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International Center for Public Policy
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Does Trade Liberalization Foster Intimate Partner Violence?

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Abstract

By exploiting an unexpected policy change in the form of drastic tariffs reduction across several industries in Peru during the 2000s we are able to causally show that in districts where industries' employment are predominantly male, trade liberalization produced an increase in physical intimate partner violence of 36 percentage points with respect to control districts in our preferred specification. We find no such difference in districts where industries' employment is predominantly female. These findings are original and consistent with several hypotheses in the social sciences. Our results are robust to falsification and placebo tests, sensitivity to initial conditions, conflation of past and current shocks, selective migration, permutation tests and input-tariffs considerations. Finally, we find considerable heterogeneity, as education and the age of first marriage appear to be key variables that correlate with our findings.

Keywords: Domestic violence, Tariff reduction, Gender-specific shocks, Differences-in-Differences, Latin America

JEL Classification Codes: O19, O24, F13, J12.

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Introduction

In recent years, international trade has expanded the focus of what used to be considered its traditional realm of research to include broader, more interdisciplinary approaches that now consider topics related to crime, child labor, mortality, gender, inequality and poverty, among others.[1] In this research, we study the impact of trade liberalization on intimate partner violence, a seemingly unrelated issue associated to trade liberalization, but one that we believe is rather plausible and consistent with recent untested hypotheses in the social sciences.[2] In order to do this, we take advantage of the exogenous variation in industries that tend to specialize geographically due to liberalization in trade, but in addition we also exploit the fact that gender composition tends to be skewed towards either males or females depending on the particular industry, something that is typically observed more acutely in developing countries. This methodological approach allows us to study whether trade liberalization in the form of a reduction in tariffs causes an undesired negative impact in terms of intimate partner violence.

We believe that this question is rather relevant, as the rates of intimate partner violence in many countries around the world are high and pervasive. According to the Pan American Health Organization (2019), physical or sexual intimate partner violence has affected more than a quarter of women at some point in their lives. Furthermore, intimate partner violence has long-term effects on women's health. Women abused by their partners are 16 percent more likely to give birth to a low-weight baby (World Health Organization,

[1] Examples are Edmonds et al. (2009, 2010); Gaddis and Pieters (2017); Autor, et al., (2018); Kis-Katos and Sparrow (2015); Pierce and Schott (2016); Dix-Carneiro and Kovak (2015, 2017), Dix-Carneiro et al., 2018, among others.
[2] Intimate partner violence is also known simply as domestic violence and it is defined as physical or psychological harm by a current or former partner or spouse. This type of violence can occur among heterosexual or same-sex couples and does not require sexual intimacy. For more information, refer to the Centers for Disease Control and Prevention website.

2013), and are associated with higher substance abuse, worse mental health and a higher incidence of chronic diseases (Coker et al 2002; Ackerson and Subramanian, 2008; Ellsberg et al., 2008). The literature shows that the more male-dominated the society, the more likely the chance that female-generated increases in household earnings will produce negative responses in male partners, which may result in intimate partner violence, either physical or emotional. Sociologists explain that this backlash may occur as a way for males to assert their power within the household as well as a way to recover their authority (Macmillan and Gartner, 1999). In this context, it is reasonable to expect that if trade liberalization impacts industries that are predominantly associated with male-related activities within-household dynamics may evolve differently compared to how liberalization may impact predominantly women-related industries. In fact, it has been shown that trade liberalization appears to have gendered effects on labor market outcomes, which may affect the relative position of partners within a household (Gaddis and Pieters, 2017). Overall, it may be argued that depending on the gender composition of industries and the magnitude of income shocks, households may be put in a becalming or stressful situation, which in a traditional male-female household may impact the probability that the latter ends up suffering from intimate partner violence (Tauchen, et al., 1991).[3]

Our research focuses on the process of trade reform in Peru during the 2000s. Specifically, between 2004 and 2011 trade openness in the country increased rapidly, substantially and unexpectedly, as tariffs in thousands of different products were drastically

[3] Not all existing theories in the literature claim that increased earning capacity in females will increase the probability of suffering from intimate partner violence. In societies that are less male-dominated increased earnings by females may improve their bargaining power and may help them improve their options outside marriage or civil union thus helping reduce intimate partner violence (Aizer, 2010; Anderberg et al., 2016).

reduced as a result of a large and drastic liberalization policy. The average tariff was reduced from 10.34 percent in 2004 to 2.96 percent in 2011. These reductions were concentrated in mostly highly protected sectors.[4] We employ a differences-in-differences approach by exploiting gender-related labor composition by industry and at the district level. For example, if male labor is mostly specialized in mining when tariffs are reduced in a particular district, they will become relatively more exposed to trade liberalization with respect to male workers in other districts. Likewise, if female labor within a district specializes in textiles when tariffs are reduced, they will become relatively more exposed to trade liberalization with respect to female workers in other districts. We compute gender-specific measures of exposure to trade liberalization, which in order to simplify we call "male exposure" and "female exposure." In addition, we control for district fixed effects, region-year fixed effects and cluster standard errors at the district level.[5]

Our findings show that trade liberalization may foster intimate partner violence in Peruvian districts where tariff cuts affected male employment the most. In such districts, the average increase in physical intimate partner violence ranges between 18.83 and 36.59 percentage points depending on the specification. Interestingly, we find no such impact in the case of districts where female exposure is more prominent. In addition, we find considerable heterogeneity and in particular, education and age of first marriage appear to

[4] Prior to the reform no sector had tariff rates of 0 percent. After implementation, around 50 percent of (six-digit level) industries became fully unprotected by tariffs.

[5] Our identification strategy resembles Autor, et al. (2018) and Shenhav (2016). The former employs gender-specific components of U.S. labor demand shocks coming from competition with China to explore whether changes in relative economic outcomes of young men versus young women affected marriage and fertility during 1990-2014. The latter exploits gender-specific Bartik shocks and gender differences in occupational choice to test their impact on relative gender earnings in U.S. states. In contrast to Bartik shocks, which are typically exploited as local labor demand shifters our identification variation comes from the reduction in tariffs.

be key variables that correlate with our findings. Women with less education and thus likely with less bargaining power appear to suffer from more physical intimate partner violence. Similarly, the impact of trade liberalization seems to be larger among women that were aged 19 or less when they first married. We are confident that our empirical findings are solid, as we undertook a long battery of tests in order to confirm whether our methods and results are robust. In particular, we find that intimate partner violence is not correlated with post-reform tariff changes, which is consistent with no pre-existing trends. In addition, we apply placebo tests that further support our findings by using a pre-reform measure of intimate partner violence. We also test whether the initial measures we used to construct our tariff exposure indices are exogenous conditional on observables. Similarly, we address the possibility of conflating short- and long-term effects. Moreover, we study whether our findings are driven by inward and outward migration patterns. We also apply permutation tests. Finally, we exploit the fact that trade liberalization does not only impact the prices of output goods, but also of intermediate inputs and consider a set of empirical tests to take their impact into account.

The rest of the paper is organized as follows. Section 2 reviews the literature. Section 3 provides institutional background. Section 4 describes the data. Section 5 explains our empirical strategy. Section 6 shows our findings, while Section 7 shows our robustness tests. Section 8 describes how our findings vary across sub-samples. Section 9 explores the association between input tariffs and violence. Finally, Section 10 concludes.

Brief Review of the Literature

The most common approach that explains the determinants of both emotional and physical intimate partner violence relates to the ability and extent to which females can

credibly change the rules of the game within the household. In particular, an important theory is the so-called "bargaining power theory" whereby both emotional and physical violence suffered by women may be understood as a function of their bargaining power with respect to men. When women improve their economic situation, they may end up having a better position within the household in terms of bargaining power, and as a result, experience less violence (Aizer, 2010; Anderberg et al., 2016). But, if the threat of ending a relationship with a partner is not credible enough, they may end up suffering from more intimate partner violence when trying to gain independence (Eswaran and Malhotra, 2011; Tauchen, et al., 1991).[6] In addition, a complementary theory that explains intimate partner violence is referred to as the "male backlash theory." It argues that changes in the economic relationship between males and females play a crucial role so that when females improve their income position within the household they may end up suffering from violence from their partners, as a reaction of males to a perceived threat in terms of a potential change in the status quo. According to this theory, violence occurs as a way for males to assert their power and recover their perceived loss in authority (Macmillan and Gartner, 1999). There is some evidence that appears to support the views above. For instance, Heath (2014) shows that the association between financial independence and intimate partner violence depend on the age at marriage and education. Similarly, Hidrobo and Fernald (2013) show that intimate partner violence is conditional on the relative education of females. Also, Panda and Agarwal (2005) show that the effect of increasing financial independence may be conditional on the initial status of females in terms of

[6] It is sometimes argued that the applicability of this theory may be somewhat limited in the context of developing countries, which tend to be more male-dominated and where social norms may not allow for females to simply withdraw from abusive marriages or civil unions (Folbre, 1994; Mabsout and Van Staveren, 2010).

property rights.

Related theories that explain intimate partner violence place less emphasis on power balance issues within the household and more on the role of actual physical interaction of males and females. One well-known idea is the theory of "exposure reduction" by which intimate partner violence is a function of how actually exposed women are with respect to men within the household. If both partners are employed, they spend less time together, which in this context implies that the probability that intimate partner violence will occur decreases (Dugan, et al., 1999). Also, actual physical interaction may serve as a way to relieve frustration caused by adverse shocks and exacerbate violence (Tauchen, et al., 1991). This view is also related to the idea that violence may occur when individuals attempt to achieve certain goals and other individuals interrupt them, which provokes potential violence in order to restore the previous situation (Feshbach, 1964).

From the perspective of our research, changes in labor markets conditions may have direct consequences on the incidence of intimate partner violence, as they help shape the distribution of bargaining power between partners. Depending on the specific industry, changes in labor markets may increase the economic rewards of one gender over the other and may exacerbate or lessen stressful conditions or change the time that partners physically interact with each other. Hence, trade liberalization may have a direct bearing on the relationship dynamics between partners in the household. For instance, the exact direction of changes in employment that results from trade liberalization is unclear. Whereas some researchers argue that regional capital adjustments and agglomeration economies may exacerbate long-run effects of trade liberalization on local labor markets,

others explain that information asymmetries may also play a role. While some researchers find that trade liberalization may reduce employment (e.g., Dix-Caneiro and Kovak, 2017; Meñezes-Filho and Muendler, 2011; Gaddis and Pieters, 2017) some others find that employment may increase (Hasan et al., 2012). Furthermore, trade liberalization may also impact the dynamics on employment flows across sectors and formality categories (Ferreira, et al., 2010) as well as the skill premia (Dix-Carneiro and Kovak, 2015; Amiti and Cameron, 2012; others).

There is some evidence that shows that trade liberalization may have gendered impacts on labor markets. Trade liberalization may increase competition (Gaddis and Pieters, 2017), which may reduce taste-based discrimination (Black and Brainerd, 2004; Ederington, et al., 2009). Also, it may foster skill-biased technical change (Juhn, et al., 2014). If trade liberalization decreases the skill premium and females are more skilled, then their bargaining power may decrease, which may affect the incidence of intimate partner violence. In addition, trade openness may help reallocate sectorial employment, which may have gendered implications if male and female workers are viewed as imperfect substitutes (Gaddis and Pieters, 2017; Galor and Weil, 1996; Do, et al., 2016).

