Female Intensity, Trade Reforms and Capital Investments in Colombian Manufacturing Industries: 1981-2000

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Abstract

We exploit a natural experiment provided by the trade liberalisation that occurred in Colombia at the beginning of the 1990s to see its possible effects on the gender composition of the workforce across manufacturing industries. In order to account for the effects of changes in capital technology, our empirical strategy controls for different types of capital stock per worker (namely, machinery, office equipment and transport equipment) within a fixed-effects instrumental-variables framework in which estimates drawn from a variety of instruments are compared. We also include a concentration index variable in order to account for changes in the degree of market power. Our findings confirm that increasing levels of trade openness in the terms of both import penetration and export orientation tend to be associated with higher shares of female employment although this effect appears to be differentiated in terms of skill level. Equally we find that manufacturing industries with higher levels of industry concentration tend to have lower female shares of jobs. Our variables for different types of the stock of capital per worker suggest that machinery and office equipment are associated with higher shares of female jobs, particularly in the white-collar workers category.

Keywords: female intensity, trade liberalisation, panel data

JEL classification: J16, J82

1. Introduction

The process of trade liberalisation in developing countries has taken place at the same time that their labour markets witnessed an increase in female labour force participation to historically unprecedented levels. The effects of trade as well as other economic policies are expected to have a differentiated effect on women due to asymmetries in the distribution of rights over economic resources, as well as segregated roles in relation to both the market economy and within the household (Fontana, 2003). Although the increase in female employment over the last decades is the result of long-term development trends pertaining to demographic and cultural change, there is also a concern in the literature to understand the effects of trade reforms and other economic policies on labour market outcomes from a gender perspective.

An increasing body of economic literature has emerged in which the interactions between trade and gender differences in the labour market have been explored. From an economic perspective, trade liberalisation might affect employment dynamics by gender in at least four different ways. First, the opportunities for increasing exports, as well as competition in the form of imported goods, have both the potential of changing gender differences in the labour market if women are concentrated in sectors more exposed to trade (Collier, 1994). Second, trade liberalisation may change the relative prices of imported technology and capital goods in developing countries. Some studies indicate strong complementarities between technology and female labour (Galor and Weil, 1996, Weinberg, 2000, Welch, 2000). Third, according to the "taste for discrimination hypothesis" formulated by Becker (1957), any policy measure towards increased competition is likely to reduce the extent of discrimination against women and ethnic minorities in the labour market. A number of empirical studies have tried to identify the effects of trade policies on the unexplained portion of the gender wage gap that can be attributed to discrimination (Artecona and Cunningham, 2002, Black and Brainerd, 2004, Oostendorp, 2009, Reilly and Vasudeva-Dutta, 2005). Fourth and last, as a counterpart to Becker's hypothesis, increasing competition arising from trade liberalisation might weaken the bargaining position of women in female-intensive industries (see: Williams and Kenison, 1996, Williams, 1987, Darity and Williams, 1985). Berik et al. (2004) found in the case of Korea and Taiwan supportive evidence for this hypothesis.

Most of this literature has focused on the effects of trade on gender wage differences while the effects on the gender composition of employment have received less attention. The experience from developed economies indicates that both trade and industrialisation are closely interrelated to the gender composition of economic activities. For instance, Goldin (quoted in Galor and Weil (1996)) indicates that the necessity for fine motor skills in textiles during the industrialisation in the United Kingdom and the United States, and more recently in the electronics industry in Asian economies, represent examples of absolute and comparative advantage of female over male labour along the pathway of economic development. However, there is still a vacuum in the existing knowledge on how trade liberalisation, as well as the industrialisation process, is affecting the gender composition of employment across manufacturing activities in developing countries.

This paper provides an empirical application to identify the effects of trade on the gender composition of employment across manufacturing industries in Colombia. In particular, we exploit a natural experiment of trade liberalisation which took place in this country at the beginning of the 1990s to assess its possible effects on the gender composition of the workforce across industrial activities. In order to account for the effects of changes in capital technology, our empirical strategy controls for different types of average stock of capital per worker (namely, machinery, office equipment and transport equipment) across manufacturing industries. We implement a panel data strategy based on fixed-effects instrumental variables (FE-IV, hereafter) in order to address potential endogeneity problems on some of the regressors. Our findings confirm that increasing levels of trade openness in the terms of both, import penetration and export orientation tend to be associated to higher shares of female employment although this effect appears to be differentiated in terms of skill level. Equally we find that manufacturing industries with higher levels of industry concentration tend to have lower female shares of jobs. Our variables for different types of the stock of capital per worker suggest that machinery and office equipment are associated with higher shares of female jobs, particularly in the white-collar workers category. The remainder of this paper is organised as follows. The following section presents the literature review and a third provides some background information for the country describing the data used for this empirical application. The fourth reports the econometric results in the light of the existing literature. The fifth and last section offers some concluding remarks.

2. Literature review

Trade theory provides some explanations for the effects of increased foreign competition on employment patterns between men and women. In particular, the Stolper-Samuelson theorem within the Heckscher-Ohlin-Samuelson trade model indicates that trade liberalisation increases the demand for, and the returns to, the most abundant factor of production. Thus, if women constitute the abundant factor in exporting industries boosted by trade, it is possible that their returns will grow faster than those of male workers and, in this way, the gender wage gap will be reduced. Wood (1991) provides one of the first studies to survey the relationship between trade and the gender composition of the labour force in developing countries. The author investigated the effects of trade on female employment ratios in manufacturing for a sample of countries and found that increasing exports to industrialised economies are associated with higher relative demand for female intensive goods from developing countries. But at the same time, Wood (1991) found that trade flows of manufacturing goods from the 'South', which to a great extent are intensive in female labour, were not associated with reductions in relative demand for female labour in manufacturing industries from developed countries.

In a more recent study, Chamarbagwala (2006) studied the effects of trade liberalisation on the gender (and skill) wage gap in India using a non parametric methodology developed by Katz and Murphy (1992). In addition to the Stolper-Samuelson proposition just mentioned above, this work tests the four "Skill Enhancing Trade (SET)" hypotheses proposed by Robbins (1996 –referenced in Chamarbagwala, 2006, see note 4) which indicate that trade liberalisation promotes, through different channels, the demand for and wages of skilled workers in developing countries. Chamarbagwala (2006) finds increasing skill premiums and diminishing gender wage gaps in India, the former being consistent with "skill-biased technical change" (*cfr.*, Acemoglu, 2002) and the latter due to a relocation process of female and male workers between traded and non-traded sectors.

From a theoretical point of view, trade liberalisation might affect employment dynamics by gender in at least four different ways. First, as long as women and men are imperfect substitutes in production, increased trade may affect the relative demand (as well as relative wages) of one gender group with respect to another. New opportunities arising from increasing exports, as well as more competition from imported goods, have the potential for both changing gender differences in the labour market if women are concentrated in sectors more exposed to trade (Collier, 1994). Second, trade liberalisation may change the relative prices of imported technology and capital goods in developing countries. For instance, the introduction of more capital intensive production processes in semi-industrialised economies might open new employment opportunities for women as physical strength becomes less relevant. In this sense, some studies indicate strong complementarities between technology and female labour (Galor and Weil, 1996, Weinberg, 2000, Welch, 2000). Third, according to the "taste for discrimination hypothesis" formulated by Becker (1957), any policy measure inducing increased competition is likely to reduce the extent of discrimination against women and ethnic minorities in the labour market. As long as gender discrimination is costly, increasing competition from imported goods and services is likely to increase competitive forces and reduce the scope for non-competitive behaviour in the form of discrimination (Artecona and Cunningham, 2002, Black and Brainerd, 2004). Fourth and lastly, as a counterpart to Becker's hypothesis, increasing competition arising from trade liberalisation might weaken the bargaining position of women in female-intensive industries (see: Williams and Kenison, 1996, Williams, 1987, Darity and Williams, 1985). Local entrepreneurs might respond to increasing imports with cost-cutting strategies to reduce labour costs and this might affect women if they are more concentrated in formerly protected industries. In what follows in this section, we review this literature with respect to these four hypothetical effects of trade on women.

2.1 Men and women as imperfect substitutes

Trade may have a differentiated effect in terms of gender because women and men may be imperfect substitutes. A recent article by Qian (2008) on the impact of tea prices and gender imbalance in China illustrates how female workers in this country have a comparative advantage in the production of that crop as "picking requires the careful plucking of whole tender leaves [which] gives adult women absolute and comparative advantages over children and men". In this case, women's comparative advantage is magnified by the fact that both the price and quality of tea leaves increases significantly with leaf tenderness. In another study for India, Rosenzweig (2004 –quoted in Duflo (2005)) documents how the choice of language instruction for boys and girls during school instruction in Mumbai entailed skill differences which became highly valuable after economic liberalisation in India over the 1990s. According to this study, low caste girls were more likely to attend English speaking schools while boys were more likely to attend Marathi-speaking schools. With the increase of service industries such as telemarketing and software as a result of economic liberalisation, the labour market returns of possessing English as a second language skill exhibited a dramatic increase. As a result, low-caste women enjoy a comparative advantage in the export-oriented service sector of Mumbai with respect to their male counterparts, with more possibilities for better wages and, to some extent, more opportunities for social mobility. Another example of imperfect substitution between men and women is provided by Goldin (quoted in Galor and Weil (1996)) who argues that the process of industrialisation is responsible for the increase in demand

(and thereby, wages) of female labour. The necessity for fine motor skills in textiles during the industrialisation in the United States and United Kingdom and, more recently in the electronics industry in Asian economies, provide examples of absolute and comparative advantage of female over male labour along the pathway of economic development.

2.2 The role of technology and women

Trade liberalisation has the potential of bringing about technological change or, at least, reconversion towards more capital-intensive production processes in semi-industrialised countries as imported machinery and equipment become cheaper. In the same vein, the increase in the number of foreign-owned firms might lead to the introduction of more capital-intensive production processes compared to local firms. In both cases, the question is whether the increase in capital per worker enhances the participation of women in the labour market.

Galor and Weil (1996) formalise a microeconomic model in which women and men are imperfect labour substitutes. The model has multiple steady-state equilibriums, one in which the economy has low capital per worker, high fertility rates, low female labour participation and low wages; at the other extreme, there is another equilibrium characterised by high capital per worker, low fertility rates and high relative female wages. The authors argue that countries might converge to a development trap of high fertility, low capital per worker and low productivity in which low female wages induce women to a low labour participation/high fertility outcome which in turn further reduces capital per worker. As the process of economic development allows some increase in the capital per worker, physical strength becomes less relevant and there is more scope for female labour participation. Increasing labour demand (and wages) for nonphysical strength skills, which can be supplied by women, entail an opportunity cost to childbearing as well as an incentive for reduced fertility. This in turn permits the accumulation of more capital per worker and this reinforces a cycle of higher demand for female labour, higher female wages, higher female labour participation and, ultimately, lower fertility.

In the case of the United States, Welch (2000) reviews the trends in relative female/male wages as well as wage inequality among men. His evidence is persuasive in favour of the hypothesis according to which women enjoy an advantage in cognitive skills. He finds that behind both the increasing trend in women's relative wages and growing income inequality among men in the United States there is a common factor: a growing demand for intellectual skills. Compared to average men, male workers at the top of the income distribution as well as women in general are relatively more intensive in such skills. Thus, the increase in demand for skilled labour shifts the income distribution in favour of these two groups. In the case of women, increasing schooling levels, as well as less frequent temporal withdrawals from the labour force due to maternity, might explain not only the improvement in female relative wages but also their higher work force share in a number of occupations.

In another study for the United States, Weinberg (2000) finds that the increasing use of computers accounts for about one half of the increase in demand for female workers, a finding that is in line with the hypothesis of imperfect substitution between female and male work noted above. He also proposes a microeconomic model in which the introduction of computers not only increases the share of female employment in a number of industries but also favours their demand in non-computer jobs by changing production processes in ways that are both less physically demanding and less hazardous. Based on his empirical findings, Weinberg (2000) concludes that a substitution process between highly skilled women and less skilled men, as previously documented in other studies, might be explained by the increase in computer use.

2.3 Trade, competition and gender discrimination

In 1957, Becker formulated an influential hypothesis in relation to labour market discrimination known as 'the taste for discrimination'. According to this hypothesis, discriminating employers and their employees are willing to sacrifice part of their income or rents in order to avoid working with people possessing some characteristics (Becker, 1957). The implication of Becker's hypothesis is, therefore, that the scope for non-competitive behaviour of firms can only be afforded through some sort of monopolistic rents which permit them to exercise their taste for discrimination against minorities in the labour market. In this sense, any policy measure towards enhanced competition should lead to the elimination of these rents and, therefore, to a reduction in the scope for costly discrimination.

There is a growing body of empirical literature in which Becker's formulation has been tested by linking trade liberalisation and gender outcomes in the labour market. This literature has focused on the effects of increased competition from trade on the magnitude of the inter-industry gender wage gaps while the effects on the gender composition of employment across economic activities have merited little attention. Two studies with a similar econometric strategy, Artecona and Cunningham (2002) and Black and Brainerd (2004), investigated the effects of increasing trade and the degree of industry concentration on the 'residual gender wage gap'.[1] The former study

[1] The residual gender wage gap is estimated as "the gender wage gap that remains after one controls for differences in education and potential labour market experience" (Black & Brainerd 2004: 544).

used data from Mexico while the latter used data from the United States. Both studies find evidence that the residual gender wage gap fell more in industries with high degree of concentration which were exposed to increased levels of foreign competition. In the same vein, Reilly and Vaseudeva (2005) investigated the relationship between trade-related measures (i.e., tariffs and imports and exports shares) on the inter-industry gender wage gap with microdata for India and found some evidence that more open sectors in that country tend to report lower levels of wage discrimination against women. In another application for Mexico, Aguayo-Téllez et al. (2010) found that trade liberalisation in this country favoured the creation of female employment in export-oriented industries at the same time that labour reallocation across sectors explains about two fifths of the increase in the female wage bill share. One of the few studies using cross sectional data is Oostendorp (2009), who investigates the effects of trade and foreign direct investment (FDI) on the gender pay gap across 161 occupations in 83 countries.[2] This study suggests that the occupational gender wage gap tends to decrease with log GDP per capita, trade and net inflows of FDI but only for richer countries while the effect on poorer countries is not statistically significant. These findings lead Oostendorp (2009) to conclude that this evidence is in line with Boserup's (1970) hypothesis according to which gender discrimination is inversely related to the level of economic development.

As noted above, the effects of trade on the gender composition of particular occupations have not yet been extensively surveyed. Most of the empirical literature has focused on the effects of trade on the gender wage gap while the implications in

[2] The dataset used in this study is the *ILO October Inquiry*, collected annually by the International Labour Organisation. It contains information on wages, earnings, and hours of work for occupations defined along the International Standard Classification of Occupations of 1968 at four digits.

terms of gender based industry segregation is yet to receive the same empirical attention. In this context, we should note Becker's assertion that

> If an individual has "taste for discrimination" he must act as if he were willing to pay something, either directly or in the form of a reduced income, to be associated with some persons instead of others (Becker, 1957: 14p.).

Here we find a segregation dimension in which discrimination not only involves a monetary cost in terms of "reduced income" but also encompasses a compositional dimension of the labour force which should be reflected in a disproportionately smaller share of women (or minority) workers in discriminating industries. In other words, as the economy becomes more liberalised, gender industry segregation should decrease in formerly protected sectors as their rents to indulge in gender discrimination shrink.

2.4 Trade and the bargaining position of women in the labour market

There are also alternative interpretations for the effects of trade on gender discrimination in the labour market. In a study for Korea and Taiwan, Berik et al. (2004) find a positive association between gender wage discrimination and increased levels of foreign competition in concentrated industries. The authors indicate their evidence supports a non-neoclassical hypothesis (see: Williams and Kenison, 1996, Williams, 1987, Darity and Williams, 1985) according to which increased levels of trade competition push employers to cost-cutting strategies that lessen the bargaining position of female and ethnic minority workers. A similar proposition is put forward by Seguino (2000) who argues that, in the case of semi-industrialised countries, "gender inequality has a positive effect on technical progress and growth" as low female wages provide a comparative advantage for export industries to succeed and earn the foreign currency to purchase imported capital goods, intermediate inputs and technology.

