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AN EMPIRICAL ANALYSIS OF THE TRADE LIBERALISATION EFFECTS ON THE GENDER WAGE GAP: THE CASE STUDY OF LATIN AMERICAN COUNTRIES

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ABSTRACT

The aim of this paper is to empirically explore the trade liberalisation effects on the gender wage gap in twelve Latin American countries between 1995 and 2014. Static and dynamic panel data techniques are used to investigate the effects on Latin American countries. As a static technique, the Fixed Effects Model presents a positive relationship between trade openness and the gender wage gap. In addition, as a control variable, it is found that female parliamentary participation has a negative correlation with the gender wage gap. On the other hand, in dynamic panel data analysis, the findings of the Pooled Mean Group Model (PMG) show that trade liberalisation leads to an increase in the gender wage gap, and the education level (HDI) is negatively correlated with the gender wage gap in the long-run. However, in this study, no significant results could be found for the concentration index of export.

1. INTRODUCTION

Over the recent decades, economic reform and trade liberalisation policies have been implemented by many countries, and the effects of these reforms on gender differences in the labour market has started to emerge as an important research subject in international economics and development studies.

Marchand (1996) claims that trade liberalisation has both exclusive and inclusive effects for females. Many theorists support the idea that global economic integration increases opportunities for females in the labour market, but the integration does not remove all barriers to women's improvement (Mears 1995; Joekes and Weston 1994; L. Meyer 2007). The intensity of female labour forces in low-skilled jobs, and gender wage gaps at significant levels, are still the important issue across the world regardless of the development level (AlAzzawi, 2013).

This study is aimed to analyse the effects of trade liberalisation on the gender wage gap in Latin American countries from 1995 to 2014. Due to the data limitations, only twelve countries could be examined. The reason for choosing this period is that since the 1990s, Latin American countries have experienced the opening trade and capital markets to foreign flows. During the trade liberalisation process, Latin American economies have deeply integrated into global markets for capital, goods and services (Ventura-Dias, 2010). A cross-country study is chosen for this paper because there are only a few studies that look at the relationship between trade and the gender wage gap on an international level, while there are many country-specific studies analysing the relationship.

The remainder of this paper proceeds as follows: the next section provides the literature review with the main titles of cross-country studies and country-specific studies; the cross-country studies part consists of four subsections that include: developed countries; economies in transition; developing countries; and the least developed countries. In section three, the data and methodology is given, including the ordering variables and test static panel data analysis, and the dynamic panel data analysis; the fourth section includes empirical results, and a conclusion is provided in the fifth section.

2. LITERATURE REVIEW

Becker has started the literature on the economics of discrimination in 1957 (Weichselbaumer and Winter-Ebmer, 2005). According the theory of Becker, if employers have a taste for discrimination, they will expose some economic cost, and the discrimination could make firms worse off. Opening a country to trade causes competition, and competition makes discrimination more costly and burdensome for employers to discriminate. The ascending competition is expected to reduce the discrimination against women and minorities in the long run (Becker,1957). Also, Wood (1991) argues that increasing international trade will extend job opportunities under favour of absorbing an increasing number of women in export-oriented industries. (Black and Brainerd, 2004; Rasekhi and Hosseinmardi, 2012; Neumayer and Soysa, 2007; Cagatay and Erturk, 2004; Becker, 1957; Busse and Spielman, 2006)

Becker has a neo-classical approach for discrimination. On the other hand, the non-neoclassical approach presumes that increasing foreign competition could reduce female workers' bargaining power and the situation may make their positions powerless. In sectors which are competitive because of low labour costs, women will be overrepresented. The sectors affected by increasing trade tend to maintain their competitiveness, and because of that intention, there will be a farther downward force on female workers' wages. The circumstance causes an increase in an unadjusted wage gap (Williams, 1987).

Heckscher-Ohlin (1991) model points to the factor-price equalisation impact of trade. Less-skilled labour abundant economies tend to specialise in less-skilled labour-intensive exports, and the demand for less-skilled labour will increase. Thus, the wages of less-skilled labour will rise relative to skilled labour (Korinek, 2005; Aguayo-Tellez, 2012). This effect also indicates a decreasing of gender wage gap in emerging economies because women usually work for lower wages and in lower-skilled jobs than men (Wolszczak-Derlacz, 2013; Korinek, 2005).

According to Heterodox Theory, international trade is a result of comparative advantage. Elson (2007) says that there is hierarchy, and the weak are dominated by the forceful. So, it does not contribute to the hypothesis of full employment, perfect competition and price adjustment, which comes with trade.

2.1.Cross-Country Studies

Oostendorp (2004) studies the impact of globalisation on the gender wage gap, focusing on occupational gender wage gap. He uses the ILO October Inquiry Database on the gender wage gap in 161 carefully identified occupations in more than 80 countries for 1983-1999. The study shows that the occupational gender wage gap tends to decline with trade and foreign direct investment (FDI) in richer countries, but there is little evidence about poorer countries. Lutz's findings (2006) support Ooestendorp's results about poorer countries. Lutz analyses the effects of trade openness and growth on gender inequality and finds a weak relationship for African countries.

Yahmed's (2012) study provides theoretical evidence for heterogeneous effects of trade openness on the wage depending on the position throughout the skill distribution. The study shows that if abilities and job loyalties are complementary, and are complementary to technological upgrading, statistical inequity based on job loyalty anticipations produce a greater gender wage gap at the top of the skill distribution. Also, the study shows that the gender wage gap is decreased at the lower part of the skill distribution because of general equilibrium effects.

Wolszczak-Derlacz (2013) analyses the relationship the impacts of domestic and foreign competition on male-female wage differentials for twelve manufacturing sectors in eighteen OECD countries between 1970 and 2005. The results show that, in general, the growth of the gender wage gap declines for all classes of skills and all countries during the period. Furthermore, concentrated industries have a greater gender wage gap than competitive industries. The rise in both import and export inflow is correlated with a lower growth of the high-skilled gender wage in concentrated industries. In contrast, trade increases the growth of the medium and low-skilled sectoral gender wage gap in less-competitive industries. In competitive industries, trade openness increases the high-skilled male-female wage differentials, while it causes decrease for medium and low-skilled.

Neumayer and Soysa (2007) analyses the relationship between globalisation and women's economics rights as a quantitative study. The results show that countries which are more open to trade have much better economic rights for women, and have lessened of forced labour. They claim that the result is valid in a global sample and in a subsample of developing countries.

Gupta (2002) refers that women are overrepresented in low-paying occupations and the situation is a major factor of women's lower relative pay in developing countries, similar to in industrialised countries. Also, the author measured market emergence by export expansion collaborated with industrialisation, and there is a significant positive relation between market emergence and gender wage gap.

Meyer (2007) finds that trade openness rises women's share of earned income using cross-sectional and panel OLS regression analysis for 55 countries for 1975 to 1998. Rasekhi and Hosseinmardi (2012) analyse the impact of globalisation on the gender wage gap for 21 developing countries between 2002 and 2007. They found a negative and significant effect between trade openness and gender wage gap.

