantial.

The following three columns show that the impact of trade exposure on fertility is driven mostly by women. First, we see that while the interaction specification in column (3) is qualitatively similar to before, quantitatively the tendency to have more children is stronger for unmarried than for all workers. Separate regressions for male and female workers in columns (4) and (5) show that unmarried women respond by giving births. One in three of unmarried women have one or more births during the sample period, which means that trade exposure raises fertility by just under one quarter (= 0.077/0.33). The coefficient for men is also positive, but only about one third in size and not significant.

We can go further by separating workers who live with a partner (co-habitating) from those workers

who have no partner (single), see columns (6) to (8) for the results. If workers lose their job and have to move to other cities to find a new job they might be exposed to new networks which could increase the likelihood of finding better matches. This may be an additional factor explaining why the response of single workers is stronger than for co-habitating or married ones, in line with the local spillovers argument above.[36] From Table 3 we see that one in three workers who can have children (they are of fertile-age in 1999) is single, and singles account for more than half of all unmarried fertile-age workers.

We see that the point estimate for men now is negative (insignificant), and the female interaction coefficient is more than five times as large as for all unmarried workers (column (6) versus column (3)). This gender difference is confirmed by performing separate specifications for men and women (columns (7) and (8)). Specifically, the coefficient in column (8) means that for single women, import exposure accounts for close to half of all childbirth (= 0.11 relative to the mean of 0.23). Turning to the overtime effect among the single men and women, Figure 3 shows the results for changing the end year of the post-shock period from 2002 until 2009. We see that the difference between men and women becomes significant at the 5% level in 2003, and after eight post-shock years the disproportionate impact on women is still positive and significant.

Overall, these results indicate not only that import competition has a sizable impact on increasing fertility but they also demonstrate that the earnings impact of rising import competition is likely to manifest itself over a long period. Single workers are almost by definition at an early stage of their lives, implying that their fertility choice will affect a relatively large part of their life.[37]

## 4.3 Import Competition and Parental Leave Taking

We now turn to the relationship between import competition and workers' decision to take parental leave. The outcome variable is a zero-one variable, indicating a spell of parental leave over the pre- or post-period.[38] While some of the leave parents take may be associated with newborn children, in principle our analysis encompasses also parental leave taken for existing children. Table 4 shows the results.

The outline of the parental leave analysis follows that of new births in the previous section. In-

[36] See Huttunen, Moen, and Salvanes (2018) on increased regional mobility after job displacement.

[37] Similar findings to Table 3 are obtained when we employ probit models that control for an extensive set of worker, firm, and partner characteristics, see Table D10 in Appendix D.

[38] This indicator variable is preferred given that women tend to take longer parental leaves than men. Parental leave may be thought of as a more incremental move towards family activities compared to birth.

terestingly, we find quite similar results.[39] This suggests that the parental leave effect of import competition is mainly driven by newborn children. First, notice that import competition does not lower parental leave take-up; if anything it causes an increase, although the coefficients in column (1) are not precisely estimated.

When we focus on workers who are unmarried as of 1999 to account for possible local spillovers, we find that exposed workers have a significantly higher likelihood of taking parental leave than workers who are not exposed to rising import competition (column (2)). Quantitatively, the coefficient of 0.059 means that the marginal impact of trade exposure is about one quarter of all parental leave taking in the sample (= 0.059/0.26). This is somewhat higher than for new childbirths (20 percent). Furthermore, we see that women are contributing to the trade-induced increase in parental leave more than men (columns (3) to (5)).

Table 4: **Parental Leave and Import Competition**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| Gender | All | All | All | Men | Women | All | Men | Women |
| Sample | | Not married | | | | Single | | |
| ImpComp | 0.035 | 0.059** | 0.044 | 0.044 | 0.067* | -0.028 | -0.028 | 0.095** |
| | (0.023) | (0.023) | (0.028) | (0.028) | (0.037) | (0.030) | (0.030) | (0.039) |
| ImpComp x Female | 0.0001 | | 0.023 | | | 0.122** | | |
| | (0.029) | | (0.046) | | | (0.048) | | |
| Observations | 10,418 | 5,784 | 5,784 | 2,808 | 2,976 | 3,160 | 1,708 | 1,452 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time FE | ✓ | | ✓ | | | ✓ | | |

**Notes:** Dependent variable is one if worker $i$ takes parental leave during period $s$, and zero otherwise. The sample in all columns is textile workers of fertile age (between 18 and 39 years old as of 1999). In columns (2) to (5) the sample is limited to workers not married as of 1999, and in columns (6) to (8) it is further limited to single workers as of 1999, that is, neither married nor co-habiting. Estimation of equation (1) in columns (2), (4), (5), (7), (8) and equation (2) in columns (1), (3) and (6) by least squares. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. $^{*}$, $^{**}$ and $^{***}$ indicate significance at the 10 %, 5% and 1% levels respectively.

As in the case of childbirth, the impact as well as the gender differential is further strengthened

[39] At the same time, the parental leave and birth data come from two independent data sources. The parental leave data come from income statistics registers while we derive the birth data from the population registers.

when we concentrate on single workers (columns (6) to (8)). Now, the coefficient estimate for men is negative and women's response is significantly different than men's. The magnitude of the gender differential is comparable to that of child birth, and the marginal impact of trade exposure is about 45 percent of all parental leave taking for single women (= 0.095 relative to a mean of 0.21). Figure G5 in the Appendix shows the overtime evolution of the impact across men and women, which is strikingly similar to the impact on births. This confirms the large impact of import exposure that we have seen for child birth in Table 3.[40]

Summarizing, exposure to rising import competition increases not only fertility but also parental leave uptake for 1999 textile workers. The increase in fertility and parental leave is driven by women, not men. In particular, it is single women who are at a relatively early stage of their lives that in the face of lower labor market opportunities shift towards child-related activities. Given that the incidence is concentrated on early-stage workers who would typically not retire from the labor market for many years, the import competition will have long-run consequences not only via its direct impact on earnings but also via the family channel because early-stage career interruptions can be damaging especially for women (Del Bono, Weber, and Winter-Ebmer 2012).

## 4.4 Getting Married in Response to a Trade Shock

Table 5 shows evidence on marriage behavior in the face of rising import competition. The sample includes all workers who are not married as of the year 1999.[41]

The first important result is that workers do not marry less due to rising import competition (columns (1) to (3)). The point estimate for men is negative but close to zero and imprecisely estimated, whereas for women we find that import exposure increases female workers' likelihood to get married. How large is the marriage impact of rising import competition? A back-of-the-envelope calculation compares the marginal effect of import competition with the average marriage probability in the sample. Rising import competition accounts for a sizable portion, 30% (= 0.058/0.19) of the overall marriage probability in the sample (based on the results for women in column 3).[42]

Next, we present results for fertile-age workers (ages 18 to 39 in year 1999) whose fertility and

[40] Supplementary results using probit models are presented in the Appendix, Table D11, and confirm these results.

[41] The marriage decision is directly relevant only for unmarried workers. Workers who in 1999 are married would have to divorce before marrying again; we analyze divorce in the following section.

[42] Similar findings are obtained when we employ probit models, see Table D12 in Appendix D.

parental leave responses we analyzed above (columns (4) to (6)). The Female interaction coefficient is positive and higher than before (compared with column (1)). The increase in marriage likelihood induced by trade disproportionately results from responses of fertile-age women. This finding is in line with our fertility and parental leave findings and underlines that marriage and child-related responses come to some extent in a bundle for these women.

It is useful to contrast our result with the finding that the overall marriage rate in the U.S. has fallen due to import competition from China (Autor, Dorn, and Hanson 2019). Our result is of interest from a policy perspective because it indicates that the institutional framework may be important in shaping family responses to import competition. The level of government transfers and insurance payments to workers who lose their jobs is greater in Denmark than in the United States. As shown in Table 2, while labor earnings are down there is no significant negative effect of import exposure on the (transfer-inclusive) personal income of Danish workers. This is likely why we do not observe a similar decline in marriage rates as estimated for the United States. In contrast, we find evidence for higher marriage rates due to substitution from labor market to family activities, in part because negative income effects are not as dominant as they are in the United States.[43]

In the final set of marriage results we focus on workers who in 1999 were single; that is, without a partner to co-habitate with (columns (7) to (9)). We see that the negative labor shock particularly induces single fertile age women, but not men, to marry. Exposure to rising import competition causes the relatively drastic change from single to married family status, not just the comparatively incremental step from co-habitation to marriage.[44] This means that the marriage response of women cannot entirely be explained by insurance motivations. Rather, these results highlight the fact that women's re-optimization after a negative labor shock is likely to have consequences for long-run career outcomes, family structure, and gender inequality.

[43] In addition, see section 6.3 for a comparison of our estimation approach with that in work on job displacement.

[44] Exploring the impact of import competition on the joint decision of marriage and childbirth, we also find a positive response by women who were single in 1999, which supports the idea that women's move toward family is motivated by fertility considerations. The results are available upon request.

Table 5: Marriage Decisions and Import Competition

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| Gender | All | Men | Women | All | Men | Women | All | Men | Women |
| Sample | | | | Fertile Age | | | Single | | |
| ImpComp | -0.019 | -0.019 | 0.058** | -0.012 | -0.012 | 0.080** | -0.026 | -0.026 | 0.094** |
| | (0.027) | (0.027) | (0.03) | (0.034) | (0.034) | (0.037) | (0.033) | (0.033) | (0.042) |
| ImpComp x Female | 0.077** | | | 0.092** | | | 0.119** | | |
| | (0.036) | | | (0.046) | | | (0.051) | | |
| Observations | 8,216 | 3,876 | 4340 | 5,784 | 2,808 | 2,976 | 3,160 | 1,708 | 1,452 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time | ✓ | | | ✓ | | | ✓ | | |

**Notes:** Dependent variable is one if worker *i* married during period *s*, and zero otherwise. Sample is all unmarried textile workers. Estimation of equation (1) in columns (2), (3), (5), (6), (8), (9) and equation (2) in columns (1), (4) and (7) by least squares. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

## 4.5 The Impact of Import Competition on Dissolution of Marriages

The final step in our analysis of family responses to trade exposure is to examine workers' likelihood of divorce. Note that being married typically means that workers are at a later stage in their lives, as reflected in their average age of about 42 years, in contrast to unmarried workers who are on average about 34 years (see Table X1). Given this age difference one would not necessarily expect that the motives of being in a marital union are the same for the two sets of workers. Table 6 shows results from estimating equations (1) and (2) when the outcome variable is divorce.[45]

We find that exposure to import competition significantly reduces divorce likelihood. The coefficient in the sample with both gender is negative at -0.03 (column (1)). On average, the divorce rate for these workers is 0.057, so the impact of trade exposure is to reduce divorce rates by more than half.[46]

[45] Analogous results using probit regressions are presented in Table D13 in the Appendix.

[46] There are a number of reasons why exposure to import competition might lead to lower divorce rates. One of them is insurance. When employment opportunities diminish due to rising import competition, an existing marital union may provide income security that not exposed workers do not need to the same extent. We return to this issue by examining the role of partner income in Table X4.

The next set of results shows that exposure to import competition leads to stronger reduction in the divorce likelihood of women. Column (2) presents results for the pooled sample, while estimates for men and women separately are shown in columns (3) and (4) respectively. The marginal impact of trade exposure on divorce for women evaluated at the average divorce rate is large (the average divorce rate for the sample underlying column (4) is 5.4 percent, and -0.040/0.054 is about a fall of 75 percent. The result that trade exposure significantly reduces divorce rates for women is in line with our findings on the shift of women to family activities in other dimensions. To the extent of our knowledge, it is also the first result on the impact of import competition on divorce behavior.

The marriage analysis above showed that women of fertile age tend to react more strongly to trade exposure than older workers. We have also seen that such workers respond strongly to rising import competition in terms of fertility and parental leave. Is it possible that planning for children plays a role for divorce decisions? In columns (5)-(8) of Table 6 we focus on divorce decisions of the relatively young, fertile-age workers (aged 18 to 39 in year 1999).

We find that while men's divorce response to rising import competition is not much affected by age (compare columns (7) and (3)), fertile-age women respond roughly twice as much to trade exposure as the average married woman (columns (4) and (8)). Fertile-age women are also behind the larger overall negative trade impact on divorce, see columns (1) and (5). These findings are consistent with women's fertility goals being a major reason for these divorce responses.

Summarizing, we show that trade exposed workers increase their family activities in several dimensions. Workers have both a higher likelihood to get marry and a lower likelihood to divorce in response to rising import competition, and women are central to the shift towards family activities. This is consistent with the fertility and parental leave behavior documented above. It suggests that the different margins of women's increased family activity are driven by the same underlying force, namely child-bearing and child-rearing.

Table 6: **Exposure to Import Competition Reduces Divorce Likelihood**

| Gender | (1) All | (2) All | (3) Men | (4) Women | (5) All | (6) All | (7) Men | (8) Women |
|---|---|---|---|---|---|---|---|---|
| | | | | | Fertile Age | | | |
| ImpComp | -0.030*** | -0.021 | -0.021 | -0.040*** | -0.057*** | -0.024 | -0.024 | -0.086*** |
| | (0.009) | (0.014) | (0.014) | (0.011) | (0.018) | (0.026) | (0.026) | (0.022) |
| ImpComp x Female | | -0.019 | | | | -0.062* | | |
| | | (0.017) | | | | (0.032) | | |
| Observations | 11,780 | 11,780 | 4,934 | 6,846 | 4,634 | 4,634 | 1,840 | 2,794 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time | | ✓ | | | | ✓ | | |

**Notes:** Dependent variable is one if worker $i$ has a divorce during period $s$, and zero otherwise. Sample is textile workers who are married as of 1999. Estimation of equation (1) in columns (1), (3), (4), (5), (7), and (8), estimation of equation (2) in columns (2) and (6). Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. $^{*}$, $^{**}$ and $^{***}$ indicate significance at the 10 %, 5% and 1% levels respectively.

# 5 Responses of all Private-Sector Workers

While analyzing the quota removal trade liberalization has a number of advantages, a concern may be that our findings do not generalize because they are based on a relatively small sample. In this section, we present economy-wide results, based on the cohort of workers who in 1999 were employed in Denmark's private sector.[47]

The impact of import competition is estimated by exploiting six-digit industry variation in the change of import penetration in Denmark. Because the change in imports from China in Denmark might be endogenous, we employ an instrumental-variables approach with geography-based transportation costs, the strength of the pre-existing distribution channels, and Chinese imports in other advanced countries (all at the six-digit level) as instrumental variables. An important identification condition of this approach is that China's export growth is primarily driven by its economic reform

[47]These workers were employed in mining, manufacturing, wholesale and retail trade, hotels and restaurants, transport, storage and communication, as well as real estate, renting and business activities. Workers initially employed in public administration, education, and health are not part of our sample as education and health sectors in Denmark are to a large extent publicly owned. At the same time, we follow all 1.6 million workers in our sample wherever they go, including to the public sectors. As before, we focus on workers who in 1999 were employed in full-time jobs.

and the global decline in trade costs.[48]

In the case of child birth, the estimation equation is the following:

$$BIRTH_i^{2000-09} = \beta_0 + \beta_1 \Delta IMP_i + \beta_2 \Delta IMP_i \; x \; Female_i + Z_i^W + Z_i^F + Z_i^P + \varepsilon_i. \tag{3}$$

The dependent variable $BIRTH_i^{2000-09}$ is an indicator variable that takes the value one if worker $i$ is the biological parent of a newborn child over the period from 2000 to 2009, and zero otherwise. On the right hand side is the change in import penetration in worker $i$'s six-digit industry from China over 1999-2009 ($\Delta IMP_i$) as well as an extensive set of worker ($Z_i^W$), firm and six-digit industry ($Z_i^F$), and partner $Z_i^P$ characteristics as of year 1999. We include two-digit occupation and industry fixed effects as well as the share of college educated workers in each six digit industry to address industry- and occupation-specific shocks, including technology shocks.[49] The term $\varepsilon_i$ is a mean zero regression residual; we allow for dependence by clustering at the industry level. Linear terms are always included when the equation has an interaction coefficient. Specifications for other family outcomes (parental leave, marriage, and divorce) and labor market outcomes are analogous to this birth equation, with the following changes: the interaction of female with change in import penetration is dropped if the sample is only women or men, and it is replaced by exposure to import competition of the spouse in the divorce equation. The spousal exposure to import competition is also instrumented similarly as the worker's own exposure using the information on the spouse's workplace. In this section we provide key second-stage results (Tables 7 and 8) while a more detailed discussion including first-stage regressions is provided in Appendix C.

Results for the impact of import competition on childbirth are shown for the close to the 1 million workers who are of fertile-age (between 18 and 39 years old in 1999) in column (1) of Table 7. We find evidence for a gender differential in trade-induced fertility decisions, with a positive and significant female interaction coefficient of about 0.3, and a coefficient for men that is small and

[48] This and other aspects of our approach are discussed in Appendix C. The other countries employed in this approach are Australia, Finland, Germany, Japan, Netherlands, New Zealand, Switzerland, and the USA. The transportation cost variable is defined as the log average of the distance from Denmark's import partners in 1996.

[49] $Z_i^W$ includes, among others, age, immigration status, marital status, children, education, occupation, wage, history of unemployment spells, labor market experience characteristics. $Z_i^F$ includes firm wage, firm size, the separation rate, in the employment in the six-digit product line of employer between 1993-1999, the share of college educated workers in the six-digit product line of employer. $Z_i^P$ includes whether the individual has a partner (if not married), the partner's age, the partner's labor earnings, an indicator whether the partner is a Danish citizen, an indicator whether the partner is employed in manufacturing, an indicator whether the partner is employed in the same six-digit product line, and an indicator whether the partner is employed in a highly trade exposed industry (95th percentile and above of trade exposure), the age difference between the partners, and an indicator whether the partner has higher earnings. The full list of variables is presented in the Appendix C.

negative. The impact of import competition on the likelihood of taking parental leave is significant and negative for men, as shown in column 2. It is less negative for women, though not significantly so. This indicates that import competition tends to decrease women's likelihood of taking parental leave if it is not associated with birth. The finding of an overall non-negative fertility effect is in line with what we obtain in the quota removal experiment (see Table 3 above). Furthermore, a gender difference, that women have a more positive fertility response in comparison to men, is also what we find for the quasi-experiment sample.

Turning to the impact of import competition on marriage, there is no evidence that import competition from China strongly affects the marriage likelihood of men (column (3)). This finding generalizes the non-negative marriage response estimated above for the quasi-experiment sample for Denmark as a whole. Comparing our results with those of Autor, Dorn, and Hanson (2019) for the US suggests that institutional differences across countries matter.