These causation links subsequently lead to reinforcing and self-fulfilling cycles of export growth, technical progress and, ultimately, economic growth. Her econometric estimates from a panel of semi-industrialised middle-income countries provide evidence of a positive relationship between gender income inequality and economic growth via two channels: (i) increased exports, technological change and growth and (ii) more investment. It should be noted that although Berik et al. (2004) and Senguino (2000) are implicitly assuming an opposite direction in the causality relationship between trade and gender wage discrimination, they ultimately concur in the notion that increasing competition arising from globalisation weakens the bargaining position of female workers in export oriented industries.

3. Background and data: trade liberalisation and labour markets in Colombia

3.3.1 Female share of jobs in manufacturing industries

As in other developing countries, Colombia has experienced a remarkable increase in female labour participation over the last decades. Between 1990 and 2004, female labour participation for the seven largest urban areas rose from 43.3% to 55.9% (Isaza et al., 2007). A number of factors have been cited in the literature to explain this trend. First, demographic change coupled with a smaller number of children per household has increased female labour participation in this country (Arango and Posada, 2002, Tenjo and Ribero, 1998). Second, increased educational levels amongst the female population have not only increased their probability of labour participation (Arango and Posada, 2002) but have also influenced female aspirations in terms of professional success (Gilbert, 1997). Lastly, the third factor is economic change (more closely associated with the reforms), which according to Farné, (cited in Gilbert, 1997) has encouraged the development of new occupations that fit both the skills and the social role of women. There is also some agreement in the Colombian literature that the growing labour force participation of secondary family members during the 1990s (mainly women) was motivated by an added worker effect exacerbated by adverse circumstances in the economy at the end of this decade (Isaza, 2002, Isaza, 2006, Santa María and Rojas, 2001, Tenjo and Ribero, 1998).[3]

Employment estimates of the female share of jobs across manufacturing industries for this empirical application are based on data from the Annual Manufacturing Survey (AMS hereafter) administered by the National Statistical Administrative Department (DANE, from its initials in Spanish). The survey can be considered as a census in the sense that it is gathered annually amongst nearly all manufacturing enterprises with more than ten workers since 1975. The economic classification under which the survey was collected from 1981 to 2000 is the International Standard Industrial Classification –ISIC, Rev. 2. Data for subsequent years were gathered using the ISIC Rev.3 which renders unfeasible comparisons with previously collected data. Figure 1a displays the total number of both, female and male workers across two broad categories, white collar and blue collar. This broad characterisation, on which we base subsequent

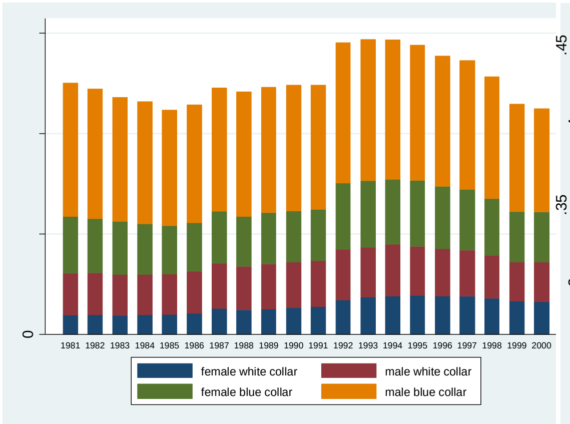
[3] It is noted that the long term trend of increasing real wages may have played an important role in the increasing female labour participation reported in urban Colombia. According to figures from Isaza *et al.* (2007), mean labour incomes rose 21.3% among men and 8.8% among women in the seven largest cities of this country between 1990 and 2004. Although it has not been found specific research on this regard for urban areas of this country, growing female earnings in the labour market may have entailed higher opportunity costs to households' fertility and, thus, increased the participation of women in the labour market. This interpretation is in line with the formulation given by Welch (2000) for the United States where the growing demand for intellectual skills explains not only the improvement in female wages but also their higher female labour force participation.

analyses, is preferred to other dis-aggregations of the labour force as the AMS was subject to changes in the questionnaires over the years analysed here regarding the classification of workers. It should be observed that other divisions of the labour force, namely by skill, hierarchical and contractual status, are not possible for the whole time period from 1981 to 2000. From the figures presented in Figure 1a, we observe a stagnation pattern in the employment dynamics of Colombian manufacturing industries for all groups analysed here where only in the case of female white collar workers is there an absolute increase in the number of jobs between the beginning of the 1980s and the end of 1990s. This sluggish pattern in employment growth could be attributed to a number of factors including an increased exit rate of plants after the introduction of trade liberalisation reforms introduced in 1990 (Eslava et al., 2009), weaker demand for Colombian manufactured goods internally due to a severe economic downturn at the end of the 1990s, as well as a less competitive position of Colombian manufacturing exports originated due to the appreciation of the Colombian currency for most of that decade (Ocampo et al., 2004). Goldberg and Pavcnik (2003) argue that labour market rigidities (rather than trade liberalisation) are also a major factor contributing to the informalisation of urban employment –and thus, the stagnation of formal employment in manufacturing firms over the 1990s.

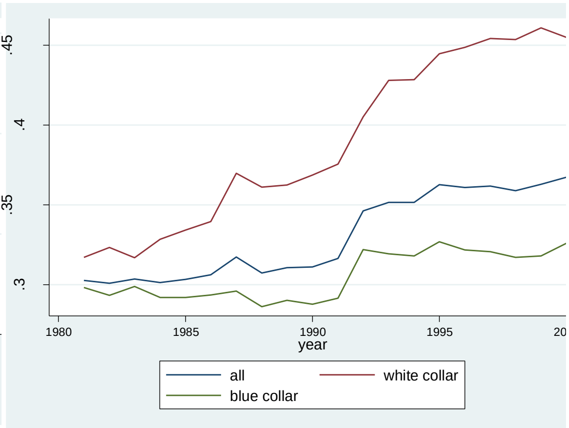
The same figures provide the basis for the calculation of the percentage of female jobs by skill level in manufacturing (see Figure 1b). They indicate that the female share of jobs for all workers rose from around 30 per cent at the beginning of the 1990s to more than 36 per cent from 1995 onwards. This increase was more pronounced amongst white collar workers as their share of female jobs increased from 31.7 percent in 1981 to 45.5 per cent in 2000 compared to a more modest rise from 29.8 per cent to 32.6 per cent in the case of blue collar workers over the same years. These trends are in line with the findings in the literature reviewed in section 3.2, above, according to which increasing female labour force participation is concomitant with the process of economic development.

Figure 1: Number of jobs and gender composition of employment across white and blue collar workers and gender in all manufacturing industries, Colombia: 1981-2000

a) Number of jobs



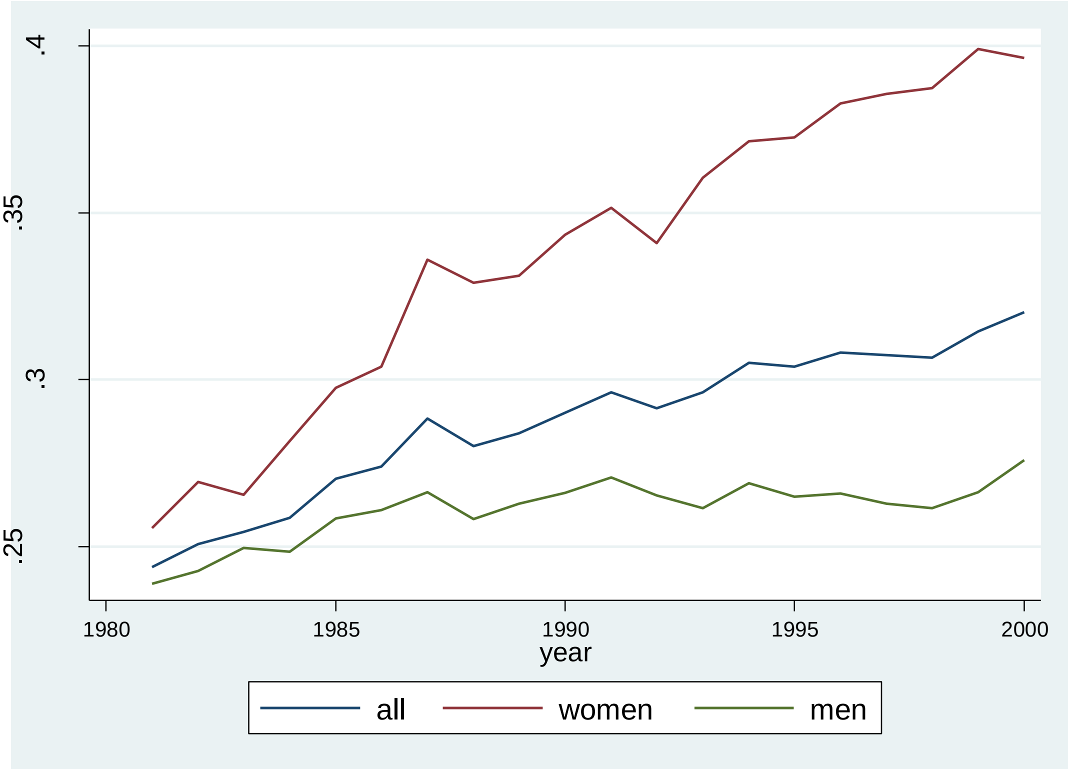
b) % of female jobs



Own estimates based on Annual Manufacturing Survey microdata.

The structure of manufacturing employment in Colombia has also experienced a structural transformation in terms of the skill composition of the labour force over the years analysed here. Employment figures from the AMS indicate that the percentage of white collar jobs has grown for both men and women although, this increase has been more pronounced amongst the latter (see Figure 2). These trends suggest that the process of economic development in Colombia has favoured a structural transformation of the manufacturing employment composition by skill level in which the increasing proportion of white collar workers is benefiting on the margin the incorporation of more women into the manufacturing labour force. This finding could be rationalised in terms of the literature reviewed in section 3.2.2, above (Galor and Weil, 1996, Welch, 2000), according to which the incorporation of technology in production processes is complementary to both, the demand of skilled workers and female labour.

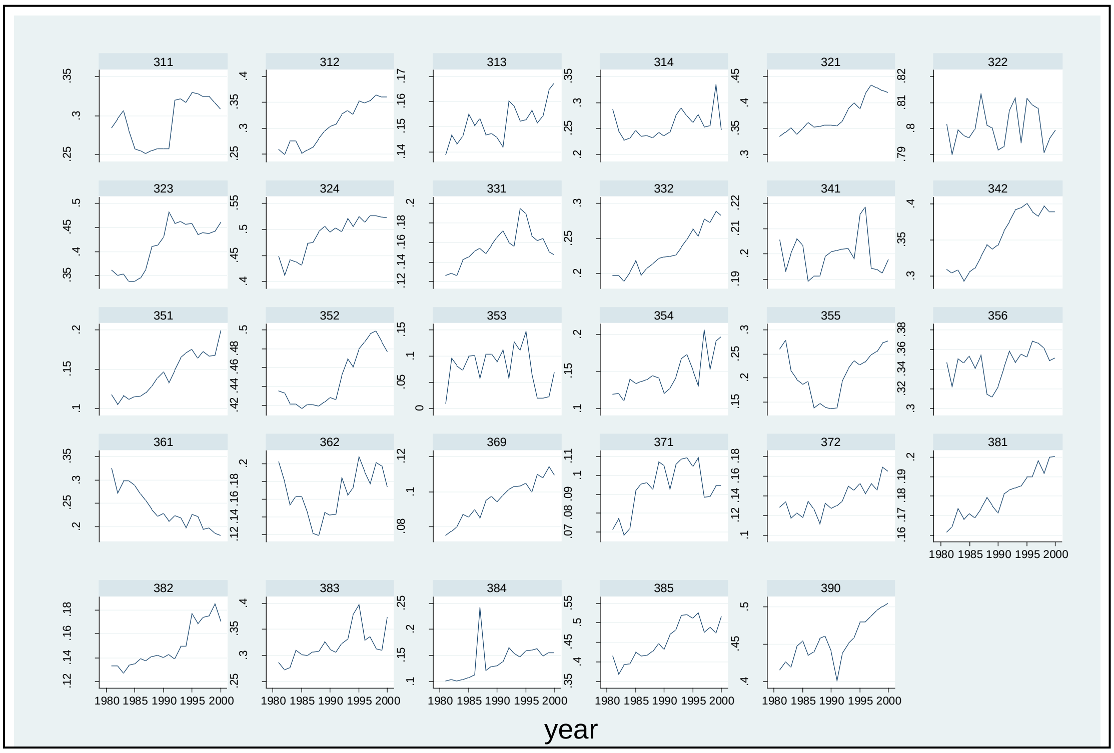
Figure 2: % white collar jobs by gender in all manufacturing industries, Colombia: 1981-2000



Own estimates based on Annual Manufacturing Survey microdata.

The increasing proportion of women reported in Figure 1b above, can also be plotted across 29 manufacturing industries using the ISIC Rev. at three digits (see Figure 3). With the exception of *353- Petroleum refineries* and *361- Pottery, china and earthenware*, all other manufacturing industries have increased the share of female workers within their labour force over these years. They indicate also that most of the industries with the highest female intensity over most years are those related to the textile-clothing-footwear production chain, this is, *322- Wearing apparel, except footwear*, *324- Footwear*, *323- Leather and products of leather* and, *321- Textiles*. These could be characterised as light industry in which production processes are intensive in both female labour and fine motor skills. Other industries have also experienced important increases in the female share of jobs. This is the case of *385- Measuring & controlling equipment*, *312- Food for animals*, and *342- Printing, publishing and allied* *industries* where most of the increment in the proportion of women workers took place in the form of more jobs into the white-collar category.

Figure 3: Proportion of female jobs across manufacturing industries, Colombia: 1981-2000



Own estimates based on Annual Manufacturing Survey microdata.

3.2 Tariffs and trade

Trade reforms in Colombia at the beginning of the 1990s evolved around two elements. The first one was the signing of trade agreements with México and Chile, on the one hand, and with the Andean countries of Venezuela, Ecuador, Peru and Bolivia, on the other. The second element was a reduction of the protective structure. According to Attanasio et al. (2004), Goldberg and Pavcnik (2005b, 2005a) and Jaramillo and Tovar (2006), one of the interesting features of Colombia is that this country did not participate in the GATT negotiations for the reduction of trade tariffs, so the level of protection was very high before the reforms. The removal of trade barriers was started in 1990 with the idea of a gradual approach over a time horizon of more than three years including the elimination of non-tariff barriers and reductions in both the number and level of import tariffs which were assumed to be complemented with a policy of exchange rate depreciation. Macroeconomic circumstances such as high inflation and a dramatic increase in the inflow of foreign capital, besides a reduction in trade flows (both, imports and exports), compounded a scenario in which Colombian authorities decided to speed up the liberalisation process. Thus, the initial liberalisation schedule for 1994 was completed in terms of non-tariff barriers and import tariffs by the end of 1991 (Edwards, 2001).