2.2. Country-Specific Studies

The classification of the countries is based on the World Economic Situation and Prospects (United Nations, 2014).

2.2.1. Developed Countries

Berik, Rodgers, and Zveglich (2004) search the effects of competition from international trade on gender wage inequality for Taiwan and Korea. The research shows that in concentrated industries, competition is positively correlated with wage discrimination against female workers, as opposed to Neoclassical Theory. Greater trade openness causes larger residual wage gaps between males and females in Taiwan. In Korea, a little attenuation in export openness is related to less gender wage discrimination in concentrated industries.

Seguino's study (2000) is based on the period (1981-1992) in which Korea and Taiwan headed towards technological upgrading of economies and further liberalisation of markets. He finds that the gender wage gap has expanded in Taiwan and has decreased in Korea. The author refers that economic remodelling in the newly industrialised economies (NIEs), the latest attempts to liberalise the economies of Taiwan and Korea and sustain export competitiveness provided a wider gender gap in the former country, but some progresses in the latter.

Black and Brainerd (2004) analyse the impact of globalisation on gender discrimination in manufacturing industries in the US. They compare the change in the gender wage gap in competitive industries and concentrated industries between 1976 and 1993 in the US. The

results indicate that trade rises wage inequality by plainly diminishing the relative wages of less-skilled workers. Also, it is clear that it reduces the ability of firms to discriminate against women.

Bøler, Javorcik and Ulltveit-Moe (2015) discuss the difference in gender wage between export-oriented firms and non-exporters for Norwegian manufacturing sectors from 1996 to 2010. The results indicate that female workers earn 24 percent less then male workers in general, but in exporting sector they earn 19 percent less than men. The authors also study the heterogeneity between exporters. It is estimated that firms exporting to a multitude of places tend to be greater and more productive than other exporters, and these firms have higher expectations from their employees. These expectations are expected to cause a wider gender gap in the firms. The findings prove that increasing numbers of export markets expands the gender wage gap in the exporting firms.

Sauré and Zoabi (2014) argue the relationship between trade expansion and female labour participation using bilateral data for the US and Mexico between 1990 and 2007. They separate the sectors as female intensive sectors and male intensive sectors, and claim that female labour participation decreases if trade extents the female intensive sectors. They say that trade integration of capital abundant economy extends the female intensive sector and narrows the male intensive sector when the female intensive sector is capital-intensive. As a result, male workers in male intensive sectors change their sector to female intensive sectors and reduce the capital-labour ratio in female intensive sectors. The marginal productivity of females falls more than men's as capital and female labour are complementary at high level. As a result, female labour participation decreases and the gender wage gap increases. They found that the trade effects of NAFTA in the United States on employment are widely compatible with their theory.

Kucera (2001) analyses the effects of trade liberalisation on gender relative wages and female labour participation in Germany and Japan between 1970 and 1996. The author finds that the female-to-male hourly wages ratio increased between the mid-1970s to 1990s in Germany. Inversely, the ratio decreased in Japan.

Leven (2008) focuses on the effect of Polish economic transition on female employees for 1990 to 2003. During the first phase of the economic transition, women had a better position in the financial sector and entrepreneurial ranks, and women took advantage of the increasing

women's movement. The findings indicate that during the first stage of transition, income has increased for women in the financial sector, and now, Polish women have the highest dividend of female entrepreneurs in any Central or West European economy. In addition to Leven's findings, other studies confirm that the gender wage gap in the Polish financial sector narrowed between 1996 and 2002 (Keane and Prasad, 2002; Grajek, 2003). In spite of the narrowing gender wage gap, men still earn a higher income than women, even in the finance and business sectors.

Jolliffe and Campos (2005) study gender differences in the labour market during the economic transition from 1986 to 1998 in Hungary. They find that the female-to-male wage ratio increased from 73 percent to 80 percent. In addition, the results show that the gender differences in log wages decreased remarkably from 0.31 to 0.19 during the transition using standard Oaxaca decomposition.

2.2.2. Economies in Transition

Kecmanovic and Barret (2011) examine the impact of economic transition on the mean gender wage gap in Serbia between 2001 and 2005. They use quantile decompositions for the study and the results show that the increase in female workers' productivity explains the fall in the gender wage gap across distribution. The mean gender wage gap decreased from 14 percent to 6 percent during the period as the residual component of the gender wage gap remained large and raised lightly during the transition. Overall, Serbian female workers' relative wages increased in the first stage of the transition.

2.2.3. Developing Countries

Juhn, Ujhelyi and Villegas-Sanchez (2013) claim that more productive firms tend to renovate their technology and get in export markets through a reduction in tariffs. New technology reduces the need for physically demanding skills. The process enhaces the relative wage and labour participation for women in blue-collar jobs, but not in white-collar jobs. The authors test their model for Mexico and find the consistent results with their model.

Artecona and Cunningham (2002) research the alteration in the gender wage gap during the trade liberalisation period in the manufacturing sector in urban Mexico. They find their results to be similar as with Dominguez-Villalobos and Brown-Grossman's (2008). The results show that trade liberalisation increased the gender wage gap in the Mexican manufacturing sector. Cunningham illustrates that the rise in gender wage gap can be

explained with the general increase in the skills premium, because the skills premium affects relatively abundant less-skilled female workers. Nonetheless, Cunningham refers that the discrimination constituent of the gender wage gap declined with competition.

Siddiqui (2009) analyses gender aspects of the impacts of trade liberalisation in Pakistan using a Computable General Equilibrium (CGE) Model. The results show that trade liberalisation increased female labour participation in unskilled jobs. Also, it increased females' real wage income more than males'. However, the study also shows that trade liberalisation increased workload and relative income poverty for relatively poor women, and it corrupted their abilities.

Menon and Rodgers (2009) explain the impact of trade reforms on the gender wage gap in Indian manufacturing industries for 1983 to 2004. The authors have found that rising trade openness in more concentrated manufacturing industries is related with increasing the residual gender wage gap in India.

After 1980, Taiwan have achieved success with export-oriented manufacturing. Berik (2000) analyses effects of the greater export-orientation on gender wage inequality in Taiwan, from 1984 to 1993, using industry-level panel data. The study refers that larger export-orientation affected both females' and males' wages negatively, but reduced gender wage inequality because male workers confront a greater wage penalty than women. The study also finds that increasing capital intensity increases both males' and females' wages, and to shift from wage to salaried jobs, raises men's wages, but decreases women's wages.

Nicita (2007) analyses the effects of the export growth in the textile and apparel sector on social well-being and poverty reduction in Madagascar. The study also has a gender perspective: in general, the effect of the export increase on women is positive, but for unskilled women, the gains were much lower. These gains come from finding jobs in the sector, rather than increasing in the wages. Yet, still female workers benefit less than male workers. Furthermore, Nicita observes gender wage differentials for unskilled individuals and finds that male workers earn about 32 percent more than female workers in textile industry.

AlAzzawi (2013) researches the expansion of gender inequality in the Egyptian manufacturing sector, and the effect of trade reform on the gender wage gap and on the female labour force. The study shows that the rise in trade liberalisation has had substantially negative effects on women's relative wages and their employment, despite the control of

public-private segmentation and occupational segmentation. Besides, female relative wages increasing in trade liberalisation widens the gender pay gap in the manufacturing sector.