Institutional Background

During the 2000s, the Peruvian economy enjoyed a very favorable external environment due to a sharp increase in commodity prices. Between 2000 and 2010, exports grew from around US$8 million to more than US$40 million and the gross domestic product per capita increased by fifty percent. The aim was to further take advantage of the favorable environment by seeking new international markets and signing free trade agreements with other countries. Interestingly, the administration at the time pursued a

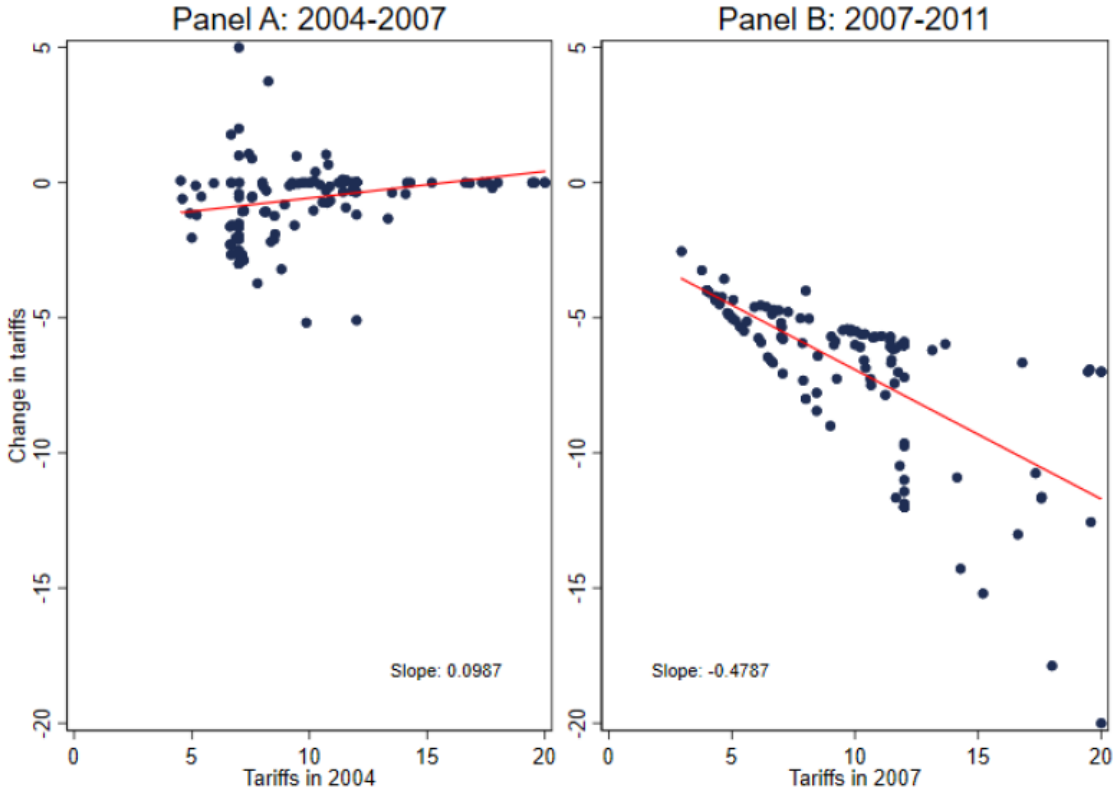
drastic and unexpected policy change with little warning. Tariffs were cut drastically and unilaterally between 2007-2008 and again between 2010-2011, albeit somewhat less dramatically. The private sector was not consulted about these actions and these actions took practically all economic agents by surprise. Even the government's own documents describe pursuing a policy of slowly reducing tariffs as a very important strategy so as to not compromise in any way the bargaining position of the Peruvian government when negotiating free trade agreements with other nations, a crucial policy objective at the time.[7]

The main reduction in tariffs occurred during 2007 and was rather large. It included nearly 5,000 different products and eliminated most tariffs and related fees. Unsurprisingly, the sectors that were the most protected were also the most affected by the reduction in tariffs. This can be seen in Figure 1. After the reform, tariff rates were simplified into two categories, 9 percent or 17 percent, and around half of the six-digit level products were assigned no tariff at all.[8]

[7] In 2011 José Luis Silva Martinot, a previous head of the most important association of exporters, said: "with free trade agreements, tariffs were going to be cut after 10 to 17 years, others after a shorter time and some others not at all. However, in the end they were all totally eliminated." Similarly, Eduardo Farah, ex-head of the National Society of Industries, said: "with these measures, the country loses bargaining power for the negotiation of future free trade agreements."

[8] Prior to the 2007 trade reform, not a single product was tariff-free.

Figure 1: Correlation between Initial Tariffs and Tariff Changes

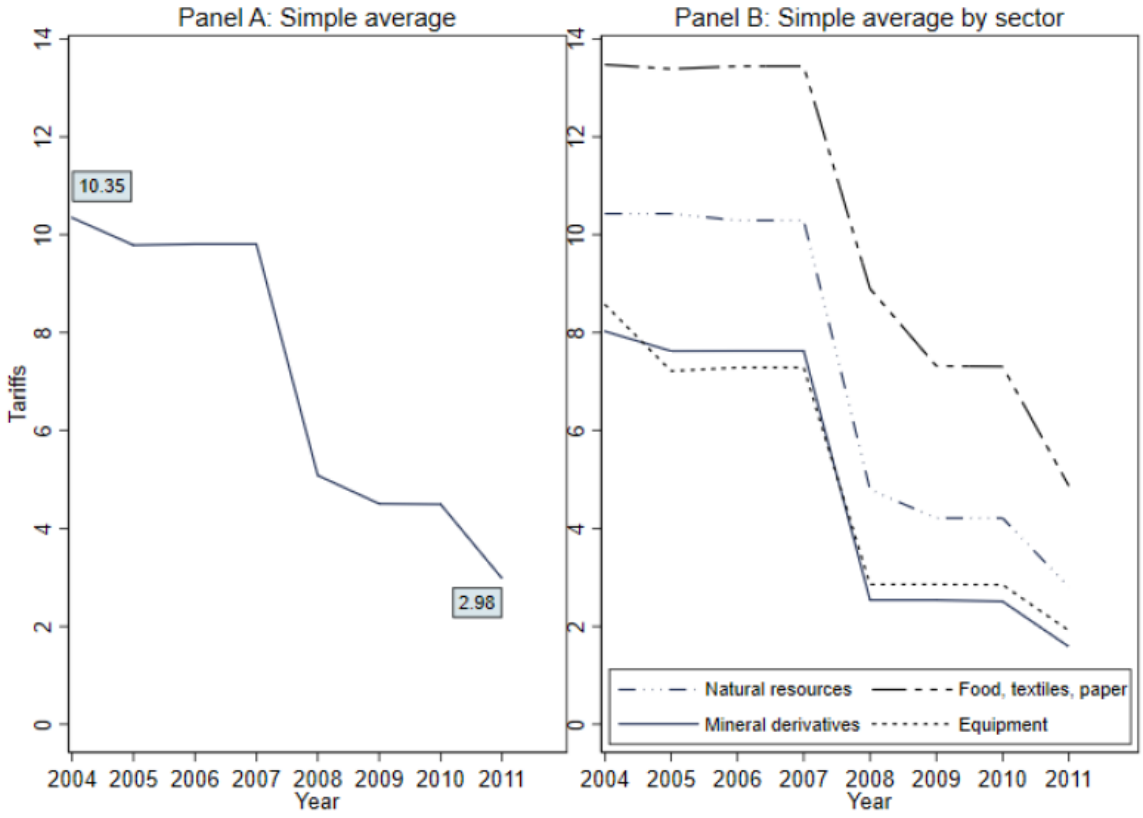


Source: World Bank TRAINS and World Bank's concordance tables

Notes: Tariffs and tariff changes were computed at the industry level using ISIC3 codes. Originally, industries were coded based on the Trade Classification Harmonized System (HS). We translated this classification into the International Standard Industrial Classification (ISIC3) using the concordance tables available online.

A second wave of tariff reductions occurred in 2010 and 2011 and while still significant, this wave was less dramatic than the first one. Figure 2 illustrates these tariff reduction waves. Panel A shows the evolution of average tariffs. On average, they decreased from 10.35 percent in 2004 to 2.98 percent in 2011. Panel B shows tariff reductions by sectors. As described in detail in the next section, our identifying variation is provided by: the differences in the timing of introduction of tariff reductions, including the magnitude of reduction, and with the district-level variation in industrial employment composition, taking into account the differential intensity in the use of male and female workers by industry.

Figure 2: MFN Tariffs, 2004-2011



Source: World Bank TRAINS and World Bank's concordance tables

Notes: Tariffs and tariff changes were computed at the industry level using ISIC3 codes. Originally, industries were coded based on the Trade Classification Harmonized System (HS). We translated this classification into the International Standard Industrial Classification (ISIC3) using the concordance tables available online.

Data

The data for this research come from several sources. We first construct a measure of exposure to trade liberalization. We use employment shares by industry per district to weigh how tariff reductions may impact each district. We employ the 1993 Peruvian Household Census[9] along with data on Most-Favored-Nation (MFN) tariffs reported by the World Bank at the six-digit level of the Trade Classification Harmonized System (HS).[10]

[9] Whereas another National Census was carried out in 2007, just when the reform was starting, we avoid using it given endogeneity concerns. In this regard, Jaeger, et al., (2018) argue that lagging the base period used to weigh tariff cuts and construct the can help with identification by minimizing the correlation between tariff changes and current demand shocks.

[10] According to the World Bank, "in current usage, MFN tariffs are what countries promise to impose on imports from

We match the industry codes reported by the World Bank (i.e. HS2007) to the industry code used in the Census (i.e. ISIC3).[11] We use data from the 1993 Census microfiles as these are the closest available to the first wave of tariff reductions in 2007. The data on intimate partner violence come from the Demographic and Health Surveys (DHS), which were conducted by the Ministry of Health between 2004 and 2011.[12] The data consist of women aged 15 to 49 who are asked if they had suffered from emotional or physical violence by a partner by any of various different means. In general, the survey contains detailed information on the characteristics of females and the incidence of intimate partner violence.

We construct a dummy variable that accounts for physical intimate partner violence and focus on females that are in a relationship, only.[13] We also compute a dummy describing emotional intimate partner violence.[14] In addition, we also employ several demographic variables available in the survey. Finally, other data collected are exports (aggregated to the 6-digit level) and foreign direct investment by industry (aggregated to the 2-digit level, the highest available), which we use as controls. Appendix B provides definitions and description of the data.

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other members of the World Trade Organization (WTO), unless the country is part of a preferential trade agreement. In practice, MFN rates are the highest and most restrictive that WTO members charge one another."

[11] We employ the concordance table provided by the World Bank here.

[12] In the survey that we employ, physical partner violence refers to women reporting that at least one act of physical violence took place during the 12 months preceding the interview, which is a stricter definition to the one provided by the Centers for Disease Control (see Footnote 2).

[13] In our survey data, a "couple" refers to couple consisting of a male and a female only.