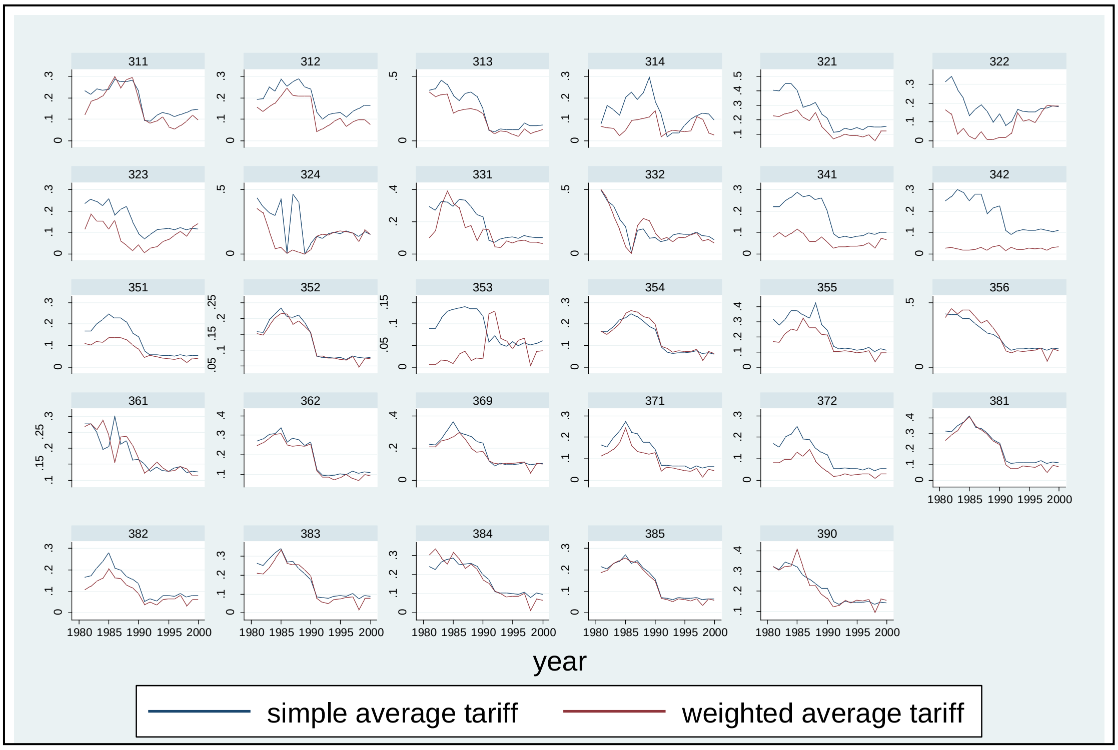
In order to measure the degree of trade openness in Colombia, we use in this empirical application a number of trade measures including import tariff data from the National Planning Department. Import tariffs were originally reported at eight-digit level according to the Nandina[4] classification. For expositional purposes of this analysis, we collapsed these data into 29 sectors defined by the ISIC Rev.2 at three-digit level in order to match it with the employment data presented in section 3.3.1, above (see Figure 4). According to these estimates, weighted average import tariffs for all manufacturing industries fell from 16.9 per cent in 1981-1984 to 6.4 per cent in 1997-2000.[5] The largest reductions on weighted tariffs over these years (all of which were more than 20 percentage points) were reported on *356- Plastic products*, *313- Beverage industries*, *384- Transport equipment*, *381- Fabricated metal products* and, *332- Furniture and fixtures*. Some studies for this country suggest that industries with a high intensity of unskilled labour were more protected before the reforms and thus, experienced the largest reductions in tariffs during the liberalisation period (Attanasio

[4] This is a harmonised trade classification for Andean countries.

[5] Weights are based on imports value in US dollars.

et al., 2004, Goldberg and Pavcnik, 2003, Goldberg and Pavcnik, 2005b, Goldberg and Pavcnik, 2005a, Jaramillo and Tovar, 2006).

Figure 4: Simple and weighted average tariffs across manufacturing industries, Colombia: 1981-2000

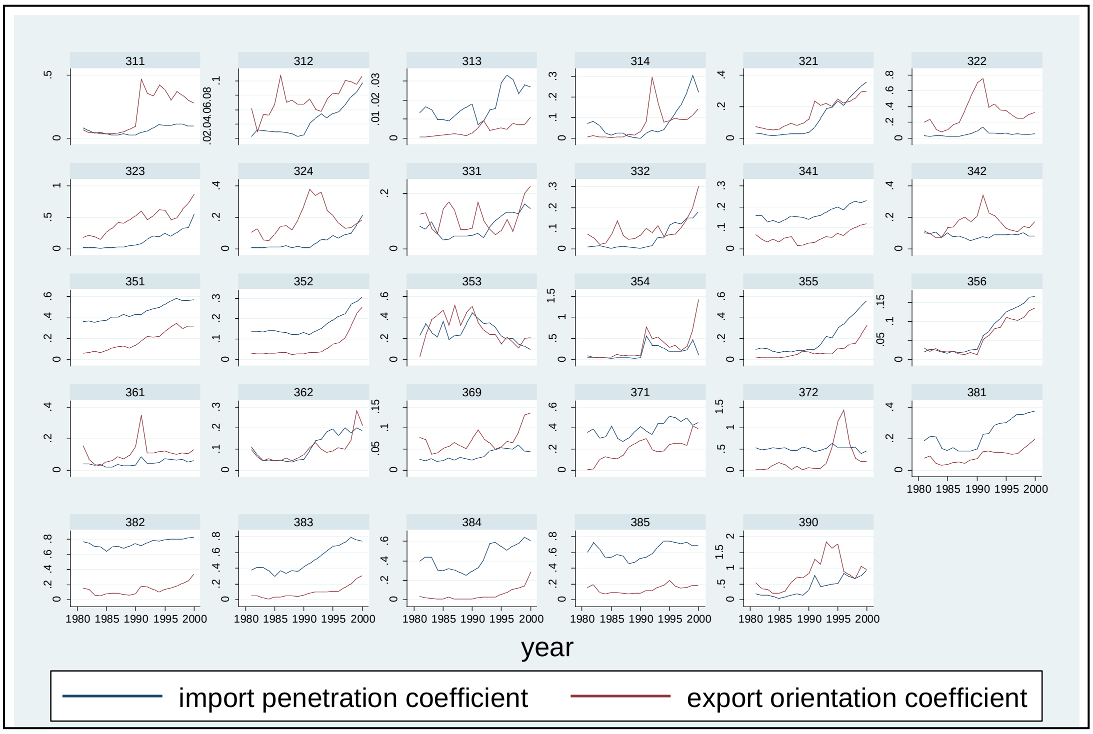


Own estimates based on tariff data from National Planning Department -DNP. Weights are based on import values in Col Pesos.

It should also be remarked that the process of tariff removal in Colombia was initiated in some industries in the early 1980s from which *332- Furniture and fixtures*, *322- Wearing apparel, except footwear* and, *321- Textiles* experienced reductions of more than ten percentage points over the pre-reform period (1985-1989) so, their reductions during the reform period (1990-1994) were more modest compared to other manufacturing industries. As a result of this process, the manufacturing industries with the lowest level of import tariffs over the post-reform period were mainly producers of intermediate goods such as *372- Non-ferrous metal basic* *industries*, *351- Industrial chemicals*, *353- Petroleum refineries*, *371- Iron and steel basic industries*, *354- Products of petroleum and coal* and, *352- Other chemical products*.

Some studies have previously used tariff data in order to assess the effects of trade policy on employment outcomes in Colombia (Attanasio et al., 2004, Goldberg and Pavcnik, 2003, Goldberg and Pavcnik, 2005b, Jaramillo and Tovar, 2006). In particular, Jaramillo and Tovar (2006) claim that tariff rates are "the most direct measure of trade policy available" in the Colombian case. But other important direct measures of trade policy such as Non-tariff barriers (NTBs hereafter), on the other hand, are only available after 1991 and, therefore, tariff rates provide a just a partial picture of trade policy in Colombia. For this reason, we focus our analysis on two commonly used indicators of trade policy, import penetration coefficient (*IPC)* and export orientation coefficient (*EOC*) that are readily available from the National Planning Department at three-digit level of the ISIC Rev.2. We believe that these measures represent superior indicators of trade policy as they display changes in trade flows, which are the ultimate objective of changes in the trade regime. The *IPC* measures the share of the domestic market in a given industry that is supplied with imports while the *EOC* indicates the percentage of domestic production in a given industry that is exported to other countries and thus, provides a crude measure of comparative advantage. The results for these trade measures are presented in Figure 5 and provide convincing evidence that most of Colombian manufacturing industries became more open in terms of both import penetration and export orientation. The *IPC* indicates that imported goods represented 18.9 per cent of the internal demand of all manufacturing goods in 1981-1985 and 32.4 per cent in 1996-2000. In general, only two out of 30 manufacturing industries examined here (*353- Petroleum refineries* and *342- Printing, publishing and allied industries*) report a reduction in this coefficient after liberalisation in 1991. The same figures indicate that the industries with the largest increments in import penetration over these years were *390- Other Manufacturing Industries*, *355- Rubber* *products*, *383- Electrical machinery apparatus, appliances*, *354- Products of petroleum and coal*, *323- Leather and products of leather* and, *321- Textiles*. In turn, *EOC* suggests that while 6.9 per cent of the domestic manufacturing product of traded goods in 1981-1984 was exported, this proportion grew to 21.3 per cent in 1996-2000. According to this coefficient, all manufacturing industries, except *353- Petroleum refineries*, became more export-oriented over these years. The largest increments in the *EOC* over this period were reported by *390- Other Manufacturing Industries*, *354- Products of petroleum and coal*, *323- Leather and products of leather* and, *372- Non-ferrous metal basic industries*, all of which experienced increases of more than 30 percentage points. It is worth to mention that *323- Leather and products of leather* and *321- Textiles*, the two sectors with the highest proportion of female workers (see 3.3.1 section, below), reported large increments in both export orientation and import penetration.

Figure 5: Import Penetration and Export Orientation coefficients across manufacturing industries, Colombia: 1981-2000



Source: National Planning Department -DNP.

3.3 Concentration, market power and trade reforms

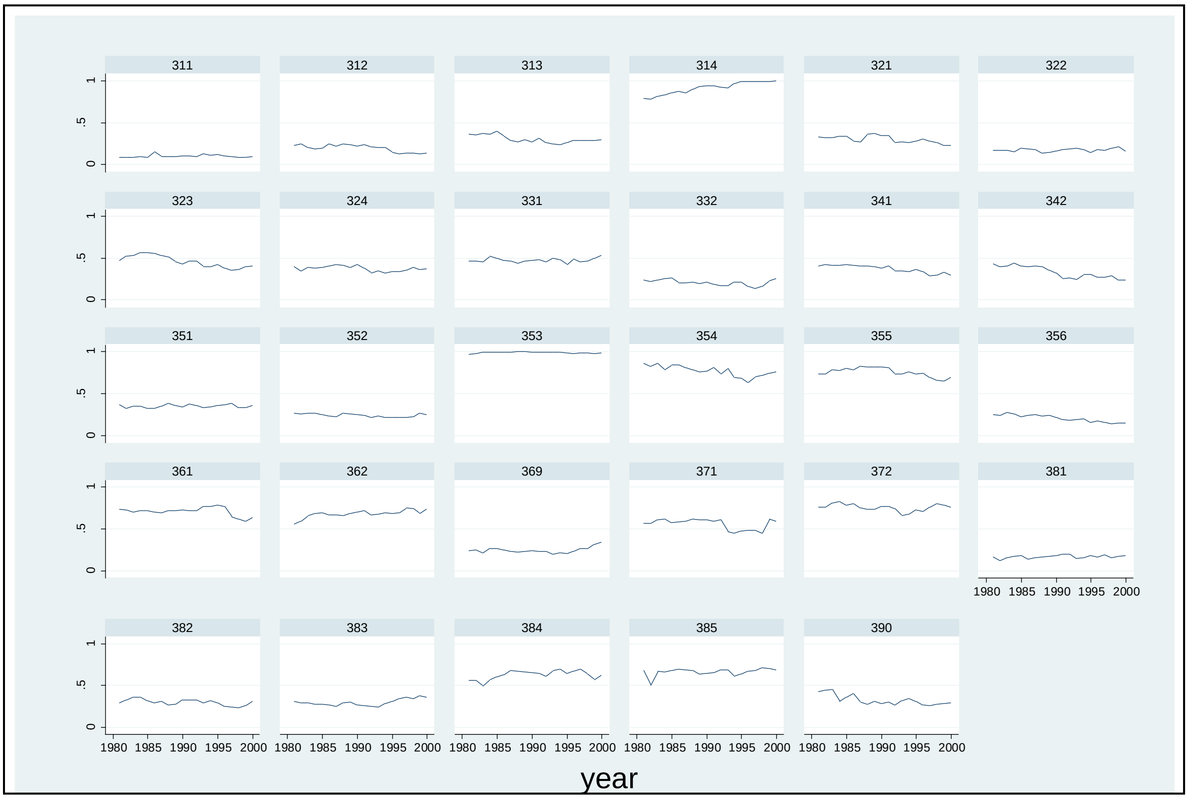
As explained in section 2.3 above, trade liberalisation has the potential to bring about more competition in the form of increased imports which, in turn, might reduce the scope for costly gender discrimination. On the other hand, section 2.4 suggests the possibility that increasing competition from imports may reinforce the bargaining position of local firms in the labour market as the number of employers is being reduced and workers have fewer options for employment within a given industry.

In order to control for the effects of market structure, we compute a conventional four-firm concentration ratio (CR_4) across industries based on the ratio between the gross product value from the four largest firms within a given industry and the total gross product value for the same industry as follows

$$CR_4 = \sum_{i=1}^{4} S_i \tag{3.1}$$

where S_i denotes the gross product share of the i firm in the total gross product of a given industry. According to this index, there has been a slight reduction in the degree of concentration along the two decades defined in this study, from an average of 0.452 in 1981 to 0.439 in 2000. Figure 6 displays this concentration ratio for each of the 29 ISIC sectors along the years defined in this study. We plotted concentration ratios on an identical scale in order to display the high degree of stability in the ranking of the most (and less) concentrated sectors. Thus, *353- Petroleum refineries*, *314- Tobacco manufactures*, *354- Products of petroleum and coal*, *372- Non-ferrous metal basic industries*, *355- Rubber products* and, *361- Pottery, china and earthenware* emerge as the most concentrated ones in which the value of production for the top four firms represents more than 70 per cent of their corresponding industry. In contrast, *311- Food products*, *381- Fabricated metal products* and, *322- Wearing apparel, except* *footwear* appear as the least concentrated industries over the years reviewed here as their concentration index ranks, on average, below 20 per cent.

Figure 6: Concentration Indices (based on Gross Product Values) across manufacturing industries, Colombia: 1981-2000



Own estimates based on Annual Manufacturing Survey microdata.

3.4 Capital equipment

The interaction of trade with employment dynamics by gender has multiple dimensions. As explained by Galor and Weil (1996), the process of economic development allows increases in the availability of capital per worker which make physical strength less relevant and, thus, may lead to increased female labour participation. Since trade liberalisation facilitates the access to imported technology, there is the possibility of significant interactions with employment dynamics by gender.