Isaza (2013) found a significant positive relationship between the rise in import penetration, and greater women shares of jobs in manufacturing industries in urban Colombia, but they refer that they could not find any significant results for the effects of increased competition, import penetration and market concentration on the expansion of the gender wage gap as their dependent variable is the female share of jobs.

2.2.4. Least Developed Countries

Hoyos, Bussolo and Núñez (2012) claim that the enlargement of the maquila sector, which is export-oriented, has provided gender equality in employment and wages in Honduras. The results show that gender discrimination is 16 percent smaller in maquila sector firms. The result conforms to Becker's theory, which is that competition decreases gender discrimination. Considering the intensity of the use of female labour, the rise in the maquila sector promoted the reduction in the gender wage gap in Honduras.

3. DATA AND METHODOLOGY

In this section, after presenting the dataset and explaining dependent variable and independent variables, the econometric methodology and the implications of it will be defined.

3.1. Data

3.1.1. Dependent Variable

To construct the gender wage gap variable, the labour datasets are taken from the CEDLAS Gender database and the ILOSTAT Earnings and Labour Cost database. Estimation of the gender wage gap alters considerably depending on the choice of time reference used in the wage statistics. For comparability purposes, nominal wages in the local currency is converted into international dollars using purchasing power parity (PPP). The PPP data is extracted from the World Bank. In this analysis, nominal hourly wages are used to calculate the gender wage gap, because taking into account weekly or monthly wages can cause distortions in the results. The reason for distortion is explained by the fact that women generally work fewer hours than men (Leaker, 2008).

3.1.2. Independent Variables

As trade-related measures, trade openness and commodity concentration are used. Trade openness is defined as the ratio of sum of exports and imports of goods and services to GDP (OECD, 2011), and the trade openness dataset is taken from the World Bank World Development Indicators. According to Becker's Theory of Discrimination, the relationship between trade openness and the gender wage gap is expected to be negative because of the rise in competition. Bhagwati (2004) refers that the international trade, especially exports, have progressively positive effects on women.

Commodity concentration – a Gini-Hirschman Index of concentration dataset is extracted from the United Nations Conference on Trade and Development (UNCTAD). Concentration index of export and import are included separately into the study to see the import and export effects on the gender wage gap in more detail. The commodity concentration, or trade dependency, indicates that the export role of nations is limited to just a few commodities. The countries with more varied exports have more opportunities to respond to fluctuations in the global economy (Meyer, 2007; Beer and Boswell, 2001). According to dependency theorists, trade dependency will have negative effects for countries. Ward (1984), Pampel and Tanaka (1986), and Miller (1999) have foreseen that higher levels of commodity concentration in the world trading system will decrease women's share in economic resources more than men's (Meyer, 2007).

"Foreign direct investment is defined as cross-border investment by a resident entity in one economy with the objective of obtaining a lasting interest in an enterprise resident in another economy" (OECD, 2013).

Foreign direct investment (FDI) provides durable and direct relationships between economies. For this reason, FDI is a key element for international economic integration. Also, it promotes the transfer of technology and know-how across countries. Many economists studied the relationship between FDI and the gender wage gap and found that FDI has an important influence on the gender wage gap because FDI brings greater wages and much better employment opportunities (Aguayo-Tellez, 2014). Oostendorp (2004) claims that globalisation can be measured using aggregate trade and FDI net inflows as a percentage of GDP. Following Oostendorp's (2004) study, FDI net inflows percentage of GDP is used in the analysis and the dataset is obtained from World Bank.

According to Human Capital Theory, education is one of the factors used to explain the gender wage gap. Education attainment variable is used in the analysis as a control variable. Rasekhi and Hosseinmardi (2012) use Human Development Index as a proxy of education level in countries. Following Rasekhi's (2012) study, HDI is used in this analysis, and the dataset is taken from United Nations Human Development Report.

The percentage of female legislators is included as a control variable. The data is measured using the proportion of females in parliament data. The dataset is taken from World Bank World Development Indicator. Social institutions, like parliaments, have a significant role on gender inequality. Odongo (2009) claims that more female legislators are more likely to lower the gender wage gap. In this case, a negative relationship is expected between the percentage of female legislators and the gender wage gap. (Rasekhi and Hosseinmardi, 2012)

3.2. Methodology

Regularly repeated observations on the same individuals necessitate panel data analysis. These individuals can be countries, firms or people (Schmidpeter, 2017b). In this study, the dataset is a balanced panel that includes 240 observations which were observed regularly between 1995 and 2014 for twelve countries (T=20, N=12).

To explain the relationship between the dependent variable and the independent variables, static and dynamic panel data techniques will be used in this study.

3.2.1. Cross-Section Dependence Test and Panel Unit Root Test

Baltagi (2007) refers that in macro panels with long time series (20-30 years), cross-sectional dependence can be more of a problem than micro panels (fewer years and large cross-sections, N>T). When the series exposes shocks, all horizontal cross-sectional units must be analysed to assess whether they are affected by shock equally. Phillips and Sul (2003) claim that if there is sufficient cross-sectional dependence in the data, and this is not considered in the estimation, the efficiency of the estimation can decrease largely. Therefore, the Breusch-Pagan (1980) LM Test is applied for the Cross-sectional Dependence Test. The Breusch-Pagan LM Test is preferred instead of the Pesaran CD Test because time dimension (T) is larger than cross-sectional dimension (N) in our model (Hoyos and Sarifidis, 2006). In this test, null hypothesis is that there is no cross-section dependency. Failing to reject the null hypothesis indicates that variables have a cross-section dependency.

In the regression analyses, data is examined using panel unit root tests to find their stationarity. The reason for applying the Unit Root Test is that our time dimension exceeds our cross-section dimension (Nowak-Lehmann, Herzer, Martinez-Zarzoso and Vollmer, 2006). According to cross-section dependence test results, we perform a Pesaran Cross-Sectional Augmented Dickey-Fuller Test (PESCDAF) as a second generation unit root test, rather than first generation Unit Root Tests. Using Peseran Cross-sectional ADF Test gives more efficient results when variables are cross-section dependent (Pesaran, 2003). Additionally, Pesaran showed that in the Monte Carlo simulations, the CADF test is valid for both N> T and T> N (Pesaran, 2003).

Rejecting the null hypothesis of Pesaran's Cross-sectional ADF Test implies that variables are stationary in panel level, and they can be shown as I (0). Failing to reject the null hypothesis indicates that variables are integrated at order 1, and can be shown as I (1) (Im, 2003).

3.2.2. Static Panel Data Analysis

Panel data models analyse individual effects and/or time effects to handle the observed or unobserved individual effects. These effects can be fixed or random and are examined by fixed or random effect models. (Park, 2011)

3.2.2.1. Fixed Effects Model

The Fixed Effects Model analyses the relationship between the dependent variable and independent variable within an individual. Each individual has their own individual characteristics that can or cannot impact the independent variables. The model presumes that the individual specific effects are correlated to independent variables. The Fixed Effects Model is used to analyse the effect of variables varying over time; the model eliminates the effect of time-invariant characteristics and the net effects of independent variables on the dependent variable can be assessed (Torres-Reyna, 2007).