[14] According to the World Health Organization, it is possible to differentiate between emotional intimate partner violence (i.e. husband humiliating his wife, threatening to harm her or to take away her children) and controlling behavior (i.e. husband trying to limit his wife's contact with her family or friends, being jealous or insisting on knowing where she is). Furthermore, whereas it is conceivable to find instances of physical intimate partner violence among adults from females to males, the overwhelming majority of cases in our country of study, Peru, occur from males to females (96 percent according to the Ministry of Women and Vulnerable Populations (2019)). In addition, whereas most of our findings using emotional intimate partner violence and even controlling behavior are consistent with the results we find for physical intimate partner violence, it is true that the first two may occur with more frequency from females to males. These results are available upon request.

Table 1 provides summary statistics. The average incidence of physical and emotional intimate partner violence in Peru is 38.7 percent and 30.2 percent, respectively. While in 2004, 43.0 percent and 32.9 percent of women reported to have suffered corresponding episodes of physical and emotional violence by 2011 these figures were of 37.6 percent and 29.0 percent, respectively. In addition, we find that other variables show a pattern that is consistent with the literature on intimate partner violence such as the age of the brides, the age difference with the partner, and the education gap between couples.[15]

Table 1: Intimate Partner Violence - Summary Statistics

Panel A: Pooled sample

Variable	Number Obs.	Mean	Std. Dev	Min	Max
Physical violence	78,276	0.387	0.487	0	1
Emotional violence	78,276	0.302	0.459	0	1
Age at marriage	78,276	19.951	4.600	10	48
Age	78,276	33.604	8.189	15	49
Partner's age	78,263	37.653	9.354	15	96
Years of educ. (YoE)	78,276	8.182	4.540	0	17
Partner's YoE	77,925	9.122	3.956	0	17
HH. head is woman	78,276	0.079	0.270	0	1
Non-Spanish	78,269	0.150	0.358	0	1
HH. size	78,276	4.856	1.818	1	19
Altitude	78,276	1517	1503	0	5037

Panel B: Average violence by survey year

Variable	2004	2006	2008	2010	2011
Physical violence	0.430	0.402	0.380	0.367	0.376
Emotional violence	0.329	0.319	0.298	0.279	0.290

Notes: The sample consists of women that were in a relationship when they were surveyed and that report whether they have ever suffered physical intimate partner violence or not. Definitions of variables are described in Appendix B.

Empirical Strategy

Our identification approach is analogous to other research on trade liberalization among others Kis-Katos and Sparrow (2015), Gaddis and Pieters (2017), Dix-Carneiro and

[15] See for instance Jensen and Thornton, (2003); Yount, et al. (2018), Mabsout and van Staveren (2010), Heath (2014); Aizer, (2010), Fiedberg and Webb (2006), Hidrobo and Fernald (2013), among others.

Kovak (2015, 2017) where the main idea is to exploit the distribution of *overall* employment within each district and across industrial sectors in order to measure how households are impacted by tariff changes. Unlike previous research, we exploit the pre-reform composition of *male* and *female* employment in addition to the time-series variation stemming from tariff changes.[16] As shown in Table A.1 the labor force is predominantly male in a significant number of industries and it is predominantly female in others.

As described above, this is relevant as trade liberalization may end up impacting males and females differently. As a simple example, consider the case where male labor is mostly specialized in mining when tariffs are reduced in a particular district. If this occurs the treated male workers will become relatively more exposed to trade liberalization with respect to unexposed male workers in other districts. Likewise, if female labor in a particular district specializes in textiles when tariffs are reduced, they will become relatively more exposed to trade liberalization with respect to unexposed female workers in other districts. Our differences-in-differences strategy is closest to Autor, et al., (2018) and Shenhav (2016). The former employs gender-specific components of the United States large labor demand shocks coming from competition with China to explore whether changes in relative economic outcomes of young men versus young women affect marriage and fertility. The latter exploits gender-specific Bartik shocks and gender differences in occupational choices to test their impact on relative gender earnings in the United States.[17]

We compute two sex-specific measures of exposure to trade liberalization for each

[16] Notice that the share of female workers per industry in 1993 is uncorrelated with tariff reductions in the period 2004-2011, as shown in Figure A.1.
[17] See also Chauvin, 2018.

district d and year t, which to simplify we simply call "male exposure" and "female exposure":

$$TrTariff^{M}_{d,t} = \sum_{i} \frac{L^{M}_{1993,i,d}}{L^{M}_{1993,d}} \times tariff_{i,t} \quad (1)$$

$$TrTariff^{F}_{d,t} = \sum_{i} \frac{L^{F}_{1993,i,d}}{L^{F}_{1993,d}} \times tariff_{i,t} \quad (2)$$

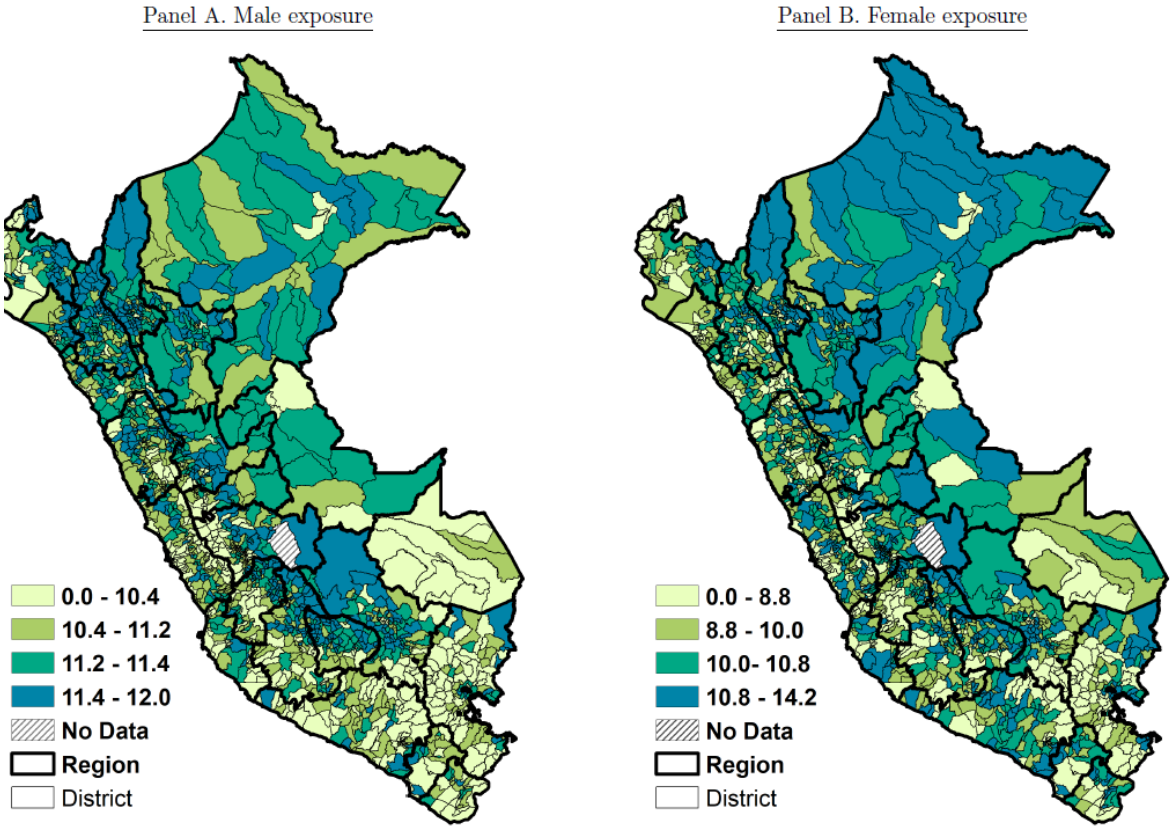
where $L^{G}_{1993,i,d}$ is the number of workers of gender $G = \{M, F\}$,[18] employed in sector i in district d in 1993, $L^{G}_{1993,d}$ is the district d's total number of workers of gender $G = \{M, F\}$ in 1993, and $tariff_{i,t}$ is the Most-Favored-Nation (MFN) tariff of industry i at year t.

Given that mechanically tariffs are assigned zeros in the non-tradable sector, districts with larger non-tradable sectors will automatically yield a lower value for both $TrTariff^{M}_{d,t}$ and $TrTariff^{F}_{d,t}$. If the size of the non-tradable sector in 1993 is correlated with any un- observed determinant of current intimate partner violence within households the resulting coefficients may be biased. For example, the size of the non-tradable sector may be correlated with the female labor force participation, which in turn may be correlated with intimate partner violence within households (Gaddis and Pieters, 2017; Aizer, 2010). Given the above, the evidence presented in this research fully excludes the non-tradable sector, which has become standard practice in the literature (Kovac, 2013).[19] Figure 3 graphically shows the variation of tariff reductions by district. The darker the district, the deeper the tariff reduction faced.

[18] M and F stands for male and female, respectively.

[19] We also exclude these four 4-digit ISIC3 industry codes, 1110, 0111, 0112, and 0121, which account for extraction of crude petroleum and natural gas; growing of cereals and other crops; growing of vegetables, horticultural specialties and nursery products; and farming of cattle, sheep, goats, horses, asses, mules and hinnies. We do this because tariff changes in these sectors were not parallel to tariffs changes in other sectors during the period prior to the first wave of tariff reductions. It should be said that we do not find any significant differences in our results when including these industries.

Figure 3: Reductions in tariff exposure, 2004-2011



Source: World Bank TRAINS, World Bank's concordance tables and the 1993 Population and Household Census.

Based on our approach above, we estimate the following reduced form:

$$y_{j,d,t} = \alpha - \beta_1 TrTariff_{d,t}^{M} - \beta_2 TrTariff_{d,t}^{W} + \alpha_d + \alpha_t + \alpha_{r,t} + [\gamma_1' X_{j,d,t} + \gamma_2' Z_{d,t} + f(W_{d,1993}, trend, \gamma_3)] + \varepsilon_{j,d,t} \quad (3)$$

where $y_{j,d,t}$ is a dummy that takes the value of 1 if woman j reports to have ever been attacked by her partner. α_d, α_t, and $\alpha_{r,t}$ are respectively district, year, and region-year fixed effects. α_d capture time-invariant heterogeneity at the district level, while αt controls for macroeconomic shocks affecting the country as a whole. We add region-year fixed effects to allow the impact of macroeconomic shocks to vary across the 25 Peruvian regions.[20] It is

[20] Peru as of 2018 was divided into 25 regions, 196 provinces and 1874 districts.

important to control for region-year fixed effects, because there may be unobserved characteristics, correlated with the 1993 initial conditions, capable of predicting labor market developments and therefore correlated with changes in intimate partner violence. Hence, by including region-year fixed effects we attenuate potential bias produced by the dynamics stemming from these initial conditions (Goldsmith-Pinkham, Sorkin and Swift, 2018). As an alternative strategy to control for these unobserved dynamics, we also include a set of initial conditions (i.e. $W_{d,1993}$) interacted with time trends as it is usually done in the literature (e.g. Edmonds et al, 2009, 2010; Gaddis and Pieters, 2017; Hasan et al, 2012; Kis-Katos and Sparrow, 2015; and Topalova, 2010.).