In contrast to the insignificant marriage response of men, the marriage likelihood of female workers increases with exposure to rising import competition (the interaction coefficient is significant at the one percent level and about six times larger than that of men). Thus, our finding that import competition increases marriage rates in the quasi-experiment sample carries over to the economy as a whole. Furthermore, we show that increased marriage rates due to import competition are driven by women, not men.

For the analysis of divorce behavior we take advantage of unique person identification across different Danish registers, in particular that for every worker $i$ we know his or her spouse, so we can match spouse's ID with the labor market database to identify spouse's six-digit industry of employment.[50] The results show that there is no significant impact of trade exposure on the divorce behavior of male workers (column (4)). At the same time, their wives' trade exposure significantly lowers the divorce likelihood. That is, for men, rather than their own exposure to import competition it is their spouse's exposure to import competition that matters. These findings for a economy-wide sample confirm our result obtained for the quasi-experiment sample that trade exposure induces female workers more strongly to stay in their marriages than male workers.

[50] This information is also available in our data for co-habitating individuals, and we utilize it in equation (3) to construct $Z_i^P$.

Table 7: **Family Responses to Rising Import Competition II**

| | (1) Birth | (2) Parental Leave | (3) Marriage | (4) Divorce |
|---|---|---|---|---|
| $\Delta IMP_i$ | -0.190* | -0.293** | 0.073 | -0.025 |
| | (0.103) | (0.122) | (0.115) | (0.076) |
| $\Delta IMP_i$ x Female | 0.314*** | 0.116 | 0.451*** | |
| | (0.092) | (0.102) | (0.108) | |
| Spouse's $\Delta IMP_i$ | | | | -0.142*** |
| | | | | (0.047) |
| Observations | 903,629 | 903,629 | 757,302 | 478,354 |
| Sample | Age 18-39 | Age 18-39 | Unmarried | Married Men |
| Two dig. industry FE | ✓ | ✓ | ✓ | ✓ |
| Two dig. occupation FE | ✓ | ✓ | ✓ | ✓ |
| Worker, firm, partner charac. | ✓ | ✓ | ✓ | ✓ |
| SW F-stat ( $\Delta IMP_i$) | 12.79 | 12.79 | 13.02 | 9.95 |
| SW F-stat ($\Delta IMP_i$x Female) | 15.08 | 15.08 | 14.41 | - |
| SW F-stat (Spouse's $\Delta IMP_i$) | 15.08 | 15.08 | 14.41 | 806.44 |
| Hansen J | 3.781 | 0.83 | 1.634 | 7.962 |
| Hansen J P-value | 0.151 | 0.66 | 0.442 | 0.093 |
| No. of Clust | 756 | 756 | 757 | 752 |

**Notes:** Dependent variable is given at top of column. Dependent variable is an indicator for that family outcome of worker $i$ over 2000-2009, and zero otherwise. Estimation by two-stage least squares; second-stage coefficients shown. S-W stands for Sanderson-Windmeijer. Robust standard errors clustered at the industry level in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table 8 shows evidence of the gender differential in labor market outcomes from rising import competition in the economy as a whole (N = 1,651,774). The results indicate that the negative impact of import competition on labor earnings is found only for women (column (1)). The linear coefficient is positive but close to zero, while the female interaction coefficient is negative and significant. This result parallels our findings for the case of the textile quota removals.

In column (2), the dependent variable is the cumulative earnings obtained from the initial six-digit industry. For both men and women, import competition leads to significant earnings reductions while they work in the exposed industry. In contrast, in the long-run exclusively women do not

recover from these earnings losses (column (1)). This result, too, mirrors our findings for the case of the textile quota removals.

Table 8: **Gender Differential in Labor Market Adjustment to Trade II**

| | (1) Earnings | (2) Earnings 1999 Job | (3) Hours | (4) Hours 1999 Job | (5) Unem-ployment | (6) Personal Income |
|---|---|---|---|---|---|---|
| $\Delta IMP_i$ | 2.946 | -10.220* | 0.968 | -8.401* | 10.070* | 3.171 |
| | (2.713) | (5.66) | (1.575) | (4.711) | (5.253) | (2.158) |
| $\Delta IMP_i$ x Female | -13.050** | 1.035 | -8.038** | 0.531 | 18.530** | 2.455 |
| | (6.055) | (2.495) | (3.518) | (1.946) | (8.835) | (6.546) |
| Observations | 1,651,774 | 1,651,777 | 1,642,413 | 1,642,413 | 1,651.77 | 1,651,757 |
| Two dig. industry FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Two dig. occupation FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Worker, firm, partner charac. | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| S-W F-stat ( $\Delta IMP_i$) | 13.45 | 13.45 | 13.44 | 13.44 | 13.45 | 13.45 |
| S-W F-stat ($\Delta IMP_i$x Female) | 14.90 | 14.90 | 14.93 | 14.93 | 14.90 | 14.90 |
| Hansen J-stat | 0.898 | 3.106 | 0.609 | 2.649 | 0.754 | 1.217 |
| Hansen J P-value | 0.638 | 0.212 | 0.737 | 0.266 | 0.686 | 0.544 |
| No. of Clusters | 761 | 761 | 761 | 761 | 761 | 761 |

**Notes:** Dependent variable is given at top of column. All variables are cumulative over the period 2000 to 2009. The earnings, hours and income variables are measured in worker $i$'s initial earnings, hours worked and income, respectively. Estimation by two-stage least squares; second-stage coefficients shown. S-W stands for Sanderson-Windmeijer. Robust standard errors clustered at the industry level in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Furthermore, in the quota removal liberalization we have seen that earnings losses were mainly driven by reduced market work. We see that the differential reduction in hours worked is also an important reason in this broader sample why the gender earnings gap increases, see column (3). In column (4) we confirm that the gender difference in hours worked rises after displacement from the exposed industry.

Furthermore, import competition significantly increases unemployment, and disproportionately so for women (column (5)).[51] Finally, Table 8 shows the impact of rising import competition

[51] In the quasi-experiment, in contrast, we do not find evidence that rising import competition leads to a gender differential in unemployment.

on the workers' personal income (labor earnings plus insurance payments plus transfers). As in the case of the quota liberalization, there is no evidence that rising import competition has significantly reduced personal income (column (6)). This is consistent with the highly effective, person-oriented policies in Denmark. Furthermore, there is no significant difference in the impact of import competition on the personal income of women, versus men.

Overall, family and labor market responses of Danish workers to rising import competition as summarized in Tables 7 and 8 are quite similar to the responses of the subset of 1999 textile workers to the MFA quota removals.

# 6 Biological Clock: On Gender Differences in the Market versus Family Response

This section lays out our biological clock hypothesis for the substantial gender difference in response to the trade shock both in the labor market and for family. We support our hypothesis by showing that the gender difference in trade adjustment costs is not driven by (1) stronger exposure of women to the shock, (2) sample composition or (3) technological shocks; and that the long-term gender gap occurs among younger workers as opposed to workers who are past their fertile period. Further evidence for the biological clock argument comes from documenting heterogeneity of female responses in demographic, education and occupation dimensions. We also discuss the role of opportunity cost, income, and substitution effects in our context. The final part of this section is devoted to methodology by comparing our approach with that in job displacement studies.

## 6.1 The Biological Clock Explanation

We argue that our findings are related to child-bearing and child-rearing activities, in which women are more involved than men. As shown above, women respond more strongly to the import shock by moving towards family activities than men, especially relatively young, fertile-age workers. Now we show that these younger women do this at the expense of successful trade adjustment and better careers in the labor market. Table 9 presents the results.

Table 9: **The Earnings Differential by Age and Stage of Life**

| Sample | (1) All | (2) Fertile Age | (3) Not Fertile Age | (4) Married | (5) Not Married | (6) Family Activity 2002-9 |
|---|---|---|---|---|---|---|
| Panel A. Dependent Variable: Earnings from all employment | | | | | | |
| ImpComp | 0.051 | 0.822 | -0.919*** | -0.491 | 0.906 | 0.731 |
| | (0.425) | (0.599) | (0.352) | (0.457) | (0.608) | (0.689) |
| ImpComp x Female | -1.274** | -1.921** | -0.483 | -0.508 | -2.440** | -2.391** |
| | (0.515) | (0.841) | (0.464) | (0.545) | (0.966) | (1.198) |
| Panel B. Dependent Variable: Earnings from employment at the 1999 firm | | | | | | |
| ImpComp | -1.134*** | -0.876** | -1.423*** | -1.421*** | -0.784** | -0.739* |
| | (0.368) | (0.380) | (0.418) | (0.418) | (0.365) | (0.436) |
| ImpComp x Female | 0.083 | -0.009 | 0.198 | 0.331 | -0.215 | -0.009 |
| | (0.278) | (0.292) | (0.369) | (0.336) | (0.318) | (0.424) |
| For both panels: | | | | | | |
| Observations | 19,526 | 10,234 | 9,292 | 11,490 | 8,036 | 4,727 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |

**Notes:** Sample is all 1999 full-time textile and clothing workers. Dependent variable in Panel A is worker *i*'s cumulative earnings. Dependent variable in Panel B is worker *i*'s cumulative earnings at the initial, 1999, textile firm. They are expressed in worker *i*'s average annual earnings over 1996-1999. Fertile age is defined as ages 18 to 39 in the year 1999. Estimation of equation (2) by least squares. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Panel A of column (1) repeats our earlier finding that women lose about 120 percent of their initial annual earnings due to import competition with China over the eight post-shock years, while men do not have significantly lower earnings. Before we turn to the role of women's biological clock it is useful to address a number of immediate issues.

**Sample Composition** One concern is that maybe the jobs in which women are disproportionately employed are more negatively affected by rising import competition than the jobs in which men typically work. If so, this might explain the difference in long-run earnings between female and male workers. However, the least squares specifications include worker fixed effects, which means only within-worker variation is exploited. Because this amounts to comparing female clerks

with male clerks, female managers with male managers, and so on, such sample differences are addressed by these fixed effects.[52]

**Technology Shocks Interacting with Gender** We have considered shocks other than import competition, in particular because they could interact with gender. Technology shocks related to computer and information communication technology were common in the 2000s, and if they affect more negatively jobs in which women are disproportionately working that might explain our result (column (1), Panel A). To address this, we have extended the specification with differential time trends across occupations, which implicitly attributes all cross-occupation variation to factors other than trade (including technology). We find that our results are robust to this, see Table X3. We have also examined the role of technology by focusing on the routine task intensity (RTI) of different occupations, finding similar results.[53] Women have disproportionately large earnings losses from rising import competition at the same time when they respond more strongly towards family than men.

**Gender Difference in the Size of the Shock** Another reason why earnings losses are concentrated on women (Panel A, column (1)) might be that male workers, for one reasons or another, are not exposed to a same extent to the trade shock as female workers. By examining the impact on the cumulative earnings obtained from the initial firm we can shed light on the size of the shock. Results are given in Panel B of Table 9 and in Figure X2. Notice that when we focus on the effect on earnings at the exposed firm, we see that both gender are significantly affected by the shock; the female interaction coefficient close to zero provides evidence that the initial earnings losses are not concentrated on women. Figure G7 in the Appendix also shows the impact on the hours worked at the initial firm and confirms that trade shock leads to displacement from the exposed firm for both gender, and women are not disproportionately impacted. The gender difference arises when we study the responses of workers to the shock after the initial displacement, in other words, as they adjust to the trade shock.[54]

[52] Another way of seeing this is to compare the estimates of gender differentials obtained from probit regressions with and without detailed four-digit occupation fixed effects. Probit results presented in Table D14 in the Appendix shows that gender differentials are similar whether or not such fixed effects are included. This provides more evidence that our results are not driven by gender sorting on particular occupations or by sample composition.

[53] It is well-established that occupations vary in how susceptible they are to the influence of automation. The RTI index (Autor, Levy, and Murnane 2003) captures this based on data from the Dictionary of Occupational Titles (today the O*NET data base). Using the RTI measure as translated to ISCO codes by Goos, Manning, and Salomons (2014), we interact the routine-task intensity of each worker's initial occupation with time to allow for differential time trends for occupations with different RTIs. The results are in line with our results here, see Table F17.

[54] By studying the timing of the family activities with respect to labor market trajectories of workers, we also show that trade-induced movement of female workers toward family happens mostly after workers are separated from their 1999 firms. See Table H18 in the Appendix for these results.

We now turn to the workers' biological clock by distinguishing younger workers in their fertile-age from older workers, see columns (2) and (3) of Table 9. Notice that the differential long-run earnings effect is much larger (and significant) for fertile-age workers compared to older workers. Because there is no major difference in the size of the trade shock between male and female fertile-age workers (see Panel B, column (2)), this supports the hypothesis that the choices of female workers subsequent to the shock, including family activities, are central to the long-run gender earnings differential. While fertile-age men do not have significantly lower long-run labor earnings, fertile-age women lose on average more than 100 percent of their initial annual earnings over the eight post-shock years (point estimate of -1.921 + 0.822= -1.099).

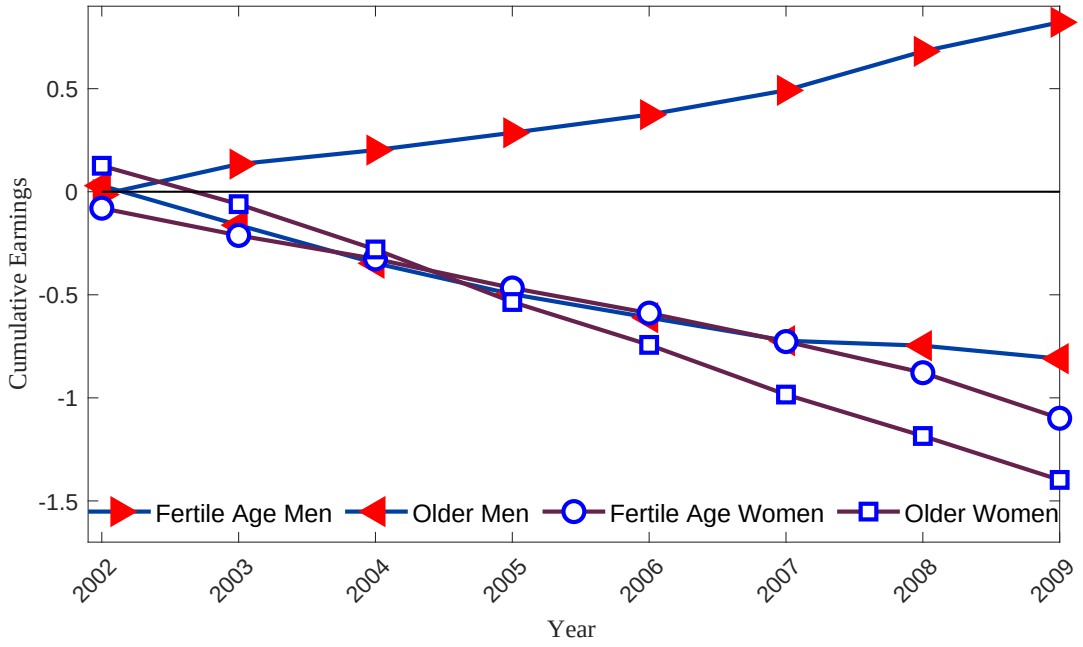


Figure 4: **The 'Missing' Earnings of Young Women**

**Notes:** Shown are *ImpComp* treatment point estimates from equation 1 with four different samples (fertile-age men, fertile-age women, not fertile-age men, not fertile-age women) and eight different endpoints of the treatment period. All regressions include worker and time fixed effects.

Figure 4 develops this striking result by showing labor earnings point estimates for a rolling end-point, 2002 to 2009, for four different sets of workers, female versus male and fertile-age versus older workers, respectively. We see that older workers tend to adjust relatively poorly to the shock, and by 2009 those workers have around one annual salary lower earnings than not exposed work-

ers. That younger workers adjust relatively well is in line with human capital theory as well as recent evidence (Dix-Carneiro 2014, Utar 2018). Importantly, it goes for both male and female workers.

In contrast, while younger male workers adjust to the shock without significant long-run earnings loss, younger female workers' earnings strikingly fall almost as much as the earnings of older female workers. Essentially, in adjusting to the shock, young age does not carry any advantage for the earnings performance of women.

A similar difference in the adjustment of female versus male workers is present when we focus on marital status instead of age, see columns (4) and (5) of Table 9. This supports our biological clock argument because unmarried workers are on average 8 years younger than married workers (see Table X1).[55] Furthermore, it is indeed women's family activity that leads exposed female workers to have a relatively weak earnings performance. This is confirmed by the final specification of Table 9 where we restrict the sample to those workers who have family activity in 2002 or later. In particular, exposed female workers in this sample lose substantially more labor earnings than the typical exposed female worker (compare columns (6) and (1), respectively).[56] Our evidence is strengthened by finding that a woman's move towards family occurs after displacement from the initial job, and typically coincides with a non-employment spell.[57]

In sum, the labor market-family choice eliminates the advantage of being "young" for women in the presence of a negative labor market shock, but not for men. The reason for this is the essence of our biological clock explanation. Birth is a female-time-intensive event by nature, and women, in contrast to men, tend to have difficulties conceiving beyond their early forties. As a consequence, a younger woman's reservation value to stay in the labor market is higher than a same-age man's reservation value. In the presence of a negative labor shock of a given size, a woman will have stronger incentives to take up family activities than a man, versus committing to a new career path with the associated investment in training. This explains both the gender difference in family responses to rising import competition and that women's labor earnings fall behind those of men as a consequence of this shift to family. The argument applies as long as the woman's age is low enough so that child-bearing and child-rearing are still playing a role, as we show in Table 9. If

[55] Consistent with the more strongly negative effect of exposure on women's earnings, Hakobyan and McLaren (2018) estimate that wage growth of exposed women in the U.S. was more reduced by the NAFTA liberalizations than that of exposed men; at the same time, they find this gender gap to be stronger for married than for single workers, not the reverse. Hakobyan and McLaren (2018) explain their finding by selective non-participation whereby higher-paid married female workers drop out of the labor force.

[56] Family activity here is defined as marriage, birth, or parental leave up-take.

[57] See Tables H18 in the Appendix.

our theory is true, the impact should be stronger the closer the women are to the end of their fertile periods. We now turn to testing this hypothesis.

## 6.2 Mechanisms for Adjustment towards Family

In this section we exploit heterogeneity in the family responses of female workers in order to shed additional light on the underlying mechanisms. We also put the biological clock argument into the context of income, substitution, and other effects that have been shown to influence fertility responses to negative labor shocks.