The Fixed Effects Model can be written as:

$$y_{it} = \alpha_i + \beta_1 x_{it} + u_{it}$$

where y_{it} is dependent variable observed for individual i in time t, α_i, is the unknown intercept for each individual, x_{it} is independent variable, β is the coefficient for the independent variable, and u_{it} is the error term.

3.2.2.2. Random Effects Model

The Random Effects Model assumes that the variation across individuals is random and uncorrelated with the independent variables, unlike the Fixed Effects Model. This assumption allows the time-invariant variables (like gender) to act as explanatory variables (Torres-Reyna, 2007; Schmidpeter, 2017b)

The Random Effects Model can be shown as:

$$y_{it} = \beta x_{it} + \alpha + u_{it} + \varepsilon_{it}$$

where y_{it} is a dependent variable observed for individual i in time t, α is the unknown intercept, x_{it} is independent variable, β is the coefficient for the independent variable and u_{it} is between entity error and ε_{it} is within entity error.

In this study, the Fixed Effects Model is selected to eliminate country differences, and Random Effects Model is employed as a justification model. Besides, The Hausman Test (1978) is performed to verify whether our decision is correct. In the Hausman Test, the null hypothesis is that the Random Effects Model is an appropriate model for the analysis; the alternative hypothesis is the Fixed Effect Model which is suitable for the study.

Following Wolszczak-Derlacz's (2013) study, we define the gender wage gap using natural logarithms of male and female wages:

$$W_{i,t} = \ln W_{m,i,t} - \ln W_{f,i,t}$$

where i denotes country, t is time, m is male and f is female.

Our Fixed Effects Model is written as:

$$W_{i,t} = \alpha_i + \beta_1 \text{Trade Openness}_{i,t} + \beta_2 \text{ConExp}_{i,t} + \beta_3 \text{ConImp}_{i,t} + \beta_4 \text{FDI}_{i,t} + \beta_5 \text{HDI}_{i,t} + \beta_6 \text{Parl}_{i,t} + u_{it}$$

where $W_{i,t}$ shows the gender wage gap, trade $openness_{i,t}$ indicates the ratio of the sum of exports and imports of goods and services to GDP, $ConExp_{i,t}$ is concentration index of export, $ConImp_{i,t}$ is concentration index of import, $FDI_{i,t}$ is foreign direct investment net flows percentage of GDP, $HDI_{i,t}$ is Human Development Index, $Parl_{i,t}$ shows the proportion of females in parliament.

3.2.3. Dynamic Panel Data Analysis

According to the results of Pesaran's CADF, we cannot reject the null hypothesis for independent variables. It means that they are not stationary in panel level. For a dependent variable, we can reject the null hypothesis; it is stationary in panel level. To obtain stationarity for non-stationary variables, taking differences and setting up the models by using the differences can be used as a one of the methods. However, taking first difference to perform static panel data techniques can cause a significant loss of information. Therefore, we use dynamic panel techniques as a different approach to comprehend the impacts of trade liberalisation on the gender wage gap.

Many authors use the Generalised Methods of Moments (GMM) Model for trade and gender wage gap studies. Oostendorp (2004) refers that using GMM would be more appropriate to analyse the impacts of globalisation on the gender wage gap, but the model is not preferred in this study because Roodman (2006) and Pesaran and Smith (1995) claim that the model can create spurious results in the panel data, which has small cross-sectional dimensions (N) and large time dimensions (T).

Thanks to the panel Autoregressive Distributed Lag (ARDL) Model, we can run stationary I (0) and non-stationary I (1) variables together on the same estimation (Pesaran and Shin, 1999) in T>N models. For this reason, panel ARDL Model is more suitable than the conventional panel co-integration tests (Samargandi, 2013). In addition, the panel ARDL Model provides both short-run and long-run estimation results. Also, the model gives results for short-term dynamics for each country. Lastly, the panel ARDL Model involves lags of dependent and independent variables, and this involvement provides consistent coefficients, even in the presence of endogeneity.

The panel ARDL Model is performed by three different estimators: the pooled mean group (PMG), mean group (MG) and dynamic fixed effects (DFE). The three estimators allow us to consider the country-specific heterogeneity issue.

3.2.3.1. Pooled Mean Group (PMG) Model

The PMG Model is a functional intermediate estimator between evaluating separate regressions that allow each coefficient and error variance to vary across groups, and orthodox fixed-effect estimators which presume that all slope coefficients and error variances are the same. The PMG estimator allows short-run coefficients, involving intercepts and error variances to vary across groups, but long-run slope coefficients are limited to homogenous countries.

3.2.3.2. Mean Group (MG) Model

Pesaran and Smith (1995) say that the Mean Group (MG) Model estimates separate regressions for each country. In the MG Model, coefficients are computed as unweighted means of the predicted coefficients for the individual countries. It does not enforce any restrictions. Mean group estimator allows for each coefficient to change and be heterogeneous both in the long-run and short-run (Samargandi, 2013).

3.2.3.3. Dynamic Fixed Effects (DFE) Model

The DFE estimator is very close to the PMG estimator. To provide equality of slope coefficients and error variances across all countries in the long run, the DFE Model imposes restrictions on the slope coefficients and error variances. Also, the DFE Model imposes limits on the speed of adjustment coefficients and the short-run coefficients to be equal too.

Our model will be estimated using these three estimators. Before running the test, the ARDL lag structure is determined performing the Lag Length Criteria Test. The common approach is to use information criteria (IC), such as Akaike IC, Hannan IC and Schwarz IC. After applying the estimators, the Hausman Test will be performed to see which estimator is appropriate. Rejecting the null hypothesis shows that the PMG Model is appropriate for the analysis. Failing to reject the null hypothesis indicates that the MG Model is suitable for the analysis.

Pesaran, Schin and Smith (1999) introduce the ARDL Model (p, q, q, . . , q) i = 1, 2, . . . ,

N, and t = 1, 2, . . . ,T model:

$$y_{it} = \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta_{ij}' x_{i,t-j} + \mu_i + \varepsilon_{it}$$

Pesaran, Schin and Smith (1999) claim that the re-parameterisation of the first equation model is suitable to work:

$$\Delta y_{it} = \phi_i y_{i,t-1} + \beta_i' x_{it} + \sum_{j=1}^{p-1} \lambda_{ij}^{*} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta_{ij}^{*\prime} \Delta x_{i,t-j} + \mu_i + \varepsilon_{it}$$

i = 1, 2, . . . , N, and t = 1, 2, . . . ,T where

$$\phi_i = -(1 - \sum_{j=1}^{p} \lambda_{ij}), \quad \beta_i = \sum_{j=0}^{q} \delta_{ij}$$

j = 1, 2 , . . , p - 1 and j = 1 , 2 , . . , q - 1

$$\lambda_{ij}^{*} = -\sum_{m=j+1}^{p} \lambda_{im}, \quad \delta_{ij}^{*} = -\sum_{m=j+1}^{q} \delta_{im}$$

where y is the gender wage gap, x_{it} (k x 1) is the vector of independent variables including trade openness, export concentration index, import concentration index, foreign direct investment, human development index and female parliament participation. λ and δ represent the short-run coefficients of lagged dependent and independent variables respectively. β are the long-run coefficients. μ_i symbolises the fixed effects; the coefficient of the lagged dependent variables, λ_{it}, are scalars; and δ_{ij} are k x 1 coefficient vectors. ϕ represents the coefficient of speed of adjustment to the long-run equilibrium. P is lag of dependent variable and q is lag of independent variables. Also, Δ indicates that $\Delta y_i = y_i - y_{i-1}$, $\Delta x_i = x_i - x_{i-1}$