We also consider a set of time-varying individual and household level covariates, $X_{j,d,t}$, which follows the standard literature of determinants of intimate partner violence (Jensen and Thornton, 2003; Yount, et al., 2018; Mabsout and van Staveren, 2010; Heath, 2014; Aizer, 2010; Hidrobo and Fernald, 2013; among others). These covariates consist of the woman *j*'s age and years of education, her partner's age and years of education, her age when she first married, a dummy whether she speaks Spanish, the household's size, and the sex of the household head. In addition, we include the altitude at where the household is located.[21] Finally, $Z_{d,t}$ is a set of time-varying district level variables that may be correlated with $TrTariff^{M}_{d,t}$ and $TrTariff^{F}_{d,t}$. This set consists of a measure of exposure to exports and to foreign direct investments as well as a measure of exposure to input tariffs. During the period of analysis exports and foreign direct investment grow exponentially and heterogeneously across sectors. Just as with $TrTariff^{M}_{d,t}$ and $TrTariff^{F}_{d,t}$, this growth may

[21] In Peru, there is a negative correlation between the altitude and access to health services and economic development.

affect a district in a particular way depending on how specialized the district is with respect to an industry. We also include a measure of exposure to input tariffs, as tariff changes may influence households not only through final output prices, but also through intermediate input prices (Edmonds, et al., 2010).[22]

The male and female exposure coefficients, β_1 and β_2, are our parameters of interest. They measure the impact of a decrease in tariff protection in industries that are more male-predominant or female-predominant on the probability of suffering increased intimate partner violence.[23] We identify β_1 and β_2 by comparing more exposed to less exposed districts hence as with any differences-in-differences framework we are not identifying the effect of trade liberalization as a whole, but its differential effect on the more exposed districts (Topalova, 2005, 2010). The assumption implicitly made in order to estimate β_k (where k is either 1 or 2) is that any unobserved district-specific time varying shock affecting the chances of suffering intimate partner violence is uncorrelated with any change in our two measures of trade exposure over time. Since our measures of trade exposure, $TrTariff^M_{d,t}$ and $TrTariff^F_{d,t}$, are simply the interaction between the 1993 initial industrial composition with the national level tariff changes, the only source of bias comes from differential time-trends in intimate partner violence correlated with both sources of variation simultaneously (Topalova, 2005, 2010).

Baseline Results

Table 2 presents our findings. Each column reports a different version of equation

[22] We follow Edmonds, et al. (2010) and use the 1993 Peruvian national input-output table, the 1993 national census and the World Bank's data on output tariffs to construct this measure. Please refer to Appendix B.
[23] Note that in equation (3) we have multiplied $TrTariff^M_{d,t}$ and $TrTariff^F_{d,t}$ by minus one to facilitate the reading of our tables. Recall that we are interested in the reduction of tariff protection, not in its increase.

(3). All regressions include district and year fixed effects as well as a set of region-year fixed effects, as was explained above. Standard errors are clustered at the district level. Column 1 shows that a unit decrease in the measure of male exposure increases violence physical in 1.78 percentage points. On the other hand, tariff cuts on industries that are predominantly female do not trigger any change in intimate partner violence. If anything, they reduce violence in 1.06 percentage points. These results are robust to the inclusion of individual and household level covariates (see column 2). When we control for time-varying controls at the district level (i.e. FDI, exports and input tariffs), the coefficient on male exposure increases in about 40 percent compared to that of column 2, and it is significant at the 1 percent level (see column 3). In contrast, the coefficient on female exposure remains statistically insignificant and considerably smaller in magnitude than the coefficient on male exposure.[24]

We believe that controlling for input tariffs increases the magnitude of the coefficient on male exposure, because tariff changes may influence households through final output prices, but also through intermediate input prices (Edmonds et al., 2010) and these impacts may be opposite in coefficient sign. When we ignore input tariffs, our measure of male exposure is conflating both effects. For instance, lower output tariffs may decrease wages as industries lose their protection. In contrast, lower input tariffs may increase wages as they may enhance productivity through access to cheaper inputs. To the extent that these opposing effects in the labor markets are transmitted to within-household

[24] We also estimate the impact of male and female exposure on emotional intimate partner violence. Our results are similar, as male exposure to trade liberalization increases emotional violence in similar magnitudes, whereas female exposure tends to decrease it. The main difference is that the negative sign in the coefficient of female exposure is larger and more robust than for physical violence. Results are shown in Table A.2. For controlling behavior, the results are also similar, but for the sake of space are not reported. They are available upon request.

dynamics, we should observe opposite effects of output and input tariffs on the incidence of intimate partner violence. Similar conclusions have been reached in the literature for other outcomes. For example, Kis-Katos and Sparrow (2015) find that decreases in output tariffs raise poverty, whereas decreases in input tariffs have the opposite effect. Similarly, Amiti and Cameron (2012) show that input tariffs reductions contributed to the closure of the industrial skill wage gap in Indonesia, whereas Dix-Carneiro and Kovak (2015) show that cuts in output tariffs modestly widened the skill wage gap in Brazil.[25]

Table 2: The Effect of Trade Liberalization on Physical Intimate Partner Violence I

	Has suffered from physical intimate partner violence			
	(1)	**(2)**	**(3)**	**(4)**
$TrTariffs^M$	0.0178	0.0202	0.0344	
	(0.0087)**	(0.0084)**	(0.0106)***	
$TrTariffs^F$	-0.0106	-0.0126	-0.0118	
	(0.0081)	(0.0080)	(0.0089)	
TrTariffs				0.0203
				(0.0089)**
Mean dep. var.	0.387	0.388	0.388	0.388
N. districts	1066	1066	1066	1066
Adjusted R^2	0.0506	0.0743	0.0744	0.0743
N	78276	77906	77906	77906
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates		X	X	X
District-level covariates			X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Finally, in order to explore the effect of the overall exposure to lower tariffs, net of input tariffs, in column (4) we estimate the relationship between intimate partner violence

[25] We explore the issue of input tariffs in detail below.

and a general measure of tariff exposure. We construct this measure in a similar fashion as before. The main difference is that here we employ the initial shares of the overall industrial employment, as it is usually done in the literature (Please refer to Appendix B). We show that for each unit decrease in the measure of tariff exposure, intimate partner violence increases by 2.03 percentage points relative to districts that were less intensively exposed.

These effects are of economic significance. According to our preferred estimates—those from column 3—a unit decrease in male exposure increases physical intimate partner violence by 3.44 percentage points. This implies that physical intimate partner violence increases by 36.59 percentage points in a district experiencing the average tariff cut on industries that are predominantly male (i.e. a change of 10.64 units between 2004 and 2011), relative to a district experiencing no change at all.[26] This is considerably larger than the national decline in intimate partner violence of about 6 percentage points (see Table 1). In the case of female exposure, the decrease in physical intimate partner violence in a district experiencing the average tariff cut ranges between 10.34 and 12.29 percentage points, albeit these decreases are not statistically different from zero.[27]

As an alternative scale of physical intimate partner violence, we compute a measure of intensity of physical violence by simply adding up all the corresponding dummy variables that refer to an episode of physical intimate partner violence in our survey. This

[26] Our lowest estimated increase in physical intimate partner violence is 18.83 percentage points, which comes from Column 1.

[27] In an on-going research piece, we calculate wages for males and females between 2004-2011 and find that a unit decrease in male exposure widens the gender wage gap by 1.8 percentage points, while a unit decrease in female exposure narrows it by 1.7 percentage points. We find qualitatively similar developments on employment rates. These results show that male and female exposure impact labor markets and household dynamics and ultimately intimate partner violence. These findings are available upon request.

measure of intensity goes from zero to nine. In addition, we also use principal components with these variables and calculate its components. We run specification (3) using all these as dependent variables. Results are shown in Table 3. We find similar results as those shown above.

Table 3: The Effect of Trade Liberalization on Physical Intimate Partner Violence II

	Violence intensity		1st principal component	
	(1)	**(2)**	**(3)**	**(4)**
$TrTariffs^M$	0.0588	0.1048	0.0216	0.0389
	(0.0344)*	(0.0460)**	(0.0126)*	(0.0168)**
$TrTariffs^F$	-0.0202	-0.0190	-0.0072	-0.0067
	(0.0331)	(0.0330)	(0.0121)	(0.0121)
Mean dep. var.	1.119	1.119	0.422	0.422
N. districts	1066	1066	1066	1066
Adjusted R^2	0.0975	0.0976	0.0992	0.0993
N	77906	77906	77906	77906
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates	X	X	X	X
District-level covariates		X		X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Robustness and Threats to Identification

Falsification and Placebo Tests for Pre-Existing Trends

Whereas recent research shows that current productivity of industries may not predict future tariffs (Baldarrago and Salinas, 2017) the possibility that pre-existing trends may be correlated with changes in the outcome being studied still needs to be addressed, as the government may endogenously protect certain industries depending on productivity.

We follow Topalova (2010) and test whether pre-existing trends in intimate partner violence are correlated with post-reform tariff changes. If tariff cuts are correlated with pre-existing trends in intimate partner violence, the coefficient β_1 should be similar whether we use pre- or post-reform data.[28] Since the first wave of tariffs cuts occurred between

[28] The same applies for β_2, although we focus on β_1 because β_2 was not statistically different from zero.

2007 and 2008 and the second wave occurred between 2010 and 2011 we use the period 2004 to 2007 as pre-reform data. We run two regressions. The first one links the 2004-2007 incidence of intimate partner violence using 2007-2010 tariff data to take advantage of the first wave of tariff reductions. The second one relates intimate partner violence with the 2008-2011 tariff data to take advantage of the second wave.[29]

The results from these regressions are reported in the first two columns in Table 4. We estimate our preferred specification, which includes individual- and district-level covariates. We can compare these results to those under column 3 in Table 2. The estimated β_1 in columns 1 and 2 is around forty and five times smaller than our baseline estimation, respectively. Both are statistically indistinguishable from zero.

We also run a placebo test exploiting one particular question from our survey data: "Has your father ever beaten your mother?" Since women from the survey are adults, this episode of physical intimate partner violence refers to a past event, long before tariffs were first cut. We can think of this variable as a pre-reform measure of intimate partner violence at the household level. Then, we run a regression between this variable and our measures of tariff cuts exposure. We report results in column 3 in Table 4. We find an estimated coefficient that is around five times smaller compared to that of our baseline estimation and statistically indistinguishable from zero, which is reassuring of our identification strategy, especially considering that this question is highly correlated with intimate partner violence.

[29] In order to clarify, in the first regression we match 2004 data on intimate partner violence with 2007 tariffs, 2005 data with 2008 tariffs, 2006 data with 2009 tariffs, and 2007 data with 2010 tariffs. For the second regression, we match the data on intimate partner violence from 2004 with 2008 tariffs, 2005 data with 2009 tariffs, and so on.