Because the fertile-age window differs across gender, and heavy investment and training to learn new skills for the new job, industry, or occupation might be incompatible with caring for a young child in which women are disproportionately involved, female workers respond differently to a negative labor shock than male workers. Moreover, the biological clock matters more for women who are closer to the end of their fertile period than for women who are far away from (or past) it. Thus, while generally fertility declines with age, if indeed the fixed fertile-age time window causes a gender difference in the response to the shock one should expect that the impact of the shock is getting stronger as women get closer to the end of their fertile period.

In the following we investigate the role of age for trade-induced fertility and parental leave responses of female workers. Table 10 shows these results for both birth and parental leave decisions (extensive margin) and also for the number of births and parental leave days (intensive margin). The age interaction coefficients for both fertility and parental leave outcomes are positive, indicating that the trade-induced family response becomes stronger as the biological clock is running out. This evidence supports our argument.

Table 10: **Fertility Response and Women's Age**

| | (1) Birth | (2) Parental Leave | (3) Log Number of Births | (4) Log Number Par'Leave Days |
|---|---|---|---|---|
| $\Delta IMP_i$ | -1.399** | -1.665** | -2.069*** | -11.380*** |
| | (0.563) | (0.674) | (0.773) | (4.268) |
| $\Delta IMP_i$ x Age | 0.044** | 0.050** | 0.067*** | 0.345** |
| | (0.018) | (0.021) | (0.025) | (0.136) |
| Observations | 398,530 | 398,530 | 398,530 | 398,530 |

**Notes:** Dependent variable is given at top of column. One is added before taking logs in columns (3) and (4). Sample is fertile-age women (ages 18 to 39 in year 1999). Estimation by two-stage least squares; second-stage coefficients shown. All regressions include the set of worker, firm and partner characteristics of section 5, Table 7. All regressions include two-digit occupation and two-digit industry fixed effects. Age, as part of the full-set of worker characteristics, is included in each regression. Robust standard errors clustered at the industry level in parentheses (744 clusters). *, ** and *** indicates significance at the 10 %, 5% and 1% levels respectively.

To better understand the age pattern of women's responses, we compare the fertility responses to rising import competition for alternative samples of textile workers that differ in their age. Figure 5 shows the female interaction point estimate from estimating equation (2) for each of these samples. The average age of the workers in 1999 is given on the horizontal axis. The first point on the left, for example, is for the sample of workers who are between 16 and 37 years old in 1999 with the average sample age of 30.1, while the right-most point is for the sample of 30 to 51 year old workers (as of 1999).[58]

Figure 5 shows that for workers who are typically around thirty years old as of 1999, there is hardly any gender difference in the fertility response to rising import competition. However, for workers that are typically around 36-37 years as of 1999, women have birth rates that are almost 2.5 percentage points higher than men. Finally, for workers that are above 40 years at the beginning of the sample period, the gender difference in the fertility response is again almost zero. The inverse-U shaped gender difference in Figure 5 provides additional support for our biological clock

[58] Throughout we employ samples with 22 cohorts each to guarantee a sample size sufficient to obtain robust results (N is above 9,000 in all samples). Note that this also removes the fertile-age sample restriction (aged 18 to 39 in year 1999).

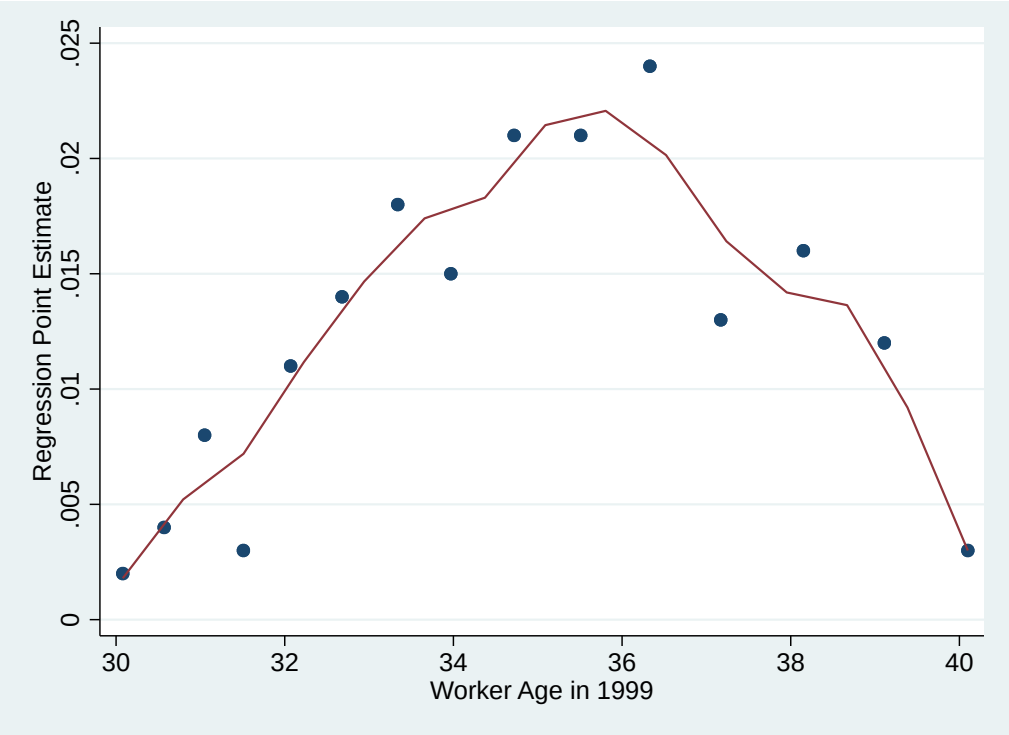


Figure 5: **Fertility Response and Age**

**Notes**: Figure shows the Female interaction coefficient from estimating equation (2) when the dependent is child birth for samples of workers that vary in their age as of 1999. Mean worker age as of 1999 is shown on the horizontal axis. Each sample of workers consists of 22 cohorts; for example, all workers who are between 25 and 46 years old in the year 1999 (mean age of 35.5), with N = 12,730.

argument. The largest gender difference is estimated for workers who are about 39 years old with the onset of rising import competition from China (in the calendar year 2002). By then, the typical female worker will have only a few of years on her biological clock left.[59]

As a consequence, if this worker experiences a trade shock or another labor market shock that induces substantial displacement, as long as the female worker plans to ever have a(another) baby she will have a strong incentive to have this baby now. The negative labor shock reduces workers' opportunity costs, providing an incentive to substitute from labor market to family activities (positive fertility response). This substitution effect, however, applies to both male and female workers, although possibly to a different extent due to household specialization. The gender difference in the workers' biological clocks is a key reason why the fertility response of women is stronger than that of men and has the pattern shown in Table 10 and Figure 5.

Of course, while the biological clock induces women to respond to a trade shock by having a

[59]Less than one in thousand US women gives birth at an age between 45 and 54, compared to 52 for the range of 35 to 39 years (Child Trends 2019).

baby, this is not the only factor affecting their behavior. Lower labor market opportunities tend to translate into lower income, and as long as children are normal goods this income effect means a lower demand for children. While earlier research has often found negative income effects, in the presence of extensive insurance and transfer policies, as in Denmark, one might expect that this static income effect is comparatively muted. The fact that we do not find import competition to reduce (insurance- and transfer-inclusive) personal income of our workers is consistent with that (see Tables 2 and 8).

While the biological clock is common to all women (vs. men), in making their fertility decisions, female workers trade off a shift towards family activities with a successful labor market adjustment–which differs across workers. To shed more light on the precise mechanisms, we examine how women's fertility decisions vary by demographic, education, occupation, and earnings profiles. Table 11 presents these results for the economy-wide sample of fertile-age female workers. Each column presents results from a separate two-stage least squares regression. We show results for the interaction of import competition with a 1999 worker characteristic in addition to the linear effects of import competition. The specification includes the 1999 worker-, firm-, and partner- characteristics described in section 5 and in Appendix C.

First, notice that the fertility response is particularly large if a woman already has one or more children as of the year 1999 (column (1)). A stronger fertility response for women with children is plausible for a number of reasons. To the extent that there are scale economies in child-rearing, the increase in costs of the first child is higher than for subsequent children. In addition, if a young child lowers the worker's ability to signal high commitment for investing into training and skill acquisition to future employers, it will be easier for women who already have a child to add another one than for a woman to have her first child.[60] We also examine whether the fertility response varies with the worker's stage of life, measured by whether the women has a partner (married or co-habitating), finding no significant influence (column (2)).[61]

The remaining specifications in Table 11 consider the roles of education, earnings, and occupation for heterogeneity in women's fertility response to the negative import competition shock. We find that college-educated women have a positive fertility response to trade shock, see column (3).[62] Statistics Denmark's IDA database also provides a hierarchical ranking of jobs within a firm; using

[60] Del Bono, Weber, and Winter-Ebmer (2012) refer to this as the impact of previous children on employability. In line with our results they find that having children is associated with a higher fertility response.

[61] Recall that using our quasi-experimental approach we find particularly strong responses for single workers. As discussed above this may be related to the geographic concentration in the textile industry.

[62] This is consistent with our age result in Table 10 because education is a major reason to postpone fertility.

this information, we identify top-level jobs and find that women in such jobs have a strong positive birth response to the trade shock (column 4, Table 11).[63] A female worker who has to consider switching into a new position at 39 years of age knows that finding and succeeding in a new position will require considerable investments. The worker also knows that more is expected from a professional, for example, than from an office clerk. If she has not yet fulfilled her fertility goals, the worker might come to the conclusion that she cannot achieve both career and fertility goals in the remaining time of her fecundity. Furthermore, because high-powered women would have to invest most heavily to succeed in their new job in the new sector, it is these women who will contribute most strongly to the positive fertility response after an adverse labor shock.[64] To be sure, there are other factors pushing towards a lower fertility response of educated women, such as that highly-educated workers tend to have more transferable human capital and a better ability to re-accumulate skills. In this respect our result of a positive fertility response indicates that the time window drawn by a woman's biological clock can be quantitatively important. The trade-off between committing to be successful in a new top-level job and devoting time for child-rearing and child-bearing before the end of their biological clock induces these women to give birth in response to trade shock.[65]

To better pin down the role of future career and time commitment to learn the new job, firm, industry, and sector-specific skills in the face of biological clock, we focus on the impact across a specific set of occupations. Confirming our findings for college educated women, and women holding top hierarchical positions before the shock, we find that women in professional occupations such as industrial engineers, or finance professionals have a positive fertility response effect due to trade (column (5)). On the other hand, the shock does not induce women in less demanding occupations, such as office clerks, to give birth (column (6)). The impact of the shock on women in service occupations is even negative (column (7)). When we focus on machine operators, a typical manufacturing occupation which constitutes the majority of textile workers (see Figure

[63]The finding of a positive fertility response to adverse labor shocks for highly-skilled, highly-paid women is the opposite of the female fertility responses to job displacement found in recent work (Del Bono, Weber, and Winter-Ebmer 2012, Huttunen and Kellokumpu 2016). These studies explain their finding that highly-skilled women reduce their fertility upon job loss by difficulties of re-establishing career paths when having a young child, in particular that a woman with a young child might not be able to keep up with the intensive training that occurs in the first few years on a new job. Our analysis shows how this interacts with a woman's biological clock.

[64]See Table C9 for evidence that trade increases manufacturing workers' likelihood of moving to the service sector.

[65]Further support is shown in Table C9 in the Appendix. Import competition generates a strong push to the service sector for both men and women (Panel D, column (2)) and also induces a significant occupational change for fertile age women (Panel C, column (4)). In these new occupations, likely less career-oriented as women face career interruptions due to child birth, women earn much less (compare columns (2) and (4) in Panel C). This contributes to our finding of an overall worse labor market adjustment of fertile age women, see column (1), Panel B.

A3 in Online Appendix), we find a significant positive fertility response. While this occupation is neither highly skilled nor highly paid, as a classic manufacturing occupation, these jobs are vanishing in Denmark and in other high-income countries in response to low-wage competition. Workers who hold these jobs move strongly to other industries in the service sector, and they are in need to build up a substantial amount of new human capital as they move to the service sector (Utar 2018; Keller and Utar 2016). Our finding of a positive fertility response points to significant industry-switching investments that these workers would have to incur. Turning to earnings, we find that birth effects tend to be increasing with earnings (column (9)). In line with our findings by occupation that do not vary one for one with pay, however, our results do not show a significant impact for the top earnings individuals. These results reveal that the trade-induced birth effect is stronger on women who face a costly adjustment to re-establish their careers either in the form of a re-investment in their skill-sets or otherwise of a strong commitment for proving themselves in a new job.

We have noted above that while rising import competition is a sizable negative shock to labor earnings (Table 9, column (1), Panel B), given the insurance and transfer policies of Denmark the lower labor earnings do not translate into significantly lower personal incomes of our workers (Tables 2 and 8). If the personal income of women does not fall, they might be able to afford a positive fertility response despite the fact that their labor earnings are substantially reduced.

In addition to the role of insurance and government transfers, the fertility response of workers might depend on the labor earnings of their spouses. In particular, exposed women who are married to a high-earning husband can likely afford a child better than women who are married to a low-earning husband. In a sense, spousal income might provide income security that influences fertility decisions beyond the role of insurance and transfer payments. Indeed, we find that only women who are (in 1999) married to a man who has substantially higher earnings respond to trade exposure by giving birth and take up parental leave (see Table X4). By showing that resources at the level of the couple matter even when a country has broad insurance and transfer policies, we confirm that income effects– tilting worker response to lower fertility–are operating as well.

The following section examines the role of methodological aspects for our results.

Table 11: **Heterogeneous Fertility Effects among Women**

| | (1) | (2) | (3) Education | (4) Within Firm Job Ranking | (5) | (6) | (7) Occupations | (8) | (9) Wage |
|---|---|---|---|---|---|---|---|---|---|
| $\Delta IMP_i$ | -0.309 | 0.063 | -0.067 | -0.031 | -0.031 | 0.006 | 0.009 | -0.097 | -0.093 |
| | (0.215) | (0.146) | (0.118) | (0.117) | (0.117) | (0.122) | (0.119) | (0.136) | (0.154) |
| $\Delta IMP_i$ x Have a child | 0.511* | | | | | | | | |
| | (0.297) | | | | | | | | |
| $\Delta IMP_i$ x Have a partner | | -0.135 | | | | | | | |
| | | (0.146) | | | | | | | |
| $\Delta IMP_i$ x College | | | 0.356*** | | | | | | |
| | | | (0.129) | | | | | | |
| $\Delta IMP_i$ x Top Ranked Positions | | | | 0.782** | | | | | |
| | | | | (0.308) | | | | | |
| $\Delta IMP_i$ x Professionals | | | | | 0.913** | | | | |
| | | | | | (0.461) | | | | |
| $\Delta IMP_i$ x Office Clerks | | | | | | -0.055 | | | |
| | | | | | | (0.115) | | | |
| $\Delta IMP_i$ x Service Occupations | | | | | | | -0.841** | | |
| | | | | | | | (0.364) | | |
| $\Delta IMP_i$ x Machine Operator | | | | | | | | 0.424* | |
| | | | | | | | | (0.252) | |
| $\Delta IMP_i$ x Earnings 4th Quartile | | | | | | | | | 0.231 |
| | | | | | | | | | (0.207) |
| $\Delta IMP_i$ x Earnings 3rd Quartile | | | | | | | | | 0.191 |
| | | | | | | | | | (0.151) |
| $\Delta IMP_i$ x Earnings 2nd Quartile | | | | | | | | | 0.039 |
| | | | | | | | | | (0.193) |
| Observations | 398,530 | 398,530 | 398,530 | 398,530 | 398,530 | 398,530 | 398,530 | 398,530 | 398,530 |
| Hansen J statistic | 1.248 | 1.184 | 0.807 | 0.772 | 0.734 | 0.932 | 1.054 | 0.609 | 0.789 |
| Hansen J Pval | 0.536 | 0.553 | 0.668 | 0.68 | 0.693 | 0.627 | 0.59 | 0.737 | 0.674 |

**Notes:** Dependent variable is Birth. Sample is fertile-age women. Estimation by two-stage least squares; second-stage coefficients shown. All regressions include the set of worker, firm and spouse characteristics employed in section 5. All regressions include two-digit occupation and two-digit industry fixed effects. Linear terms of each interaction variable is included in the respective regression if not already among the set of controls. Robust standard errors clustered at the industry level (744) in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

## 6.3 Comparing the Quota Liberalization and Job Displacement Approaches

Another reason why we find relatively strong evidence for positive family responses after a negative labor shock might lie in our empirical approach. In order to see whether a plausibly exogenous shock that leads to lay-offs and plant closures influences our results, we implement an identification strategy closer to that in the job displacement (JD) literature to compare it with our quota liberalization approach (MFA for short).

In particular, we re-define a textile worker as treated if he or she has a period of unemployment anytime during the years 2002 to 2005. The year 2002 is the time when the uncertainty regarding China's WTO accession fully resolved and the first year that textile workers were exposed to import competition from China due to the lifting of quotas for China, while 2005 is the year of the fourth and final phase of quota liberalizations. We emphasize that the following analysis of JD is limited in a number of ways, and a comprehensive analysis is left to future work. This includes the fact that job displacement is not always associated with unemployment (see the analysis of Del Bono, Weber, and Winter-Ebmer 2015). At the same time, we believe that our analysis captures the most salient forces.

Table 12 reports two sets of results, corresponding to the two approaches, referred to as JD and MFA. For both sets of results we employ the same samples, and estimate equation (2). The only difference between the JD and the MFA results is the definition of worker treatment. Table 12 shows results for four family responses (birth, parental leave, marriage, and divorce), as well as on cumulative earnings. The MFA results presented in Table 12 are the same as in earlier tables.[66]

Beginning with the family responses of workers (columns 1 to 8), we see that irrespective of the definition of treatment there is evidence for a gender difference. For example, the female interaction coefficients for JD and MFA are 0.130 and 0.128 in the case of births (columns (1) and (2), respectively). The gender difference is not always estimated to be so small, and the gender difference with the JD approach is not always significant. At the same time, given that the female interaction point estimate is always positive there is evidence that women move more strongly towards family than men using either approach.[67]

Moving to the level (and sign) of the worker response, there is a clear difference between the JD and MFA approaches. In particular, for all family responses, we estimate a negative impact

[66] Birth results are from Table 3, Parental Leave from Table 4, Marriage from Table 5, Divorce from Table 6, and Earnings from Table 2.

[67] The interaction coefficient is negative for divorce because here the shift towards family means a negative value.

of exposure for men using the JD approach (positive for divorce), and in three out of four cases significantly so (columns (1), (3), and (5)). Furthermore, we find that women's family responses tend to be negative according to the JD approach.[68]

In sum, while we find broadly the same gender difference with either the JD or the MFA approach, family responses are strikingly more negative according to the JD approach than if we employ the the MFA approach. In particular, the family response of female workers tends to be negative using the JD approach, whereas it is positive using the MFA approach, and often significantly so.