4. EMPIRICAL RESULTS

In this section, we examine the empirical findings using panel data techniques in two main parts. In the first part, static panel data analysis is performed using the Random Effects Model and the Fixed Effects Model. In the second one, panel ARDL estimators are applied as a dynamic panel data analysis with the Pooled Mean Group Model (PMG), Mean Group Model (MG) and Dynamic Fixed Effects Model (DFE). As mentioned before, the analysis in this

section is based on the effects of trade liberalisation on the gender wage gap in twelve Latin American countries between 1995 and 2014.

4.1. Cross-Section Dependence Test

As the first step of the analysis, the Cross-section Dependence Test is used to understand whether the horizontal cross-sectional units are affected from shocks equally, and to decide which unit root test is more appropriate for our study.

Table 1
Cross-Section Dependence Test

Variables	Breusch-Pagan LM	Pesaran CD
Gender Wage Gap	0.0000***	0.0000***
Concentration Index(Export)	0.0000***	0.0006***
FDI	0.0000***	0.0813*
Trade Openness	0.0000***	0.0000***
Concentration Index (Import)	0.0000***	0.0000***
HDI	0.0000***	0.0000***
Female Parliament Participation	0.0000***	0.0000***

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

Both the Breusch-Pagan LM Test and Pesaran CD Test are run for comparability purposes, but for our study, the Breusch-Pagan LM Test is considered (T>N). Table 1 shows that null hypothesis is rejected strongly at the 1% significance level, and there is a cross-section dependency for all variables. The results mean that countries are interdependent (Sarafidis and Wansbeek, 2010). Therefore, as we mentioned in the methodology part, Pesaran Cross-sectional ADF Test is used as a second generation Unit Root Test.

4.2. Unit Root Test

Since our panel data spans twenty years, the variables are likely to have unit root (Nelson and Plosser, 1982). To test the order of integration of variables, the Pesaran Cross-sectional Augmented Dicky-Fuller Unit Root Test is employed for static and dynamic panel anlaysis. Although order of integration is not important for the panel ARDL Model, the PESCADF Test is performed to make sure no series exceeds I (1) order of integration in panel ARDL – otherwise the model cannot be run.

Table 2
Pesaran Cross-Sectional Augmented Dicky-Fuller Unit Root Test

Variables	In level	First Difference
Gender Wage Gap	**0.019*****	0.0000***
Concentration Index(Export)	0.116	0.0000***
FDI	0.461	0.0000***
Trade Openness	0.350	0.0000***
Concentration Index (Import)	0.293	0.0000***
HDI	0.825	0.0000***
Female Parliament Participation	0.615	0.0000***

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

Table 2 presents that for the level series, the PESCADF Test rejects the null hypothesis of a unit root for the gender wage gap with the significance level of 1%. For all independent variables, the results fail to reject the null hypothesis. It means that they have unit root. For this reason, the first difference level is controlled, and the PESCADF Unit Root Test rejects the null hypothesis for all independent variables at the 1% significance level. In brief, it can be claimed that all the series are non-stationary in level, except gender wage gap; these series are stationary at first difference level (I (1)).

4.3. Results of Static Panel Data Analysis

According to unit test results, independent variables are not integrated in level; therefore, the method of taking differences of these variables is used to employ static models. In other words, if y_t is a non-stationary vector, the model necessitates to establish the model using Δy_t. To avoid the spurious regression, first differences are taken for all variables (Shrestha and Bhatta, 2017). The Random Effects Model is employed firstly and then the Fixed Effects Model is used in this study.

Table 3
Panel FGLS-Random Effects Test

Variables	Coefficients	t-stat	p-value
ΔTrade Openness	-.001107	1.609	0.108
ΔFDI	.002734	1.212	0.226
ΔConcentration Index (Export)	-.156753	-1.012	0.312

ΔConcentration Index (Import)	-.126917	-0.488	0.625
ΔHDI	.001877	-0.095	0.924
ΔFemale Parliament Participation	**-.002406**	**-3.310**	**0.063***
C	-.003305	-0.629	0.529

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

The Random Effects Model is employed as a justification model and to be able to run the Hausman Test in our analysis. As can be seen in Table 3, all independent variables, except female parliament participation, are insignificant. For the female parliament participation variable, p-value equals 0.063, which is significant at 10% significance level, and the variable is negatively correlated with the gender wage gap.

Table 4

Hausman Test		
Prob	:	**0.0190**

After running the Random Effects Model, the Hausman Test is employed to decide which model is proper for this analysis. As can be observed in Table 4, the null hypothesis is rejected at 5% significance level. More clearly, the Fixed Effects Model is accepted as an appropriate model by using the Hausman Test.

Table 5

Fixed Effects Model

Variables	**Coefficients**	**t-stat**	**p-value**
ΔTrade Openness	**.001216**	**1.750**	**0.082***
ΔFDI	.002691	1.189	0.236
ΔConcentration Index(Export)	-.182767	-1.163	0.246
ΔConcentration Index (Import)	-.109616	-0.416	0.678
ΔHDI	-.001997	0.099	0.921
ΔFemale Parliament Participation	**-.002486**	**-1.896**	**0.059***
C	-.003226	-0.611	0.542

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

The Fixed Effects Model is applied to control country differences between Latin American countries, and result of the Hausman Test has supported our approach. According to Table 5, four explanatory variables are insignificant. The p-value of the female parliament participation variable is 0.059, which is significant at only 10% significance level. The coefficient of this variable is -.002486, and negatively related to the gender wage gap. Trade openness p-value equals 0.082 and it is significant at 10% significance level, and the variable impacts the gender wage gap positively.

In conclusion, two different static panel data models are run to examine the impacts of trade liberalisation on the gender wage gap, and there are only two significant variables which affect the gender wage gap. The female parliament participation variable is significant and has a negative impact on the gender wage gap in both models, which is consistent with Odongo's (2009) findings. Trade openness is significant in only the Fixed Effects Model. The effects of trade openness will be explained below with the dynamic panel data analysis results.