Table 4: Falsification and Placebo Tests, Physical Intimate Partner Violence

	Falsification: PIPV (2004-2007)		Placebo: PIPV between parents
	(1)	**(2)**	**(3)**
TrTariffs^{M} (2007-2010)	0.0007		
	(0.0070)		
TrTariffs^{F} (2007-2010)	0.0014		
	(0.0054)		
TrTariffs^{M} (2008-2011)		0.0062	
		(0.0079)	
TrTariffs^{F} (2008-2011)		0.0004	
		(0.0066)	
TrTariffs^{M} (2004-2011)		0.0075	
			(0.0106)
TrTariffs^{F} (2004-2011)			-0.0131
			(0.0080)
Mean dep. var.	0.405	0.405	0.482
N. districts	595	595	1066
Adjusted R^2	0.103	0.103	0.0510
N	29330	29330	73733
District and year FE	X	X	X
Region-year FE	X	X	X
Individual-level covariates	X	X	X
District-level covariates	X	X	X

Notes: PIPV stands for physical intimate partner violence

Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Sensitivity to Initial Conditions

Whereas we employ a more general estimator it may be viewed as part of the family of shift-share identification instruments and in particular Bartik estimators. Goldsmith-Pinkhman, et al., (2019) show that there are two set of alternate identification conditions. If the number of industries is fixed, we require, conditional on observables, exogenous initial employment shares—those employed to construct the Bartik shock. This

is because the two-stage least square estimator is equivalent to a generalized method of moments estimator using the initial shares as instruments.[30] When the number of industries goes to infinity along with the number of locations, what matters is whether the tariff cuts are uncorrelated with the bias stemming from the initial shares. If this is the case, the presence of a large number of shocks causes the bias to average out. This latter identification condition seems to be more relevant in our context since tariff cuts are a consequence of an unexpected and massive trade reform and are not correlated with pre-existing trends of our outcome of interest.

We now assume that tariff cuts are correlated with the potential bias generated by the initial shares, even though it seems to not be the case. We test whether the initial shares used to construct our measures of tariff exposure are exogenous conditional on observables. If so, controlling for different sets of initial conditions should not affect our estimates. We run specification (3) testing for four different sets of initial conditions, $W_{d,1993}$, interacted with quadratic linear trends, instead of including region-year fixed effects. The first set of initial conditions consists of the share of employment destined to agriculture and fishing, mining, manufacture, and construction as well as the number of female and male workers. The aim is to capture the broad employment structure of each district, which may be correlated with household dynamics. The second set of variables consists of the share of individuals with complete primary, high school, and post-secondary education. We consider this set as Goldsmith-Pinkhman, et al., (2019) show that in Autor, et al. (2013) the industries driving identification are located in more educated areas.

[30] One caveat is that not every share should be exogenous. Goldsmith-Pinkhman, et al., (2019) show that in practice just a small number of industries tends to account for a large portion of the identifying variation.

The third set of variables is related to household dynamics and social norms: the share of individuals that live together, the share of individuals that are Catholics, and the share of individuals that are Evangelists, and where other beliefs is the omitted category. To control for social norms at the district level we use Spanish speakers, female, younger than 18, aged 18 to 40, aged 40 to 65; older than 65 is the omitted category. Finally, the fourth set is linked to the structure of the labor market. These are the share of employment, the share of female workers and the share of workers employed in small and medium firms. Table 5 describes the results, which are robust in all specifications. This suggests that either our specification is already partialling out the potential bias generated by the initial shares or tariff cuts are uncorrelated with this bias.[31]

Table 5: Sensitivity to Initial Conditions

	Has suffered from physical intimate partner violence				
	(1)	**(2)**	**(3)**	**(4)**	**(5)**
$TrTariffs^M$	0.0260	0.0242	0.0247	0.0259	0.0285
	(0.0081)***	(0.0071)***	(0.0073)***	(0.0074)***	(0.0082)***
$TrTariffs^F$	-0.0010	0.0036	0.0021	0.0010	0.0006
	(0.0078)	(0.0076)	(0.0075)	(0.0075)	(0.0076)
Mean dep. var.	0.388	0.388	0.388	0.388	0.388
N. districts	1066	1066	1066	1066	1066
Adjusted R^2	0.0736	0.0735	0.0740	0.0735	0.0743
N	77906	77906	77906	77906	77906
District and year FE	X	X	X	X	X
Region-year FE					
Individual-level covariates	X	X	X	X	X
District-level covariates	X	X	X	X	X
Initial conditions interacted with quadratic trends:					
Initial shares	X				X
Education		X			X
Household			X		X
Labor				X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

Initial shares: This set includes the share of employment destined to agriculture and fishing, mining, manufacture and construction. It also considers the total number of female and male workers. Education: the

[31] The findings are similar if we consider year dummies instead of quadratic trends.

share of individuals with complete primary, with complete high-school, and with complete post-secondary. Household: this set includes the share of individuals that live together, the share of individuals that are Catholics, and the share of individuals that are Evangelical Christians. We also consider the share of the population that speaks Spanish, the share of the population that is female, and the share of people younger than 18, aged between 18 and 40, and aged between 40 and 65. Labor: the share of overall employment, the share of workers that are female, and the share of workers employed in small and medium enterprises. We interact these variables with quadratic trends. **The set of individual-level covariates** includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located**. The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Conflating Past and Current Shocks

In recent years shift-share instruments have been criticized. Jaeger, et al., (2018) argue that if it takes time for markets to adjust, shift-share instruments may conflate short-term responses and long-term effects. In this situation they suggest adding lagged measures of the instrument. However, to be able distinguish between short- and long-term effects, the variation of the instrument across time periods should be independent enough. In our context, the composition of industries affected by tariff reductions and their magnitudes vary across time, which is reflected in the fact the autocorrelation across districts once we condition on district fixed effects is low. Following Jaeger, et al. (2018) we calculate the serial correlation of the first difference of our variables of interest. Results are shown in Table A.3 and Table A.4. Compared to Jaeger et al. (2018) our serial correlations seem to be low. As such, we control for dynamic responses by adding lagged measures of exposure and include five lags.[32] Results are shown in Table 6. Overall, β_1, remains almost unchanged if we add these lags. At most, it increases by 20 percent.

[32] Jaeger, et al., (2018) estimate mid-to-long-term impacts of immigration inflows employing data from different decades. We focus on short-term impacts as we exploit year-to-year changes It is reasonable to expect more persistence as it takes time for markets to adjust.

Table 6: Controlling for Previous Shocks

	Has suffered from physical intimate partner violence					
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
-TrTariffs^M	0.0334	0.0409	0.0416	0.0401	0.0399	0.0397
	(0.0106)***	(0.0115)***	(0.0116)***	(0.0117)***	(0.0117)***	(0.0117)***
-TrTariffs^F	-0.0118	-0.0115	-0.0121	-0.0107	-0.0107	-0.0105
	(0.0082)	(0.0092)	(0.0093)	(0.0093)	(0.0093)	(0.0093)
L1. -TrTariffs^M		-0.0106	-0.0086	-0.0087	-0.0088	-0.0089
		(0.0089)	(0.0091)	(0.0091)	(0.0091)	(0.0093)
L1. -TrTariffs^F		-0.0028	-0.0047	-0.0050	-0.0049	-0.0049
		(0.0088)	(0.0091)	(0.0091)	(0.0091)	(0.0092)
L2. -TrTariffs^M			-0.0078	-0.0113	-0.0120	-0.0121
			(0.0097)	(0.0115)	(0.0120)	(0.0120)
L2. -TrTariffs^F			0.0055	0.0078	0.0079	0.0080
			(0.0092)	(0.0097)	(0.0098)	(0.0097)
L3. -TrTariffs^M				-0.0228	-0.0270	-0.0276
				(0.0408)	(0.0454)	(0.0457)
L3. -TrTariffs^F				0.0276	0.0287	0.0301
				(0.0335)	(0.0338)	(0.0340)
L4. -TrTariffs^M					0.0053	0.0015
					(0.0192)	(0.0247)
L4. -TrTariffs^F					-0.0101	-0.0096
					(0.0192)	(0.0201)
L5. -TrTariffs^M						-0.0027
						(0.0143)
L5. -TrTariffs^F						-0.0019
						(0.0122)
Mean dep. var.	0.388	0.388	0.388	0.388	0.388	0.388
N. districts	1066	1066	1066	1066	1066	1066
Adjusted R^2	0.0744	0.0744	0.0744	0.0744	0.0744	0.0744
N	77906	77906	77906	77906	77906	77906
District and year FE	X	X	X	X	X	X
Region-year FE	X	X	X	X		X
Individual-level covariates	X	X	X	X	X	X
District-level covariates	X	X	X	X	X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Selective Migration

Selective migration may bias our results as it may affect the composition of victims between highly- and lowly-affected areas. For instance, if females that were already victims before liberalization migrate to highly affected areas, we will observe that trade liberalization is associated with a higher prevalence of violence. The opposite is true if female victims migrate from highly- to lowly-exposed areas. This is because our dependent variable asks about past episodes of violence, including those that happened before liberalization.

However, given that we exploit year-to-year changes in tariff exposure, migration may not be a problem. This is especially true if we consider that the short- and medium-term migration rates seem to be low. In fact, only 5.10 percent and 16.4 percent of our sample have changed their residence in the last year and in the last five years, which means that the share of people reallocating from one district to another is even lower. In fact, according to 2007 Census data, in the last five years, the inter-district migration rate of females between 15 and 49 years old was of 16.1 percent (and 15.9 percent in 2017). Moreover, migration does not appear to be related to male nor female exposure. In Table 7 we estimate equation (3) using dummies indicating if individuals have changed their residence as dependent variables. These are: (i) a dummy variable that takes the value of 1 whether the female *j* has at least changed her residence once during her lifetime, M_{ever}; (ii) a dummy whether she has changed her residence at least once since 1991, M_{1991}; (iii) a dummy whether she has changed her residence at least once in the last five years, M_{5yrs}; and (iv) a dummy whether she has changed her residence at least once in the last year, M_{1yr}. Our results show that male and female exposures are not statistically associated with the

probability of changing residence. If anything, male exposure seems to be negatively associated with long-term migration and positively correlated with short-term migration. Hence, endogenous sorting does not seem to be a problem in our setting. This goes in line with Dix-Carneiro et al. (2015) as they show that migration may play a limited role as an adjustment mechanism in Brazil.

Table 7: The Effect of Trade Liberalization on Migration

	Has changed residence at least once ...			
	... ever	... since 1991	... in the last 5 years	... in the last year
$TrTariffs^M$	-0.0105 -	0.0047	0.0035	0.0008
	(0.0130)	(0.0123)	(0.0090)	(0.0046)
$TrTariffs^F$	-0.0013	0.0002	0.0036	0.0028
	(0.0102)	(0.0095)	(0.0071)	(0.0034)
Mean dep. var.	0.559	0.420	0.164	0.051
N. districts	1066	1066	1066	1066
Adjusted R^2	0.140	0.147	0.116	0.0571
N	77891	77891	77891	77891
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates	X	X	X	X
District-level covariates	X	X	X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

We also evaluate if the effect of male exposure is larger on the sample of migrants compared with nonmigrants. If female victims are migrating from districts in which male employment was hit harder by liberalization, we would be underestimating the effect of male exposure and the effect on the sample of nonmigrants should be larger. If female victims are migrating into affected districts, we would be overstating the effect of male

exposure and the effect on the sample of nonmigrants should be smaller.[33] To carry out this exercise, we estimate the following equation:

$$\begin{aligned} y_{j,d,t} &= \alpha - \beta_1 TrTariff^M_{d,t} - \beta_2 TrTariff^W_{d,t} + \alpha_d + \alpha_t + \alpha_{r,t} \\ &+ \delta_0 M_{i,t} - \delta_1 [M_{i,t} \times TrTariff^M_{d,t}] - \delta_2 [M_{i,t} \times TrTariff^F_{d,t}] \\ &+ [\gamma_1' X_{j,d,t} + \gamma_2' Z_{d,t}] + \varepsilon_{j,d,t} \end{aligned} \quad (4)$$

where $M_{i,t}$ is one of the dummies defined above. The coefficients δ_1 and δ_2 measure the difference in the effect of trade liberalization between migrants and non-migrants. We show the results of estimating equation (4) in Table 8. The row labeled 'Test Male' shows the p-value of testing the null hypothesis: $\beta_1 + \delta_1 = 0$. Likewise, the row labeled 'Test Female' shows the p-value of testing the hypothesis: $\beta_2 + \delta_2 = 0$. Table 8 shows that the effect of trade liberalization is similar among those that have changed their residence and those that have not. The impact seems to be lower among those that have changed their residence at least once, which implies that we may be underestimating the effect of trade liberalization. However, it should be pointed out that the differences are not statistically significant.