The main difference in the response levels between the two approaches appears to be that unemployed workers (both female and male) exhibit a negative family response, a factor that is relatively strongly reflected in the JD results. Consistent with this, we see that the JD approach leads to a much more negative earnings impact estimate than the MFA approach, see columns (9) and (10) of Table 12. Moreover, there is no significant gender difference in the earnings impact according to the JD approach. It appears that once a worker is unemployed the current and future earnings outlook is so bleak that the biological clock mechanism or other reasons for a positive fertility response carry little weight.

This finding is consistent with the previous section, which showed that with MFA identification income effects continue to play a role for workers' responses. At the same time, it appears that by conditioning on unemployment one obtains a considerably larger negative income effects with the JD approach than with our MFA approach.

We emphasize that one should not draw overly strong inferences from this back-of-the-envelope comparison.[69] A limitation is that in our JD approach we define a worker to be treated if he or she exhibits any unemployment spell during the years 2002 to 2005, whereas other JD analyses often focus on plant closure or mass layoff.[70] At the same time, mass layoff due to plant closure is also an important reason for unemployment among the textile workers exposed to rising import competition (Utar 2014), and we believe that the results reported in Table 12 capture key factors.[71]

[68] Marriage is the exception; the female marriage response is essentially zero with the JD approach.

[69] For example, it is possible that rising import competition triggers a stronger sectoral shift from manufacturing to services than shocks analyzed in the job displacement literature. If so, the loss of sector-specific human capital might be greater with a trade shock, and correspondingly it might provide relatively strong incentives to move to family.

[70] Data on plant closures is not available to us at this point.

[71] Furthermore, we have experimented with alternative versions of the JD approach, such as relying on unemployment in either 2002 or 2003 instead of the longer period of 2002-2005, finding similar results.

Table 12: **Comparison with Job Displacement Approach**

| | (1) Birth | (2) Birth | (3) ParL | (4) ParL | (5) Marriage | (6) Marriage | (7) Divorce | (8) Divorce | (9) Earn'gs | (10) Earn'gs |
|---|---|---|---|---|---|---|---|---|---|---|
| | JD | MFA | JD | MFA | JD | MFA | JD | MFA | JD | MFA |
| ImpComp | -0.219*** | -0.019 | -0.106*** | -0.028 | -0.113*** | -0.012 | 0.071 | -0.024 | -3.490*** | 0.051 |
| | (0.043) | (0.037) | (0.034) | (0.030) | (0.041) | (0.034) | (0.071) | (0.026) | (0.341) | (0.425) |
| ImpComp x Female | 0.130* | 0.128* | 0.072 | 0.122** | 0.119** | 0.092** | -0.022 | -0.062* | -0.448 | -1.274** |
| | (0.070) | (0.055) | (0.059) | (0.048) | (0.057) | (0.046) | (0.076) | (0.032) | (0.455) | (0.515) |
| Observations | 3,160 | 3,160 | 3,160 | 3,160 | 5,784 | 5,784 | 4,634 | 4,634 | 19,526 | 19,526 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |

**Notes:** Dependent variable is given on top of column. Sample is textile workers; single for birth and parental leave; unmarried fertile-age for marriage, and married for divorce. Estimation of equation 2. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

# 7 Concluding Remarks

Using population register data on all marriages, divorces and births together with employer-employee matched data from Denmark, we show that a labor market shock in the form of intensified import competition with China has a significant impact on gender inequality through its effect on the family-market work balance. Generally, single workers exposed to import competition have a higher likelihood to marry, have children, and take parental leave, while married workers do not divorce their spouses as often as similar non-exposed workers. Strikingly, even though the negative earnings impact of the shock is similar for men and women, the shift to family is largely accounted for by women, not men, and correspondingly the negative long-run earnings impact of import competition on women is much higher than for men. These results carry over to the Danish economy at large.

We show that the gender differential in the workers' family-market work adjustment is neither because women are employed in jobs at particular firms, industries, or occupations, nor because women suffer more strongly from concurrent shocks such as technology. Instead, the strong family response with its corresponding gender earnings inequality is due to a woman's biological clock. Because women are unlikely to conceive beyond their early forties, fertile-age women have a higher reservation value to stay in the labor market than men, and a negative labor demand shock

due to trade exposure induces a woman to move towards family more than it induces a man. Combined with the fact that having and caring for a young child is female-intensive, making investments to succeed in a new career before the biological clock runs out is difficult, and may induce women to focus on family instead of 'having it all'. Support for this comes from finding that there is no gender differential for workers past their fertile age, and that the move towards family is particularly strong among women who would have to make the highest investments into new careers and are closest to the end of their fertile periods. The shock on labor market opportunities through trade is but one labor market shock where the biological clock differences between women and men may have important implications for gender inequality, household specialization, and family structure.

Our paper shows how otherwise gender-neutral labor market shocks can generate strongly gender-specific results by interacting with a woman's biological clock. Our findings demonstrate that the gender-gap in the labor market can arise even when education, occupation, or preference differences between men and women, as well as discriminatory practices and policies in labor markets, are limited.

This paper also provides evidence that globalization can have a strong impact on earnings inequality because women and men do not substitute family work for market work in the same way due to a higher reservation value to stay in the labor market induced by their biology. In the absence of effective insurance and government transfer systems, the negative effect on earnings may outweigh any substitution from labor market to family, while in a setting with substantial family-oriented support systems as in Denmark, the shift of women towards family activities–with the corresponding implications for gender inequality– is stronger. There is clearly a need for more economic research on the importance of the market-family margin.

# Additional Figures and Tables

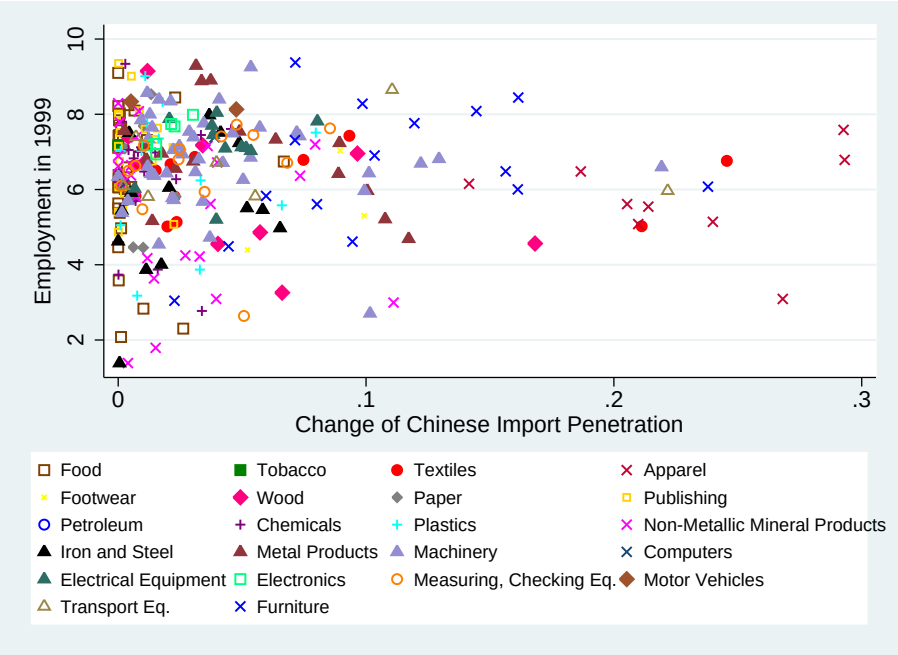


Figure X1: **Change in Chinese Import Penetration**

**Notes**: Figure shows the change in Chinese import penetration across six-digit industries. Each two-digit industry is given the same color and symbol. The vertical axis shows the logarithm of the total number of employees in 1999.

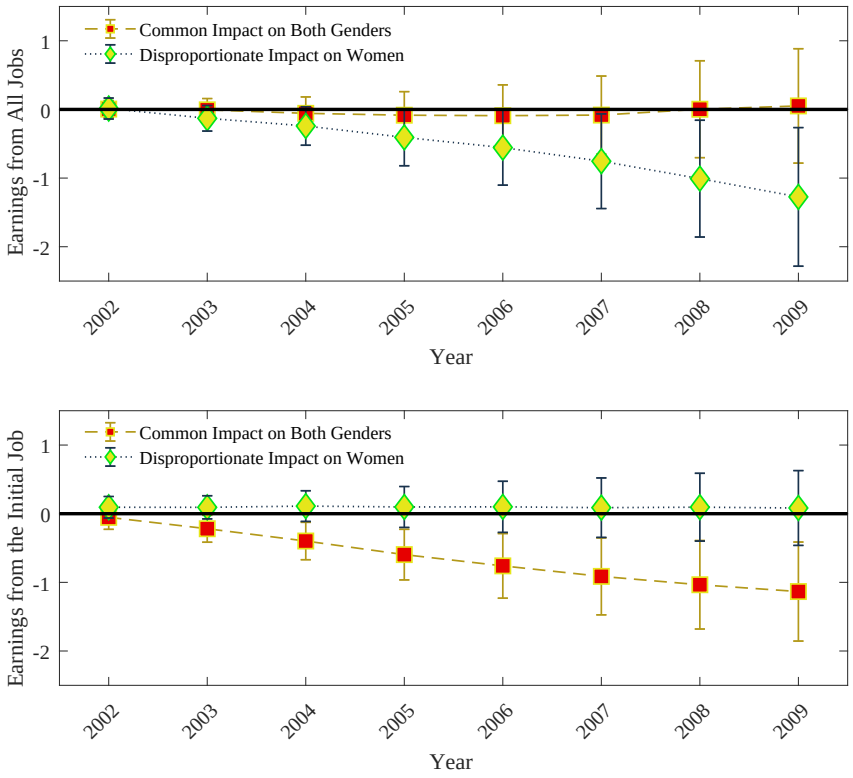


Figure X2: **Gender Difference in Trade Adjustment Costs**

**Notes**: Figure draws the coefficient estimates obtained from estimating equation (2) on a rolling end-year basis. Dependent variable in top panel is cumulative earnings from all jobs, in bottom panel cumulative earnings in 1999 job (bottom). Squares indicate the coefficient estimates of ImpComp (Post x Exposure), circles indicate the coefficient estimates of ImpComp x Female (Post x Exposure x Female). Shown is the 95% confidence interval.

Table X1: **Characteristics By Gender, Family, and Treatment Status**

| | Treated Mean | Not Treated Mean | Diff | t-stat |
|---|---|---|---|---|
| **Panel A. Women** | N = 3,069 | N = 2,524 | | |
| Age | 39.28 | 39.21 | 0.07 | 0.26 |
| Hourly Wage | 134.86 | 134.28 | 0.58 | 0.49 |
| **Panel B. Married Women** | N = 1,889 | N = 1,534 | | |
| Age | 42.18 | 41.90 | 0.28 | 0.93 |
| Hourly Wage | 136.02 | 135.14 | 0.88 | 0.57 |
| Partner's Log Income | 12.50 | 12.47 | 0.04 | 2.22 |
| **Panel C. Unmarried Women** | N = 1,180 | N = 990 | | |
| Age | 34.64 | 35.05 | -0.41 | -0.94 |
| Hourly Wage | 133.02 | 132.95 | 0.06 | 0.04 |
| Partner's Log Income | 12.41 | 12.39 | 0.01 | 0.41 |
| **Panel D. Men** | N = 1,674 | N = 2,731 | | |
| Age | 39.07 | 39.24 | -0.18 | -0.57 |
| Hourly Wage | 189.64 | 181.71 | 7.93 | 2.50 |
| **Panel E. Married Men** | N = 974 | N = 1,493 | | |
| Age | 43.01 | 43.16 | -0.15 | -0.43 |
| Hourly Wage | 206.98 | 193.68 | 13.31 | 2.86 |
| Partner's Log Income | 12.14 | 12.15 | -0.01 | -0.40 |
| **Panel F. Unmarried Men** | N = 700 | N = 1,238 | | |
| Age | 33.58 | 34.52 | -0.94 | -2.11 |
| Hourly Wage | 165.51 | 167.28 | -1.77 | -0.46 |
| Partner's Log Income | 12.06 | 12.12 | -0.06 | -1.95 |

**Notes**: Table shows averages of 1999 worker characteristics. See the text for definition of treatment. Partner characteristics in the case of unmarried workers are for co-habitant.

Table X2: **Worker Characteristics in Economy-wide Sample**

| | | | **Men** | | **Women** | |
|---|---|---|---|---|---|---|
| | **Mean** | **SD** | **Mean** | **SD** | **Mean** | **SD** |
| | **N=1,651,774** | | **N=915,702** | | **N=736,072** | |
| Age | 38.194 | 9.868 | 38.039 | 10.051 | 38.387 | 9.632 |
| Immigrant | 0.041 | 0.199 | 0.044 | 0.206 | 0.038 | 0.190 |
| Labor Market Experience | 14.440 | 5.807 | 14.498 | 5.898 | 14.368 | 5.691 |
| Married | 0.543 | 0.498 | 0.521 | 0.500 | 0.570 | 0.495 |
| Number of Children | 1.337 | 1.152 | 1.264 | 1.179 | 1.428 | 1.112 |
| Log Earnings | 12.255 | 0.669 | 12.362 | 0.665 | 12.128 | 0.626 |
| College | 0.280 | 0.449 | 0.246 | 0.431 | 0.322 | 0.467 |
| Vocational Educated | 0.411 | 0.492 | 0.437 | 0.496 | 0.380 | 0.485 |
| Birth Event | 0.046 | 0.208 | 0.048 | 0.215 | 0.042 | 0.201 |
| Divorce Event | 0.008 | 0.088 | 0.007 | 0.085 | 0.008 | 0.092 |
| Marriage Event | 0.026 | 0.159 | 0.026 | 0.158 | 0.026 | 0.160 |
| Managers | 0.037 | 0.188 | 0.051 | 0.219 | 0.020 | 0.139 |
| Professionals | 0.143 | 0.350 | 0.147 | 0.354 | 0.138 | 0.345 |
| Office Workers | 0.129 | 0.335 | 0.059 | 0.235 | 0.216 | 0.412 |
| Machine Operators | 0.056 | 0.230 | 0.061 | 0.240 | 0.049 | 0.217 |

**Notes**: Summary statistics for all full-time, private sector workers as of the year 1999. Immigrant is indicator variable for first and second generation immigrants. Occupation categories follow one-digit ISCO classification except Machine Operators (ISCO = 82). Birth, Divorce, and Marriage outcome variables are indicators on whether a worker fathers or mothers a newborn baby, divorces, or marries in a given year, respectively.

Table X3: **Occupation-specific Time Trends**

| | (1) Earnings | (2) Earnings 1999 Job | (3) Birth Event | (4) Parental Leave | (5) Marriage | (6) Divorce |
|---|---|---|---|---|---|---|
| Imp Comp | -0.180 | -1.213*** | -0.021 | -0.028 | -0.031 | -0.020 |
| | (0.391) | (0.366) | (0.039) | (0.031) | (0.028) | (0.014) |
| ImpComp x Female | -0.874* | 0.144 | 0.127** | 0.123** | 0.096*** | -0.020 |
| | (0.479) | (0.274) | (0.057) | (0.049) | (0.035) | (0.017) |
| Female x Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Occupation x Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Worker FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| N | 19,526 | 19,526 | 3,160 | 3,160 | 8,216 | 11,780 |

**Notes**: Dependent variables are listed at top of column; Earnings is the cumulative earnings over 1999-2001 and 2002-2009 periods; Earnings 1999 Job are cumulative earnings obtained from the initial, 1999, firm over 1999-2001 and 2002-2009 period and analogously for all dependent variables. Least squares estimation of equation (2) augmented with occupation-specific fixed effects, for managers, professionals and technicians, office workers, production operators, craft workers and labourers, respectively. Estimation samples in columns 3 and 4 are single fertile-age workers, in column 5 unmarried workers and in column 6 married workers, all as of year 1999. Robust standard errors clustered at the 1999 firm-level. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table X4: **Family Responses of Married Workers and Spouse Earnings**

| Gender | (1) **Women** Birth | (2) **Men** Birth | (3) **Women** Parental Leave | (4) **Men** Parental Leave |
|---|---|---|---|---|
| $\Delta IMP_i$ | -0.037 | -0.322* | -0.209 | -0.343 |
| | (0.186) | (0.169) | (0.194) | (0.227) |
| $\Delta IMP_i$ x High Spouse Earnings | 0.247* | -0.247 | 0.323** | -0.286 |
| | (0.144) | (0.287) | (0.139) | (0.275) |
| Observations | 178,528 | 177,352 | 178,528 | 177,352 |
| Two-digit Industry FE | ✓ | ✓ | ✓ | ✓ |
| Two-digit Occupation FE | ✓ | ✓ | ✓ | ✓ |

**Notes:** Dependent variable is given at top of column. Sample is all married fertile-age workers as of 1999. High spouse earnings is defined as at least 130 percent of the worker's own salary in 1999. All regressions include the set of control variables $Z_i^W$, $Z_i^F$, and $Z_i^P$ described in section 5. Estimation by two-stage least squares; second-stage coefficients shown. Robust standard errors clustered at the industry level in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

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# Appendix:

# "Globalization, Gender and the Family"

Wolfgang Keller and Hale Utar

May 3, 2020

# Contents (Appendix)

# A Trade Liberalization in Textiles and Clothing

## A.1 The Multi Fibre Arrangement System

When the General Agreement on Tariffs and Trade (GATT) was signed in 1948, world trade in textile and clothing was excluded from the agreement. Instead, trade in textiles and clothing was governed by bilateral agreements. As the number of agreements grew, the Multi-fibre Arrangement (MFA) was introduced in 1974 to govern the world trade in textile and clothing. The European Union negotiated most (MFA) quotas for the bloc of countries as a whole, and since 1993 any member state-specific restrictions were removed and the quotas started to be managed at the EU level. In 1995 the Agreement on Textiles and Clothing (ATC) replaced the MFA and made provisions for phasing it out in four steps over a period of 10 years. The liberalizations occurred at the beginning of the years 1995, 1998, 2002 and 2005. Based on the volume of imports in 1990, quotas were eliminated on 16% of 1990 imports at the beginning of 1995, an additional 17% at the beginning of 1998, 18% at the beginning of 2002, and the remaining 49% at the beginning of 2005. Due to the surge of Chinese imports in the first few months of 2005, the EU renegotiated the quotas with China, with the result that they agreed on additional quotas on certain products until 2008 (the so-called "Bra War"). We exclude these categories from our treatment classification.