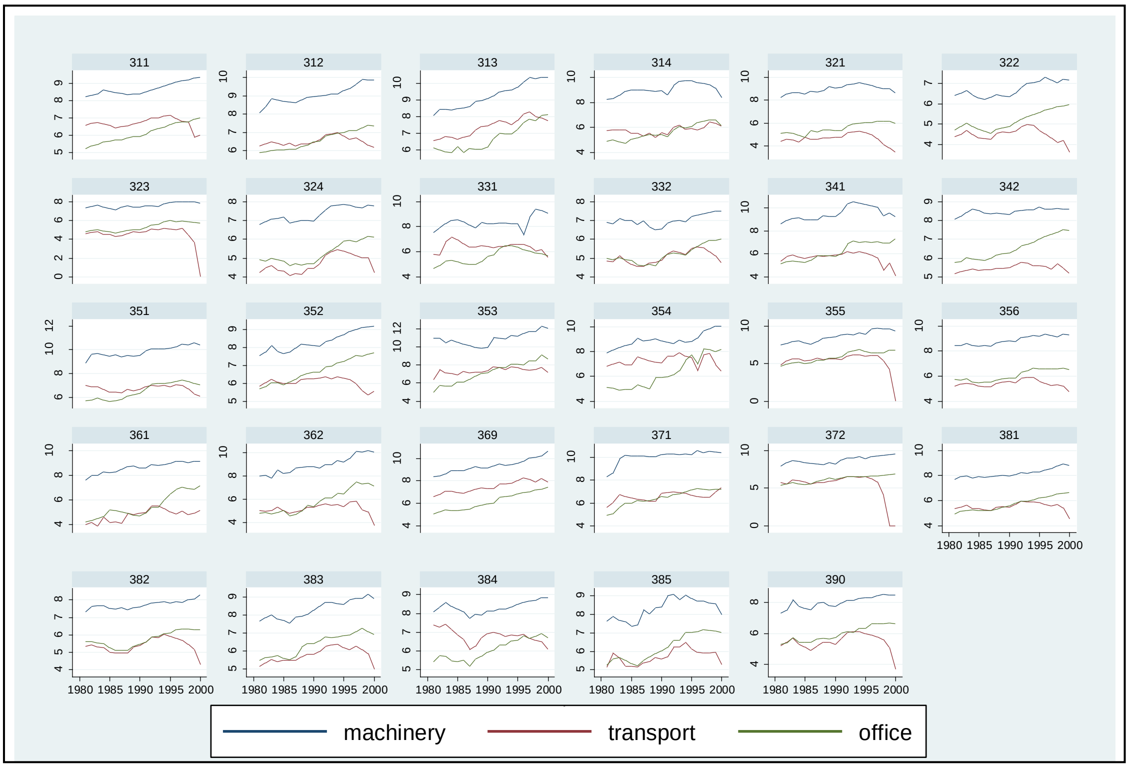
In order to test these possible relationships between employment dynamics by gender and trade, we also investigated the changes in capital investment across manufacturing industries. For this purpose, we computed the stock of three different types of capital over the fiscal year using AMS microdata. These are (i) machinery and equipment, (ii) transport equipment, and (iii) office equipment. In order to control for scale differences, we computed separately capital stocks per worker in natural logarithms expressed in constant 1999 Colombian Pesos. . Capital stocks were estimated using a perpetual inventories approach according to the following expression:

$$K_{it} = K_{it-1} + I_{it} + (K_{it-1} + I_{it}) * D_i \qquad (3.2)$$

where K denotes the capital stock of industry i at the beginning of year t, I represents the gross investment of industry i and, D depicts the observed depreciation rate of industry i estimated by Pombo (1999) at the ISIC Rev.2, 3-digit level industries. Figure 7 displays our estimates for the logarithm of the capital stock per worker across the 29 manufacturing industries defined in this study from 1981 to 2000. Capital stocks per worker of both machinery equipment and office equipment reported net increases between 1981-1985 and 1996-2000 for all manufacturing industries reviewed here. Contrastingly, transport equipment per worker reported net increases only in 14 out of 29 manufacturing sectors over the same time period. The largest increases in the stock of machinery equipment per worker between 1981-1985 and 1996-2000 were reported by *313- Beverage industries*, *362- Glass and glass products*, *355- Rubber products* and, *369- Other non-metallic mineral products*. In the case of transport equipment, the largest increases were found in *313- Beverage industries*, *369- Other non-metallic mineral products*, *361- Pottery, china and earthenware*, *324- Footwear* and, *371- Iron and steel basic industries*. Finally, the largest increases in office equipment per worker were recorded by *354- Products of petroleum and coal*, *353- Petroleum refineries*, *362- Glass and glass products*, *361- Pottery, china and earthenware*, *369-*

Other non-metallic mineral products and*, 313- Beverage industries*. From this, it is evident that *313- Beverage industries* was the most dynamic sector in terms of investments of all three types of capital equipment reviewed here, followed by *362- Glass and glass products*, a complementary sector of the former. A similar remark could be made for industries dedicated to the production of non-metallic mineral manufactures such as *361- Pottery, china and earthenware*, and *369- Other non-metallic mineral products* with some of the largest increments in their stock of the three types of capital per worker examined here.

Figure 7: Capital Equipment (Machinery, Transport and Office) per Worker across manufacturing industries, Colombia: 1981-2000



Own estimates based on Annual Manufacturing Survey microdata.

4. Econometric analysis

4.1 Methodology

In order to explain the effects of trade policy on the gender composition of the workforce across manufacturing industries, we implement different panel data models including fixed-effects instrumental variables (FE-IV). As technological changes are also likely to affect the share of female jobs across manufacturing industries over a time span of two decades, our empirical strategy also incorporates the three explanatory variables for the capital stock per worker (in logarithms) explained above in section 3.3.4, namely, machinery equipment, transport equipment and, office equipment. In addition, we control for the effects of changes in market structure with the inclusion of a concentration index based on expression 3.1 in Section 3.3.3, above.

The FE-IV approach adopted here is based on an individual industry effects model

$$y_{it} = x'_{it}\beta + \alpha_i + \varepsilon_{it} \qquad (3.3)$$

where y_{it} represents the female share of jobs in industry i at time t, x'_{it} is a set of explanatory variables and β depicts the coefficients to be estimated. The structure of the error component in (3.3) assumes the existence of unobserved time-invariant factors across the cross-section units depicted by α_i plus a conventional random component ε_{it}. Provided the existence of adequate instruments, z_{it}, FE-IV provide consistent estimates of β even in cases where the regressors contained in x_{it} are correlated with the random component ε_{it}. The key characteristic of such instruments is that they are uncorrelated to the error term ε_{it} so,

$$E(\varepsilon_{it} | \alpha_i, z_{i1}, \dots, z_{it}, \dots, z_{iT}) = 0. \qquad (3.4)$$

Under the assumption that (3.2) is upheld by the data, FE-IV provides consistent estimates. As it is normally the case with panel data, if the assumptions for the idiosyncratic error term notably $\varepsilon_{it} \sim (0, \sigma_\varepsilon^2)$ are not satisfied, conventionally computed standard errors are inaccurate. According to Cameron and Trivedi (2009), this assumption can be relaxed by the use of standard errors that allow for intergroup correlation. This is achieved with the estimation of a variance-covariance matrix that is adjusted with a clustered sandwich estimator. Chapter 8 of Angrist and Pischke (2008) describe this and other procedures for robust covariance matrix estimation in panel data applications whose observations are correlated within groups.[6] The estimation of FE-IV models presented in this application is performed using the ***xtivreg2*** Stata command developed by Schaffer and Stillman (2010) which allows for this type of cluster-robust standard errors. In the case of models without instruments, cluster-robust standard errors can be estimated with the conventional ***xtreg*** Stata command.

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[6] Chapter 10 in Cameron and Trivedi (2009) provides also a review of different estimates for the variance-covariance matrix including the cluster-robust procedure. More formally, the cluster-robust standard errors procedure implemented in this application is a generalization of White's (1980) procedure for the estimation of a robust covariance matrix of the following form:

$$\hat{\Omega}_{cl} = (X'X)^{-1}\left(\sum_g X_g \hat{\Psi}_g X_g\right)(X'X)^{-1}$$

where $\hat{\Psi}_g = a\hat{e}_g\hat{e}'_g = a\begin{bmatrix} \hat{e}_{1g}^2 & \cdots & \hat{e}_{1g}\hat{e}_{n_g g} \\ \vdots & \ddots & \vdots \\ \hat{e}_{1g}\hat{e}_{n_g g} & \cdots & \hat{e}_{n_g g}^2 \end{bmatrix}$,

X_g is the matrix of regressors for g groups, $\hat{e}_{ig}$ are the estimated residuals clustered around g groups of data and a is a factor adjustment which makes a degrees of freedom correction. See ANGRIST, J. D. & PISCHKE, J.-S. 2008. *Mostly Harmless Econometrics: An Empiricist's Companion,* Princeton, New Jersey.: 312-313p.

4.2 Results

As a departure point, Table 1 describes the variables included in the models presented in this section while Table 2 reports their variance decomposition of them. All variables have no missing values and are within the expected range. To facilitate interpretation and estimation under different methods, all our variables are continuous measures within the 0 to 1 range, except in the case of capital per worker variables as they are expressed in logs in Colombian Pesos at constant 1999 prices. For all variables but the log of office equipment per worker variable (*lnkpw_office*), most of the variation occurs *between* manufacturing industries rather than *within* manufacturing industries.

Table 1 Variable definitions

label	variable	definition
femshare	female share of jobs: all workers	female share of jobs in industry i at time t amongst all workers
wc_femshare	female share of jobs: white-collar workers	female share of jobs in industry i at time t amongst white collar workers
bc_femshare	female share of jobs: blue-collar workers	female share of jobs in industry i at time t amongst blue collar workers
ipc	import penetration coefficient	$ipc_{it} = \frac{M_{it}}{Y_{it} + M_{it} - X_{it}}$ where Y, M and X denote, respectively, the gross product, imports and exports of industry *i* at time *t*.
eoc	export orientation coefficient	$eoc_{it} = \frac{X_{it}}{Y_{it}}$ where X and Y denote, respectively, exports and the gross product of industry *i* at time *t*.
CIGP	Concentration index	See expression (3.1) in text and details on it.
lnkpw_mach	ln(capital equipment per worker: machinery)	See expression (3.2) in text and details on it.
lnkpw_trans	ln(capital equipment per worker: transport)	
lnkpw_office	ln(capital equipment per worker: office equipment)	

To begin with, we want to test whether there is a relationship between the female share of jobs, on the one hand, and two selected trade variables on the other. The trade variables are the import penetration coefficient –*ipc* and the export orientation coefficient –*eoc*. These models are presented in Tables 3 and 4, from top to bottom, for *all workers*, *white-collar workers* and *blue collar workers*. All the reported specifications use clustered-robust standard errors as described in the preceding section.

Table 2: Panel summary statistics: within and between variation

Variable		Mean	Std. Dev.	Min	Max	Observations
isic	overall	349.897	25.064	311	390	N = 580
	between		25.486	311	390	n = 29
	within		0.000	349.897	349.897	T = 20
year	overall	1990.5	5.771	1981.0	2000.0	N = 580
	between		0.000	1990.5	1990.5	n = 29
	within		5.771	1981.0	2000.0	T = 20
femshare	overall	0.2701	0.1598	0.0096	0.8135	N = 580
	between		0.1596	0.0785	0.8007	n = 29
	within		0.0296	0.1824	0.3727	T = 20
wc_femshare	overall	0.3761	0.1019	0.0364	0.6704	N = 580
	between		0.0861	0.1597	0.5987	n = 29
	within		0.0567	0.2528	0.6486	T = 20
bc_femshare	overall	0.2295	0.1889	0.0032	0.8697	N = 580
	between		0.1899	0.0224	0.8516	n = 29
	within		0.0281	0.1200	0.3487	T = 20
CIGP	overall	0.4429	0.2462	0.0836	0.9990	N = 580
	between		0.2463	0.0985	0.9894	n = 29
	within		0.0442	0.2799	0.5702	T = 20
ipc	overall	0.2189	0.2218	0.0005	0.9456	N = 580
	between		0.2023	0.0176	0.7511	n = 29
	within		0.0980	-0.1337	0.7527	T = 20
eoc	overall	0.1717	0.2237	0.0006	1.8409	N = 580
	between		0.1636	0.0041	0.8421	n = 29
	within		0.1555	-0.4653	1.3417	T = 20
lnkpw_mach	overall	8.6275	0.9976	6.2089	12.3023	N = 580
	between		0.8888	6.7036	10.9253	n = 29
	within		0.4807	6.8639	10.0045	T = 20
lnkpw_trans	overall	5.8333	1.0772	0.0000	8.2848	N = 580
	between		0.9121	4.4789	7.4737	n = 29
	within		0.5964	0.5224	7.0625	T = 20
lnkpw_office	overall	6.0298	0.8434	4.2137	9.1494	N = 580
	between		0.5027	5.1884	7.1754	n = 29
	within		0.6833	3.8663	8.0038	T = 20

In Table 3, Column 1 reports pooled OLS regression estimates featuring only *ipc* as a regressor. The coefficients for manufacturing employment disaggregated by broad skill types are poorly determined as their statistical significance lies outside the 10 per cent level. However, there is a remarkable gain in efficiency as well as an increase in the magnitude of the *ipc* coefficient when we control for fixed effects using the (within) FE estimator in Column 2. In this case we find a positive and well determined relationship between import penetration and the female share of jobs; the size of the coefficients suggests that this effect is stronger amongst *white collar workers*. This relationship is confirmed in Column 3 for *all workers* and *white collar workers* when we include a trend variable while it turns out statistically insignificant for blue-collar workers. We also check in Column 4 whether this relationship holds when we lag the trade variable as the presumed effects of import penetration in manufacturing industries on their female share of jobs might exhibit some persistence over time. The estimates in Column 4 are quite similar in terms of both size and statistical significance to those from the FE with no trend in Column 2. The inclusion of a time-trend variable in addition to the lagged *ipc* variable in Column 5 yields a sizeable reduction in the size of the coefficients while standard errors are slightly larger so the statistical significance is consequently reduced, particularly amongst *blue collar workers*. Finally, Column 6 features coefficients based on a first-difference estimator. As the variables are in differences while the *ipc* variable is lagged one period, there is a reduction in the number of observations with respect to the FE models based on the mean-difference estimator in Columns 2 and 3. First differencing reduces the size of the coefficients dramatically and they are well determined only when the dependent variable is the female share of jobs for *all workers*.

Table 3 Female share equations, trade variable: import penetration coefficient (ipc)

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	**OLS**	**FE**	**FE+trend**	**FE: ipc_{t-1}**	**FE: ipc_{t-1} + trend**	**Differences: D.Y = f(D.ipc_{t-1})**
			All workers			
ipc	-0.0799	0.1445***	0.0728**			
	(0.1192)	(0.0306)	(0.0339)			
trend			0.0021***		0.0023***	
			(0.0007)		(0.0007)	
L.ipc				0.1499***	0.0694**	
				(0.0299)	(0.0334)	
LD.ipc						0.0334***
						(0.0088)
Constant	0.2876***	0.2384***	0.2320***	0.2390***	0.2304***	
	(0.0423)	(0.0067)	(0.0050)	(0.0064)	(0.0050)	
			White-collar workers			
ipc	0.0515	0.3334***	0.1075**			
	(0.0583)	(0.0556)	(0.0472)			
trend			0.0066***		0.0071***	
			(0.0008)		(0.0009)	
L.ipc				0.3219***	0.0784	
				(0.0553)	(0.0504)	
LD.ipc						0.0022
						(0.0365)
Constant	0.3648***	0.3031***	0.2828***	0.3109***	0.2848***	
	(0.0226)	(0.0122)	(0.0092)	(0.0118)	(0.0104)	
			Blue-collar workers			
ipc	-0.1323	0.0673**	0.0682			
	(0.1428)	(0.0310)	(0.0433)			
trend			-0.0000		0.0002	
			(0.0008)		(0.0008)	
L.ipc				0.0729**	0.0647	
				(0.0315)	(0.0427)	
LD.ipc						0.0220
						(0.0175)
Constant	0.2585***	0.2148***	0.2149***	0.2136***	0.2127***	
	(0.0499)	(0.0068)	(0.0069)	(0.0067)	(0.0069)	
Observations	580	580	580	551	551	522

Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Table 4 provides a similar set of econometric results with respect to those commented above but this time the trade variable is represented by the export orientation coefficient –*eoc*. Both OLS and FE estimates in Columns 1 and 2 indicate that manufacturing industries with higher levels of export orientation tend to have larger shares of female jobs. The coefficients for the *eoc* variable are statistically significant at the 1 per cent level for all dis-aggregated measures of the labour force. With the inclusion of a time trend variable in Columns 2 and 3 the *eoc* coefficient still yields a positive coefficient in all cases but the size and the statistical significance is drastically reduced. A similar outcome is observed in Columns 4 and 5 with the incorporation of a one-lag version for this explanatory variable either with or without a trend control. The first-differenced results reported in Column 6 suggest that changes in export orientation might be positively associated with changes in the female share of jobs in the case of *all workers* and *blue collar workers* while they exert no independent effect amongst *white collar workers*. Notwithstanding, this positive effect amongst the *blue collar workers* is statistically significant only at the 10 per cent level.