4.4. Results of Dynamic Panel Data Analysis

Optimal Lag Length Test

Table 6

Var Lag Order Selection Criteria

Lag	LogL	LR	FPE	AIC	SC	HQ
0	-1659.020198	NA	0.081206	17.35438	17.47314	17.40248
1	-331.0512163	2545.274	**1.33e-07***	**4.031784***	**4.981886***	**4.416582***
2	-298.0611566	60.82542	1.57e-07	4.198554	5.979996	4.920051
3	-260.3808924	**66.72547***	1.78e-07	4.316468	6.929250	5.374664
4	-227.2764515	56.20858	2.12e-07	4.482046	7.926169	5.876941

To decide the appropriate number of lags required for the panel ARDL Model, the VAR Lag Order Selection Criteria Test is employed, choosing maximum lag order as four. Then, Lag 1 is chosen by considering the FPE, AIC, SC and HQ results. The mechanism of lag order selection criteria is based on the selection of the minimum values of columns as demonstrated with stars in Table 6.

4.4.1. Panel ARDL Model

Before the interpretations of the variables significance, the error correction coefficient (EC) should be considered. The error correction coefficient is showed in the short-run estimation part, and expected to be negative and statistically significant. Table 7 shows that error correction coefficient equals -.6042077. It is highly significant and has a negative sign; this indicates that all variables are co-integrated. More clearly, there is a stable long-run relationship between all the variables. Also, the coefficient shows that the deviation from the long-run equilibrium level of (gender wage gap) the current period is straightened by 60.42% in the next period to bring offset (Dritsakis, 2011).

Table 7 presents the results of the Pooled Mean Group Model estimations using ARDL with Lag 1 structure. The table indicates the long-run and short-run coefficients between gender wage gap and the independent variables, and the speed of adjustment error correction coefficient (EC). In the long-run, as can be seen in the results, trade openness is significant at 1% significance level. Contrary to Becker's Discrimination Theory (1957), it is found that trade openness has a positive impact on the gender wage gap. Oostedorp (2004) refers that he could not find any clear effects of FDI on the gender wage gap for developing/poorer countries, but our results present that FDI is significant at 10% significance level, and has a positive impact on the gender wage gap.

The concentration index of import has the highest coefficient between variables, and is statistically significant at 1% significance level. The coefficient shows that a rise in concentration index of import increases the gender wage gap. Unlike the import results, we could not find any significant results in our models for the concentration index of export despite of Gupta's (2002) findings.

HDI is highly significant and has a negative effect, while FDI, trade openness and the concentration index of import have positive impacts on the gender wage gap. It means that HDI helps to decrease the gender wage gap in the long-run. In the short-run, only HDI is statistically significant, and interestingly, positively related with the gender wage gap.

Table 7

Pooled Mean Group Regression

Long-run Estimations			
Variables	**Coefficient**	**t-statistics**	**p-value**
Trade Openness	**.001679**	**4.76**	**0.000*****
FDI	**.003739**	**1.74**	**0.082***
Concentration Index (Export)	.103821	1.07	0.283
Concentration Index (Import)	**.454091**	**2.87**	**0.004*****
HDI	**-.017371**	**-3.37**	**0.001*****
Female Parliament Participation	.000660	0.73	0.468

Short-run Estimations			
Variables	Coefficient	t-statistics	p-value
EC	**-.6042077**	**-4.74**	**0.000*****
ΔTrade Openness	.0000324	0.02	0.981
ΔFDI	.0023294	0.68	0.494
ΔConcentration Index (Export)	-.0402267	-0.15	0.879
ΔConcentration Index (Import)	-.1490408	-0.25	0.804
ΔHDI	**.0712623**	**3.35**	**0.001*****
ΔFemale Parliament Participation	-.0013304	-0.54	0.591
Constant	.0485281	1.08	0.278

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

To confirm which estimator is appropriate for our panel dataset, the Hausman Test is performed. Rejecting the null hypothesis indicates that the mean group estimator is proper for the analysis. According to the result, the null hypothesis cannot be rejected, which means that pooled mean group estimator is appropriate for our model.

Table 8

Hausman Test		
Chi-square stat	:	11.390
Prob	:	**0.0771**

After running the Pooled Mean Group Model, the mean group and dynamic fixed effects

estimators are applied to see the differences in results. According to Table 9, only HDI is a significant variable, and is significant at 5% significance level, and negatively correlated with the gender wage gap in the long run, but positively correlated with the gender wage gap in the short-run. As is seen from Table 10, the DFE model gives a similar result for HDI in the long-run but the HDI coefficient is weaker than MG model's coefficient. The error correction is highly significant and negative in both the MG and DFE Models. Results of the DFE Model demonstrate that in the short-run, trade openness is significant at 5 per cent significance level and has a positive impact on the gender wage gap. In addition to trade openness, female parliament participation is highly significant and has a negative effect on the dependent variable.

Table 9

Mean Group Estimation: Error Correction Form

Long-run Estimations			
Variables	**Coefficient**	**t-statistics**	**p-value**
Trade Openness	.0022774	0.56	0.577
FDI	-.0015249	-0.11	0.916
Concentration Index (Export)	-.5305003	-0.70	0.486
Concentration Index (Import)	-.0113166	-0.01	0.989
HDI	**-.077882**	**-2.21**	**0.027****
Female Parliament Participation	-.0018813	-0.57	0.569

Short-run Estimations			
Variables	**Coefficient**	**t-statistics**	**p-value**
EC	**-1.03701**	**-7.31**	**0.000*****
ΔTrade Openness	-.001075	-0.53	0.599
ΔFDI	.001894	0.34	0.731
ΔConcentration Index (Export)	.459844	0.68	0.497
ΔConcentration Index (Import)	-.539505	-0.88	0.379
ΔHDI	**.115113**	**3.93**	**0.000*****
ΔFemale Parliament Participation	.000636	0.19	0.852
Constant	.779744	2.88	0.004

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

Table 10

Dynamic Fixed Effects Regression: Estimated Error Correction Form

Long-run Estimations			
Variables	**Coefficient**	**t-statistics**	**p-value**
Trade Openness	-.0010138	-1.30	0.193
FDI	.0009376	0.24	0.808
Concentration Index (Export)	.1061202	0.67	0.505
Concentration Index (Import)	.0182143	0.06	0.951
HDI	**-.026642**	**-2.41**	**0.016****
Female Parliament Participation	.0003759	0.26	0.798

Short-run Estimations			
Variables	**Coefficient**	**t-statistics**	**p-value**
EC	**-.6139673**	**-9.82**	**0.000*****
ΔTrade Openness	**.0012631**	**2.02**	**0.043****
ΔFDI	.0018895	0.84	0.400
ΔConcentration Index (Export)	-.2131916	-1.53	0.126
ΔConcentration Index (Import)	-.1636933	-0.67	0.501
ΔHDI	.0029731	0.17	0.865
ΔFemale Parliament Participation	**-.0036053**	**-2.86**	**0.004*****
Constant	.2956749	3.20	0.001

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

As mentioned before, the panel ARDL Model provides the short-run dynamics for each country. The three biggest Latin American economies (by GDP (PPP)) are chosen to examine, and the last nine countries can be seen in the Appendix B. As can be seen in Table 11, trade openness is significant at 1% significance level only for Argentina. It has a positive effect on the gap, which means that the rise in trade openness increases the gender wage gap. For Brazil and Mexico, only female parliament participation is significant at 5% and 1% significance levels, respectively. The coefficients of this variable for both countries are negatives and shows that female parliament participation declines the gender wage gap in Brazil and Mexico. Other variables are insignificant for both three countries.