Between 1986 and 1994 the EU executed MFA quotas towards 19 countries. These were Argentina, Brazil, China, Czechoslovakia, Hong Kong, Hungary, India, Indonesia, the Republic of Korea, Macao, Malaysia, Pakistan, Peru, Philippines, Poland, Romania, Singapore, Sri Lanka and Thailand. Under the later ATC system, the selection of MFA products to be integrated into the normal WTO system was left to the decision of the importing country. The EU started its phasing-out process by integrating mainly products on MFA categories with no quotas towards WTO members. The same approach was chosen by the USA. During the first two phases, the EU integrated 34 MFA categories, but only very few existing quotas with respect to WTO members.

During the same time the EU also liberalized quotas on a bilateral basis for neighboring countries in Eastern Europe (Europe Agreements) and the Mediterranean area. Among the list of 19 countries above, the Czech Republic, Slovakia, Hungary, Poland, and Romania already had established quota free access to the European market before 1999. In 1997 about 70% of the total EU import value of textiles and clothing was imported without any quantitative restrictions, while the other 30% was imported under quota. Among the 81 categories for which EU quotas existed, only 18 were utilized at an average of more than 70% between 1996 and 1998. The exporting countries that

had the highest quota utilization rates were China, India, Pakistan and Indonesia. In 1998, China's share of textiles and apparel imports of Denmark was a little over 10% compared to 2.8%, 0.7% and 1.3% respectively for India, Pakistan and Indonesia. By 2007 China's share reached 26%, while the respective shares of India, Pakistan and Indonesia were 6%, 1%, and 0.5% ).

## A.2 Textile Quotas

The *Systeme Integre de Gestion de Licenses* (SIGL) database provides categories of textile and clothing products that are subject to trade quotas in the European Union for a particular year. We employ this data to identify firms in Denmark that are affected by the quota removals on Chinese exports following that country's entry into the WTO. The quota categories are administrative descriptions of quota products that do not follow standard statistical product classifications. For one, the quotas have a varying degree of coverage; for example, the quota category *Gloves, mittens and mitts, knitted or crocheted* covers nine products at the 8-digit Common Nomenclature (CN) level, while the category *Woven fabrics of synthetic filament yarn obtained from strip or the like of polyethylene or polypropylene, less than 3 m wide* corresponds to a single 8-digit CN product. Quota categories include both textile and clothing products. Also, a given category does not necessarily cover a technologically or materially homogeneous group of products, nor does it have to be comprehensive. For example, *ramie bedspreads* are covered by the quota restriction for China while *cotton bedspreads* are not, and *Brasseries of all types of textile material* is covered, in contrast to *Corselettes of all types of textile materials*. The source of the correspondence between quota categories and eight-digit products is Utar (2014), and it is available from the author.[72]

## A.3 Importance of China's Entry into the WTO

How important was China's entry into the WTO in 2002 compared with earlier and later liberalization episodes (1995 and 1998, versus 2005, respectively)? The European Union kept a relatively open trade policy in the textile and clothing sector throughout the 1990s with respect to other developing countries except for some 'sensitive MFA quota categories' which were mostly the subject of the 2005 (Phase IV) quota abolishment. For example, developing countries subject to the MFA quotas, such as India, Indonesia, Pakistan, Thailand, did not experience any quota removal as part

[72]If the quota product manufactured by one firm would be a close substitute to the non-quota product manufactured by another firm, this would mean that some workers we classify as not exposed are in fact exposed; if this contamination would be quantitatively important our analysis would underestimate the causal effect of the shock.

of Phase II (1998). For Indonesia all active quotas imposed were subject to Phase IV abolishment except 2 quotas (categories 21 and 33) which were subject to Phase III and were removed in 2002. Similarly, for India no quotas were in place that were subject to Phase I and II removal. There were only 2 quota categories that were subject to the Phase III (categories 24 and 27) and they were removed in 2002. The remaining 15 categories were removed in 2005 (source: SIGL). The quotas imposed to these countries were mostly subject to Phase IV removal and were removed in 2005. In contrast, China was strongly restricted by the MFA quotas before it entered the WTO.

The EU has no textile quotas for the least developed countries. For example, Bangladesh was benefiting from the General System of Preferences, and no textile quotas were imposed on Bangladesh throughout the sample period.

Argentina, Brazil, Macao and Pakistan had 1 category, Hong Kong had 4 and South Korea had 6 categories removed in Phase I and II. The highest utilization rate among these quotas removed under the Phase I or II was 49.6 % for category 100 from Korea. This category was not subject to quota for any other country. Giving the overall share imports from these countries and the differences of quota categories imposed across these countries, it is difficult to disentangle the impact of Phase I and II removal from the general liberalization in the textiles and clothing industry.

## A.4 The Timing of the Trade Shock

There are two natural possibilities for dating the timing of the trade liberalization, the end of the MFA system (last step in 2005) and China's entry into the WTO (in 2002). While the two events are related to each other, our empirical strategy exploits the expiration of the MFA quotas for China due to its WTO membership. A key reason for that is that the abolishment of the MFA quotas were scheduled in 1995 and therefore there was no uncertainty associated with its timing. However, China was not able to benefit from these quota removals as long as the country was not a member of the WTO. The shock that matters for the difference-in-difference estimation strategy comes from uncertainty regarding China's accession to the WTO, as well as its timing.

Employing the year 2002 as the beginning of the treatment period implies that we do not distinguish between phases I to III (in 2002) and phase IV (in 2005) of the MFA liberalizations. If the liberalizations in phase IV were the most comprehensive, perhaps because advanced countries tended to leave the quotas that were most strongly binding for the latest liberalization phase, our analysis might underestimate the trade liberalization effects. At the same time, Utar (2014) docu-

ments strong overlap between firms producing quota products that are subject to the 2002 and the 2005 removal for China and shows that firms strongly reacted to the removal of quotas already in 2002 by downsizing even if more of their products were subject to the 2005 removal.[73] As a consequence, it is challenging to estimate the effects of the 2002 and 2005 liberalizations separately. This has been shown in Utar (2014); the following reports a number of key results from that paper for convenience.

The variable $Q2002_j$ is an indicator variable that is 1 if firm $j$ produces a quota good as of year 1999 which is subject to the 2002 removal for China. Similarly, $Q2005_j$ is 1 if firm $j$ produces a quota good as of year 1999 which is subject to the 2005 removal for China. Further, $Post02_t$ is an indicator for years 2002 and later, while $Post05_t$ denotes the years 2005 and later. The following equation is estimated at the firm-level for the period 1999-2007:

$$\ln Y_{jt} = \alpha_0 + \alpha_1 Q2002_j x Post02_t + \alpha_2 Q2005_j x Post05_t + \delta_j + \tau_t + \varepsilon_{jt} \quad \text{(A1)}$$

In equation A1 $Y_{jt}$ denotes one of several firm-level outcome variables, $\delta_j$ denotes firm fixed effects and $\tau_t$ denotes year fixed effects. The results are reported in Table A1. We see that the reduction in sales is stronger in response to the 2005 removal yet employment responds more to the 2002 removal (columns (1), (3) and (4)). Column (5) show that employment for less educated workers drops 16% annually in response to the 2002 removal even when one controls for the impact of the 2005 removal. The impact of the 2002 removal on workers with vocational education on textile production (machine operators) is even stronger. The annual reduction is estimated to be 20% (column (6)).

Overall, the results in Table A1 provide evidence for substantial employment effects for both the 2002 and the 2005 quota liberalizations that would be difficult to fully separate with any research design. Therefore we adopt the approach of treating all post-2002 years as treatment period.

[73] Among 191 firms producing products subject to the 2005 removal 97 of them were also producing products subject to the 2002 removal.

Table A1: **The 2002 versus 2005 Quota Removals: Firm-level Evidence**

| | (1) Sales | (2) Value Added | (3) Employment | (4) Full-time Employment | (5) Employees w/ High School Education | (6) Employees w/ Textile Production Education |
|---|---|---|---|---|---|---|
| Q2002xPost02 | -0.075 | -0.081 | -0.123*** | -0.146** | -0.164*** | -0.201*** |
| | (0.064) | (0.061) | (0.059) | (0.057) | (0.053) | (0.046) |
| Q2005xPost05 | -0.158*** | -0.187*** | -0.081 | -0.125** | -0.152*** | -0.049 |
| | (0.059) | (0.067) | (0.054) | (0.059) | (0.046) | (0.037) |
| Firm FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year FEs✓ | ✓ | ✓ | ✓ | ✓ | ✓ | |
| N | 4,555 | 4,536 | 4,503 | 4,545 | 4134 | 4,134 |

Source: Utar (2014). The estimation sample is annual observations for textile and clothing firms over 1999-2007.All dependent variables in logs. Full-time employment is number of full-time equivalent employees. Robust standard errors are clustered at the firm level. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

## A.5 Possible Pre-Trends and Placebo Results

We begin by showing summary statistics on textile worker characteristics by exposure for the year 1995, four years before the beginning of our sample period.

Table A2: **1995 Worker Characteristics by Exposure to Import Competition**

| | Treated N = 4,667 Average | Control N = 5,163 Average | Difference | t-stat |
|---|---|---|---|---|
| Age | 35.37 | 35.43 | -0.06 | -0.29 |
| Labor Market Experience | 11.35 | 10.96 | 0.39 | 3.71 |
| Log Annual Earnings | 11.99 | 12.01 | -0.03 | -1.45 |
| Married | 0.54 | 0.53 | 0.01 | 0.86 |
| Marriage Event | 0.02 | 0.03 | 0 | -1 |
| Divorce Event | 0.01 | 0.01 | 0 | 0.56 |
| Birth Event | 0.05 | 0.05 | 0 | -0.87 |

**Notes**: Shown are averages of the 1995 characteristics of workers treated and not treated (control) to rising import competition from China. Treated workers are those whose firm in 1999 manufactured in Denmark a product protected by a quota that would be removed with China's entry into the WTO, control otherwise. Log earnings is measured in 2000 Danish Kroner.

As Table A2 indicates, the differences are limited.

The following analysis checks for possible pre-trends by following the 1999 textile workers back to the year 1990 for a number of placebo exercises. We apply the same specification as in the text and estimate an equation analogous to equation (1) for the sample period 1990-1999. As the pre-shock period, we employ the period 1990-94. The (placebo) treatment period is assumed to be 1995-99 as this is the period that covers the phase I and II removals for developing countries other than China, as well as the removal of quotas for the Eastern European countries. Table A3 shows labor market and income results, separately for men (Panel A) and women (Panel B). Table A4 reports in addition earnings and income results separately for married and unmarried workers, as well as evidence on three different family outcomes (birth, marriage, and divorce).

Beginning with Table A3, notice that none of the coefficients are significantly different from zero at standard levels, neither for men nor for women. This is what one would expect in the absence of major pre-trends. Turning to the results in Table A4 here we also see that none of the estimated coefficients are significantly different from zero.

We have also interacted the exposure variable with year dummies instead of one variable for the entire period. Using annual worker-level data over 1990-1999, we estimate the following equation:

$$Y_{it} = \alpha_0 + \alpha_1 Exposure_i x Year_t + \iota + \tau + \varepsilon_{it}. \tag{A2}$$

In this equation, $\iota$ and $\tau$ denote worker and year fixed effects, respectively. $Exposure_i$ is defined as before, and we interact this variable with year indicators starting from 1995.[74] Results are shown in Table A5.

The results confirm that there are no major pre-trends for the labor market measures. Further, despite the occasional weakly significant coefficients we conclude that there are no major pre-trends for family variables either. Furthermore, Figure A1 presents the pre-sample analysis for the family variables separately for women and men. We find no evidence on pre-existing differential trends between treatment and control groups. Finally, we also combine our pre-sample information for 1990-99 with the first years of our sample and estimate the above equation. Figure A2 shows the evolution of annual impact of import competition over 1995-2003 for the probability of birth. We see that the fertility response of workers starts with the first year of the shock, but not before. Based on these findings, we can rule out the possibility of major pre-existing trends in our analysis.

[74] This year-by-year difference-in-differences specification is likely to suffer from an over-estimation bias as described by Bertrand, Duflo, and Mullainathan (2004).

Table A3: **Potential Pre-Trends I: 1990-1999**

| | (1) Earnings | (2) Personal Income | (3) Hours Worked | (4) Hourly Wage | (5) Unemployment |
|---|---|---|---|---|---|
| Panel A. Men | | | | | |
| $Exposure_{i,99} * Post95_s$ | 0.009 | 0.019 | -0.009 | 0.017 | -0.085 |
| | (0.033) | (0.028) | (0.014) | (0.020) | (0.107) |
| N | 8,248 | 8,248 | 7,964 | 7,964 | 8,248 |
| Panel B. Women | | | | | |
| $Exposure_{i,99} * Post95_s$ | 0.013 | -0.012 | 0.015 | -0.002 | -0.052 |
| | (0.028) | (0.025) | (0.015) | (0.014) | (0.117) |
| N | 10,374 | 10,374 | 9,850 | 9,850 | 10,374 |

**Notes:** Dependent variables on top of column. All variables are expressed in logs. They are the average annual value of earnings, personal income, hours worked, hourly wage and the unemployment index, respectively. Unemployment index takes the value of one when no unemployment is recorded in a given year, and ranges to 1001 which indicates unemployment for the whole duration of year. E.g., the value 501 indicates a half year of unemployment. Averages are taken across the pre- and post-1995 periods, namely 1990-1994 and 1995-1999. Estimation by least squares. All specifications include worker and time fixed effects. Robust standard errors clustered at the 1999 firm in parentheses.

Table A4: **Potential Pre-Trends–Subsample Analysis**

| | (1) Earnings | (2) Personal Income | (3) Divorce | (4) Marriage | (5) Birth |
|---|---|---|---|---|---|
| Panel A. Men | | | | | |
| $Exposure_{i,99} * Post95_s$ | 0.003 | 0.009 | 0.003 | 0.013 | 0.006 |
| | (0.024) | (0.019) | (0.007) | (0.014) | (0.018) |
| N | 8,550 | 8,542 | 8,550 | 8,550 | 8,550 |
| Panel B. Women | | | | | |
| $Exposure_{i,99} * Post95_s$ | 0.024 | -0.007 | -0.003 | 0.012 | 0.017 |
| | (0.027) | (0.013) | (0.006) | (0.013) | (0.016) |
| N | 10,954 | 10,946 | 10,954 | 10,954 | 10,954 |
| Panel C. Married Workers as of 1999 | | | | | |
| $Exposure_{i,99} * Post95_s$ | -0.014 | 0.020 | 0.003 | 0.029 | 0.005 |
| | (0.032) | (0.025) | (0.007) | (0.023) | (0.027) |
| $Exposure_{i,99} * Post95_s * Woman_i$ | 0.042 | -0.028 | -0.002 | -0.017 | 0.007 |
| | (0.039) | (0.025) | (0.008) | (0.029) | (0.034) |
| N | 11,548 | 11,548 | 11,548 | 11,548 | 11,548 |
| Panel D. Unmarried Workers as of 1999 | | | | | |
| $Exposure_{i,99} * Post95_s$ | 0.042 | 0.010 | 0.006 | -0.011 | 0.014 |
| | (0.032) | (0.021) | (0.013) | (0.009) | (0.020) |
| $Exposure_{i,99} * Post95_s * Woman_i$ | -0.021 | -0.012 | -0.012 | 0.022 | 0.012 |
| | (0.054) | (0.023) | (0.019) | 0.014 | 0.031 |
| N | 7,956 | 7,940 | 7,956 | 7,956 | 7,956 |

**Notes:** Dependent variables at the top of the column. Estimation by least squares. All specifications include worker and time fixed effects and a constant. Regressions in Panels C and D also include $Post95_s * Woman_i$ but this is omitted from the table. Earnings is the average earnings over 1990-1994 and 1995-1999 normalized by the worker's own 1999 earnings. Similarly, Personal Income is the average personal income across the pre- and post-1995 period normalized by the worker's own personal income as of year 1999. Divorce, Marriage, and Birth variables take 1 if the individual has an event of divorce, marriage, or birth (fathering or mothering a new born child) over the periods, 1990-1994 and 1995-1999, and zero otherwise. Robust standard errors clustered at the 1999 firm in parentheses.

Table A5: **Potential Pre-Trends Using Annual Data**

| | (1) Earnings | (2) Income | (3) Hours | (4) Hourly Wage | (5) Unemp-loyment | (6) Divorce | (7) Marriage | (8) Birth |
|---|---|---|---|---|---|---|---|---|
| Exposure x Y95 | 0.004 | 0.009 | 0.012 | -0.010 | -0.014 | 0.003* | -0.005 | -0.002 |
| | (0.017) | (0.012) | (0.011) | (0.009) | (0.088) | (0.002) | (0.004) | (0.005) |
| Exposure x Y96 | 0.002 | 0.008 | 0.000 | -0.006 | 0.039 | -0.001 | 0.003 | -0.001 |
| | (0.019) | (0.015) | (0.012) | (0.010) | (0.100) | (0.002) | (0.004) | (0.005) |
| Exposure x Y97 | 0.011 | -0.001 | 0.011 | -0.006 | -0.020 | 0.001 | 0.005* | 0.011* |
| | (0.025) | (0.018) | (0.013) | (0.010) | (0.084) | (0.002) | (0.003) | (0.005) |
| Exposure x Y98 | 0.005 | -0.004 | 0.002 | 0.002 | -0.029 | 0.001 | 0.005 | 0.003 |
| | (0.026) | (0.021) | (0.013) | (0.014) | (0.101) | (0.002) | (0.004) | (0.006) |
| Exposure x Y99 | 0.024 | 0.011 | 0.003 | 0.004 | 0.080 | 0.000 | 0.003 | -0.002 |
| | (0.032) | (0.025) | (0.016) | (0.014) | (0.093) | (0.002) | (0.003) | (0.005) |
| N | 84,227 | 84,227 | 80,548 | 80,548 | 84,227 | 84,227 | 84,227 | 84,227 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Year FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |

**Notes:** Dependent variables are given at the top of the column. Estimation by least squares in the annual sample of 1990-1999. Labor market variables in logs. Divorce, Marriage, and Birth variables take one if the individual has an event of divorce, marriage, or birth (fathering or mothering a newborn child) in a year, and zero otherwise. Robust standard errors clustered at the 1999 firm in parentheses.