The preceding findings from models featuring only one explanatory trade variable (plus a time trend in some cases) deserve some reflection. Estimates from the FE models using the mean-difference estimator suggest that manufacturing industries with high levels of both import penetration and export orientation tend to have a larger share of jobs occupied by women. The use of the first-difference estimator yields slightly less convincing evidence in favour of trade as a positive explanation for the growing proportion of female jobs in manufacturing industries. At best, these results suggests that the effects of increased trade in the gender composition of employment of manufacturing industries in urban Colombia are unevenly distributed across the two categories of jobs defined in this study. While changes in import penetration might be associated with a larger share of female jobs amongst *white collar workers*, changes in export orientation might be associated with increasing shares of jobs amongst *blue collar workers*. More importantly, the poor significance of the trade coefficients in some specifications suggests that other variables may have played a role in the incorporation of women in manufacturing. So far, we have implicitly assumed that the trade variables are uncorrelated to the error term ε_{it}. In other words, we have not dealt yet with any potential endogeneity problems that may contaminate these estimates.

Table 4 Female share equations, trade variable: export orientation coefficient (eoc)

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	**OLS**	**FE**	**FE+trend**	**FE: eoc$_{t-1}$**	**FE: eoc$_{t-1}$ + trend**	**Differences: D.Y = f(D.eoc$_{t-1}$)**
			All workers			
eoc	0.1960**	0.0594***	0.0178			
	(0.0730)	(0.0212)	(0.0152)			
trend			0.0026***		0.0028***	
			(0.0006)		(0.0006)	
L.ipc				0.0641**	0.0224	
				(0.0241)	(0.0168)	
LD.ipc						0.0190**
						(0.0093)
Constant	0.2364***	0.2599***	0.2395***	0.2604***	0.2369***	
	(0.0246)	(0.0036)	(0.0052)	(0.0040)	(0.0057)	
			White-collar workers			
eoc	0.1594***	0.1396***	0.0211			
	(0.0409)	(0.0309)	(0.0192)			
trend			0.0075***		0.0075***	
			(0.0008)		(0.0008)	
L.ipc				0.1419***	0.0283*	
				(0.0335)	(0.0149)	
LD.ipc						0.0076
						(0.0124)
Constant	0.3487***	0.3521***	0.2942***	0.3562***	0.2920***	
	(0.0147)	(0.0053)	(0.0072)	(0.0055)	(0.0084)	
			Blue-collar workers			
eoc	0.1940**	0.0201	0.0122			
	(0.0877)	(0.0178)	(0.0180)			
trend			0.0005		0.0007	
			(0.0006)		(0.0006)	
L.ipc				0.0269	0.0167	
				(0.0206)	(0.0197)	
LD.ipc						0.0241*
						(0.0122)
Constant	0.1962***	0.2260***	0.2221***	0.2248***	0.2190***	
	(0.0293)	(0.0031)	(0.0061)	(0.0034)	(0.0066)	
Observations	580	580	580	551	551	522

Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

For these reasons, we now implement the FE-IV approach by incorporating additional explanatory variables in our modelling strategy, namely, a concentration index of the gross product described in section 3.3.3 (*CIGP*), and the three measures of the stock of capital equipment per worker detailed on section 3.3.4 (*lnkpw_mach*, *lnkpw_trans* and, *lnkpw_office* –see Table 1 for definitions). Under this framework, we control for endogeneity problems through the use of instruments for both trade measures already incorporated in the models presented in Tables 3 and 4 and the concentration index variable (*CIGP*) discussed in Section 3.3, above. We base our decision on which variables to instrument on a version of the Hausman test of endogenous regressors developed in Stata™ by Schaffer and Stillman (2010) that is robust to violations of conditional homoskedasticity. The results for this test, under different specifications, are presented in the Statistical Appendix 1 of this paper (see Tables A1 and A2); they indicate that the null hypothesis that a given set of regressors is exogenous can be safely rejected in the case of the concentration index variable (*CIGP*) and the two trade measures (*ipc* and *eoc*).[7] Thus, we instrumented *CIGP* with the logarithm of the number of firms, *ipc* with average tariffs (see section 3.3.2, above) and, *eoc* with a conventional relative trade balance measure (*RTB*) constructed as follows:

$$RTB_{it} = \frac{X_{it} - M_{it}}{X_{it} + M_{it}} \qquad (3.5)$$

where X_{it} and M_{it} denote the exports and imports, respectively, from industry *i* at time *t*.

The rationality for the use of these instruments is justified not only on the grounds that they are highly correlated to the endogenous variables (we test formally this below) but also on their theoretical validity. In the case of the import penetration, we argue that average tariffs represent an appropriate instrument measure of trade policy as they are aimed at moderating import flows. On this it should be mentioned that some empirical applications dealing with the effects of trade on labour market outcomes in Colombia have directly relied on tariffs as a proxy measure of trade policy (Attanasio et al., 2004, Goldberg and Pavcnik, 2003, Goldberg and Pavcnik, 2005b, Jaramillo and

[7] See notes at Tables 3.A1 and 3.A2 for details on the structure of this test.

Tovar, 2006).[8] We believe that using tariffs instead of import penetration as a variable to control for the impact of trade policy on the labour market is problematic as it omits the effects of other trade policy measures such as import licences and import quotas. Contrastingly, import penetration provides an outcome measure of the effects of trade policy on the competitive environment in which local firms have to operate. Tariffs instead provide a good instrument for import penetration as they embody a trade policy measure aimed specifically at moderating import flows into the domestic economy. In the case of the export orientation coefficient, we use a relative trade balance measure described in expression (3.5) as it represents a reasonable estimate of the competitive position of manufacturing industries with rich variation across sectors and over time. We also instrument the concentration index of gross product (CIGP) variable with the natural logarithm of the corresponding number of firms for each combination of industries and years based on the assumption that more competitive industries (i.e., with a lower concentration index) have, on average, a larger number of firms.

In Table 5 we test formally the association between the endogenous regressors and the selected instruments incorporated in subsequent FE-IV models presented below. According to these results, we can reasonably be confident that our instruments are highly correlated with the endogenous regressors not only in terms of the FE within estimator (see Columns 1, 3 and 5) but also in terms of the first-differences specification (see Columns 2, 4 and 6). As in other models presented along this paper, the standard errors reported in Table 5 are robust for cluster correlation. On these

[8] On these papers, Attanasio et al (2004) use tariffs at the beginning of the 1980s interacted world coffee prices as instruments for tariffs while Goldberg and Pavcnik (2005b) perform an identical strategy. Jaramillo and Tovar (2006) also use tariffs at the beginning of the 1980s interacted with annual exchange rates.

results we verify a negative association between import penetration (*ipc*) and average tariffs (*a_tariffs*) as can be seen in the regression coefficients in Columns 1 and 2 which are statistically significant at the one per cent level in the case of the FE estimator and, at the five percent level in the case of the first-differences estimator. We confirm also a negative association between the concentration index of gross product (*CIGP*) and the natural logarithm of the number of plants (*ln_noplants*) as can be inferred from the estimated coefficients in Columns 3 and 4 of Table 5. Lastly, we corroborate a positive relationship with statistically significant coefficients at the one per cent level between export orientation (*eoc*) and the relative trade balance measure (*rtb*) presented in expression (3.5), above.

Table 5: Testing the relevance of instruments: fixed-effects and first-differences estimates

VARIABLES	**(1)** **ipc**	**(2)** **D.ipc**	**(3)** **CIGP**	**(4)** **D.CIGP**	**(5)** **eoc**	**(6)** **D.eoc**
a_tariffs	-0.5688***					
	(0.1193)					
D.a_tariffs		-0.1196**				
		(0.0593)				
ln_noplants			-0.1075**			
			(0.0399)			
D.ln_noplants				-0.1022***		
				(0.0240)		
rtb					0.2237**	
					(0.0816)	
D.rtb						0.2995***
						(0.0824)
Constant	0.3189***	0.0073***	0.9762***	-0.0007***	0.1933***	0.0096***
	(0.0210)	(0.0024)	(0.1978)	(0.0000)	(0.0079)	(0.0005)
Observations	580	551	580	551	580	551
R-squared	0.2586	0.0078	0.1702	0.0620	0.1223	0.2061
Number of isic	29	29	29	29	29	29

Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1

Notes: (1) features *ipc* as a dependent variable against average tariffs (*a_tariffs*) as a single explanatory variable while (2) features the same variables in differences. (3) features *CIGP* as a dependent variable against the logarithm of the number of firms (*ln_noplants*) as a single explanatory variable while (4) features the same variables in differences. (5) features *eoc* as a dependent variable with the relative trade balance (*rtb*) as a single explanatory variable while (6) features the same variables in differences.

Results for our FE-IV estimates for the effects of import penetration on the female share of jobs are presented in Table 6. In order to check the robustesness of our FE-IV estimates, we also estimate the same female share equations with instruments derived from their lagged values. Standard errors for FE-IV models presented on Table 6 are robust for cluster serial autocorrelation (see Section 3.4.1, above). To further check these results, we present in the Statistical Appendix 2, estimates using the Generalised Method of Moments approach developed by Arellano and Bover (1995) and Blundell and Bond (1998).

As a natural reference point, Column 1 on Table 6 shows conventional FE with no instrumental variables. The trade variable, *ipc*, shows well determined coefficients for *all workers*, *white collar workers* and *blue collar workers* pointing towards a positive relationship between import penetration and the female share of jobs, a finding that confirms our previous results from Table 3. The use of instruments presented under different specifications in Columns 2 to 7confirm this result for both, *all workers* and *white collar workers*. In the case of *blue collar workers*, the choice of instruments affects the statistical significance of this variable and this casts some doubt on the effects of import penetration in the female share of jobs amongst this category. Results for the *ipc* variable using the linear dynamic panel data procedure presented in the Statistical Appendix 2 confirm that its effect on the female share of jobs is both negative and statistically different from zero only in the case of *white collar workers*. These results suggest that import penetration has a differentiated effect in the female share of jobs across the labour force categories defined in this study pointing suggesting that some of the presumably positive effects of increased import penetration tend to favour the insertion of women mainly into the *white collar workers* category.

Table 6 Fixed-effects IV estimation of female share equations; trade variable: import penetration coefficient (ipc)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					All workers				
ipc	0.0892***	0.1600***	0.1247***	0.1752***	0.1505***	0.1698**	0.1102***	0.1708**	0.1065***
	(0.0128)	(0.0551)	(0.0157)	(0.0362)	(0.0144)	(0.0742)	(0.0165)	(0.0701)	(0.0166)
CIGP	-0.0787***	-0.4516***	-0.1096***	-0.4496***	-0.1221***	-0.4519***	-0.0961***	-0.4490***	-0.0968***
	(0.0249)	(0.0951)	(0.0368)	(0.0957)	(0.0377)	(0.0961)	(0.0368)	(0.0946)	(0.0368)
lnkpw_mach	0.0005	0.0019	0.0119***					0.0026	0.0030
	(0.0038)	(0.0048)	(0.0027)					(0.0047)	(0.0039)
lnkpw_trans	-0.0033*			-0.0016	-0.0011			-0.0018	-0.0029
	(0.0018)			(0.0022)	(0.0018)			(0.0026)	(0.0018)
lnkpw_office	0.0124***					0.0001	0.0106***	-0.0011	0.0095***
	(0.0028)					(0.0052)	(0.0021)	(0.0055)	(0.0030)
					White-collar workers				
ipc	0.1626***	0.6532***	0.2336***	0.7203***	0.3390***	0.5719***	0.1691***	0.5461***	0.1666***
	(0.0196)	(0.1008)	(0.0258)	(0.0701)	(0.0256)	(0.1230)	(0.0257)	(0.1129)	(0.0261)
CIGP	-0.0621	0.1353	-0.1097*	0.1177	-0.1551**	0.1330	-0.0520	0.1256	-0.0542
	(0.0382)	(0.1740)	(0.0606)	(0.1856)	(0.0673)	(0.1593)	(0.0575)	(0.1524)	(0.0575)
lnkpw_mach	0.0075	0.0155*	0.0467***					0.0041	0.0073
	(0.0058)	(0.0087)	(0.0045)					(0.0076)	(0.0062)
lnkpw_trans	-0.0018			0.0101**	0.0054			0.0057	-0.0017
	(0.0028)			(0.0043)	(0.0033)			(0.0042)	(0.0028)
lnkpw_office	0.0412***					0.0183**	0.0434***	0.0169*	0.0398***
	(0.0043)					(0.0086)	(0.0032)	(0.0089)	(0.0047)

Table 6 (Continuation)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					Blue-collar workers				
ipc	0.0721***	0.0579	0.0863***	0.0033	0.0782***	0.0914	0.0868***	0.0959	0.0883***
	(0.0141)	(0.0560)	(0.0168)	(0.0366)	(0.0150)	(0.0753)	(0.0179)	(0.0714)	(0.0182)
CIGP	-0.0416	-0.3885***	-0.0253	-0.3848***	-0.0212	-0.3869***	-0.0270	-0.3893***	-0.0260
	(0.0275)	(0.0967)	(0.0395)	(0.0967)	(0.0394)	(0.0975)	(0.0401)	(0.0964)	(0.0402)
lnkpw_mach	-0.0067	-0.0097**	-0.0038					-0.0045	-0.0034
	(0.0041)	(0.0049)	(0.0029)					(0.0048)	(0.0043)
lnkpw_trans	0.0006			-0.0013	0.0006			0.0011	0.0011
	(0.0020)			(0.0023)	(0.0019)			(0.0026)	(0.0020)
lnkpw_office	0.0012					-0.0096*	-0.0023	-0.0076	-0.0006
	(0.0031)					(0.0053)	(0.0022)	(0.0056)	(0.0033)
Observations	580	580	551	580	551	580	551	580	551
Instruments									
- tariffs		Yes		yes		yes		yes	
- ln(number of plants)		Yes		yes		yes		yes	
- One lag			yes		Yes		yes		yes

Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Constant omitted. Column (1) displays conventional FE with no instrumental variables. Columns (2) to (9) display FE-IV estimates; see bottom of table for chosen instruments. Import penetration coefficient (ipc) instrumented with either average tariffs or its own lag. Concentration index based on gross product (CIGP) instrumented with either the natural logarith of the number of firms or its own lag. FE-IV with cluster-robust stantandard errors estimated with the **xtivreg2** Stata command developed by Schaffer and Stillman (2010).

Models presented in Table 6 also investigate the effects of other variables commented on in the literature review in Section 3.2 above. Our measure of the degree of market concentration (*CIGP*) discussed in Section 3.3.3, above, is negative and statistically significant at the one per cent level for *all workers* and *blue collar workers* and performs poorly in the case of *white collar workers*. Results from dynamic panel data presented in Appendix 2 also suggest that the degree of market concentration is inversely correlated with the female share of jobs for all employment groupings analysed here, with well determined coefficients in most cases. Overall, the econometric evidence presented in both the main text and Appendix 2 is in line with the segregation dimension implicit in Becker's hypothesis of labour market discrimination in the sense that increased levels of market competition should erode monopolistic rents to discriminate against women. Although we do not have any evidence of reduced gender discrimination, we do observe that more competitive industries tend to have, on average, higher female shares of jobs. At least, this is what we would expect according to Becker's hypothesis in terms of the gender composition of the labour force as a result of increasing competition. In any case, we remain agnostic on whether this inverse relationship between market concentration and the female share of jobs across manufacturing industries is in any extent related to lower levels of gender discrimination. The same could be said regarding the results for the *ipc* variable commented above which could be rationalised in terms of the increased levels of market competition induced by increasing import penetration.