Table 11

Pooled Mean Group Regression

	Argentina		**Brazil**		**Mexico**	
	Coefficient	**p-value**	**Coefficient**	**p-value**	**Coefficient**	**p-value**
ΔTrade Openness	**.0045744**	**0.002*****	.00203	0.282	-.0010124	0.840
ΔFDI	-.0050843	0.159	.0040587	0.272	.0024725	0.813
ΔExport	-.3437108	0.726	.2857773	0.386	1.892908	0.097
ΔImport	-.3999316	0.729	.5028823	0.271	3.436606	0.131
ΔHDI	.0499418	0.125	.0138123	0.304	.0897796	0.380
ΔParliament	.0021288	0.640	**.0126614**	**0.018****	**.008422**	**0.008*****

*, **, and *** denotes the rejection of the null hypothesis at 10%, 5% and 1% levels, respectively.

In summary, comparing the results of the Fixed Effects (static) Model and the Pooled Mean Group (dynamic) Model shows that there is only one common significant variable, which is trade openness. In both models, the coefficients are close to each other, and trade openness affects the gender wage gap positively, which means that a rise in trade openness enhances the gender wage gap. Also, it is provided by the DFE model that trade openness is significant in the short-run. In the panel ARDL estimators, only HDI is a common, statistically significant variable in the long-run. The panel ARDL tables provide that HDI decreases the gender wage gap in all three models in the long-run; but interestingly, in the short-run the PMG and MG Models show that HDI is positively correlated with the dependent variable.

5. CONCLUSION

This paper empirically studied the effects of trade liberalisation on the gender wage gap in twelve Latin American countries between 1995 and 2014, using static and dynamic panel data models. The Fixed Effects Model and the panel ARDL-Pooled Mean Group Model have been employed as a baseline model; and the Random Effects Model, Mean Group Model and Dynamic Fixed Effects Model have been performed as justification models. The main findings can be summarised as follows. First, by controlling the country differences using the Fixed Effects Model, it has been found that trade openness enhances the gender wage gap, while other trade-related variables have no significant effects. In addition, it has been deduced that female parliament participation, which has been included proxy of percentages of female

legislators, narrows the gender wage gap and the results of the Random Effects Model and Dynamic Fixed Effects Model short-run estimations have supported this outcome.

Second, error correction coefficient in the pooled mean group shows that all variables are co-integrated in the long-run. In the "The Economics of Discrimination" (1957) study, Becker claims that trade openness brings competition, and competition leads to a decrease in the gender wage gap, but in contrast to Becker's theory, we found that trade openness expands the gender wage gap in the long-run, and has no impact in the short-run. However, the Dynamic Fixed Effect Model shows that trade openness has a statistically significant and positive effect on the gender wage gap in the short-run. As one of the first global studies examining the effects of globalisation on the gender wage gap, Oostendorp (2004) could not find any significant effects of FDI on the gender wage gap in developing countries. Conversely, in our analysis, the PMG Model has provided that FDI has a significant and positive impact on the gender wage gap in the long-run. In addition to FDI results, it has been found that the concentration index of import has a highly significant and positive effect in the PMG Model. It can be interpreted as a higher concentration index of import and FDI widen the gender wage gap. As a control variable, the Human Development Index (HDI), which is a proxy of education level, has been included the analysis, and all three of the panel ARDL models show it is highly significant and negatively correlated in the long-run, which means that the higher level of education lowers the gender wage gap. Nevertheless, both the PMG and MG model present that the HDI variable has a highly positive impact on the gender wage gap in the long-run, unexpectedly.

In this study, the concentration index of export, which can be called commodity concentration or trade dependency, has been used as one of the proxies of trade liberalisation variables. Ward (1984), Pampel and Tanaka (1986), and Miller (1999) state that higher trade dependency brings negative outcomes for countries, and decreases women's share in economic resources more than men's; but no significant effects of the concentration index of exports on the gender wage gap in both static and dynamic models could be found.

According to the results, our findings are compatible with the non-neoclassical approach, which assumes that rising trade openness increases the foreign competition, and this circumstance can cause expansiveness in the gender wage gap (Williams, 1987).

For further studies, the Heckscher-Ohlin model can be used as a different approach to understand the effects of the trade liberalisation on the gender wage gap at different skill levels in Latin America. Korinek (2005) and Wolszczak-Derlacz, (2013) claim that in emerging economies, females generally work in low-skilled jobs and work for lower wages than men; the Heckscher-Ohlin model assumes, lower-skilled labour abundant economies tend to specialize in low-skilled, labour-intensive exports, and this tendency increases the demand for low-skilled labour. All in all, the rising demand for low-skilled labour leads to increases in the low-skilled workforce's wages, and this circumstance decreases the gender wage gap in developing countries.

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APPENDIX A

LIST OF COUNTRIES

ARGENTINA

BOLIVIA

BRAZIL

COSTA RICA

ECUADOR

EL SALVADOR

HONDURAS

MEXICO

PANAMA

PARAGUAY

PERU

URUGUAY

APPENDIX B

ARGENTINA

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-1.187147	.2505164	-4.74	0.000	-1.67815	-.6961436
Tradeopennes D1.	.0045744	.0014951	3.06	0.002	.0016441	.0075047
FDIofGDP D1.	-.0050843	.0036108	-1.41	0.159	-.0121613	.0019927
Export D1.	-.3437108	.9794373	-0.35	0.726	-2.263373	1.575951
Import D1.	-.399932	1.156115	-0.35	0.729	-2.665876	1.866012
HDI D1.	.0499418	.0325433	1.53	0.125	-.0138418	.1137255
Parliament D1.	.0021288	.0045451	0.47	0.640	-.0067794	.011037
_cons	.1807795	.0930582	1.94	0.052	-.0016113	.3631702

BOLIVIA

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.2787367	.1535595	-1.82	0.069	-.5797078	.0222343
Tradeopennes D1.	.0058285	.003023	1.93	0.054	-.0000964	.0117535
FDIofGDP D1.	.0094285	.0049489	1.91	0.057	-.0002712	.0191283
Export D1.	.6270371	.5262106	1.19	0.233	-.4043166	1.658391
Import D1.	-3.60002	1.225062	-2.94	0.003	-6.001098	-1.198943
HDI D1.	-.0637391	.0452974	-1.41	0.159	-.1525204	.0250422
Parliament D1.	-.0037921	.0019489	-1.95	0.052	-.0076118	.0000275
_cons	.0451258	.0419694	1.08	0.282	-.0371327	.1273843

BRAZIL

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.519355	.1047193	-4.96	0.000	-.7246011	-.314109
Tradeopennes D1.	.00203	.0018884	1.07	0.282	-.0016713	.0057312
FDIofGDP D1.	.0040587	.0036958	1.10	0.272	-.003185	.0113024
Export D1.	.2857775	.3293969	0.87	0.386	-.3598285	.9313835
Import D1.	.5028822	.4568016	1.10	0.271	-.3924325	1.398197
HDI D1.	.0138123	.0134493	1.03	0.304	-.0125478	.0401725
Parliament D1.	.0126614	.0053593	2.36	0.018	.0021573	.0231654
_cons	.1560083	.053834	2.90	0.004	.0504956	.2615211