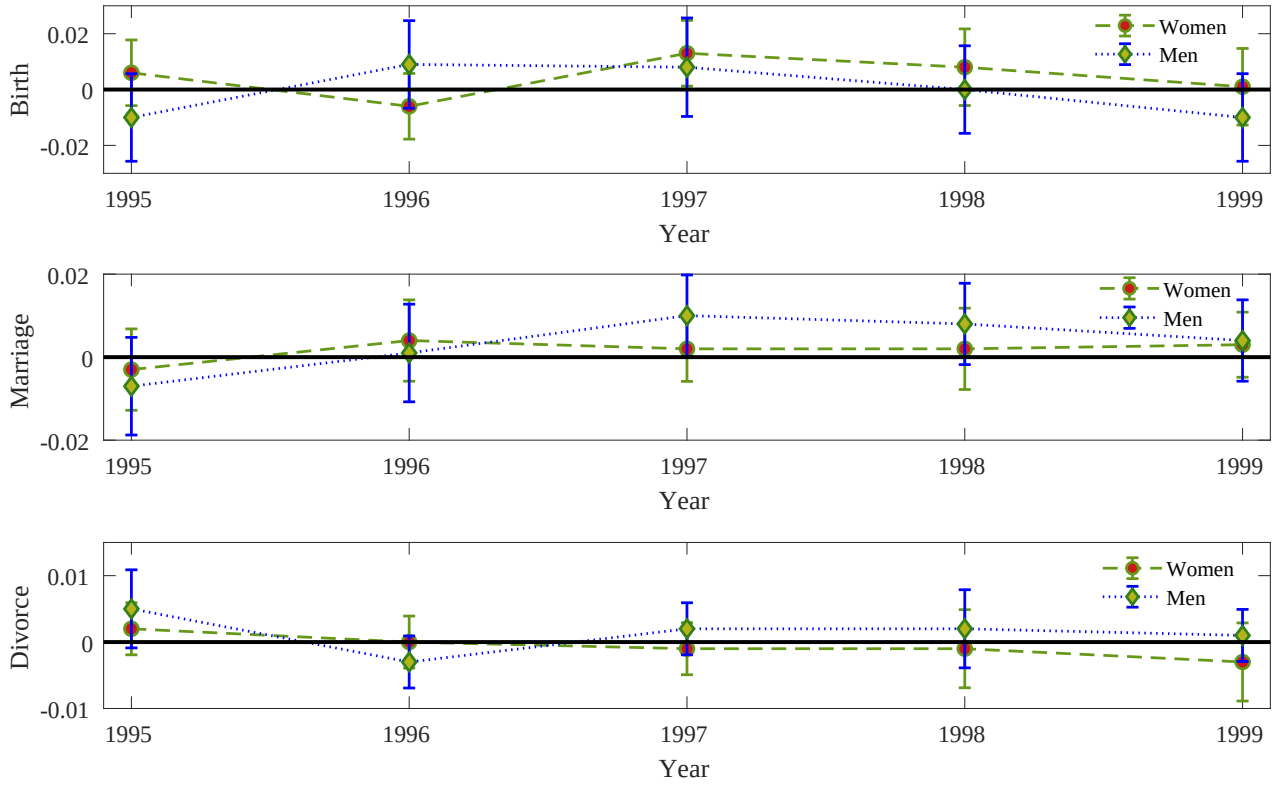


Figure A1: **Family Activities in the Pre-Sample Period by Gender**

**Notes**: Figure draws the difference-in-differences coefficient estimates obtained from estimating equation A2 for each gender in the annual sample of 1990-1999. Robust standard errors clustered at the 1999 firm. Shown is the 95% confidence interval.

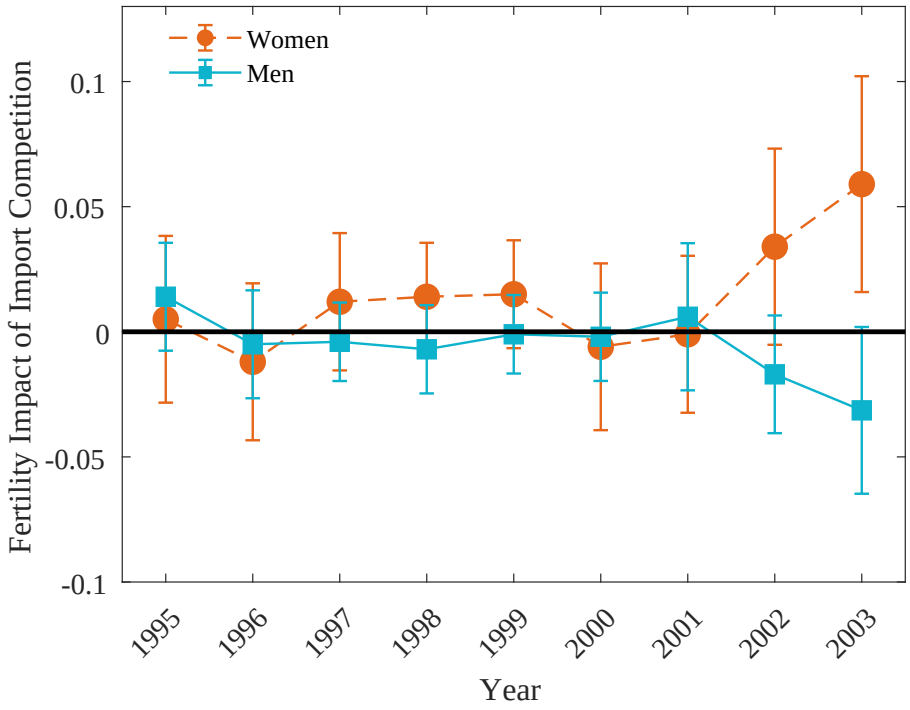


Figure A2: **The Evolution of Annual Fertility Effect for 1995-2003**

**Notes**: Figure draws the difference-in-differences coefficient estimates obtained from estimating equation A2 over 1990-2003 when the outcome variable is fathering or mothering a newborn baby. Robust standard errors clustered at the 1999 firm. Shown is the 95% confidence interval.

## A.6 Textile and Clothing Workers: Occupation and Sector by 2009

Figure A3 shows the occupational distribution of female textile workers by treatment status. We see that the differences are limited.

Turning to the sectoral movement of 1999 textile workers, Figure A4 shows that 38 percent of workers not exposed to rising import competition are still in manufacturing by 2009, while 31 percent have moved to the services sector. This confirms the general trend of a shift of employment away from manufacturing towards services.[75] At the same time, Figure A4 shows that of the set of exposed workers, 44 percent are employed in the service sector by 2009, while only 26 percent have still a manufacturing jobs. This difference suggests that rising import competition has sped up structural change for exposed workers. If manufacturing firms exposed to new import competition have shut down, displacing their workers, or they have scaled down their production, the rate at which exposed workers seek to find jobs in services will be relatively high. In line with this, note that the disproportional shift of exposed workers into services is virtually the same size as their

[75] Other factors that may explain this shift towards services are the relocation of manufacturing jobs to other countries and relatively high rates of labor-saving technological change in manufacturing.

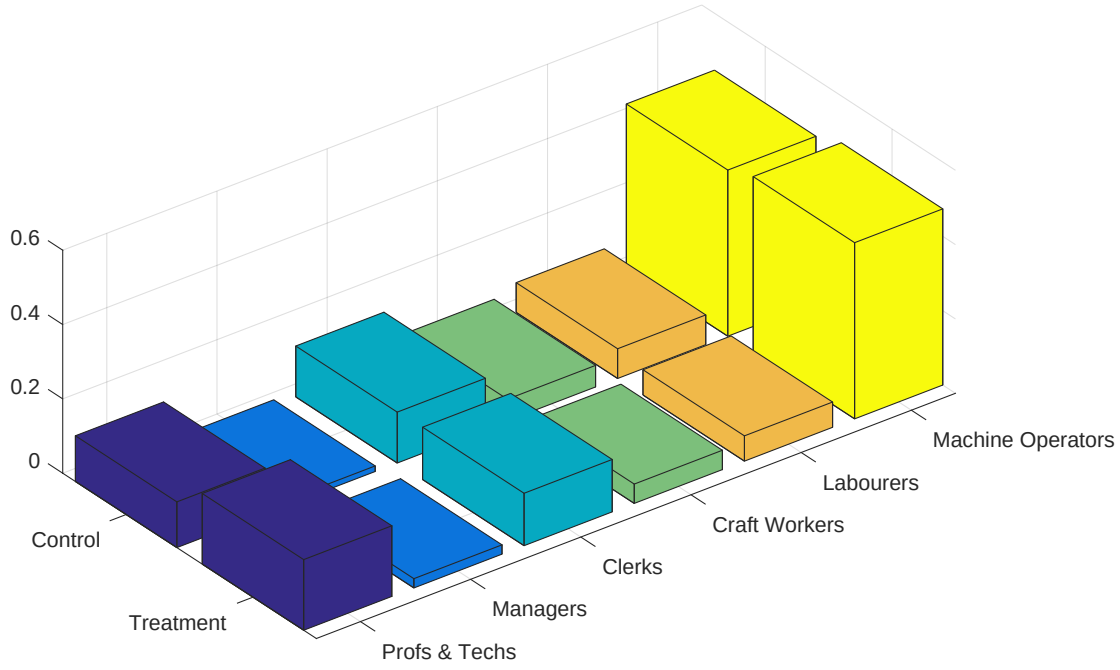


Figure A3: **Occupational Distribution of Female Workers in 1999**

Shown here is the distribution of female textile workers across major occupations by exposure to import competition in 1999.

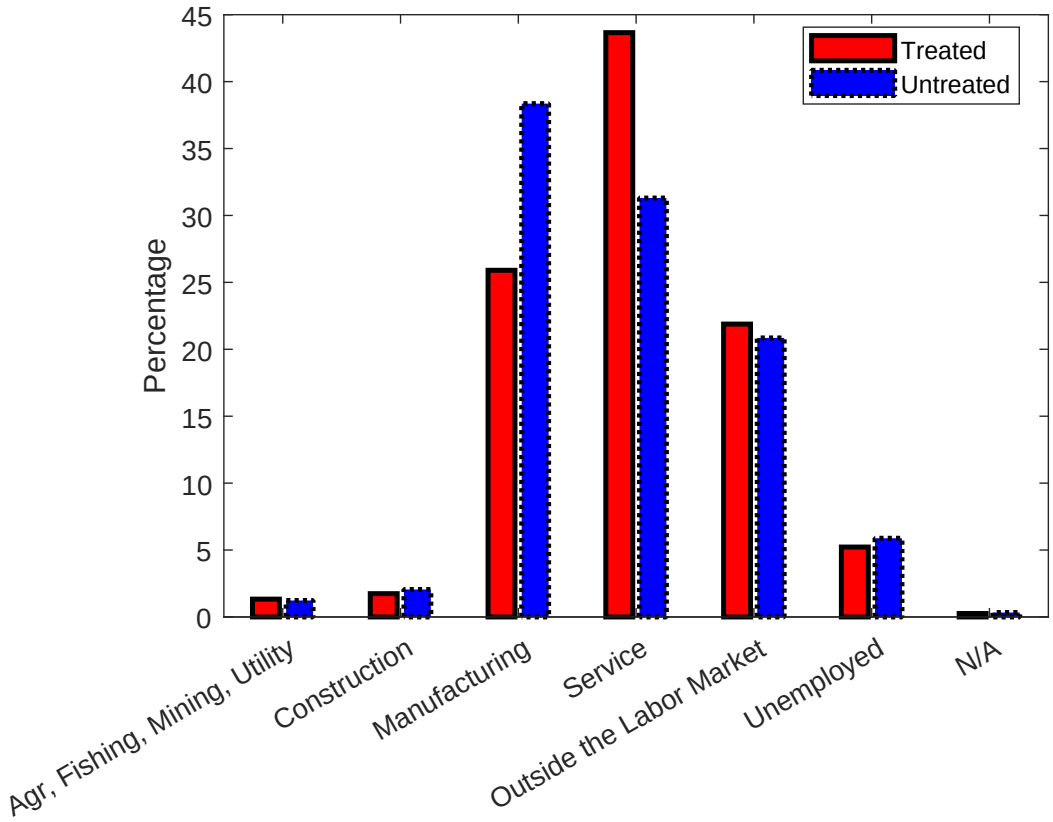


Figure A4: **Sectoral Distribution of Workers in 2009**

All regressions include worker fixed effects and the post-WTO accession period indicator.

lower tendency of staying in manufacturing (12, versus 13 percentage points, respectively). While Figure A4 shows that exposed workers are somewhat more likely to be out of the labor force than not exposed workers, overall Figure A4 suggests that the most important influence of trade exposure appears to be on the shift from manufacturing to services.

# B Treatment Variable: Revenue Share instead of Indicator

This section presents results obtained using an alternative, continuous treatment definition. Instead of a simple indicator whether a firm in 1999 manufactures any products for which the quotas will be liberalized in 2002 or 2005, we employ the revenue share of these products in 1999. Results are shown in Tables B6 and B7.

Table B6: **Import Competition and Fertility, Parental Leave, and Marriage Responses using Continuous Treatment Definition**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| Gender | | Women | Men | | Women | Men | | Women | Men |
| Dep Var. | Birth | Birth | Birth | ParL | ParL | ParL | Marriage | Marriage | Marriage |
| PostxRevShare | -0.139 | 0.302** | -0.139 | -0.151 | 0.346** | -0.151 | -0.228** | 0.193 | -0.228** |
| | -(0.119) | (0.145) | (0.119) | (0.097) | (0.134) | (0.097) | (0.111) | (0.145) | (0.111) |
| PostxRevSharexFemale | 0.441** | | | 0.497*** | | | 0.421** | | |
| | (0.189) | | | (0.165) | | | (0.179) | | |
| Observations | 3,160 | 1,452 | 1,708 | 3,160 | 1,452 | 1,708 | 3,160 | 1,452 | 1,708 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time FE | ✓ | | | ✓ | | | ✓ | | |

**Notes**: Dependent variable at top of column. Sample period is 1999-2009. Sample includes all fertile age (18-39) single textile workers as of 1999. ParL stands for parental leave. Robust standard errors are clustered at the 1999 firm level. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table B7: **Divorce and Earnings Effects using Continuous Treatment Definition**

| | Married Workers | | | All Workers | |
|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) |
| Gender | All | Women | Men | All | All |
| Dep. Var. | Divorce | Divorce | Divorce | Earnings | Earnings 1999 Firm |
| PostxRevShare | -0.081 | -0.112*** | -0.081 | 0.756 | -4.715*** |
| | (0.052) | (0.038) | (0.052) | -1.669 | -1.241 |
| PostxRevSharexFemale | -0.031 | | | -5.233*** | 0.589 |
| | (0.063) | | | (1.879) | (0.939) |
| Observations | 11,780 | 6,846 | 4,934 | 19,526 | 19,526 |
| Worker FE | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ |
| Female x Time FE | ✓ | | | ✓ | ✓ |

**Notes**: Sample period is 1999-2009. Robust standard errors are clustered at the 1999 firm level. *, ** and *** indicate significance at the 10%, 5% and 1% levels respectively.

Comparing the results of Tables B6 and B7 with the corresponding results obtained with the indicator treatment definition shown in the text indicates that the two sets of results are similar.

# C Analysis of the Economy-wide Labor Force

In this section we provide additional details on our analysis for the economy-wide sample discussed in section 5. The workers in this sample were, in 1999, employed in a wide range of industries, including mining, manufacturing, wholesale and retail trade, hotels and restaurants, transport, storage and communication, as well as real estate, renting and business activities. Sectors that are not included as initial employment of workers in the sample are public administration, education, health, and a wide range of small personal and social service providers. Following Keller and Utar (2019) we estimate the impact of import competition from China by employing six-digit NACE industry (or product line) variation in the change of Chinese import penetration in Denmark. We control for two-digit industry fixed effects to avoid broad industry variations that are likely to be convoluted with technological changes.

In the case of earnings the estimation equation is as follows:

$$Earnings_i^{2000-09} = \beta_0 + \beta_1 \Delta IMP_i + Z_i^W + Z_i^F + Z_i^P + \varepsilon_i. \tag{C3}$$

The dependent variable $Earnings_i^{2000-09}$ is the cumulative earnings over 2000-2009 measured in worker $i$'s 1999 annual earnings. On the right hand side we have the change in import penetration from China ($\Delta IMP_i$), as well as measures of worker ($Z_i^W$), firm and product-line ($Z_i^F$) and partner $Z_i^P$ characteristics as of year 1999.

The vector $Z_i^W$ includes gender, age, the interaction between gender and age, immigration status, marital status (married indicator, widow indicator, an indicator whether an individual has ever been in any form of homosexual union), the number of children, the squared number of children, education (college dummy, vocational education dummy, at most high-school diploma dummy), occupation (two-digit ISCO fixed effects), the logarithm of the hourly wage, the history of unemployment spells, an indicator whether the individual is a union member, and finally the worker's labor market experience as measured by the number of years in the labor market.

The vector $Z_i^F$ includes the average wage in the firm, the size of the firm as measured by the full-time equivalent number of employees, the separation rate of workers from the original firm between 1998-1999, pre-trends in the employment in the six-digit product line of employer between 1993-1999, the share of college educated workers in the six-digit product line of employer, and two-digit industry (NACE) fixed effects.

The set $Z_i^P$ includes partner characteristics, which include an indicator whether the individual has a partner (if not married), the partner's age, the partner's labor earnings, an indicator whether the partner is a Danish citizen, an indicator whether the partner is employed in manufacturing, an indicator whether the partner is employed in the same six-digit product line, and an indicator whether the partner is employed in a highly trade exposed industry (95th percentile and above of trade exposure), the age difference between the partners, and an indicator whether the partner has higher earnings. All $Z_i^W$, $Z_i^F$ , and $Z_i^P$ characteristics, if they are not explicitly indicated otherwise, are of the year 1999. Specifications for marriage, divorce, parental leave, fertility and other labor market outcomes are analogous to this earnings equation.

## C.1 Instrumental Variables and First-stage Regressions

Because the change in imports from China across industries might be endogenous, we employ an instrumental-variables approach. We construct our measure of Chinese import competition by developing a mapping between the international trade data at the eight-digit product level from Statistics Denmark's UHDI database and Denmark's six-digit industry classification, DB93. Our mapping follows the match between Combined Nomenclature (CN) and Classification of Products by Activity (CPA) available at Eurostat's RAMON database. We adapt this according to Danish industrial production using Statistics Denmark's VARES database. The mapping between trade (CN and Harmonized System, HS) and production data (DB93) is created separately for the three CN/HS versions, CN/HS-1996, CN/HS-1999 and CN/HS-2009. To construct Danish consumption figures at the six-digit DB93 level, we employ data on exports and imports from the UHDI database together with manufacturing revenue obtained from Statistics Denmark's FIRE database.