The results in Table 6 also feature the effects of the stock of capital investments per worker (in natural logarithms of Col Pesos of 1999) under the three categories discussed in Section 3.3.4, above. Columns 2 to 7 display the effects of these variables one by one using either average tariffs + the number of firms in logs (Columns 2, 4 and 6) or lagged values (Columns 3, 5 and 7) as instruments for both, the trade variable (*ipc*) and the concentration index variable (*CIGP*). It is worth reiterating that we could not find evidence indicating the necessity to instrument our capital equipment variables based on the version of the Hausman test of endogenous regressors explained above (see Appendix 3.1). Compared to the baseline specification with no instruments (Column 1), only our office equipment variable (*lnkpw_office*) is statistically significant for *all workers* and *white collar workers* while it tends to perform poorly for *blue collar workers*. This relationship is confirmed by our dynamic panel data estimates presented in Table A2.1 in Appendix 2. On this we should recall the discussion presented in Section 3.2.2 according to which increases in the availability of capital per worker enhances the participation of women in the labour market. In particular, the positive relationship between the capital stock of office equipment and the female share of jobs observed in our results is consistent with the hypothesis supported by some of the studies reviewed above (Galor and Weil, 1996, Weinberg, 2000, Welch, 2000) which suggest that women enjoy a comparative advantage in cognitive skills.

The fact that the estimates for our office equipment variable is not statistically significant for blue collar workers (a result that is also confirmed by dynamic panel data estimates in Table A2.1 in Appendix 2) might in a way be interpreted as a confirmation that the investments in office equipment are complementary to skilled female labour which tend to be concentrated in the *white collar category*. This interpretation is, to some extent, in line with the formulation proposed by Weinberg (2000) who argues that, in the case of the United States, a substitution process between highly skilled women and less skilled men might be explained by the increase in computers use which, on the margin, tends to favour the former. Figures presented in Figure 1a, above, suggest that this phenomenon might also be happening in Colombian manufacturing industries as female *white collar workers* were the only group of the labour force which shows an absolute increase of employment levels between 1981 and 2000. Contrastingly, male *blue collar workers* were the group with the largest reduction in manufacturing employment over these years in both absolute and relative terms.

In regard to the other two capital equipment variables reported in Table 6, we observe less clear cut results. Coefficients for the machinery equipment variable (*lnkpw_mach*) in Columns 2 and 3 suggest a positive and statistically significant relationship in the case of *white collar workers* when this variable is instrumented with its lagged values but this result turns out statistically insignificant when all capital regressors are simultaneously included in the model as can be seen in Column 7. Our dynamic panel data estimates presented in Table A2.1 of Appendix 2 indicate that this relationship is well determined only for *all workers*. In the same vein, the transport equipment variable shows up statistically significant at the 5 per cent level only in the case of *white collar workers* when we instrument both *ipc* and *CIGP* with average tariffs and the log of the number of firms, respectively (see Column 4 in Table 6). This finding is also confirmed by our dynamic panel data estimates from Table A2.1 in Appendix 2. When we switch our IV strategy to lagged values, our results presented in the main text indicate that this coefficient is not statistically different from zero. As in the case of the machinery equipment, the transport equipment variable loses its statistical significance when all capital regressors are simultaneously included in our FE-IV models in Table 6, a result that is also confirmed by our dynamic panel data estimates in Table A2.1 in Appendix 2.

Table 7 Fixed-effects IV estimation of female share equations; trade variable: export orientation coefficient (eoc)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					All workers				
eoc	0.0292***	-0.0015	0.0503***	0.0162	0.0675***	0.0023	0.0430***	-0.0026	0.0427***
	(0.0075)	(0.0242)	(0.0106)	(0.0229)	(0.0106)	(0.0230)	(0.0103)	(0.0233)	(0.0103)
CIGP	-0.0946***	-0.5540***	-0.1410***	-0.6239***	-0.1727***	-0.5210***	-0.1124***	-0.5240***	-0.1104***
	(0.0255)	(0.0774)	(0.0387)	(0.0795)	(0.0405)	(0.0794)	(0.0385)	(0.0792)	(0.0384)
lnkpw_mach	0.0003	0.0142***	0.0177***					0.0041	0.0022
	(0.0039)	(0.0039)	(0.0027)					(0.0048)	(0.0041)
lnkpw_trans	-0.0048***			-0.0032	-0.0026			-0.0051**	-0.0046**
	(0.0018)			(0.0025)	(0.0020)			(0.0023)	(0.0018)
lnkpw_office	0.0168***					0.0113***	0.0155***	0.0100**	0.0148***
	(0.0028)					(0.0029)	(0.0019)	(0.0040)	(0.0029)
					White-collar workers				
eoc	0.0543***	-0.1012***	0.1029***	-0.0050	0.1564***	-0.0850***	0.0786***	-0.0955***	0.0780***
	(0.0116)	(0.0383)	(0.0179)	(0.0386)	(0.0197)	(0.0326)	(0.0163)	(0.0334)	(0.0163)
CIGP	-0.0908**	-0.4268***	-0.1606**	-0.7453***	-0.2641***	-0.2339**	-0.0671	-0.2370**	-0.0665
	(0.0395)	(0.1223)	(0.0655)	(0.1338)	(0.0756)	(0.1127)	(0.0608)	(0.1139)	(0.0608)
lnkpw_mach	0.0069	0.0729***	0.0569***					0.0123*	0.0056
	(0.0060)	(0.0062)	(0.0046)					(0.0069)	(0.0065)
lnkpw_trans	-0.0045			0.0036	0.0019			-0.0057*	-0.0042
	(0.0028)			(0.0042)	(0.0037)			(0.0033)	(0.0029)
lnkpw_office	0.0492***					0.0610***	0.0502***	0.0555***	0.0476***
	(0.0043)					(0.0042)	(0.0030)	(0.0057)	(0.0047)

Table 7 (continuation)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
					Blue-collar workers				
eoc	0.0156*	-0.0170	0.0286***	-0.0226	0.0294***	-0.0191	0.0261**	-0.0176	0.0264**
	(0.0082)	(0.0243)	(0.0110)	(0.0215)	(0.0105)	(0.0239)	(0.0111)	(0.0242)	(0.0111)
CIGP	-0.0575**	-0.4505***	-0.0526	-0.4350***	-0.0541	-0.4535***	-0.0460	-0.4542***	-0.0444
	(0.0279)	(0.0776)	(0.0405)	(0.0744)	(0.0401)	(0.0825)	(0.0413)	(0.0824)	(0.0414)
lnkpw_mach	-0.0066	-0.0041	0.0008					-0.0030	-0.0036
	(0.0042)	(0.0039)	(0.0028)					(0.0050)	(0.0044)
lnkpw_trans	-0.0006			-0.0013	-0.0001			-0.0009	-0.0004
	(0.0020)			(0.0023)	(0.0020)			(0.0024)	(0.0020)
lnkpw_office	0.0052*					-0.0025	0.0021	-0.0008	0.0041
	(0.0031)					(0.0031)	(0.0020)	(0.0041)	(0.0032)
Observations	580	580	551	580	551	580	551	580	551
Instruments									
- relative trade balance		Yes		yes		yes		yes	
- ln(number of plants)		Yes		yes		yes		yes	
- One lag			yes		yes		yes		yes

Cluster-robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. Constant omitted. Column (1) displays conventional FE with no instrumental variables. Columns (2) to (9) display FE-IV estimates; see bottom of table for chosen instruments. Export orientation coefficient (eoc) instrumented with either relative trade balance or its own lag. Concentration index based on gross product (CIGP) instrumented with either the natural logarith of the number of firms or its own lag. FE-IV with cluster-robust stantandard errors estimated with the **xtivreg2** Stata command developed by Schaffer (2010).

The econometric results presented in Table 7 are intended to investigate the effects of an alternative trade variable, the export orientation coefficient –*eoc*. In this case, increased levels of trade in the form of export orientation tend to be statistically different from zero in a number of specifications. Nevertheless, the sign of the coefficient proves to be sensitive to the choice of instruments in this case. When we base our IV strategy on lagged values of endogenous regressors, the coefficient for *eoc* is both, positive and statistically significant at the one per cent level in all specifications (and for all breakdowns of the manufacturing employment) analysed here. This result is well supported by our dynamic panel data estimates presented in Table A2.2, Appendix 2, particularly in the case of blue collar workers. To a lesser extent, our first differences estimates from Table 4 point to a similar relationship. The simultaneous use of tariffs and the number of firms (in logs) as instruments (in Columns 2, 4, 6 and 8) yields less convincing results indicating that the *eoc* coefficient turns either statistically insignificant (for *all* and *blue collar workers*) or negative (for *white collar workers*). Overall, these results suggest that export orientation in manufacturing industries may be associated to larger shares of female workers in employment and some of the coefficients imply that this effect might be stronger amongst *blue collar workers*, a result that is somehow evident by comparing estimates from FE-IV and dynamic panel data. In the case of *white collar workers*, estimates from different methods provide a less coherent picture in terms of sign, size of coefficients and, statistical significance. From a conservative point of view, these results prove inconclusive in the case of *white collar workers* while they also suggest that export orientation might be associated to higher female shares of jobs in the *blue collar* category as indicated by the majority of our FE-IV and dynamic panel data estimates presented in Table 7 and Appendix 2, respectively.

The results in Table 7 also reveal the effects of other variables on the female share of jobs in manufacturing industries. In the case of our market concentration variable

(*CIGP*), there is strong evidence of its negative association with the female share of jobs for all labour force groupings analysed here. Coefficients for this variable are well determined in most cases, particularly for *all workers* and *white collar workers* where their statistical significance lies at the one per cent level in most cases. We observe also that the size and statistical significance of the coefficients tend to be reduced by the joint use of tariffs and the number of firms in logs as instruments of, respectively, *eoc* and *CIGP*. The negative association between the female share of jobs and our market concentration variable is better supported by our dynamic panel data results from Table A2.2 in Appendix 2 in which this variable is well determined in all cases.

Regarding our stock of capital per worker variables, results from Table 7 also indicate that both machinery (*lnkpw_mach*) and office equipment (*lnkpw_office*) exhibit a positive association with the female share of jobs for *all workers* and *white collar workers*. These results are equally confirmed by our dynamic panel data results from Table A2.2 in Appendix 2, according to which the coefficients for these two variables are well determined for the same the labour force groupings. In contrast, our measure of the stock of transport equipment per worker (*lnkpw_trans*) tends to be statistically insignificant in most cases, except for *all workers* when it is included simultaneously with the other two capital per worker variables just mentioned above.[9] We also find

[9] It should be highlighted that the sign of the coefficient for this variable in this case is negative. This result is just partially replicated by our dynamic panel data coefficients reported in Table A3.2.2 of the Statistical Appendix 3.2 where this variable appears statistically significant only for *blue collar workers*. These results suggest that manufacturing industries with a high intensity in the use of transport equipment tend to have lower proportions of jobs occupied by women, an interpretation that might be plausible if we take into account that occupations related to the operation of transport equipment tend to be performed almost exclusively by men in urban Colombia. This interpretation is supported by the household survey microdata used in this empirical application according to which around 98 per cent of those working as

strong evidence that the same three capital measures are uncorrelated with the female share of jobs amongst *blue collar workers*, as indicated by their coefficients in all specifications for this labour group. Compared to our dynamic panel estimates from Table A2.2 in Appendix 2, evidence of a relationship between the female share of jobs and the stock of capital equipment can only be confirmed in the case of *white collar workers* for machinery equipment and office equipment variables.

5. Concluding remarks

This paper provides new evidence on the relationship between trade reforms and employment outcomes by gender with an empirical application to Colombian manufacturing industries. Given some data limitations discussed below, our empirical approach had to innovate by looking at the effects of trade liberalisation on the gender composition of employment in manufacturing industries. Although the evidence presented in this paper does not formally test whether women are more (or less) discriminated in the labour market, our empirical results suggests that trade liberalisation, as well as some of the structural transformations in terms of the degree of market competition and the capital intensity of economic activities, are somehow related to the gender composition of employment in Colombian manufacturing industries.

We found convincing evidence that increased levels of import penetration are positively associated with higher female shares of jobs in manufacturing industries. Different econometric techniques presented in this paper point towards a similar

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"Transport Equipment Operators" between 1984 and 2004 are men, indicating that this occupation ranks as one of the most segregated in the labour market of this country.

conclusion and they indicate that this effect was probably stronger amongst *white collar workers*. Increasing levels of export orientation suggest a similar pattern although this result might be even more pronounced in the case of *blue collar workers*. Likewise, we found persuasive evidence that higher levels of market concentration as measured by our concentration index of gross product (*CIGP*) are negatively associated with the female share of jobs in manufacturing industries, indicating that more competitive environments are more likely to incorporate larger shares of female employment. So far, this is what we expected to find from the literature review presented in section 3.2.3 in relation to the segregation dimension implicit in Becker's hypothesis of labour market discrimination. As our dependent variable is the female share of jobs, we remain agnostic as to whether the effects of increased competition, either in the form of import penetration or in the form of market concentration, have any effect on the extent of gender pay discrimination. These results, however, suggest that increasing levels of competition are associated with higher shares of female employment and this is the type of result we would expect to encounter on the gender composition of the labour force if increased trade has an effect on gender discrimination. We should stress that increasing levels of female employment in manufacturing industries could occur with or without improvements in the bargaining position of women in the labour market. On this we should remember that higher levels of trade might also push employers to cost-cutting strategies that lessen the bargaining position of women, as suggested by the study of Berik et al. (Berik et al., 2004) for east Asian economies discussed in Section 3.2.4, above. In that Section, we reviewed also a study by Seguino (2000) who argues that low female wages might encourage the hiring of women workers in export oriented industries. For all these reasons, our findings are only suggestive of some of the positive effects of trade on gender differences in the labour market and further research is needed to establish whether the participation of women in Colombian manufacturing industries was accompanied by a reduction of gender discrimination.

We could also verify some complementarities between female labour and the use of some types of capital equipment. Our estimates under different panel data techniques are suggestive that the increasing use of office equipment is concomitant with higher shares of female employment in manufacturing industries of urban Colombia. This result was robust even in cases were other types of capital equipment were simultaneously controlled for. These findings provide further support of the hypothesis that the increasing use of technology favours the incursion of women in the labour market as they enjoy a comparative advantage in cognitive skills (Galor and Weil, 1996, Weinberg, 2000, Welch, 2000). This finding is further supported by the fact that the presumably positive effect derived from the increasing use of office equipment is econometrically stronger amongst the *white collar* group where the most qualified women tend to be concentrated. In the same vein, the fact that increasing female shares of jobs are also positively associated with increases in the use of machinery equipment suggests that the growing demand for intellectual skills not only favours the relative demand for female labour but also that this might entail an incentive for fertility decline as implied by Galor and Weil (1996).

We attempted to reconcile results from different econometric techniques, including FE-IV. The appropriateness of instruments and their validity in terms of both economic and statistical theory was assessed by comparing results drawn from different econometric techniques. The use of different methods to verify the relationships between the female share of jobs and some variables related to the economic development process provides a sound basis for statistical inference. We were fortunate to verify that many of these relationships we robust to the use of different instruments. From an empirical point of view, we believe that the results outlined along this paper are well supported by a number of methods pointing in the same direction.

The findings encountered along this paper also provide an interesting picture from an economic development perspective. To some extent, the evidence presented here is suggestive that the incorporation of women in manufacturing industries is concomitant not only with increased levels of trade, but also with capital intensification (in terms of machinery and office equipment) of productive processes in a number of industries. As a whole, our findings are consistent with Boserup's (1970) hypothesis according to which gender discrimination is inversely related to the level of economic development. This claim, however, deserves some qualifications as we could observe along this paper that these effects are highly differentiated across the labour market groupings defined in our data. In this sense, the selection of women into the *white collar workers* category appears to be more successful than the case of the blue collar workers category and this differentiated pattern appears to be biased in favour of the most qualified (and, presumably, better off) women.