[33] We should note that positive or negative selection into migration could also affect the magnitude of the effect of male exposure. The direction of this bias is difficult to know a priori.

Table 8: The Effect of Trade Liberalization by Migration Status

	Has suffered from physical intimate partner violence			
	(1)	**(2)**	**(3)**	**(4)**
$-TrTariffs^{M}$	0.0367	0.0357	0.0352	0.0347
	(0.0110)***	(0.0108)***	(0.0108)***	(0.0106)***
$-TrTariffs^{F}$	-0.0136	-0.0131	-0.0124	-0.0120
	(0.0086)	(0.0084)	(0.0083)	(0.0082)
$M_{ever} \times -TrTariffs^{M}$	-0.0041			
	(0.0037)			
$M_{ever} \times -TrTariffs^{F}$	0.0036			
	(0.0043)			
$M_{1991} \times -TrTariffs^{M}$		-0.0032		
		(0.0037)		
$M_{1991} \times -TrTariffs^{F}$		0.0032		
		(0.0041)		
$M_{5yrs} \times -TrTariffs^{M}$			-0.0048	
			(0.0045)	
$M_{5yrs} \times -TrTariffs^{F}$			0.0047	
			(0.0049)	
$M_{1yr} \times -TrTariffs^{M}$				-0.0055
				(0.0079)
$M_{1yr} \times -TrTariffs^{F}$				0.0067
				(0.0088)
Mean dep. var. (M=0)	0.362	0.384	0.393	0.391
Mean dep. var. (M=1)	0.408	0.392	0.359	0.330
Test Men	0.003	0.003	0.008	0.0319
Test Women	0.244	0.258	0.416	0.684
N. districts	1066	1066	1066	1066
Adjusted R^2	0.0748	0.0744	0.0745	0.0747
N	77891	77891	77891	77891
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates	X	X	X	X
District-level covariates	X	X	X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Permutation Tests

We carry out a number of permutation tests to show that our main results are not driven by spurious effects caused by (i) trends in intimate partner violence and (ii) time invariant cross-sectional patterns across districts. We randomize our sample to generate false data that we use to re-estimate equation (3). These placebo tests are useful to check whether our model is mis-specified and to calculate exact empirical p-values (Hsiang and Jina, 2014). We randomize the vector [$TrTariff^{M}_{d,t}$ and $TrTariff^{F}_{d,t}$] 2,000 times, without replacement and holding everything else fixed. We re-estimate equation (3) each time (the same equation estimated for column 3 in Table 2).[34]

Following Hsiang and Hina (2014), we conduct this randomization in two ways. First, we randomize the cross-sectional structure between districts. That is, we randomly re-assign each district's complete history of male and female exposure to tariff cuts to another district while preserving the ordering of years. Since this preserves the time structure within the data, this exercise serves to test whether national and regional trends are generating spurious correlations. Second, we randomize the time structure within districts. Put it differently, we randomly re-order each district's time-series of male and female exposure while keeping them assigned to the original district. Since this preserves the cross-sectional structure of the data and only alters its time structure, this exercise serves to test whether time invariant cross-sectional patterns across highly and lowly exposed districts are generating spurious results.

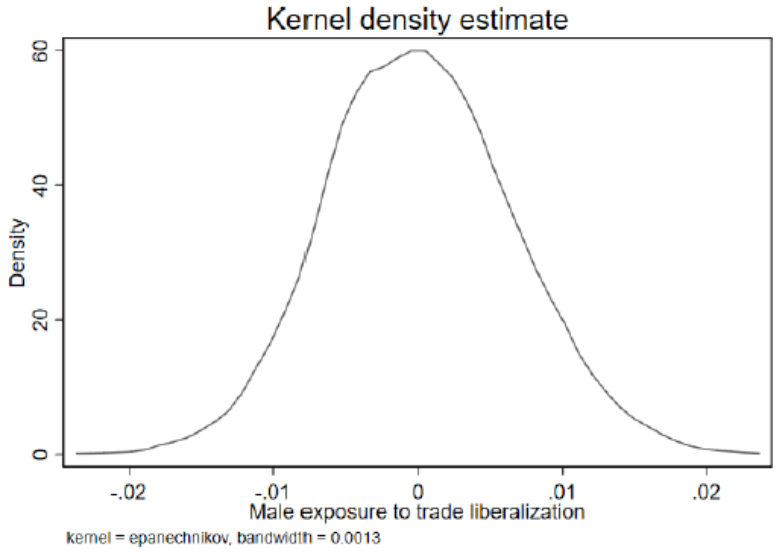
Figure 4 shows the results for β_1. We do not show the results for β_2 because it is not

[34] i.e. including region-year fixed effects and time-varying district level covariates. Our results are similar if we do not include them.

statistically different from zero. We confirm that both randomization procedures give two distributions properly centered at zero. Furthermore, both empirical p-values are below 0.001, validating our asymptotic confidence intervals (i.e. using clustered standard errors at the district level).

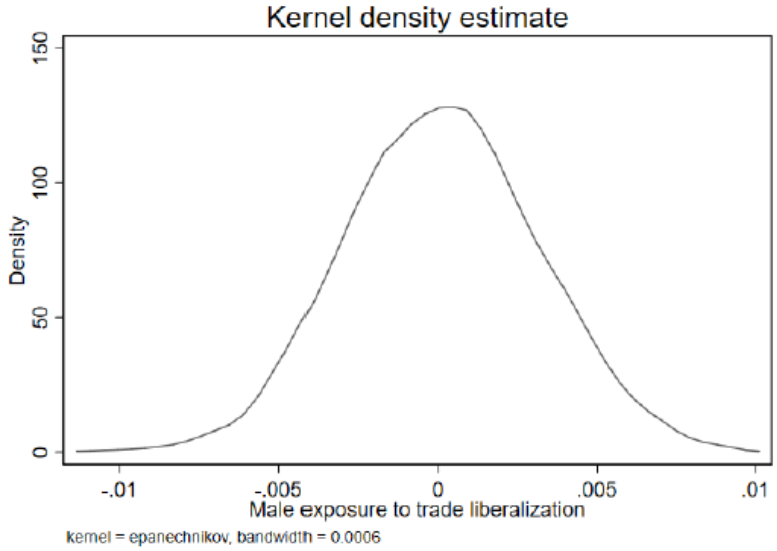
Figure 4: Empirical Distribution of Coefficients for

Panel A: Randomizing cross-sectional structure between districts



Exact p-value: 0.000

Panel B: Randomizing time structure within districts



Exact p-value: 0.000

Notes: Distribution of point estimates for $-TrTariff^{M}_{d,t}$ based on equation 3. Regressions include district and years fixed effects; region-year fixed effects; and the sets of individual- and district-level covariates. Each distribution is constructed by repeating the randomization and estimation procedure 2,000 times.

Heterogeneous Effects

We discuss whether the effect of trade liberalization is larger on different female sub-samples, including females that: (i) are younger than their partners (Friedberg and

Webb, 2006; Mabsout and van Staveren, 2010), (ii) are very young when they first married (Jensen and Thornton, 2003; Heath, 2014; Yount, et al., 2018), (iii) have little education (Heath, 2014; Aizer, 2010), and (iv) are less educated than their partners (Mabsout and van Staveren, 2010; Hidrobo and Fenald, 2013; Aizer, 2010). Each of these categories is associated with a higher probability of suffering intimate partner violence according to the literature.

In Table 9, we show the results of estimating the following equation:

$$\begin{aligned} y_{j,d,t} &= \alpha - \beta_1 TrTariff_{d,t}^{M} - \beta_2 TrTariff_{d,t}^{W} + \alpha_d + \alpha_t + \alpha_{r,t} \\ &+ \delta_0 D_{i,t} - \delta_1 [D_{i,t} \times TrTariff_{d,t}^{M}] - \delta_2 [D_{i,t} \times TrTariff_{d,t}^{F}] \\ &+ [\gamma_1^{'} X_{j,d,t} + \gamma_2^{'} Z_{d,t}] + \varepsilon_{j,d,t} \end{aligned} \quad (5)$$

where $D_{i,t}$ can be: (i) a dummy variable that takes the value of 1 whether the woman j is older than her partner (see column 1), (ii) a dummy whether she was at least 19 years old when she first married (see column 2),[35] (iii) a dummy whether she has completed high-school (see column 3), and (iv) a dummy whether she is more educated than her partner (see column 4). The coefficients δ_1 and δ_2 measure the heterogeneous effect of trade liberalization. The row labeled 'Test Male' shows the p-value of testing the null hypothesis: $\beta_1 + \delta_1 = 0$. Likewise, the row labeled 'Test Female' shows the p-value of testing the hypothesis: $\beta_2 + \delta_2 = 0$.

Although not all the interactions with male exposure are statistically significant, the negative signs of the coefficients suggest that increases in intimate partner violence are smaller among females that ex-ante were well positioned in their household. Females that

[35] 19 years old is the median age of first marriage in our sample.

are older than their partners (column 1), that are aged 19 or more when they first married (column 2), or that have at least completed high school (column 3), experience smaller increases in intimate partner violence. Similarly, given that most of the interactions with female exposure have a positive sign, decreases in violence are smaller among females that are well positioned within their household. The reverse side of this result is that decreases in violence were *larger* among females with a worse bargaining position, meaning that reducing tariff protection in female industries may actually decrease violence among certain female sub-populations.