Imports from China in eight other high-income countries are employed as an instrumental variable in the following way:

$$\Delta HIP_j^{CH} = \frac{OM_{j,2009}^{CH} - OM_{j,1999}^{CH}}{C_{j,1996}} ,$$

where $OM_{j,t}^{CH}$ is the total value of imports in the corresponding industry $j$ in the eight high-income countries at year $t$. The countries are Australia, Finland, Germany, Japan, the Netherlands, New Zealand, Switzerland, and the United States. Data for the European countries comes from Eurostat and is available in the eight-digit CN classification. Data for the non-European countries comes

from the United Nation's COMTRADE database in the six-digit Harmonized System classification.

We employ two additional instrumental variables that can be viewed as structural measures of market openness in the pre-trade shock period. They are, first, the logarithm of the weighted average distance to the source countries of the goods Denmark imported in 1996 in worker $i$'s initial industry of employment. We denote this measure by $Open_j^{Dist}$.[76] The higher the value of $Open_j^{Dist}$ indicates lower transportation costs, and higher susceptibility of the industry to the import shock. Hence we expect a positive relationship between $\Delta IMP_i$ and $Open_j^{Dist}$. The second measure is the fraction of retail trade firms in all importing firms in worker $i$'s six-digit industry in 1996. This measure, denoted by, $Open_j^{DistrCH}$, proxies the openness of the pre-existing distribution channels that makes the industry more vulnerable to exogenous supply shocks from China.[77] Hence we expect a positive relationship between $\Delta IMP_i$ and $Open_j^{DistrCH}$. Table C8 show the first-stage coefficients across different subsamples that we use in our analysis, and confirms that the instruments are strongly correlated and have the expected signs.

[76] Employing bilateral distance data from CEPII http://www.cepii.fr/CEPII/en/bdd_modele/presentation.asp?id=8, we weight these distances by the import value shares at the six-digit industry level.

[77] The data comes from the FIRE database.

Table C8: **First-Stage Coefficients**

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Full Sample | | Fertile Age Sample | | Married Men Sample | |
| | $\Delta IMP_i$ | $\Delta IMP_i$xFemale | $\Delta IMP_i$ | $\Delta IMP_i$xFemale | $\Delta IMP_i$ | Spouse's $\Delta IMP_i$ |
| $\Delta HIP_j^{CH}$ | 0.002*** | -0.0018** | 0.002*** | -0.001* | 0.002*** | 0.000 |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.001) | (0.000) |
| $Open_j^{Dist}$ | 0.015*** | 0.006*** | 0.015*** | 0.006*** | 0.014*** | 0.001 |
| | (0.005) | (0.002) | (0.005) | (0.002) | (0.005) | (0.000) |
| $Open_j^{DistrCH}$ | 0.118** | 0.078*** | 0.114* | 0.073*** | 0.081 | -0.016*** |
| | (0.058) | (0.029) | (0.058) | (0.026) | (0.058) | (0.003) |
| $\Delta HIP_j^{CH}$xFemale | -0.0001 | 0.003*** | -0.000 | 0.003*** | | |
| | (0.000) | (0.001) | (0.000) | (0.001) | | |
| Spouse's $\Delta HIP_j^{CH}$ | | | | | -0.000 | 0.002*** |
| | | | | | (0.000) | (0.000) |
| Spouse's $Open_j^{Dist}$ | | | | | 0.000 | 0.001*** |
| | | | | | (0.000) | (0.000) |
| Spouse's $Open_j^{DistrCH}$ | | | | | 0.001 | 0.346*** |
| | | | | | (0.005) | (0.014) |
| Worker Variables | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Firm Variables | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Product-Line Variables | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Partner Variables | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Two-digit Occupation FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Two-digit Industry FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Sanderson-Windmeijer F | 13.45 | 14.9 | 12.81 | 15.16 | 9.995 | 806.44 |
| Observation | 1,671,774 | 1,671,774 | 900,846 | 900,846 | 478,354 | 478,354 |

**Notes:** All regressions include the full-set of $Z_i^W$, $Z_i^F$, and $Z_i^P$. Robust standard errors clustered at the level of the six-digit industry in parentheses (761 clusters). *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

## C.2 Gender Differential in Adjustment Costs among Fertile Age Workers

Table C9 presents the results for the family and labor market adjustment among the fertile age workers (N = 903,629). In Panel A we focus on family adjustment, whereas Panels B, C, and D provides evidence of various aspects of the workers' labor market adjustment. These results show that import competition generates a strong push to the service sector for both men and women (Panel D, column (2)). As the shock induces women to move disproportionately towards family (Panel A), it also induces a significant occupational change for fertile age women (Panel C, column (4)). In these new occupations, likely less career-oriented as women face career interruptions due to child birth, women earn much less (compare columns (2) and (4) in Panel C). This contributes to our finding of an overall worse labor market adjustment of fertile age women, see column (1),

Panel B.

Table C9: **Family and Labor Market Adjustment of Fertile Age Workers**

| **Panel A** | (1) Birth | (2) log No. of Births | (3) Parental Leave | (4) Marriage |
|---|---|---|---|---|
| $\Delta IMP_i$ | -0.190* | -0.173 | -0.293** | -0.053 |
| | (0.103) | (0.114) | (0.122) | (0.094) |
| $\Delta IMP_i$xFemale | 0.314*** | 0.307*** | 0.116 | 0.495*** |
| | (0.092) | (0.090) | (0.102) | (0.099) |
| **Panel B.** | Earnings | Earnings Initial 6-dig Ind | Employment | Employment Initial 6-dig Ind |
| $\Delta IMP_i$ | 1.613 | -12.310* | -0.542 | -8.650** |
| | (3.612) | (6.572) | (0.748) | (4.250) |
| $\Delta IMP_i$xFemale | -13.080* | 1.305 | -3.177** | 0.808 |
| | (7.426) | (3.015) | (1.452) | (1.995) |
| **Panel C.** | Earnings Initial Occupation | Earnings Different Occupation | Employment Same Occupation | Employment Diff Occupation |
| $\Delta IMP_i$ | -0.555 | 2.167 | -2.078 | 1.536 |
| | (4.710) | (5.364) | (3.381) | (3.158) |
| $\Delta IMP_i$xFemale | -14.820* | 1.735 | -10.030** | 6.856** |
| | (6.572) | (2.940) | (3.904) | (2.677) |
| **Panel D.** | Years in Manufacturing | Years in Service Sector | Outside Labor Mkt | Unemployment |
| $\Delta IMP_i$ | -5.423 | 6.143*** | 0.174 | 9.118* |
| | (3.578) | (3.121) | (0.341) | (4.659) |
| $\Delta IMP_i$xFemale | -6.743*** | 2.109 | 1.188* | 16.570* |
| | (1.938) | (1.805) | (0.673) | (9.449) |

**Notes:** Sample is all fertile-age workers (between 18 and 39 years old in 1999 N = 903,629). Dependent variables at top of column. All dependent variables are cumulative over 2000-2009. Initial 6-dig Ind stands for the worker's 6-digit industry in the year 1999 and initial occupation is the worker's 1-dig ISCO in 1999. Estimation analogous to equation (3) by two-stage least squares; second-stage coefficients shown. All regressions include the set of $Z_i^W$, $Z_i^F$, and $Z_i^P$ shown in Table 7. Robust standard errors clustered at the level of the six-digit industry (756 clusters) in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels, respectively.

# D Probit Estimation Results for Family Responses

This section presents probit results for child births and parental leave, marriage and divorce that complement Tables 3, 4, 5 and 6 in the text.

Changes in family status and the number of children are relatively rare, discrete events, and consequently a natural estimation method is also probit regression. Family outcome $X_{is}$ of worker $i$ in period $s$ is specified as follows:

$$X_{is} = f(\beta_1 Exposure_{i,99} * Post_s + \beta_2 Post_s + \beta_3 Exposure_{i,99} + \beta' W_{i,99} + \varepsilon_{is}), \quad s = 0,1\ , \qquad \text{(D4)}$$

where the vector $W_{i,99}$ contains 1999 characteristics of worker $i$, including on $i$'s firm and partner (if applicable). [78] $Post_s$ captures the influence of aggregate trends affecting all workers. Of key interest is $\beta_1$ which estimates whether exposed workers show different outcomes compared to observationally similar non-exposed workers, relative to pre-shock years. We also allow for correlation within a group of workers employed by the same firm in 1999 and cluster standard errors by each worker's 1999 firm. For ease of exposition, we denote the difference-in-difference term $Exposure_{i,99} * Post_s$ by $ImpComp_{is}$, mnemonic for rising import competition.

The results of these probit regressions are presented in Tables D10, D11, D12, and D13 and they confirm the corresponding findings obtained using least squares estimation with worker fixed effects in Tables 3, 4, 5 and 6 in the text.

Probit regressions allow us as well to illustrate the role of occupational sorting for our results. Table D14 compares results for two sets of probit regressions, one without and one with four-digit fixed effects for a worker's 1999 occupation.[79] If sorting between occupations is important for gender differences, the inclusion of occupational fixed effects should lead to substantially smaller gender differentials because four-digit fixed effects eliminate a substantial part of the between variation.

Table D14 shows that the results with fixed effects are quite similar to those without occupational

[78] This includes the worker's age, gender, the number of children, an indicator for first or second generation immigrant status, an indicator for being married, an indicator for being a single parent, education (whether the most attained education is high-school degree), the average wage of the worker's firm, the separation rate of the worker's firm, the trade exposure status of the partner, and education of partner (whether the most attained education is high-school degree). The separation rate of the worker's workplace is defined as the percentage of employees that are not employed in the workplace in year 1999 with respect to 1998.

[79] These fixed effects would not be identified in our least-squares specification with worker fixed effects.

Table D10: **Import Competition and Births - Probit Results**

| Gender | (1) All | (2) All | (3) All | (4) Men | (5) Women | (6) All | (7) All | (8) Men | (9) Women |
|---|---|---|---|---|---|---|---|---|---|
| | | Not married | | | | Single (Not Married and Not Co-habiting) | | | |
| ImpComp | 0.04 | 0.164** | 0.137 | 0.156 | 0.208* | 0.162 | 0.021 | -0.079 | 0.411** |
| | (0.081) | (0.077) | (0.094) | (0.11) | (0.111) | (0.121) | (0.141) | (0.177) | (0.167) |
| Marg. Effect | 0.013 | 0.058 | 0.048 | 0.053 | 0.075 | 0.042 | 0.054 | -0.018 | 0.117 |
| ImpCompxFemale | 0.031 | | 0.035 | | | | 0.247* | | |
| | (0.082) | | (0.101) | | | | (0.133) | | |
| Marg. Effect | 0.01 | | 0.012 | | | | 0.067 | | |
| Observations | 9,864 | 5,749 | 5,749 | 2,779 | 2,970 | 3,144 | 3,144 | 1,695 | 1,449 |
| Worker, firm, partner vars | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Obs Prob | 0.278 | 0.31 | 0.31 | 0.29 | 0.328 | 0.21 | 0.21 | 0.194 | 0.228 |
| Pseudo R-Sq | 0.092 | 0.09 | 0.09 | 0.086 | 0.1 | 0.166 | 0.168 | 0.165 | 0.185 |

**Notes:** Dependent variable is one if worker *i* has a newborn child during period *s*, and zero otherwise. Estimation by probit regression. The sample in column (1) is textile workers of fertile age (between 18-39 as of 1999). "M" is Men, "W" is Women. The sample in columns (2) to (5) is workers not married as of 1999, in columns (6) to (9) workers neither married nor co-habitating as of 1999. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table D11: **Import Competition and Parental Leave - Probit Results**

| | (1) All | (2) All | (3) All | (4) Men | (5) Women | (6) All | (7) All | (8) Men | (9) Women |
|---|---|---|---|---|---|---|---|---|---|
| | | Not married | | | | Not Married and Not Co-habiting | | | |
| ImpComp(PostxExposure) | 0.03 | 0.148** | 0.041 | 0.165 | 0.17 | 0.096 | -0.117 | -0.187 | 0.303* |
| | (0.078) | (0.073) | (0.089) | (0.106) | (0.11) | (0.123) | (0.142) | (0.182) | (0.164) |
| Marg. Effect | 0.009 | 0.048 | 0.013 | 0.045 | 0.06 | 0.021 | -0.024 | -0.029 | 0.083 |
| ImpCompxFemale | 0.051 | | 0.082 | | | | 0.295** | | |
| | (0.086) | | (0.105) | | | | (0.137) | | |
| Marg. Effect | 0.015 | | 0.026 | | | | 0.07 | | |
| Observations | 9,864 | 5,749 | 5,749 | 2,779 | 2,970 | 3,144 | 3,144 | 1,695 | 1,449 |
| Worker, firm, partner vars | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FE | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Obs Prob. | 0.241 | 0.258 | 0.258 | 0.193 | 0.32 | 0.159 | 0.159 | 0.12 | 0.205 |
| Pseudo R-Sq | 0.072 | 0.062 | 0.069 | 0.054 | 0.07 | 0.107 | 0.12 | 0.107 | 0.128 |

**Notes:** Dependent variable is one if worker *i* takes parental leave during period *s*, and zero otherwise. Estimation by probit regression. The sample in column (1) is textile workers of fertile age (between 18-39 as of 1999). "M" is Men, "W" is Women. The sample in columns (2) to (5) is workers not married as of 1999, in columns (6) to (9) workers neither married nor co-habitating as of 1999. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table D12: **Import Competition and Marriage - Probit Results**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | | | | Fertile Age | | | Fertile Age Not married, and not cohabiting | | |
| Gender | All | Men | Women | All | Men | Women | All | Men | Women |
| ImpComp(PostxExposure) | -0.059 | -0.088 | 0.188 | -0.008 | -0.054 | 0.255** | -0.01 | -0.04 | 0.378* |
| | (0.094) | (0.112) | (0.114) | (0.108) | (0.129) | (0.13) | (0.16) | (0.186) | (0.229) |
| Marg. Effect | -0.014 | -0.021 | 0.048 | -0.002 | -0.015 | 0.078 | -0.002 | -0.008 | 0.084 |
| ImpCompxFemale | 0.221*** | | | 0.22** | | | 0.343** | | |
| | (0.084) | | | (0.096) | | | (0.134) | | |
| Marg. Effect | 0.059 | | | 0.067 | | | 0.078 | | |
| Observations | 8,166 | 3,838 | 4,328 | 5,749 | 2,779 | 2,970 | 3,144 | 1,695 | 1,449 |
| Probit | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Worker, firm, partner vars | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Obs Prob | 0.19 | 0.185 | 0.194 | 0.226 | 0.213 | 0.238 | 0.157 | 0.146 | 0.17 |
| Pseudo R-Sq | 0.097 | 0.084 | 0.115 | 0.081 | 0.078 | 0.088 | 0.147 | 0.133 | 0.177 |

**Notes:** Dependent variable is one if worker *i* gets married during period *s*, and zero otherwise. Estimation by probit regression. Only workers who are not married as of 1999 included in the sample. "M" is Men, "W" is Women. The fertile age refers to workers who are between 18 and 39 years old as of 1999. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table D13: **Import Competition and Divorce - Probit Results**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | | | | | | Fertile Age | | |
| VARIABLES | All | All | Men | Women | All | All | Men | Women |
| ImpComp(PostxExposure) | -0.283*** | -0.243* | -0.212 | -0.391*** | -0.466*** | -0.390** | -0.276 | -0.681*** |
| | (0.104) | (0.124) | (0.146) | (0.146) | (0.131) | (0.168) | (0.199) | (0.191) |
| | -0.018 | -0.016 | -0.016 | -0.021 | -0.053 | -0.045 | -0.033 | -0.070 |
| ImpCompxFemale | | -0.120 | | | | -0.192 | | |
| | | (0.103) | | | | (0.139) | | |
| | | -0.008 | | | | -0.023 | | |
| Observations | 10,287 | 10,287 | 4,008 | 6,279 | 4,115 | 4,115 | 1,478 | 2,637 |
| Worker, firm, partner vars | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Sample Prob | 0.058 | 0.058 | 0.06 | 0.056 | 0.098 | 0.098 | 0.099 | 0.098 |
| Pseudo R2 | 0.173 | 0.174 | 0.153 | 0.195 | 0.13 | 0.133 | 0.134 | 0.144 |

**Notes:** Dependent variable is one if worker *i* divorces during period *s*, and zero otherwise. Estimation by probit regression. Only workers who are married as of 1999 included in the sample. The fertile age refers to workers who are between 18 and 39 years old as of 1999. Robust standard errors clustered at the level of workers' initial firm are in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

Table D14: **Gender Differential in Family Response to Exposure: Within or Between Occupations?**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Marriage | | Birth | | Parental Leave | | Divorce | |
| ImpComp | -0.020 | -0.029 | -0.012 | -0.040 | -0.104 | -0.128 | -0.102 | -0.121 |
| | (0.0945) | (0.0957) | (0.136) | (0.145) | (0.142) | (0.151) | (0.112) | (0.117) |
| ImpComp x Female | 0.153* | 0.184* | 0.321** | 0.330** | 0.320** | 0.316** | -0.188* | -0.196* |
| | (0.0922) | (0.0946) | (0.140) | (0.149) | (0.139) | (0.149) | (0.0970) | (0.104) |
| Observations | 8,163 | 8,008 | 3,283 | 3,103 | 3,283 | 3,045 | 11,703 | 10,941 |
| All controls | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Period fixed effects | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Four-digit Occupation FEs | - | ✓ | - | ✓ | - | ✓ | - | ✓ |

**Notes:** Dependent variables are equal to one if worker in period s has married (columns 1-2), or has mothered/fathered a newborn baby (columns (3-4), or has taken a parental leave (columns (5-6), or has divorced (columns 7-8), zero otherwise. Sample: all unmarried workers as of 1999 (columns 1-2), all fertile age single workers as of 1999 (columns 3-6), all married workers as of 1999 (columns 7-8). Estimation by probit with period fixed effects (equation 1). All specifications include individual, partner, firm characteristics. Columns 2, 4, 6, 8 include in addition four-digit occupation fixed effects. Robust standard errors clustered at the 1999 firm level in parentheses. *, ** and *** indicate significance at the 10 %, 5% and 1% levels respectively.

fixed effects. This is consistent with Goldin's (2014) observation that within-occupation effects, rather than between-occupation sorting, are key for explaining the gender wage gap.

# E Alternative Definitions of Fertile Age

Emphasizing fertility considerations in women's trade shock responses, the analysis so far has contrasted younger with older workers employing a common fertile-age threshold of 18-39 (as of 1999). In this section we consider alternative thresholds. First, we consider gender specific fertile-ages by defining women who are 36 years old or less and men who are 45 years or less, both as of 1999, as fertile-age. We also show results for an alternative common definition of fertile age by examining workers who are between twenty and forty years old as of 1999.

Table E15 provides results on family responses of workers with the gender-specific definition of fertile age. The results show that women react to exposure with more family activities (more birth, parental leave, and marriage, less divorce), in contrast to men.