Finally, it should be remarked that this investigation in its present state could be further developed in a number of ways. As explained above, it would be desirable to verify the effects of trade on gender wage differences. Employment data used in this paper come from the Annual Manufacturing Survey which does not provide disaggregated information on wages and labour costs by gender. This limitation in the availability of data rendered impossible the further investigation of the effects of trade policies on labour market outcomes from a gender perspective, particularly in regard to the paramount issue of wage differences. An alternative to this problem would be to use household survey microdata, which are available in Colombia on a regular basis since 1984. Based on statistical analyses not presented in this study, we found that this type of data has some limitations in terms of the accuracy in the recording of the information related to the ISIC codes to describe the economic activity of household respondents in the workforce, which is based on the supply side of the labour market, as opposed to the Annual Manufacturing Survey. Therefore, our attempts to verify a relationship between trade measures and gender wage gaps were inconclusive using household survey data but we believe this issue remains an important avenue for further research.

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Appendix 3.1

Table A1a Fixed-effects IV estimation of female share equations, trade variable: import penetration coefficient (ipc) –Endogeneity test of endogenous regressors

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
						All workers						
ipc	0.3163***	0.2887***	0.3375***	0.0662***	0.0782***	0.0595***	0.1600***	0.1752***	0.1698**	0.4281	0.1183***	0.0921***
	(0.0488)	(0.0276)	(0.0729)	(0.0169)	(0.0173)	(0.0168)	(0.0551)	(0.0362)	(0.0742)	(0.5911)	(0.0230)	(0.0221)
CIGP	-0.0046	-0.0121	-0.0032	-0.5102***	-0.5645***	-0.4990***	-0.4516***	-0.4496***	-0.4519***	-0.2955	-0.1166***	-0.0842***
	(0.0365)	(0.0327)	(0.0388)	(0.0832)	(0.0858)	(0.0857)	(0.0951)	(0.0957)	(0.0961)	(0.3754)	(0.0427)	(0.0286)
lnkpw_mach	-0.0042			0.0090***			0.0019			-0.1666		
	(0.0048)			(0.0031)			(0.0048)			(0.3242)		
lnkpw_trans		0.0003			-0.0025			-0.0016			-0.0346	
		(0.0021)			(0.0023)			(0.0022)			(0.0258)	
lnkpw_office			-0.0048			0.0075***			0.0001			0.0111**
			(0.0054)			(0.0024)			(0.0052)			(0.0053)
Endogeneity test of endogenous regressors												
χ² (1 or 2):	30.704	55.951	20.004	44.259	53.755	42.196	61.707	85.728	51.935	2.933	2.703	0.195
P-val =	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0868	0.1002	0.6592
						White-collar workers						
Ipc	0.6273***	0.7110***	0.5430***	0.2153***	0.2889***	0.1664***	0.6532***	0.7203***	0.5719***	0.2857	0.2878***	0.0900***
	(0.0853)	(0.0549)	(0.1120)	(0.0224)	(0.0251)	(0.0208)	(0.1008)	(0.0701)	(0.1230)	(0.3198)	(0.0326)	(0.0331)
CIGP	0.0614	0.0818	0.0555	-0.1385	-0.3930***	-0.0400	0.1353	0.1177	0.1330	-0.1820	-0.0981	-0.0239
	(0.0638)	(0.0651)	(0.0596)	(0.1101)	(0.1246)	(0.1063)	(0.1740)	(0.1856)	(0.1593)	(0.2031)	(0.0605)	(0.0428)
lnkpw_mach	0.0166**			0.0489***			0.0155*			0.0050		
	(0.0084)			(0.0041)			(0.0087)			(0.1754)		
lnkpw_trans		0.0100**			0.0062*			0.0101**			0.0081	
		(0.0042)			(0.0034)			(0.0043)			(0.0365)	
lnkpw_office			0.0191**			0.0453***			0.0183**			0.0658***
			(0.0083)			(0.0030)			(0.0086)			(0.0079)
Endogeneity test of endogenous regressors												
χ² (1 or 2):	44.896	111.032	20.326	0.084	4.839	0.038	46.380	111.640	21.804	0.085	0.150	6.178
P-val =	0.0000	0.0000	0.0000	0.7723	0.0278	0.8458	0.0000	0.0000	0.0000	0.7709	0.6988	0.0129

Table A1a (continuation)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
						Blue-collar workers						
ipc	0.1966***	0.0983***	0.2387***	0.0375**	0.0252	0.0424**	0.0579	0.0033	0.0914	0.4767	0.0508**	0.0934***
	(0.0461)	(0.0256)	(0.0697)	(0.0171)	(0.0163)	(0.0173)	(0.0560)	(0.0366)	(0.0753)	(0.7382)	(0.0242)	(0.0247)
CIGP	0.0081	-0.0183	0.0071	-0.4013***	-0.3588***	-0.4078***	-0.3885***	-0.3848***	-0.3869***	-0.3113	-0.0754*	-0.0658**
	(0.0344)	(0.0303)	(0.0371)	(0.0842)	(0.0809)	(0.0880)	(0.0967)	(0.0967)	(0.0975)	(0.4687)	(0.0449)	(0.0320)
lnkpw_mach	-0.0151***			-0.0082***			-0.0097**			-0.2336		
	(0.0045)			(0.0031)			(0.0049)			(0.4049)		
lnkpw_trans		0.0003			-0.0011			-0.0013			-0.0379	
		(0.0020)			(0.0022)			(0.0023)			(0.0271)	
lnkpw_office			-0.0140***			-0.0063***			-0.0096*			-0.0091
			(0.0052)			(0.0025)			(0.0053)			(0.0059)
Endogeneity test of endogenous regressors												
χ^2 (1 or 2):	9.183	2.558	7.708	27.542	22.804	25.879	31.23	22.879	28.768	3.991	1.933	1.311
P-val =	0.0024	0.1097	0.0055	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0457	0.1644	0.2523

Clustered robust standard errors in parentheses estimated with the **xtivreg2** Stata command developed by Schaffer (2010). The endogeneity test incorporated in this command is robust to heteroskedasticity and is compatible with clusted-robust standard errors. The test statistic is distributed as a Chi-squared with degrees of freedom equal to the number of tested regressor and is defined as a difference between two Sargan-Hansen tests from two models, one where the concerning variables are treated as endogenous and another where these variables are treated as exogenous.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$. N = 580 and t = 20 in all cases.

(1) to (3): import penetration coefficient (ipc) is instrumented with average tariffs.

(4) to (6): concentration index based on gross product (CIGP) instrumented with the natural logarithm of the number of firms.

(7) to (9): both, ipc and CIGP instrumented as described above.

(10) to (12): capital per worker variables instrumented with the natural logarithms of net investment per worker.

Table A1b Fixed-effects IV estimation of female share equations (pooled capital variables), trade variable: import penetration coefficient (ipc) - Endogeneity test of endogenous regressors

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	All workers				**White-collar workers**				**Blue-collar workers**			
ipc	0.3360***	0.0556***	0.1708**	-4.1142	0.5166***	0.1649***	0.5461***	2.7724	0.2443***	0.0424**	0.0959	-5.7164
	(0.0686)	(0.0169)	(0.0701)	(139.8278)	(0.1025)	(0.0210)	(0.1129)	(93.3871)	(0.0659)	(0.0175)	(0.0714)	(193.8449)
CIGP	-0.0033	-0.4971***	-0.4490***	3.4727	0.0460	-0.0337	0.1256	-2.3203	0.0110	-0.4117***	-0.3893***	4.8274
	(0.0381)	(0.0846)	(0.0946)	(118.7512)	(0.0569)	(0.1054)	(0.1524)	(79.3106)	(0.0366)	(0.0875)	(0.0964)	(164.6262)
lnkpw_mach	-0.0013	0.0035	0.0026	-4.6197	0.0048	0.0073	0.0041	3.0729	-0.0080*	-0.0041	-0.0045	-6.4093
	(0.0049)	(0.0046)	(0.0047)	(155.5796)	(0.0073)	(0.0058)	(0.0076)	(103.9073)	(0.0047)	(0.0048)	(0.0048)	(215.6818)
lnkpw_trans	0.0015	-0.0040*	-0.0018	4.0828	0.0051	-0.0017	0.0057	-2.7432	0.0040	0.0001	0.0011	5.6662
	(0.0027)	(0.0022)	(0.0026)	(137.9452)	(0.0040)	(0.0028)	(0.0042)	(92.1298)	(0.0026)	(0.0023)	(0.0026)	(191.2351)
lnkpw_office	-0.0042	0.0063*	-0.0011	1.6270	0.0175**	0.0416***	0.0169*	-0.9845	-0.0103*	-0.0041	-0.0076	2.2281
	(0.0058)	(0.0036)	(0.0055)	(54.0780)	(0.0086)	(0.0045)	(0.0089)	(36.1172)	(0.0055)	(0.0038)	(0.0056)	(74.9689)
Endogeneity test of endogenous regressors												
χ^2 (1 or 2):	23.099	42.619	54.813	4.331	20.239	0.084	21.878	15.844	9.230	27.374	31.239	3.797
P-val =	0.0000	0.0000	0.0000	0.2279	0.0000	0.7722	0.0000	0.0012	0.0024	0.0000	0.0000	0.2842

Clustered robust standard errors in parentheses estimated with the **xtivreg2** Stata command developed by Schaffer (2010). The endogeneity test incorporated in this command is robust to heteroskedasticity and is compatible with clusted-robust standard errors. The test statistic is distributed as a Chi-squared with degrees of freedom equal to the number of tested regressor and is defined as a difference between two Sargan-Hansen tests from two models, one where the concerning variables are treated as endogenous and another where these variables are treated as exogenous.

*** p<0.01, ** p<0.05, * p<0.1. N = 580 and t = 20 in all cases.

(1), (5), (9): import penetration coefficient (ipc) instrumented with average tariffs.

(2), (6), (10): concentration index based on gross product (CIGP) instrumented with the natural logarithm of the number of firms.

(3), (7), (11): both, ipc and CIGP instrumented as described above.

(4), (8), (12): capital per worker variables instrumented with the natural logarithms of net investment per worker.

Table A2a Fixed-effects IV estimation of female share equations, trade variable: export orientation coefficient (eoc)
Endogeneity test of endogenous regressors

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
						All workers						
eoc	-0.0562**	-0.0321	-0.0469**	0.0194**	0.0276***	0.0168*	-0.0015	0.0162	0.0023	-0.3140	0.0682***	0.0324***
	(0.0231)	(0.0217)	(0.0210)	(0.0097)	(0.0103)	(0.0094)	(0.0242)	(0.0229)	(0.0230)	(1.1172)	(0.0193)	(0.0100)
CIGP	-0.1633***	-0.2002***	-0.1234***	-0.5222***	-0.6004***	-0.5001***	-0.5540***	-0.6239***	-0.5210***	0.6267	-0.1398***	-0.1157***
	(0.0310)	(0.0325)	(0.0288)	(0.0844)	(0.0863)	(0.0871)	(0.0774)	(0.0795)	(0.0794)	(2.3912)	(0.0467)	(0.0322)
lnkpw_mach	0.0262***			0.0126***			0.0142***			0.4888		
	(0.0034)			(0.0031)			(0.0039)			(1.5157)		
lnkpw_trans		-0.0017			-0.0033			-0.0032			-0.0500*	
		(0.0022)			(0.0024)			(0.0025)			(0.0303)	
lnkpw_office			0.0217***			0.0106***			0.0113***			0.0103*
			(0.0023)			(0.0024)			(0.0029)			(0.0056)
Endogeneity test of endogenous regressors												
χ² (1 or 2):	23.458	21.458	18.684	37.928	49.759	36.211	60.205	70.14	54.523	6.003	4.82	1.741
P-val =	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0143	0.0281	0.1870
						White-collar workers						
Eoc	-0.1246***	-0.0487	-0.0960***	0.0737***	0.1123***	0.0576***	-0.1012***	-0.0050	-0.0850***	-0.1992	0.1404***	0.0363**
	(0.0410)	(0.0415)	(0.0341)	(0.0134)	(0.0160)	(0.0120)	(0.0383)	(0.0386)	(0.0326)	(0.9146)	(0.0284)	(0.0146)
CIGP	-0.2596***	-0.3615***	-0.1457***	-0.1614	-0.5048***	-0.0277	-0.4268***	-0.7453***	-0.2339**	0.4298	-0.1752**	-0.0529
	(0.0549)	(0.0621)	(0.0467)	(0.1170)	(0.1340)	(0.1116)	(0.1223)	(0.1338)	(0.1127)	(1.9577)	(0.0687)	(0.0469)
lnkpw_mach	0.0780***			0.0598***			0.0729***			0.4373		
	(0.0060)			(0.0043)			(0.0062)			(1.2409)		
lnkpw_trans		0.0049			0.0035			0.0036			-0.0331	
		(0.0041)			(0.0038)			(0.0042)			(0.0446)	
lnkpw_office			0.0633***			0.0533***			0.0610***			0.0647***
			(0.0038)			(0.0030)			(0.0042)			(0.0082)
Endogeneity test of endogenous regressors												
χ² (1 or 2):	39.13	26.556	30.496	0.001	4.579	0.339	39.154	30.773	30.882	1.668	1.833	1.694
P-val =	0.0000	0.0000	0.0000	0.9766	0.0324	0.5606	0.0000	0.0000	0.0000	0.1965	0.1758	0.1931

Table A2a (continuation)

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
							Blue-collar workers					
eoc	-0.0653***	-0.0592***	-0.0641***	0.0030	-0.0006	0.0032	-0.0170	-0.0226	-0.0191	-0.3561	0.0305	0.0242**
	(0.0237)	(0.0216)	(0.0227)	(0.0098)	(0.0095)	(0.0097)	(0.0243)	(0.0215)	(0.0239)	(1.1910)	(0.0187)	(0.0112)
CIGP	-0.1051***	-0.1141***	-0.0896***	-0.4202***	-0.3898***	-0.4213***	-0.4505***	-0.4350***	-0.4535***	0.7178	-0.0844*	-0.1009***
	(0.0318)	(0.0323)	(0.0311)	(0.0851)	(0.0795)	(0.0898)	(0.0776)	(0.0744)	(0.0825)	(2.5492)	(0.0451)	(0.0359)
lnkpw_mach	0.0066*			-0.0055*			-0.0041			0.4995		
	(0.0035)			(0.0031)			(0.0039)			(1.6158)		
lnkpw_trans		-0.0001			-0.0013			-0.0013			-0.0444	
		(0.0022)			(0.0022)			(0.0023)			(0.0293)	
lnkpw_office			0.0070***			-0.0037			-0.0025			-0.0093
			(0.0025)			(0.0024)			(0.0031)			(0.0063)
Endogeneity test of endogenous regressors												
χ^2 (1 or 2):	16.911	17.08	16.642	26.243	24.782	24.216	42.321	41.182	40.571	6.034	2.546	3.265
P-val =	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0000	0.0140	0.1106	0.0708

Clustered robust standard errors in parentheses estimated with the **xtivreg2** Stata command developed by Schaffer (2010). The endogeneity test incorporated in this command is robust to heteroskedasticity and is compatible with clusted-robust standard errors. The test statistic is distributed as a Chi-squared with degrees of freedom equal to the number of tested regressor and is defined as a difference between two Sargan-Hansen tests from two models, one where the concerning variables are treated as endogenous and another where these variables are treated as exogenous.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$. N = 580 and t = 20 in all cases.

(1) to (3): export orientation coefficient (eoc) instrumented with relative trade balance (see text for details).

(4) to (6): concentration index based on gross product (CIGP) instrumented with the natural logarithm of the number of firms.

(7) to (9): both, ipc and CIGP instrumented as described above.

(10) to (12): capital per worker variables instrumented with the natural logarithms of net investment per worker for each type of capital: machinery, transport and office equipment.