COSTA RICA

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.9754669	.1916428	-5.09	0.000	-1.35108	-.5998539
Tradeopennes D1.	-.0056763	.0029554	-1.92	0.055	-.0114688	.0001162
FDIofGDP D1.	.0104105	.016331	0.64	0.524	-.0215978	.0424187
Export D1.	-.0965248	.1995108	-0.48	0.629	-.4875588	.2945093
Import D1.	-1.141065	.423063	-2.70	0.007	-1.970253	-.3118762
HDI D1.	.2224737	.2114488	1.05	0.293	-.1919582	.6369057
Parliament D1.	-.0031555	.0022501	-1.40	0.161	-.0075656	.0012545
_cons	-.0790179	.0741892	-1.07	0.287	-.224426	.0663903

ECUADOR

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.2655914	.131477	-2.02	0.043	-.5232816	-.0079012
Tradeopennes D1.	.0042064	.002278	1.85	0.065	-.0002584	.0086712
FDIofGDP D1.	.0015751	.0076085	0.21	0.836	-.0133374	.0164876
Export D1.	-.6694196	.3490595	-1.92	0.055	-1.353564	.0147245
Import D1.	-2.076233	.6376288	-3.26	0.001	-3.325962	-.8265034
HDI D1.	.0474877	.066761	0.71	0.477	-.0833614	.1783369
Parliament D1.	-.0016737	.002174	-0.77	0.441	-.0059347	.0025873
_cons	.0480937	.0314992	1.53	0.127	-.0136435	.1098309

EL SALVADOR

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.2821652	.1410696	-2.00	0.045	-.5586565	-.0056738
Tradeopennes D1.	.0033134	.0022927	1.45	0.148	-.0011802	.0078069
FDIofGDP D1.	.0040716	.0045343	0.90	0.369	-.0048155	.0129587
Export D1.	.2748862	.5026469	0.55	0.584	-.7102836	1.260056
Import D1.	-1.965787	1.136596	-1.73	0.084	-4.193473	.2619
HDI D1.	-.0083534	.0570566	-0.15	0.884	-.1201823	.1034754
Parliament D1.	-.002914	.0046902	-0.62	0.534	-.0121066	.0062786
_cons	.0222382	.0275989	0.81	0.420	-.0318546	.076331

HONDURAS

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.3418332	.1293161	-2.64	0.008	-.5952881	-.0883784
Tradeopennes D1.	-.0008774	.0026689	-0.33	0.742	-.0061084	.0043537
FDIofGDP D1.	.0262104	.0163999	1.60	0.110	-.0059327	.0583536
Export D1.	.1371155	1.174473	0.12	0.907	-2.164809	2.43904
Import D1.	.4203087	1.97928	0.21	0.832	-3.45901	4.299627
HDI D1.	.1032068	.1525908	0.68	0.499	-.1958658	.4022793
Parliament D1.	-.010493	.004334	-2.42	0.015	-.0189875	-.0019986
_cons	-.0625313	.0388549	-1.61	0.108	-.1386854	.0136229

MEXICO

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.1564484	.1780213	-0.88	0.379	-.5053637	.192467
Tradeopennes D1.	-.0010124	.0050246	-0.20	0.840	-.0108604	.0088356
FDIofGDP D1.	.0024725	.0104398	0.24	0.813	-.0179892	.0229343
Export D1.	1.892908	1.142186	1.66	0.097	-.3457357	4.131552
Import D1.	3.436606	2.277492	1.51	0.131	-1.027196	7.900407
HDI D1.	.0897796	.1023052	0.88	0.380	-.1107349	.2902942
Parliament D1.	.0084222	.0031563	2.67	0.008	.0022359	.0146084
_cons	.0032986	.0448489	0.07	0.941	-.0846036	.0912007

PANAMA

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-1.068237	.255721	-4.18	0.000	-1.569441	-.567033
Tradeopennes D1.	-.0018543	.0006622	-2.80	0.005	-.0031522	-.0005565
FDIofGDP D1.	.0000713	.0015962	0.04	0.964	-.0030572	.0031998
Export D1.	.1909782	.163104	1.17	0.242	-.1286998	.5106561
Import D1.	-.2175662	.2234899	-0.97	0.330	-.6555983	.2204659
HDI D1.	.1102433	.0663855	1.66	0.097	-.0198698	.2403564
Parliament D1.	.0048321	.0017562	2.75	0.006	.00139	.0082742
_cons	-.2252297	.1136328	-1.98	0.047	-.4479459	-.0025136

PARAGUAY

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.3470995	.2925772	-1.19	0.235	-.9205404	.2263414
Tradeopennes D1.	.0039167	.0021271	1.84	0.066	-.0002524	.0080857
FDIofGDP D1.	-.0213936	.0163571	-1.31	0.191	-.053453	.0106658
Export D1.	-.8285229	.3690734	-2.24	0.025	-1.551893	-.1051524
Import D1.	-.8327077	1.309051	-0.64	0.525	-3.3984	1.732984
HDI D1.	.1434602	.0650071	2.21	0.027	.0160486	.2708718
Parliament D1.	-.0205312	.0113843	-1.80	0.071	-.0428441	.0017816
_cons	.021749	.0342251	0.64	0.525	-.0453311	.088829

PERU

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-1.458404	.224501	-6.50	0.000	-1.898418	-1.01839
Tradeopennes D1.	-.0099062	.0034098	-2.91	0.004	-.0165893	-.003223
FDIofGDP D1.	-.0114719	.0063832	-1.80	0.072	-.0239827	.001039
Export D1.	-1.94832	.5739066	-3.39	0.001	-3.073156	-.8234838
Import D1.	3.44308	1.394075	2.47	0.014	.7107435	6.175416
HDI D1.	.0779949	.0387974	2.01	0.044	.0019535	.1540363
Parliament D1.	-.0016684	.0028482	-0.59	0.558	-.0072508	.0039139
_cons	.405971	.1125025	3.61	0.000	.1854701	.6264718

URUGUAY

D.genderwagegap	Coef.	Std. Err.	z	P>z	[95% Conf.	Interval]
ec	-.3700087	.1073019	-3.45	0.001	-.5803165	-.1597009
Tradeopennes D1.	-.0041544	.0013272	-3.13	0.002	-.0067558	-.0015531
FDIofGDP D1.	.0076041	.0037313	2.04	0.042	.0002908	.0149174
Export D1.	-.0049252	.2868119	-0.02	0.986	-.5670663	.5572158
Import D1.	.6419448	.2450491	2.62	0.009	.1616573	1.122232
HDI D1.	.0688403	.0215437	3.20	0.001	.0266154	.1110652
Parliament D1.	.0002188	.0024458	0.09	0.929	-.0045749	.0050124
_cons	.0658521	.0445035	1.48	0.139	-.0213731	.1530772