Results for the alternative non-gender specific definition of fertile-age are presented in Table E16; they turn out to be similar. Overall, our results are not sensitive to alternative ways of defining fertile-age.

Table E15: **Family Responses with Alternative Definition of Fertile Age I**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Birth | | Parental Leave | | Marriage | | Divorce | |
| | Men | Women | Men | Women | Men | Women | Men | Women |
| ImpComp | -0.009 | 0.135*** | -0.021 | 0.123*** | -0.021 | 0.067* | -0.027 | -0.097*** |
| | (0.034) | (0.043) | (0.026) | (0.042) | (0.03) | (0.039) | (0.02) | (0.025) |
| Worker FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 1,682 | 1,470 | 1,682 | 1,470 | 2,804 | 3,024 | 2,002 | 2,964 |
| R-squared | 0.605 | 0.602 | 0.592 | 0.611 | 0.45 | 0.427 | 0.492 | 0.505 |

**Notes:** Age limit of the sample: Male workers who are less than 46 years old as of 1999, female workers who are less than 37 years old as of 1999. The sample period: 1999-2009. Dependent variable given at top of column. Estimation by least squares with time and worker fixed effects. Sample in columns (1) to (4) is single, in columns (5) and (6) unmarried, and in columns (7) and (8) married workers, all as of 1999. Robust standard errors clustered at the level of workers' initial firm are in parentheses. *, ** and *** indicate significance at the 10%, 5% and 1% levels respectively.

Table E16: **Family Responses with Alternative Definition of Fertile Age II**

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Birth | | Parental Leave | | Marriage | | Divorce | |
| | Men | Women | Men | Women | Men | Women | Men | Women |
| ImpComp | -0.015 | 0.104*** | -0.023 | 0.091** | -0.013 | 0.076** | -0.026 | -0.087*** |
| | (0.038) | (0.04) | (0.03) | (0.038) | (0.033) | (0.036) | (0.025) | (0.022) |
| Worker FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Observations | 1,682 | 1,470 | 1,682 | 1,470 | 2,804 | 3,024 | 2,002 | 2,964 |
| R-squared | 0.605 | 0.602 | 0.592 | 0.611 | 0.45 | 0.427 | 0.492 | 0.505 |

**Notes:** Age limit of the sample: Male and female workers between 20 and 40 years old as of 1999. The sample period: 1999-2009. Dependent variable given at top of column. Estimation by least squares with time and worker fixed effects. Sample in columns (1) to (4) is single, in columns (5) and (6) unmarried, and in columns (7) and (8) married workers, all as of 1999. Robust standard errors clustered at the level of workers' initial firm are in parentheses. $^{*}$, $^{**}$ and $^{***}$ indicate significance at the 10%, 5% and 1% levels respectively.

To summarize, our main results are robust to employing alternative (and gender-neutral) restrictions on age to distinguish younger from older workers.

# F Technology as an Alternative Shock

Table F17: **Trends for Occupations with Different Routine-Task Intensity**

| | (1) Earnings | (2) Earnings 1999 Job | (3) Birth Event | (4) Parental Leave | (5) Marriage | (6) Divorce |
|---|---|---|---|---|---|---|
| Imp Comp | 0.306 | -1.239*** | -0.028 | -0.028 | -0.037 | -0.016 |
| | (0.470) | (0.384) | (0.042) | (0.035) | (0.030) | (0.016) |
| ImpComp x Female | -1.166** | 0.102 | 0.144** | 0.143** | 0.097** | -0.024 |
| | (0.545) | (0.290) | (0.064) | (0.056) | (0.039) | (0.019) |
| Female x Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| RTI x Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Worker FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| Time FEs | ✓ | ✓ | ✓ | ✓ | ✓ | ✓ |
| N | 16,552 | 16,552 | 2,468 | 2,468 | 6,752 | 10,196 |

**Notes**: Dependent variables are listed at top of columns; Earnings is the cumulative earnings over 1999-2001 and 2002-2009 periods; Earnings 1999 Job are the cumulative earnings obtained from the initial, 1999, firm over 1999-2001 and 2002-2009 period and analogously for all dependent variables. Least squares estimation of equation (2) augmented with additional fixed effects. All regressions include workers' initial two-digit ISCO occupations' routine-task intensity (RTI) x Time. Estimation samples in columns 3 and 4 include single fertile age workers, in column 5 unmarried workers and in column 6 married workers as of year 1999. Robust standard errors clustered at the level of the 1999 firm. $^{*}$, $^{**}$ and $^{***}$ indicate significance at the 10 %, 5% and 1% levels respectively.

The results indicate that our results are robust to allowing for differential time trends for occupations that have different routine task intensities. This provides evidence that the impact of rising import competition is separate from any effect that technological change might have.

# G Evolution over Time: Key Responses of Textile Workers

In this section we provide the additional results on the over-time responses of workers in reacting to rising import competition. We present figures showing the gender differential in parental leave and marriage behavior over time. Furthermore, we show that there is no major difference between male and female workers in how they are affected in their 1999 job, confirming that the initial

impact of rising import competition –at any time horizon– was similar for both gender. In the following figures, the year on the horizontal axis is the last year included in the analysis.

Figure G5 shows the evolution of the gender differential in parental leave uptake.

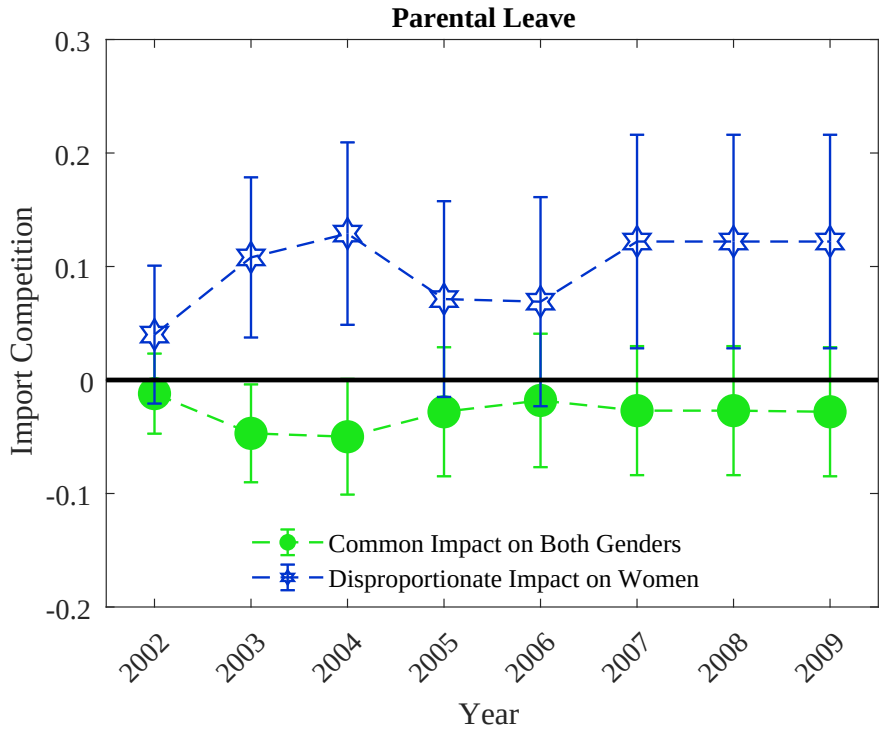


Figure G5: **Impact of Import Competition on Parental Leave over Time**

**Notes**: Figure shows the coefficient estimates obtained from estimating equation (2) on a rolling end-year basis when the dependent is parental leave take defined over the rolling periods. Blue stars indicate the coefficient estimate of ImpComp (Post x Exposure), green circles indicate the coefficient estimate of ImpComp x Female (Post x Exposure x Female). Shown is the 95% confidence interval.

Figure G6 shows the over-time effect of rising import competition on the marriage behavior workers who are single in 1999.

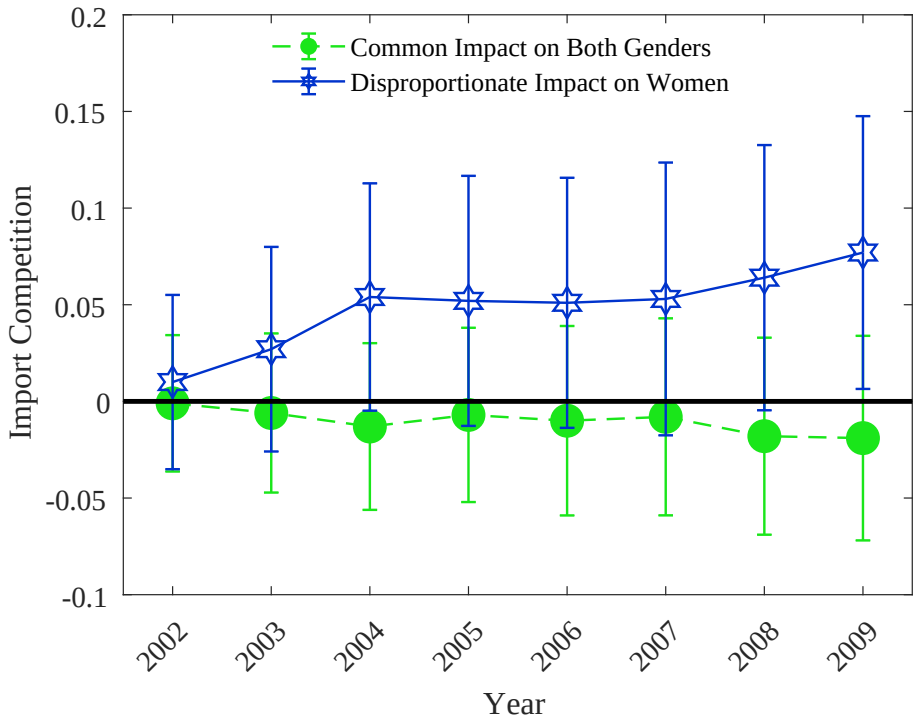


Figure G6: **Impact of Import Competition on Formation of Marriages over Time**

**Notes**: Figure draws the coefficient estimates obtained from estimating equation (2) on a rolling end-year basis when the dependent is getting married defined over the periods and the sample consists of textile workers who are not married as of 1999. Blue stars indicate the coefficient estimate of ImpComp (Post x Exposure), green circles indicate the coefficient estimate of ImpComp x Female (Post x Exposure x Female). Shown is the 95% confidence interval.

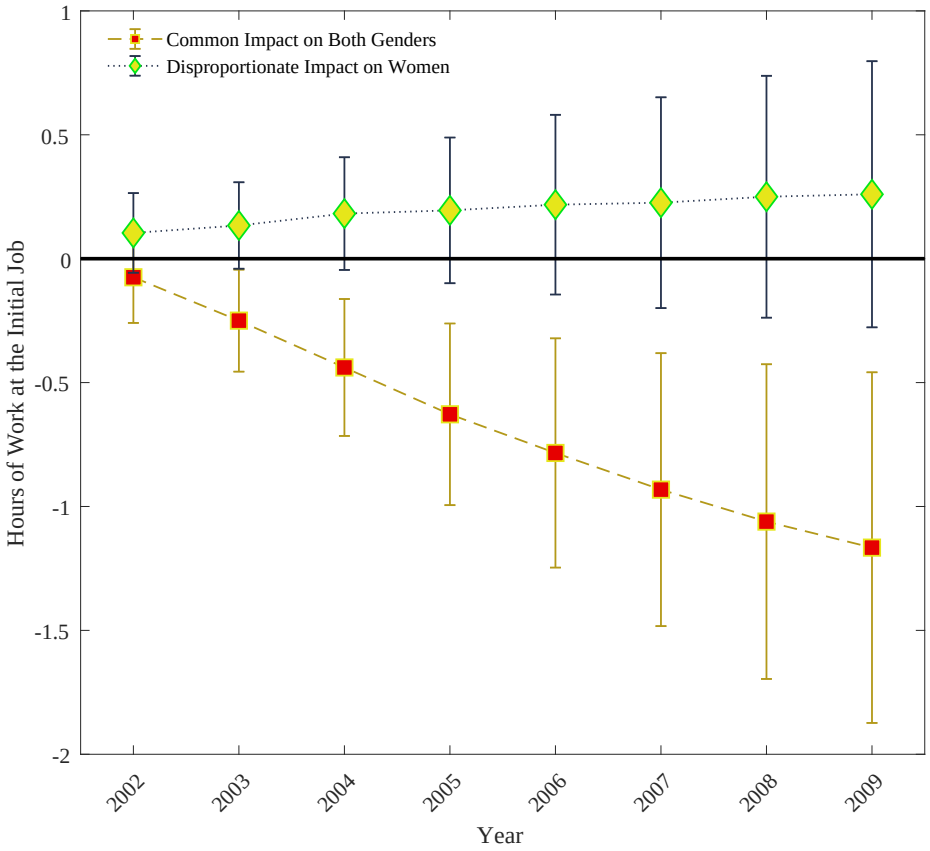


Figure G7: **Impact of Import Competition on Hours Worked at the Initial Firm over Time**

**Notes**: Figure draws the coefficient estimates obtained from estimating equation (2) on a rolling end-year basis when the dependent is cumulative hours worked at the initial job (measured in workers' own annual hours worked in 1999) over each period. Red squares indicate the coefficient estimates of ImpComp (Post x Exposure), green diamonds indicate the coefficient estimates of ImpComp x Female (Post x Exposure x Female). Shown is the 95% confidence interval.

Figure G7 shows how the gender differential due to import competition in the cumulative hours worked at the initial, 1999 firm evolved over the period 2002 to 2009. It confirms that the shock leads to significant displacement from the exposed workplace but this effect is not disproportionately felt among women.

# H Family Responses: Labor Market Status and Timing

Further evidence for the role of the biological constraint can be provided by examining the type of labor market positions that induce exposed workers to make a shift towards family activities. We exploit the timing of family activities and the workers' labor market positions at the time of these activities using our rich data. Table H18 shows the results. Each cell in Table H18 reports the difference-in-difference coefficient (and robust standard error), *ImpComp*, based on a least-squares regression with worker and period fixed effects. Results for Birth, Parental Leave, Marriage, and Divorce are shown by the four broad columns, separately for men and for women.

The first row of Table H18 shows the overall family response, irrespective of the worker's labor market position; this repeats results from the earlier Tables 3 to 6 for convenience.[80] The following two rows distinguish family responses while being employed in the original 1999 textile job from family responses after the worker has left the 1999 textile job. Finally, the lower rows distinguish two specific labor market positions after departing from the original textile job, namely (1) Outside of the labor force and (2) Unemployed. Our interest lies in which of these labor market positions, if any, is closely related to the worker's take-up of family activities.[81]

We begin with the family outcomes while the worker still works at the initial firm. The results show that trade exposure rarely generates a pro-family response while the workers are employed at the initial firm, neither for men nor for women (row 2). The coefficients tend to be small and insignificant.[82] In sharp contrast, import competition often triggers pro-family choices once a worker loses employment at their initial firm, especially for women. The results show that exposed women are induced to take pro-family action in terms of all four outcomes (row 3). Taken together, this establishes that *change* of labor market position is correlated with exposure-induced family responses.

However, this does not necessarily constitute evidence in support of Becker's (1973) hypothesis that labor market and family activities are jointly determined. Perhaps trade exposure matters because by moving to a new job workers make a new set of acquaintances, and the increase in family activities is the consequence? The final rows of Table H18 show that there seems to be

[80] The birth and parental leave results are columns (4) and (5) of Tables 3 and 4, respectively, while marriage results are from columns (2) and (3) of Table 5 and divorce results are from Table 6, columns (4) and (5).

[81] Recall that the labor market position of an individual is recorded every year in late November, while the definitions of the family outcomes cover the whole calendar year. It is thus in principle possible, for example, that in a given year an unemployed worker has taken parental leave from his job.

[82] The exception to this is the female divorce response, however, the corresponding divorce point estimate for men is similar in magnitude so it does not help to explain the gender differential.

more to it than meeting new people. Rather, we see that women make often pro-family decisions out of relatively weak positions in the labor market. For example, a relatively large share of trade-induced new births occurs when the women is outside of the labor force (coefficient of 0.043, column (2)). The same cannot be said for men (coefficient of virtually zero, column (1)). Similarly, exposure-induced parental leave uptake for women who are unemployed or out of the labor force is important, whereas this is not the case for men (rows 4 and 5, columns (3) and (4)). Women who are outside of the labor market are also marrying due to rising import competition, in contrast to men (columns (5), (6), row 4).

Overall, this analysis shows that women tend to move towards family after they are displaced, and the family activities coincide with a relatively weak labor market position.

Table H18: **Family Responses to Trade Exposure across Labor Market Positions**

| | Birth | | Parental Leave | | Marriage | | Divorce | |
|---|---|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| | Women | Men | Women | Men | Women | Men | Women | Men |
| Any Labor Market Position | 0.077** | 0.053 | 0.067* | 0.044 | 0.058** | -0.019 | -0.040*** | -0.021 |
| | (0.037) | (0.035) | (0.037) | (0.028) | (0.029) | (0.027) | (0.011) | (0.014) |
| At the Initial Job | 0.007 | 0.008 | 0.008 | 0.023 | 0.000 | -0.037 | -0.014** | -0.013 |
| | (0.025) | (0.028) | (0.028) | (0.022) | (0.018) | (0.021) | (0.006) | (0.010) |
| After Leaving the Initial Job | 0.093*** | 0.039 | 0.099*** | 0.013 | 0.057 | 0.018 | -0.026*** | -0.008 |
| | (0.031) | (0.029) | (0.033) | (0.023) | (0.022) | (0.022) | (0.010) | (0.012) |
| Of which: | | | | | | | | |
| Unemployed | 0.005 | -0.011 | 0.020 | -0.007 | -0.001 | -0.004 | -0.004 | -0.003 |
| | (0.006) | (0.007) | (0.013) | (0.004) | (0.005) | (0.004) | (0.003) | (0.004) |
| Out of Labor Force | 0.041** | 0.001 | 0.034** | -0.004 | 0.014* | 0.002 | -0.006 | 0.002 |
| | (0.017) | (0.007) | (0.016) | (0.003) | (0.008) | (0.005) | (0.003) | (0.004) |

**Notes:** Each cell gives a least squares coefficient and standard error estimate of *ImpComp* obtained from the estimation of equation (1), including worker and period fixed effects and a constant. For the sample sizes, see Tables 3 to 6. The sample of workers in columns (1) to (4) is single fertile age workers, in columns (5)-(6) is unmarried, and in columns (7)-(8) married workers. Robust standard errors clustered at the level of workers' 1999 firm are in parentheses. $^{*}$, $^{**}$ and $^{***}$ indicate significance at the 10 %, 5% and 1% levels respectively.