Table A2b Fixed-effects IV estimation of female share equations (pooled capital variables), trade variable: export orientation coefficient (eoc). Endogeneity test of endogenous regressors

VARIABLES	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
	All workers				**White-collar workers**				**Blue-collar workers**			
eoc	-0.0514**	0.0158*	-0.0026	-0.1514	-0.1058***	0.0565***	-0.0955***	0.1776	-0.0624***	0.0034	-0.0176	-0.2380
	(0.0214)	(0.0094)	(0.0233)	(0.6215)	(0.0350)	(0.0121)	(0.0334)	(0.4718)	(0.0230)	(0.0097)	(0.0242)	(0.8677)
CIGP	-0.1259***	-0.4980***	-0.5240***	0.3708	-0.1528***	-0.0230	-0.2370**	-0.2691	-0.0878***	-0.4247***	-0.4542***	0.5318
	(0.0290)	(0.0859)	(0.0792)	(1.4861)	(0.0474)	(0.1105)	(0.1139)	(1.1282)	(0.0311)	(0.0893)	(0.0824)	(2.0749)
lnkpw_mach	0.0029	0.0034	0.0041	-0.5158	0.0121*	0.0064	0.0123*	0.3742	-0.0042	-0.0038	-0.0030	-0.7318
	(0.0043)	(0.0047)	(0.0048)	(1.8036)	(0.0070)	(0.0060)	(0.0069)	(1.3692)	(0.0046)	(0.0049)	(0.0050)	(2.5182)
lnkpw_trans	-0.0055***	-0.0049**	-0.0051**	0.4066	-0.0058*	-0.0045	-0.0057*	-0.3161	-0.0013	-0.0007	-0.0009	0.5767
	(0.0020)	(0.0022)	(0.0023)	(1.4670)	(0.0033)	(0.0028)	(0.0033)	(1.1136)	(0.0022)	(0.0023)	(0.0024)	(2.0482)
lnkpw_office	0.0212***	0.0094**	0.0100**	0.2087	0.0579***	0.0504***	0.0555***	-0.0601	0.0095***	-0.0015	-0.0008	0.2689
	(0.0033)	(0.0037)	(0.0040)	(0.6700)	(0.0053)	(0.0047)	(0.0057)	(0.5087)	(0.0035)	(0.0038)	(0.0041)	(0.9355)
Endogeneity test of endogenous regressors												
χ^2 (1 or 2):	20.164	36.878	56.300	3.870	33.093	0.434	33.638	11.723	15.767	25.511	40.733	4.127
P-val =	0.0000	0.0000	0.0000	0.2759	0.0000	0.5102	0.0000	0.0084	0.0001	0.0000	0.0000	0.2481

Clustered robust standard errors in parentheses estimated with the **xtivreg2** Stata command developed by Schaffer (2010). The endogeneity test incorporated in this command is robust to heteroskedasticity and is compatible with clusted-robust standard errors. The test statistic is distributed as a Chi-squared with degrees of freedom equal to the number of tested regressor and is defined as a difference between two Sargan-Hansen tests from two models, one where the concerning variables are treated as endogenous and another where these variables are treated as exogenous.

*** p<0.01, ** p<0.05, * p<0.1. N = 580 and t = 20 in all cases.

(1), (5), (9): export orientation coefficient (eoc) instrumented with relative trade balance (see text for details).

(2), (6), (10): concentration index based on gross product (CIGP) instrumented with the natural logarithm of the number of firms.

(3), (7), (11): both, eoc and CIGP instrumented as described above.

(4), (8), (12): capital per worker variables instrumented with the natural logarithms of net investment per worker.

Appendix 2

The discussion about the validity of instruments in the context panel data has been widely documented in the literature. In order to have a yardstick of comparison for our FE-IV estimates, we also implement a dynamic panel data system strategy based on the GMM developed by Arellano and Bover (1995) and Blundell and Bond (1998). This GMM procedure consists of a simultaneous estimation of two equations, one in levels and another in differences with a set of instruments used in each equation. In principle, the general model can be expressed as:

$$y_{it} = \delta y_{i,t-1} + x'_{i,t}\beta + \rho_i + \varepsilon_{i,t} \, . \qquad \text{(A1)}$$

where $y_{i,t}$ represents the share of female jobs in the total number of jobs of industry *i* at year *t*, $x_{i,t}$ is a set of explanatory variables (in this case, a trade variable, a concentration index and, a capital stock per worker measure in logs – either machinery, transport or office equipment), ρ_i depicts a vector of industry fixed effects and, $\varepsilon_{i,t}$ is an *i.i.d.* random component. First differencing of (A1) allows the elimination of the industry fixed effects as follows,

$$y_{i,t} - y_{i,t-1} = \delta(y_{i,t-1} - y_{i,t-2}) + \beta(x_{i,t} - x_{i,t-1}) + (\varepsilon_{i,t} - \varepsilon_{i,t-1}). \qquad \text{(A2)}$$

In this specification, the choice of instruments aimed to solve endogeneity problems amongst the explanatory variables is performed in such a way that present realisations on the explanatory variables are influenced by past realisations of the dependent variable. Thus, instead of assuming strict ortogonality in the regressors, a less restrictive assumption of *weak exogeneity* is adopted. Under the two assumptions of (i) no serial autocorrelation in the residuals and, (ii) *weak exogeneity*, the following moment conditions apply:

$$E[y_{i,t-s}(\varepsilon_{i,t} - \varepsilon_{i,t-1})] = 0 \qquad \forall \qquad s \geq 2; t = 3, \ldots, T. \qquad \text{(A3)}$$

$$E\left[x_{i,t-s}\left(\varepsilon_{i,t}-\varepsilon_{i,t-1}\right)\right]=0 \qquad \forall \qquad s\geq 2; t=3,\dots,T. \qquad \text{(A4)}$$

Moment conditions (A3) and (A4) represent the basis for the GMM estimator of differences. This differences estimator is, however, characterised by low asymptotic precision and small sample biases and that is why it should be complemented with the regression equation in levels. Furthermore, when the lagged dependent and explanatory variables are persistent over time they represent weak instruments for the regression equation in differences (Blundell and Bond, 1998). According to Griliches and Hausman (1986), another problem is that the differences estimator is biased due to decreasing signal-to-noise ratios. For all of this, Arellano and Bover (1995) system estimator reduces potential biases by incorporating simultaneously the estimation of equations (A1) and (A2). Industry-specific effects at this stage ought to be controlled with instrumental variables for which lagged differences represent adequate instruments for the regression in levels. Even though, industry-specific effects may be correlated with right-hand side variables, there is no correlation between them when they are expressed in differences. Under these circumstances, the following stationarity property should hold,

$$E[y_{i,t+p}\cdot\rho_i]=E[y_{i,t+q}\cdot\rho_i]\,; \quad E[x_{i,t+p}\cdot\rho_i]=E[x_{i,t+q}\cdot\rho_i]\,; \quad \forall\, p \text{ and } q \quad \text{(A5)}$$

from which the additional moment conditions for this part of the system are given by

$$E\left[\left(y_{i,t-s}-y_{i,t-s-1}\right)\times\left(\rho_i+\varepsilon_{i,t}\right)\right]=0 \qquad for\ s=1, \qquad \text{(A6)}$$

$$E\left[\left(x_{i,t-s}-x_{i,t-s-1}\right)\times\left(\rho_i+\varepsilon_{i,t}\right)\right]=0 \qquad for\ s=1. \qquad \text{(A7)}$$

Conditions (A3) to (A7) provide the basis for the GMM procedure to generate consistent estimates of the parameters of interest in which the weighting matrix can be any symmetric, positive definite matrix (Arellano and Bover, 1995). From these matrices, the most efficient GMM estimator is generated by applying the weighting matrix based on the variance-covariance matrix for the moment conditions. Consistency of this GMM estimator relies on whether the validity of the lagged explanatory variables as adequate instruments holds in practice. According to Arellano and Bond (1991) and Arellano and Bover (1995), two tests can be implemented to verify the validity of such instruments, the Sargan test for over-identifying restrictions and, the second-order serial correlation test. The former is expressed as follows:

$$s = \hat{\epsilon}' Z \left(\frac{1}{N} \sum_{i=1}^{N} Z_i' \hat{\epsilon}_i \hat{\epsilon}_i' Z_i\right)^{-1} Z_i' \hat{\epsilon} \qquad \text{(A8)}$$

where $\hat{\epsilon}_{jt}$ are the estimated residuals and Z represents the set of valid instruments in the differenced equation. Under the null hypothesis that instruments are exogenous, S follows a χ^2_{m-r} distribution where $m - r$ is the number of instruments minus the number of exogenous variables. The Sargan test evaluates the overall validity of the instruments by assessing the sample analogue of the moment conditions used in the estimation process in which failure to reject the null hypothesis gives support to our model.

The second test examines the hypothesis of no serial correlation in the error term. In particular, we test whether the residuals from the regression in differences are first- and second-order serially correlated. Following Arellano and Bond (1991) and Arellano and Bover (1995), when this test fails to reject the null hypothesis of no second-order serial correlation, we conclude that the original error term is serially uncorrelated in accordance to the moment conditions set above.

Results for the GMM procedure outlined above are presented in Tables A2.1 and A2.2 in this Appendix. Models presented in Table A2.1 feature *ipc* as the trade explanatory variable whereas models in Table A2.2 use *eoc* as a trade variable. The layout of results in these two tables is divided along the breakdowns of the labour force outlined along this paper, namely, all workers (Columns 1 to 4), white collar workers (Columns 5 to 8) and, blue collar workers (Columns 9 to 12). For each of the labour force breakdowns, the three capital variables are introduced, first, one by one and then, simultaneously. Arellano-Bond test for first and second order serial correlation is presented at the bottom ob tables, followed by the Sargan test for the overall validity of the instruments. These results are used for reference to comment other models in the main text.

Table A2.1: Arellano-Bover/Blundell-Bond linear dynamic panel-data estimates: female share of jobs across manufacturing industries, 1981-2000. Trade variable: import penetration coefficient (ipc)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
VARIABLES	***All workers***				***White-collar workers***				***Blue-collar workers***			
Lagged Dep. Var.	0.7854***	0.7900***	0.7740***	0.7729***	0.6146***	0.6424***	0.5291***	0.5146***	0.8217***	0.8159***	0.8211***	0.8161***
	(0.0304)	(0.0304)	(0.0310)	(0.0316)	(0.0297)	(0.0295)	(0.0313)	(0.0322)	(0.0285)	(0.0282)	(0.0283)	(0.0286)
ipc	0.0076	0.0207**	0.0014	0.0006	0.0771***	0.1182***	0.0430**	0.0485**	0.0165	0.0119	0.0216*	0.0173
	(0.0114)	(0.0098)	(0.0128)	(0.0129)	(0.0210)	(0.0197)	(0.0206)	(0.0209)	(0.0115)	(0.0102)	(0.0127)	(0.0129)
CIGP	-0.0622***	-0.0587***	-0.0581***	-0.0599***	-0.1850***	-0.1519***	-0.1754***	-0.1592***	-0.0353**	-0.0332**	-0.0374**	-0.0355**
	(0.0147)	(0.0146)	(0.0145)	(0.0147)	(0.0194)	(0.0179)	(0.0170)	(0.0189)	(0.0148)	(0.0146)	(0.0149)	(0.0150)
lnkpw_mach	0.0054**			0.0026	0.0175***			0.0117**	-0.0017			0.0014
	(0.0024)			(0.0037)	(0.0038)			(0.0057)	(0.0025)			(0.0041)
lnkpw_trans		-0.0011		-0.0018		0.0061**		0.0013		-0.0039**		-0.0036**
		(0.0016)		(0.0016)		(0.0028)		(0.0027)		(0.0017)		(0.0017)
lnkpw_office			0.0049**	0.0036			0.0248***	0.0320***			-0.0024	-0.0023
			(0.0021)	(0.0033)			(0.0031)	(0.0048)			(0.0021)	(0.0036)
Constant	0.0392*	0.0861***	0.0588***	0.0555**	0.0637**	0.1454***	0.0995***	0.1465***	0.0675***	0.0768***	0.0675***	0.0769***
	(0.0219)	(0.0154)	(0.0149)	(0.0245)	(0.0286)	(0.0205)	(0.0164)	(0.0320)	(0.0242)	(0.0142)	(0.0162)	(0.0269)
Observations	551	551	551	551	551	551	551	551	551	551	551	551
Arellano-Bond test (p-values)												
First order	0.0012	0.0009	0.0013	0.0012	0.0024	0.0021	0.0033	0.0031	0.0022	0.0027	0.0022	0.0024
Second order	0.1996	0.2191	0.213	0.2049	0.225	0.2269	0.2338	0.2398	0.4506	0.4876	0.466	0.4681
***Sargan test:* χ^2(188)**	92.541	93.108	90.995	90.109	25.817	26.229	26.245	25.799	50.367	49.493	49.828	49.172
p-values	0.0520	0.0479	0.0647	0.0731	0.2135	0.1978	0.1972	0.2142	0.1050	0.1211	0.1147	0.1274

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table A2.2: Arellano-Bover/Blundell-Bond linear dynamic panel-data estimates: female share of jobs across manufacturing industries, 1981-2000. Trade variable: export orientation coefficient (eoc)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
VARIABLES	***All workers***				***White-collar workers***				***Blue-collar workers***			
Lagged Dep. Var.	0.7865***	0.7968***	0.7726***	0.7711***	0.6472***	0.7014***	0.5382***	0.5206***	0.8306***	0.8215***	0.8327***	0.8231***
	(0.0299)	(0.0297)	(0.0310)	(0.0317)	(0.0291)	(0.0286)	(0.0314)	(0.0326)	(0.0273)	(0.0273)	(0.0274)	(0.0279)
eoc	0.0069	0.0139**	0.0062	0.0053	0.0165	0.0387***	0.0102	0.0154	0.0144*	0.0122*	0.0152**	0.0136*
	(0.0072)	(0.0066)	(0.0072)	(0.0073)	(0.0132)	(0.0130)	(0.0123)	(0.0124)	(0.0076)	(0.0070)	(0.0077)	(0.0077)
CIGP	-0.0624***	-0.0577***	-0.0597***	-0.0612***	-0.1958***	-0.1537***	-0.1792***	-0.1626***	-0.0328**	-0.0313**	-0.0337**	-0.0323**
	(0.0147)	(0.0146)	(0.0144)	(0.0146)	(0.0195)	(0.0186)	(0.0171)	(0.0190)	(0.0147)	(0.0146)	(0.0148)	(0.0148)
lnkpw_mach	0.0053**			0.0023	0.0216***			0.0117**	-0.0017			0.0002
	(0.0023)			(0.0037)	(0.0038)			(0.0058)	(0.0024)			(0.0041)
lnkpw_trans		-0.0012		-0.0019		0.0041		-0.0000		-0.0041**		-0.0039**
		(0.0016)		(0.0016)		(0.0029)		(0.0027)		(0.0017)		(0.0017)
lnkpw_office			0.0044**	0.0034			0.0271***	0.0346***			-0.0017	-0.0010
			(0.0018)	(0.0029)			(0.0029)	(0.0047)			(0.0019)	(0.0032)
Constant	0.0401*	0.0867***	0.0620***	0.0603**	0.0344	0.1552***	0.0915***	0.1459***	0.0657***	0.0763***	0.0610***	0.0798***
	(0.0210)	(0.0156)	(0.0139)	(0.0250)	(0.0288)	(0.0218)	(0.0163)	(0.0332)	(0.0232)	(0.0143)	(0.0147)	(0.0271)
Number of obs.	551	551	551	551	551	551	551	551	551	551	551	551
Arellano-Bond test (p-values)												
First order	0.0014	0.0012	0.0015	0.0014	0.0035	0.0031	0.0043	0.004	0.0032	0.0032	0.0032	0.0032
Second order	0.1996	0.2022	0.2154	0.2087	0.2281	0.24	0.2358	0.2469	0.4407	0.4717	0.446	0.4562
***Sargan test:* χ^2(188)**	91.778	92.128	91.251	19.214	27.426	27.947	27.849	27.565	50.229	48.909	49.404	49.028
p-values	0.0580	0.0552	0.0624	0.5714	0.1572	0.1417	0.1445	0.1529	0.1074	0.1329	0.1228	0.1304

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.