Table 9: Heterogeneous Effects on Physical Intimate Partner Violence

	Has suffered from physical intimate partner violence			
	(1)	**(2)**	**(3)**	**(4)**
$-TrTariffs^M$	0.0359	0.0410	0.0394	0.0357
	(0.0108)***	(0.0111)***	(0.0106)***	(0.0106)***
$-TrTariffs^F$	-0.0130	-0.0183	-0.0178	-0.0124
	(0.0082)	(0.0086)**	(0.0086)**	(0.0085)
Older than partner × $-TrTariffs^M$	-0.0080			
	(0.0048)*			
Older than partner × $-TrTariffs^F$	0.0071			
	(0.0055)			
≥ 19 when marries × $-TrTariffs^M$		-0.0129		
		(0.0037)***		
≥ 19 when marries × $-TrTariffs^F$		0.0124		
		(0.0040)***		
High school × $-TrTariffs^M$			-0.0179	
			(0.0048)***	
High school × $-TrTariffs^F$			0.0162	
			(0.0053)***	
More educated × $-TrTariffs^M$				-0.0034
				(0.0037)
More educated × $-TrTariffs^F$				0.0012
				(0.0041)
Mean dep. var. (D=0)	0.386	0.421	0.407	0.393
Mean dep. var. (D=1)	0.397	0.360	0.359	0.382
Test Men	0.012	0.008	0.056	0.003
Test Women	0.535	0.475	0.859	0.184
N. districts	1066	1066	1066	1066

Adjusted R^2	0.0745	0.0748	0.0749	0.0745
N	77906	77906	77906	77906
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates	X	X	X	X
District-level covariates	X	X	X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Sex-specific Exposure to Input Tariffs

When overall tariffs are reduced both prices of output goods and intermediate inputs are affected. In contrast with output prices, reductions in prices on inputs may reduce wages and employment, as they tend to enhance productivity through lower marginal costs. To what extent these opposing effects are transmitted to households dynamics is an empirical question. Analogous conjectures have been tested in the case of other outcomes. Kis-Katos and Sparrow (2015) find that decreases in output tariffs raise poverty, whereas decreases in input have the opposite effect. Similarly, Amiti and Cameron (2012) show that input tariff reductions contribute to the closure of the industrial skill wage gap in Indonesia, whereas Dix-Carneiro and Kovak (2015) show that cuts in output tariffs modestly widened the skill wage gap in Brazil.

We calculate two additional measures of exposure to input tariffs by exploiting the pre- reform composition of male and female employment to weight input tariffs accordingly.[36] We estimate the following specification, which is analogous to equation (3):

[36] Please see Appendix B.

$$
\begin{aligned}
y_{j,d,t} &= \alpha - \beta_1 TrTariff^M_{d,t} - \beta_2 TrTariff^W_{d,t} - \rho_1 InputTariff^M_{d,t} - \rho_2 InputTariff^W_{d,t} \\
&+ \alpha_d + \alpha_t + \alpha_{r,t} + [\gamma_1^{'} X_{j,d,t} + \gamma_2^{'} Z_{d,t}] + \varepsilon_{j,d,t} \qquad (6)
\end{aligned}
$$

where ρ_1 and ρ_2 measure the impact of a decrease in input tariffs affecting male and female industries on the probability of suffering increased physical intimate partner violence. This time, the set of time-variant district-level covariates, $Z_{d,t}$, does not consider input tariffs as they have already been included. Table 10 shows our results. For ease of comparison, under column 1 we show the same estimation exhibited under column 3 of Table 2, including the coefficient on the overall measure of input tariff exposure, which we used as a district-level covariate before.

Column 1 suggests that reductions in input tariffs are negatively associated with increases in intimate partner violence, however this relationship is not statistically significant. In columns 2 to 4, we disaggregate overall input tariffs by sex-predominance industry. Column 4 is our preferred specification as it includes individual- and district-level covariates. Male exposure to reductions in output tariffs is still associated with increases in intimate partner violence, but female exposure is now statistically significant at 5 percent, which suggests that output tariff cuts on female-predominant industries may decrease violence within households. For input tariffs, we find that larger reductions in male exposure to input tariffs decrease violence, whereas larger cuts in female exposure may increase it, albeit these increases are not statistically different from zero. We reach to the same conclusions if we ignore district-level covariates (column 3) or individual-level covariates (column 2). These results confirm our conjecture that the effects of output tariffs and input tariffs should be of opposite sign, which goes in line with findings in the literature.

Table 10: The Effect of Lower Input Tariffs on Physical Intimate Partner Violence

	Has suffered from physical intimate partner violence			
	(1)	**(2)**	**(3)**	**(4)**
-$TrTariffs^M$	0.0344	0.0317	0.0342	0.0377
	(0.0106)***	(0.0109)***	(0.0107)***	(0.0109)***
-$TrTariffs^F$	-0.0118	-0.0196	-0.0222	-0.0210
	(0.0082)	(0.0108)*	(0.0105)**	(0.0106)**
-*InputTrTariffs*	-0.0546			
	(0.0333)			
-*Input* $TrTariffs^M$		-0.0823	-0.0826	-0.0790
		(0.0330)**	(0.0334)**	(0.0338)**
-*Input* $TrTariffs^F$		0.0516	0.0548	0.0466
		(0.0409)	(0.0399)	(0.0422)
Mean dep. var. (D=0)	0.388	0.387	0.388	0.388
N. districts	1066	1066	1066	1066
Adjusted R^2	0.0744	0.0507	0.0744	0.0744
N	77906	78276	77906	77906
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates	X		X	X
District-level covariates	X			X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. **The set of district-level covariates** includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Concluding Remarks

In this research we ask whether trade liberalization may impact household dynamics in such a way that the end result may be an increase in physical intimate partner violence. This question, one that has not been addressed before is rather relevant as countless people around the world suffer from physical violence on a daily basis. In order to deal with it we exploit an unexpected tariff reduction across several industries in Peru during the 2000s we are able to causally show that in districts where industries'

employment are predominantly male, trade liberalization produced an increase in physical intimate partner violence that ranges from 18.8 percentage points to around 36.6 percentage points with respect to control districts. Interestingly, we find no such statistically significant impact in the case of districts where female-predominant industries were more exposed to tariff reductions.

Our results are robust to falsification and placebo tests, sensitivity to initial conditions, conflation of past and current shocks, selective migration, permutation tests and input-tariffs considerations. Also, we find considerable heterogeneity, as education and the age of first marriage appear to be key variables that correlate with our findings. Finally, we also exploit the fact that when overall tariffs are reduced both prices of output goods and intermediate inputs are affected. In contrast with output prices, reductions in prices on inputs may increase wages and employment, as they tend to enhance productivity through lower marginal costs. In the data, we find that these opposing effects may transmit to household dynamics as male and female exposure to input tariff cuts are associated with decreases and increases in physical intimate partner violence respectively. Changes in labor markets conditions appear to have a direct bearing on the incidence of intimate partner violence, as consistent with current theories in the social sciences, they help shape the distribution of bargaining power between partners. It appears that depending on the specific industry, changes in labor markets may increase the relative economic rewards of one gender with respect to the other, but at the same time change and even compromise the psychological balance within household members.

Our work contributes to the broader social impacts of globalization and reinforce recent research showing that negative impacts on the labor market or other parts of

deindustrialization can lead to adverse marriage market impacts, fertility, and, in our case, on intimate partner violence in a wide range of settings. From a policy perspective, our findings demonstrate that sometimes, sensible economic policies can have negative, unexpected repercussions. They also provide an opportunity to policymakers to pursue proactive policy measures in order to help prevent or alleviate this issue. Two specific measures that governments may find useful are educational messages via traditional and social media and an increase in peer awareness on the typical red flags associated with intimate partner violence.

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Appendix A: Additional Tables and Figures

Table A.1: Top 10 Traded Industries by...:

Panel A: ...the share of male workers

Industry (ISIC3 Group)	Share of male workers
Manufacture of furniture (361)	0.962
Quarrying of stone, sand and clay (141)	0.958
Manufacture of structural metal products, tanks, reservoirs and steam generators (281)	0.953
Mining of uranium and thorium ores (120)	0.952
Manufacture of products of wood, cork, straw and plaiting materials (202)	0.950
Fishing, aquaculture and service activities incidental to fishing (050)	0.949
Mining of non-ferrous metal ores, except uranium and thorium ores (132)	0.946
Manufacture of aircraft and spacecraft (353)	0.945
Mining and quarrying n.e.c. (142)	0.938
Manufacture of other fabricated metal products; metal working service activities (289)	0.936

Panel B: ...the share of female workers

Industry (ISIC3 Group)	Share of female workers
Manufacture of knitted and crocheted fabrics and articles (173)	0.679
Other service activities (930)	0.614
Manufacture of wearing apparel, except fur apparel (181)	0.567
Extraction and agglomeration of peat (103)	0.500
Farming of animals (012)	0.426
Manufacture of coke oven products (231)	0.375
Manufacture of optical instruments and photographic equipment (332)	0.372
Manufacture of other textiles (172)	0.337
Manufacture of other chemical products (242)	0.311
Spinning, weaving and finishing of textiles (171)	0.283

Source: 1993 Population and Household Census

Table A.2: The Effect of Trade Liberalization on Emotional Violence

	Has ever suffered from emotional violence			
	(1)	(2)	(3)	(4)
$-TrTariffs^{M}$	0.0216	0.0237	0.0216	
	(0.0087)***	(0.0084)***	(0.0106)**	
$-\ TrTariffs^{F}$	-0.0188	-0.0202	-0.0217	
	(0.0078)**	(0.0077)***	(0.0077)***	
$-TrTariffs$				-0.0006
				(0.0083)
Mean dep. var.	0.302	0.302	0.302	0.302
N. districts	1066	1066	1066	1066
Adjusted R^2	0.0501	0.0664	0.0664	0.0663
N	78305	77935	77935	77935
District and year FE	X	X	X	X
Region-year FE	X	X	X	X
Individual-level covariates		X	X	X
District-level covariates			X	X

Notes: Standard Errors clustered at the district level. *** Significant at the 1 percent level. ** Significant at the 5 percent level. * Significant at the 10 percent level.

The set of individual-level covariates includes the age and years of education of both partners, the age of the female when first married, a dummy that captures whether the female speaks Spanish, the size of the household, the sex of the household head, and the altitude at which the household is located. The set of district-level covariates includes a measure of exposure to foreign direct investment, a measure of exposure to exports, and a measure of exposure to input tariffs. Details on the construction of these variables can be found in Appendix B.

Table A.3: Autocorrelation of $\Delta TrTariff^{M}$

Years	11-10	10-09	09-08	08-07	07-06	06-05	05-04	04-03	03-02	02-01	01-00	00-99	99-98
2011-2010	1.00												
2010-2009	-0.43	1.00											
2009-2008	0.58	-0.57	1.00										
2008-2007	0.55	0.57	-0.82	1.00									
2007-2006	.	.	.	.	.								
2006-2005	-0.17	-0.32	-0.04	-0.15	.	1.00							
2005-2004	-0.54	0.51	-0.79	0.89	.	-0.19	1.00						
2004-2003	0.55	-0.71	0.81	-0.87	.	0.31	-0.92	1.00					
2003-2002	-0.55	0.71	-0.81	0.87	.	-0.31	0.92	-1.00	1.00				
2002-2001	-0.54	0.60	-0.75	0.83	.	-0.27	0.79	-0.86	0.86	1.00			
2001-2000	0.69	-0.56	0.83	-0.78	.	0.24	-0.80	0.81	-0.81	-0.76	1.00		
2000-1999	-0.79	0.58	-0.93	0.79	.	0.03	0.75	-0.78	0.77	0.72	-0.93	1.00	
1999-1998	0.31	-0.63	0.46	-0.47	.	0.34	-0.48	0.68	-0.68	-0.57	0.45	-0.45	1.00

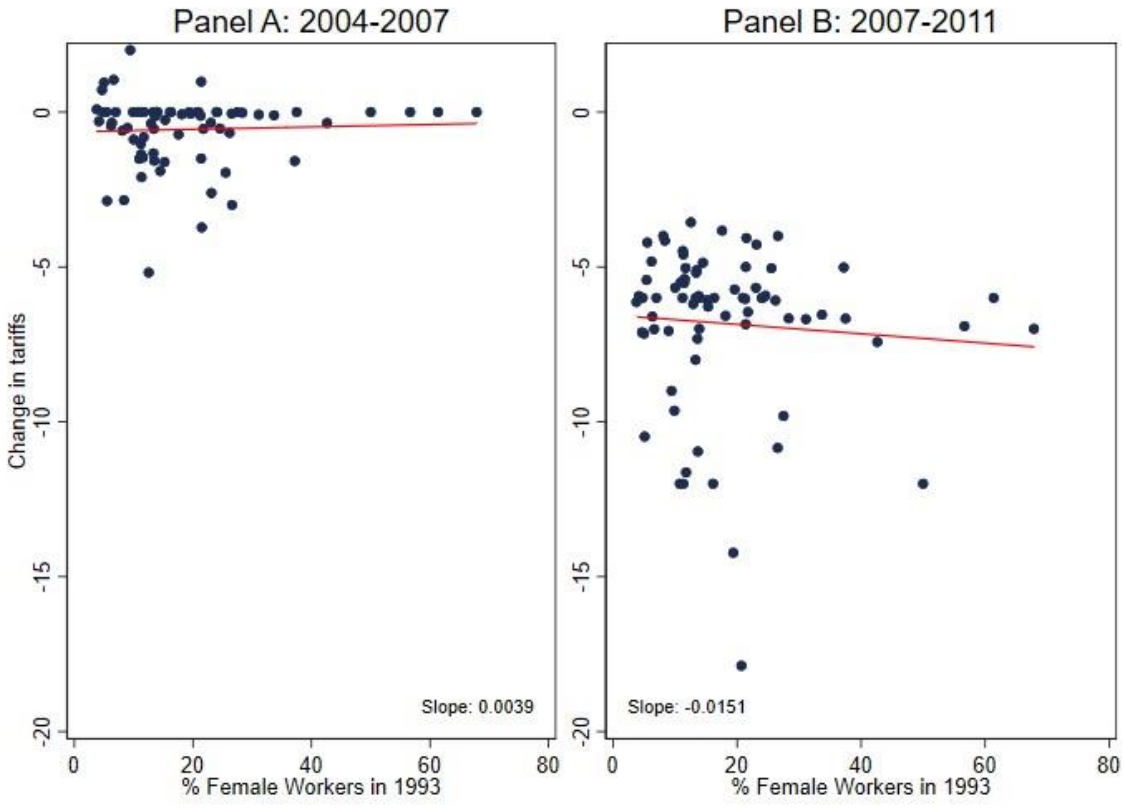
Source: World Bank TRAINS, 1993 Population and Household Census, World Bank's Concordance Table

Table A.4: Autocorrelation of Δ $TrTariff^F$

Years	11-10	10-09	09-08	08-07	07-06	06-05	05-04	04-03	03-02	02-01	01-00	00-99	99-98
2011-2010	1.00												
2010-2009	-0.02	1.00											
2009-2008	0.18	-0.29	1.00										
2008-2007	-0.35	0.15	-0.54	1.00									
2007-2006	.	.	.	.	.								
2006-2005	-0.12	-0.15	-0.05	0.01	.	1.00							
2005-2004	-0.58	0.05	-0.56	0.80	.	-0.01	1.00						
2004-2003	0.54	-0.22	0.63	-0.79	.	0.09	-0.95	1.00					
2003-2002	-0.54	0.22	-0.63	0.79	.	-0.09	0.95	-0.99	1.00				
2002-2001	-0.51	0.15	-0.54	0.68	.	-0.12	0.77	-0.82	0.82	1.00			
2001-2000	0.14	-0.25	0.97	-0.48	.	0.01	-0.51	0.59	-0.59	-0.51	1.00		
2000-1999	-0.22	0.29	-0.99	0.53	.	0.06	0.53	-0.60	0.61	0.53	-0.98	1.00	
1999-1998	0.01	-0.14	0.08	-0.03	.	0.02	-0.02	0.08	-0.08	-0.07	0.07	-0.08	1.00

Source: World Bank TRAINS, 1993 Population and Household Census, World Bank's Concordance Table

Figure A.1: Correlation between the share of female workers per industry in 1993 and tariff changes



Source: World Bank TRAINS, World Bank's concordance tables, and 1993 Population and Household Census

Notes: Tariff changes were computed at the industry level using ISIC3 codes. Originally, industries were coded based on the Trade Classification Harmonized System (HS). We translated this classification into the International Standard Industrial Classification (ISIC3) using the concordance tables available online.

Appendix B: Data

B.1 Individual-level outcomes

Physical Intimate Partner Violence

Dummy: Takes the value of one for women that have ever been (i) pushed, shook or thrown something at, (ii) slapped or arm twisted, (iii) punched with fist or something harmful, (iv) kicked or dragged, (v) strangled or burnt, (vi) threatened with a knife/gun or other weapon, (vii) attacked with knife/gun or other weapon, (viii) forced to have sex when not wanted, and (ix) forced to make other sexual acts when not wanted, by her spouse. This variable is at the individual level and comes from the DHS surveys. We only consider women that were in a relationship when they were surveyed.

Intensity: Using the dummy variables described above we compute a measure of intensity by adding them together. Hence, this measure goes from 0 to 9 and its average value is of 1.12 with a standard deviation of 1.76.

Principal component: Using the dummy variables described above we compute the first component from a principal component analysis, which accounts 41 percent of the total variance. Its average value is of 0.42 and has a standard deviation of 0.65.

Emotional Intimate Partner Violence:

Dummy: Takes the value of one for women that have ever been (i) humiliated, (ii) threatened with harm, and (iii) threatened by her spouse. This variable is at the individual level and comes from the DHS surveys. We only consider women that were in a relationship when they were surveyed.

Intensity: Using the dummy variables described above we compute a measure of intensity

by adding them together. Hence, this measure goes from 0 to and its average value is of 0.50 with a standard deviation of 0.88.

Principal component: Using the dummy variables described above we compute the first component from a principal component analysis, which accounts 63 percent of the total variance. Its average value is of 0.29 and has a standard deviation of 0.51.

B.2 Individual-level controls

Summary statistics are shown in Table 1. "Age at marriage" is the age in years of first marriage. "Age" is the age in years. "Partner's age" is the age in years of each woman's partner. "Years of educ. (YoE)" is the education in years. "Partner's YoE" is the education in years of each woman's partner. "HH. head is women" is a dummy indicating if the household head is female. "Non-Spanish" is a dummy that indicates whether a particular woman speaks Quechua, Aymara or any other language different from Spanish. "HH. size" is the number of individuals living in each woman's household. "Altitude" is the meters over the sea level at which the household is located. Source: DHS.

B.3 District-level variables

Exposure to Tariff Changes: For district d at year t we construct the following measure of exposure:

$$TrTariff_{d,t} = \sum_{i}^{I} \frac{L_{1993,i,d}}{L_{1993,d}} \times tariff_{i,t} \qquad (7)$$

where $L_{1993,i,d}$ is the number of workers in sector i in district d in 1993, $L_{1993,d}$ is the district d's total number of workers in 1993, and $tariff_{i,t}$ is the Most-Favored- Nation (MFN) tariff of industry i at year t. To compute this variable, we exclude the services sector altogether, as this has become standard practice in the literature. Finally, since the Census industry codes use the

International Standard Industrial classification (ISIC 3) aggregated at the 3-digit level, whereas tariff data use the Trade Classification Harmonized System (HS), we convert HS codes into ISIC3 codes using the concordance tables available at the World Bank's website. This means that we are able to distinguish between I = 76 different industries.

Exposure to Tariff Changes by Sex: Refer to Section 5. In addition, since the Census industry codes use the International Standard Industrial classification (ISIC 3) aggregated at the 3-digit level, whereas tariff data use the Trade Classification Harmonized System (HS), we convert HS codes into ISIC3 codes using the concordance tables available at the World Bank's website. This means that we are able to distinguish between 76 different industries.

Input Tariffs: We follow Edmonds, et al., (2010) and use the 1993 Peruvian national input-output table, the 1993 national census, and MFN tariffs to construct this variable. For each industry *i*, we create an input tariff for that industry as the weighted average of tariffs on goods used for production in industry *i* (which is between parenthesis in equation 5). Such weights were constructed using industry *j*'s share of industry *i*'s total input cost, which we call $sc_{j,i,1993}$. Then, the district input tariff is computed by weighting industry *i*'s input tariff by *i*'s employment share in the district in 1993:

$$InputTariff_{d,t} = \sum_{i}^{I} \frac{L_{i,d,1993}}{L_{d,1993}} \left(\sum_{j}^{J} sc_{j,i,1993} \times tariff_{j,t}\right) \quad (8)$$

We do not exclude the services sector when computing the input tariff of industry *i*, $(\sum_{j}^{J} sc_{j,i,1993} \times tariff_{j,t})$ to reflect the fact that some services may be used in the production of output goods. However, we do not consider them for the set of output industries I. Once we account for the industries considered in I and in J, we are able to distinguish between 32 different industries. This is because the Peruvian input-output table features 45 sectors, hence we had to

work at that level of aggregation.

Input Tariffs by sex: We compute the following measures of exposure:

$$InputTariff_{d,t}^{G} = \sum_{i}^{I} \frac{L_{i,d,1993}^{G}}{L_{d,1993}^{G}} \left(\sum_{j}^{J} sc_{j,i,1993} \times tariff_{j,t}\right) \quad (9)$$

where $G = \{M, F\}$, M stands for male, and F stands for female. To construct (9) we apply the same considerations as in the computation of (8) above.

Foreign Direct Investments: $FDI_{d,t} = \sum_{i}^{I} w_{i,J(i),d} \times FDI_{J,t}$, where the employment in production sector i in district d as a share of total employment in the district is defined as $w_{i,J(i),d} \equiv L_{i,J(i),d,1993}/L_{d,1993}$. $FDI_{J,t}$ is the total foreign direct investments destined to sector J (sector codes aggregated to 2-digits). This data was compiled from the Private Investment Promotion Agency (Pro Inversión) and it distinguishes between 14 different sectors. We drop the services sector when computing $w_{i,J(i),d}$.

Exports: $Exports_{d,t} = \sum_{i} w_{i,d} \times Exports_{i,t}$, where $w_{i,d} \equiv L_{i,d}/L_{d}$ is the employment in production sector i in district d as a share of total employment in the district. $Exports_{i,t}$ is the total value of exports made by firms in sector i. This data was compiled from the World Bank's TRAINS Data. We drop the services sector when computing $w_{i,d}$. Since the Census industry codes use the International Standard Industrial classification (ISIC 3) aggregated at the 3-digit level, whereas exports data use the Trade Classification Harmonized System (HS), we convert HS codes into ISIC3 codes using the concordance tables available at the World Bank's website. This means that we can distinguish between $I = 76$ different industries.